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Sources of Cycles in Japan, 1975–1987*

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A simple real model is used to decompose movements of aggregate inventories and output in Japan during 1975 to 1987 to three components, one due to cost shocks, one due to demand shocks, and one due to shocks from abroad. Cost shocks are estimated to account for about one tenth of the movement in GNP, one half of the movement in inventories. Most of the remaining movement in GNP is due to demand shocks, in inventories to shocks from abroad. Confidence intervals around these point estimates are, however, very large. *J. Japan. Int. Econ.*, March 1992, **6**(1), pp. 71–98. Department of Economics, University of Wisconsin, Madison, Wisconsin 53706. © 1992 Academic Press, Inc.

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I. INTRODUCTION

A number of recent studies have considered the sources of fluctuations in aggregate output in the United States, breaking down these fluctuations into components due to cost and demand. A survey of several recent contributions can be found in Blanchard (1989). There seems, however, to be very little work that constructs similar breakdowns for other countries. Apart from the intrinsic interest in sources of fluctuations in other countries, such work could, in addition, shed light on theories of the business cycle that purport to explain fluctuations in market economies in general.

This paper uses a real model to consider sources of fluctuations in Japan, 1975 : 1 to 1987 : 4. The model is based on that in West (1990c), which

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derived dynamic linear aggregate demand and aggregate supply curves in a simple closed economy, general equilibrium inventory model. To account for the open economy aspects of Japan's economy, I simply add trade-weighted indices of real foreign output and the real exchange rate to the aggregate demand curve derived in West (1990c) and assume that these two variables evolve according to an exogenous bivariate vector autoregression. Estimates of the reduced form implied by the resulting aggregate demand and supply equations are then used to break down fluctuations in real aggregate inventories and output into components due to cost (a real unobservable that shifts the aggregate supply curve), demand (a real unobservable that shifts the aggregate demand curve), and the element of shocks to foreign output and the real exchange rate that is uncorrelated with cost and demand shocks.

The basic identifying assumption is suggested by a simple production smoothing model such as that in Holt *et al.* (1960). In such a model, increases in demand will tend to cause inventories to fall and output to rise, while decreases will tend to cause the opposite pattern; increases in cost will tend to cause both inventories and output to fall, decreases the opposite. The pattern of comovements of inventories and outputs thus reflects sources of fluctuations in a straightforward manner, with procyclical inventory movements resulting from cost shocks and countercyclical movements from demand shocks.

Such a simple pattern might not prevail in a model with empirically plausible dynamics. But even in a dynamic model the pattern of comovements can be used to identify cost and demand shocks (West, 1990c). In the present open economy context, one can also infer the additional effect of the component of movements in foreign output and the exchange rate that is uncorrelated with the unobservables that shift the Japanese aggregate demand and supply curves.

I turn to Japan for two reasons. The first is that Japanese inventory data apear to be unusually good. As in the United States, but unlike some other major industrialized countries (e.g., Germany, OECD, 1981), the figures on inventory investment are benchmarked against survey data every month. And these survey data appear to be unusually good (OECD, 1979). Second, Japan is distinct among the G7 countries in that its aggregate inventory movements are not markedly procyclical (West, 1990a,b). Given the model used, Japanese data therefore seemed especially likely to yield results that would provide an interesting contrast with U.S. data.

As one might predict given the model, I do indeed find that the component of inventory and GNP movements due to costs is less in Japan than in the United States. At business cycle horizons, about 10% of the variance of changes in GNP and 60% of the variance of inventory

investment are due to cost shocks. Most of the remaining variance in GNP is due to demand shocks, and in inventories is due to shocks from abroad.

By contrast, in West (1990c) I found that about 50% of the movement in U.S. GNP and over 90% of the movement in U.S. inventories are due to costs. The relatively low figures for costs in Japan are, at least superficially, inconsistent with real business cycle models in which the major source of output fluctuations is shocks to costs.

A by-product of the analysis is a decomposition of movements in foreign output and the real exchange rate. Foreign output tends to be driven largely by shocks uncorrelated with the unobservable cost and demand shocks that shift Japan's aggregate supply and demand curves, suggesting that there is a large nation-specific element to output fluctuations. The real exchange rate appears to be driven more by shocks to cost than to demand, suggesting (again!) that the standard monetary model of the exchange rate is not very satisfactory.

However, for at least three reasons the results should be interpreted with caution. First, point estimates are extraordinarily imprecise, perhaps because the sample is relatively short (12 years). The 95% confidence interval around the variance decompositions just described sometimes include values less than 10 and more than 90, indicating that essentially no reasonable hypothesis is markedly inconsistent with the data. Thus it would be hasty to draw any firm conclusions from the point estimates. Second, while most of the parameter estimates are plausible, those summarizing the effects of foreign output and the real exchange rate on aggregate demand perhaps are not, in that they imply larger short-run than long-run elasticities. This again underscores the tentative character of the conclusions. Finally, the model is purely real, excluding by assumption any possible role for nominal shocks and nominal rigidities. Readers who accord a prominent role to nominal factors might therefore consider it premature to place much weight on the results. In this connection, it should be emphasized that the aggregate demand and aggregate supply curves are not the usual textbook ones, since the latter usually combine nominal and real variables.

One final introductory remark: A more or less traditional production-smoothing model is the starting point for my model. Much has been made of the Japanese "just-in-time" system of inventory management, which seems to be in wide use in some manufacturing industries (see, e.g., Cusumano, 1985, on the automobile industry), and which is inconsistent with the basic assumptions of production-smoothing models. But such inventory management is still consistent with the possibility that the aggregate economy is led by an invisible hand to act as if it is minimizing costs that increase with the level of production and inventories in a fashion

assumed by production-smoothing models. Thus I do not consider a production-smoothing model a priori unsuited for analysis of the Japanese economy. Indeed, in related work (West, 1990b), I have found that in Japanese manufacturing, where just-in-time practices seem to be of the greatest importance, comovements of inventories and sales are similar to those of the aggregate data studied here, and the empirical results presented below suggest that my model does a tolerable job of explaining movements in these data.

Part II presents the model; Part III presents empirical results. Part IV concludes. The Appendix contains some algebraic details. An additional appendix available on request has some results omitted from the paper to save space.

II. Model

The model is an open economy version of the closed economy model in West (1990c). It is a real model in which the goods market in isolation determines output, inventories, and the ratio of output price to the wage. This price feeds into the asset market to determine the exchange rate, nominal price level, and so on. The model derived below determines the real goods market equilibrium, but leaves asset markets and determinations of nominal quantities unspecified. One could (but I do not) add a nominal side with a specification such as is in Flood and Hodrick (1985).[1]

The goods market equilibrium is determined as follows. In inverse form, the aggregate demand curve is

$$\begin{aligned} P_t &= -g_{0S}S_t + G_0'X_t + \cdots + G_n'X_{t-n} + U_{dt} \\ &\equiv -g_{0S}S_t + G(L)'X_t + U_{dt}, \qquad G_i \equiv (g_{if}, g_{ie})', \qquad X_t \equiv (Q_{ft}, e_t)'. \end{aligned} \tag{1}$$

In (1), P_t is the ratio of output price to the wage rate (see West, 1990c), S_t is real aggregate final sales, X_t is a vector of variables that shift the aggregate demand curve, $X_t \equiv (Q_{ft}, e_t)'$, Q_{ft} is a trade-weighted index of foreign output measured in foreign currency units, e_t is a trade-weighted index of the real exchange rate (yen/foreign currency), and U_{dt} is an unobservable shock to aggregate demand that captures shifts in preferences and policy. The g_{if}'s, g_{ie}'s, and g_{0S} are parameters, with g_{0S} positive.

[1] The real business cycle assumption that the goods market in isolation determines output was suggested by Flood and Hodrick (1985), who make the slightly less stringent assumption that the asset market does feed into the goods market, but only by revealing information about the value of unobserved goods market disturbances.

In (1) and throughout, constant and trend terms that are allowed in the empirical work are omitted here for simplicity.

S_t in equation (1) is the sum of (a) domestic sales of domestically produced goods and (b) net exports. Underlying this equation are equations determining each of these two components, such as (a) domestic final sales $= -(1/g_{0S})P_t +$ demand shock (see West, 1990c, for a derivation in a closed economy), and (b) net exports $= (g_{0f}/g_{0S})Q_{ft} + \cdots + (g_{nf}/g_{0S})Q_{ft-n} + (g_{0e}/g_{0S})e_t + \cdots + (g_{ne}/g_{0S})e_{t-n}$. Thus, foreign output Q_{ft} and the real exchange rate e_t appear in the demand curve in accordance with the standard notion that an increase in foreign output or depreciation of the real exchange rate increases net exports and thus domestic output. Since net exports are part of final sales S_t, the natural specification in a model that distinguishes between sales and production is to have S_t increase with increases in foreign output and depreciation of the real exchange rate. The distributed lags on Q_{ft} and e_t allow these to affect domestic demand with a delay.[2] (The length of the distributed lag is the same for both variables since this was suggested by the empirical work; in principle, the lag lengths might be different.)

An aggregate supply curve is derived from a version of Holt *et al.* (1960). The representative firm has per period costs:

$$C_t = g_{0Q}Q_t^2 + g_{0H}H_t^2 + 2U_{ct}(hH_t + Q_t). \tag{2}$$

In (2), Q_t is real production (GNP), H_t is real aggregate inventories, and U_{ct} is a cost shock that captures shifts in technology. The identity

$$Q_t = S_t + \Delta H_t \tag{3}$$

holds. The g_{if} are parameters ($g_{0Q} > 0$, $g_{0H} > 0$) with $h \geq 0$ a parameter that measures the cost shock's impact on inventory storage costs relative to its impact on production costs.

The first term in (2) reflects increasing costs to production and can be considered an approximation to an arbitrary convex cost function. The inventory term $g_{0H}H_t^2$ captures increasing marginal storage and holding

[2] It follows from West (1990c) that one could derive (1) in a general equilibrium model, in which consumers maximize the present value of a utility function that depends on leisure and consumption, by defining $G(L)'X_t + U_{dt}$ as a shock to preferences that makes consumption more desirable relative to leisure. But this seems to strain the notion of a preference disturbance.

costs. See West (1986) for additional discussion of this and the other term in the cost function.[3]

Let the representative competitive firm maximize expected discounted profits, using a constant discount rate b, $0 < b < 1$. (An observationally equivalent model results if one assumes monopoly or an oligopoly with a Nash equilibrium (Eichenbaum, 1984).) Let E_t be mathematical expectations (linear projections) conditional on period t information. The objective for the firm is to choose inventories and production to maximize expected present discounted profits:

$$\max_{T\to\infty} \lim E_t \sum_{j=0}^{T} b^j [P_{t+j} S_{t+j} - C_{t+j}].$$

Let c_t expected present discounted costs, $c_t \equiv E_t \sum_{j=0}^{\infty} b^j C_{t+j}$. After using the identity $S_t = Q_t - \Delta H_t$ to substitute out for S_t, differentiation of this objective function with respect to H_t and Q_t leads to two first-order conditions:

$$-P_t + bE_t P_{t+1} = \partial c_t/\partial H_t = 2g_{0H} H_t + 2hU_{ct}, \tag{4a}$$

$$P_t = \partial c_t/\partial Q_t = 2g_{0Q} Q_t + 2U_{ct} \tag{4b}$$

Equation (4a) states that the firm puts goods into inventory until the marginal storage cost $\partial c_t/\partial H_t$ equals the expected excess of discounted revenue over the opportunity cost of not selling today. Equation (4b) states that the firm equates marginal production cost to price.

The four equations (1), (2), (4a), and (4b) interact to determine equilibrium P_t, Q_t, S_t, and H_t, conditional on the shocks U_{ct} and U_{dt}, and the demand shift variables Q_{ft} and e_t. To solve how the shocks and shift variables interact to determine Q_t and H_t, begin by using (1) and then (3) to eliminate P_t and then S_t from (4). This leaves two dynamic first-order conditions in the two variables H_t and Q_t, written in vector form as

$$E_t[bA_1' Y_{t+1} + A_0 Y_t + A_1 Y_{t-1} + B_0 G(L)' X_t + bB_1 G(L)' X_{t+1} + D_0 U_t + bD_1 U_{t+1}] = 0. \tag{5}$$

In (5), Y_t is the (2×1) vector $(H_t, Q_t)'$, U_t is the (2×1) vector $(U_{ct}$,

[3] A cost of changing production often is included in (2), as is a cost of having inventories deviate from a target level proportional to expected sales (e.g., West, 1990c). These are excluded in the main part of the analysis here because, as explained below, the empirical work gives no indication of such costs being present. Results of some experimentation with a model that nonetheless allows a nonzero target level are reported below.

TABLE II
LAG LENGTHS IN ΔY_t VECTOR AUTOREGRESSION

A. Lags of ΔQ_{ft} and Δe_t in ΔY_t vector autoregression

Row	No. of lags		Lags tested equal to zero		Degrees of freedom	χ^2	p value
	ΔQ_{ft}	Δe_t	ΔQ_{ft}	Δe_t			
(1)	1	1	1	1	4	1.73	0.786
(2)	1	1	1		2	0.78	0.675
(3)	1	1		1	2	1.15	0.563
(4)	4	4	1–4	1–4	16	47.43	0.000
(5)	4	4	2–4	2–4	12	45.92	0.000
(6)	4	4	1	1	4	13.12	0.011
(7)	4	4	1–4		8	25.57	0.001
(8)	4	4	2–4		6	23.95	0.001
(9)	4	4	1		2	2.30	0.317
(10)	4	4		1–4	8	31.34	0.000
(11)	4	4		2–4	6	29.18	0.000
(12)	4	4		1	2	11.43	0.003

B. Lags of ΔH_t and ΔQ_t in ΔY_t vector autoregression

Row	No. of lags		Lags tested equal to zero		Degrees of freedom	χ^2	p value
	ΔH_t	ΔQ_t	ΔH_t	ΔQ_t			
(1)	1	0	1		2	21.24	0.000
(2)	1	1		1	2	2.74	0.253
(3)	2	0	2		2	1.17	0.357
(4)	2	1	2	1	4	3.59	0.467
(5)	4	0	2–4		6	2.44	0.875
(6)	1	4		1–4	8	12.41	0.135
(7)	4	4	2–4	1–4	14	15.20	0.367

Note. The "No. of lags" in A is the variable "n" in equations (1) and (7a). In A, all equations included a constant and one lag of ΔH_t. In B, all equations include a constant and four lags of ΔQ_{ft} and Δe_t.

lines (5) to (7) suggest that a univariate forecast is likely to be good as a bivariate forecast. I thus modeled Δe_t as a univariate AR(1) as well.

In sum, in the notation of equation (7), $p = 1$ and Φ_1 is a diagonal matrix. Since the data seem to suggest that four lags of ΔQ_{ft} and e_t are appropriate in the ΔY_t equation (Table II), the implication is that $n = 4$ in the demand curve (1).

I have modeled ΔQ_{ft} and Δe_t as evolving according to an exogenously specified vector autoregression. Although I did not spell this out in the discussion above, the solution of the model implicitly involves forecasting ΔQ_{ft} and Δe_t from this vector autoregression. If ΔQ_t and/or ΔH_t Granger

where the ΔX_t autoregression is repeated for convenience. In (7), Π is a (2×2) matrix whose second column is zero, and whose first column depends on b, g_{0Q}, g_{0S}, and g_{0H}. The Γ_i are 2×2 matrices that depend not only on b and the g_{ij} but also on the Φ_i. The (2×1) disturbance V_{yt} is a linear combination of the innovations u_{ct}, u_{dt}, and V_{xt}; its variance–covariance matrix depends on all the parameters of the system.

This paper uses (7) to decompose fluctuations in Japanese GNP, Q_t, into components due to cost, demand, and a residual that remains after the innovation V_{xt} is projected onto cost and demand. (Thus, common elements of foreign and Japanese cost and demand shocks are attributed to U_{ct} and U_{dt}, with the foreign residual reflecting only the idiosyncratic element of foreign cost and demand shocks.) See the Appendix for algebraic details.

III. Empirical Results

Described below in turn are data, identification and estimation, determination of the number of unit roots and selection of lag lengths, and empirical results.

A. *Data*

All the data used were real, quarterly and seasonally adjusted, 1974 : 1 to 1987 : 4; after some lags were used for initial conditions, the sample remaining for estimation spanned the 51 quarters from 1975 : 2 to 1987 : 4. I begin in 1974 : 1 because many authors have found a distinct change in the Japanese economy around the time of the first OPEC shock, with some authors (Komiya and Yasui, 1984; Yoshikawa and Ohtake, 1987) suggesting in particular that the export market became a strong force in Japanese business cycles at about this time.

The units for the inventory and GNP data are billions of 1980 yen. GNP and aggregate private inventory investment came from the OECD's *Main Economic Indicators* (MEI), as supplied on diskettes by VAR Econometrics. A series for the level of inventories (used only in the tests for unit roots and in Fig. 1) was constructed by combining the 1980 : 4 figure for the stock (the Japanese Economic Planning Agency's *Economic Statistics Annual*, 1987, p. 340) with the MEI figure for inventory investment.[5]

[5] The Economic Planning Agency's figure for the stock apparently includes some public inventories such as stocks of rice and thus does not match up exactly with my series for private inventory investment. But this just means that my series for the level is off by a constant (namely, the level of public inventories in 1980 : 4), and adding a constant to the entire series has no effect on unit root tests.

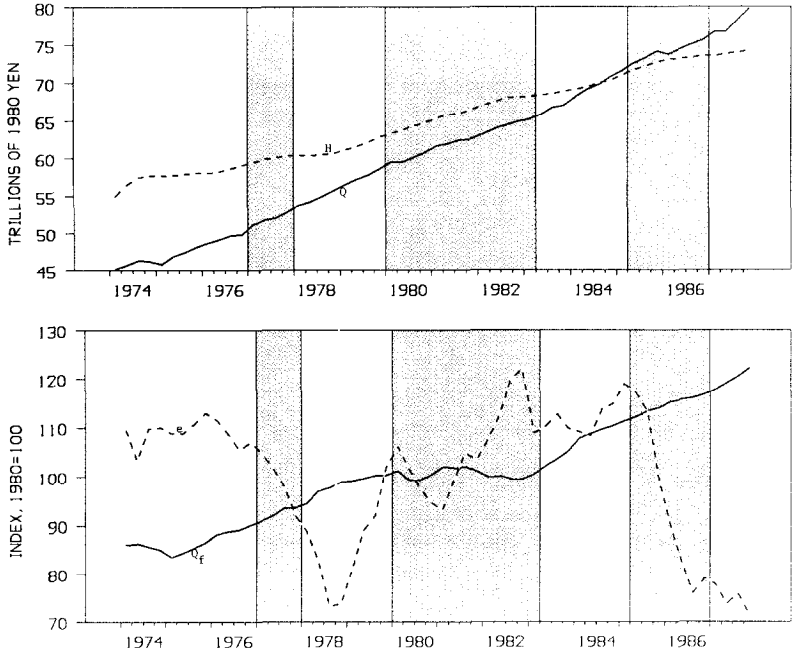

FIG. 1. Basic data.

As in Goldstein and Khan (1978) and Noland (1989), foreign output Q_{ft} was a trade-weighted index of the GNP of some of Japan's major trading partners. Data on GNP and exchange rates for the sample period were available from MEI for Australia, Canada, Germany, the United Kingdom, Italy, and the United States. Value (nominal yen) data on exports to these countries, 1982–1987, was obtained from the Bank of Japan's *Monthly Statistics of Japan* (March 1984, pp. 77–81). Weights were computed according to each country's share of total exports over that period: Australia, 0.06; Germany, 0.09; Canada, 0.05; the United Kingdom, 0.07; Italy, 0.01; and the United States, 0.72. Q_{ft} was then constructed by weighting indices of real GNP (1980 = 100), and e_t by weighting indices of the average real quarterly yen/foreign currency exchange rate (1980 = 100).

The set of six countries includes four of Japan's five largest export markets over 1982–1987 (South Korea is the exception). But since these six only account for about 35 to 40% of Japan's exports in the early as well as the later part of the sample (Allen, 1981, p. 162), this is still a rather noisy measure.

Plots of the data are given in Fig. 1. The intervals between peaks and troughs in the reference cycle, as determined by the Japanese Economic

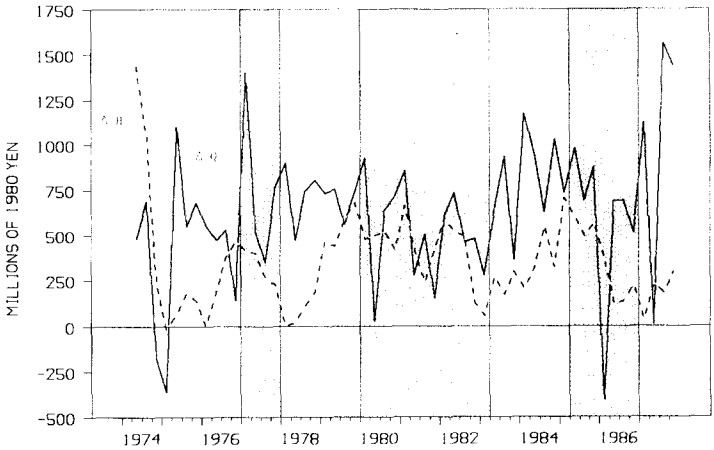

FIG. 2. Basic data—differences.

Planning Agency, are noted by shaded areas. As one would expect, given the large weight on the United States in the construction of the indices, the bottom panel of Fig. 1 indicates that it is not misleading to think of Q_{ft} as U.S. GNP and e_t as the real yen/dollar exchange rate: one sees in the path for e_t, for example, the post-1985 appreciation of the yen relative to the dollar.

Figure 2 plots the first difference of Japanese GNP and inventories. Two comments are of interest. First, it is evident that what the Economic Planning Agency identifies as a peak to trough slowdown rarely involves negative growth in GNP. Second, although the picture suggested to at least one reader some mild procyclicality of inventory movements, any such suggestion is more apparent than real. The correlation between ΔQ_t and ΔH_t is only 0.005. The comparable correlation for the United States for the same sample period is 0.35; as noted in the introduction, Japanese inventory movements are notably less procyclical than those in the United States or the other G7 countries.

B. Identification and Estimation

To determine whether unit roots are present and to select the lag lengths n and p in (7), OLS or seemingly unrelated regressions and some standard asymptotic tests were used as described below. Once these were chosen, seemingly unrelated regressions were used to estimate both unrestricted and restricted versions of (7) and confidence intervals were constructed by bootstrapping as described below. The restricted version imposed the cross-equation restrictions implied by the model. These are written out

in detail in the Appendix. A brief overview of details useful in interpreting the results follows.

A discount rate $b = 0.98$ was imposed a priori. In related research (West, 1986) I have found that the results are remarkably insensitive to the exact value of the discount rate chosen, so I did no experimentation with other values of b.

It follows from (5) and the formulas given in the Appendix that the g_{ij} parameters are identified only up to a normalization: doubling all the g_{ij} results in identical first-order conditions, apart from a rescaling of the disturbances. Consistent with West (1990c), the normalization chosen was $g_{0S} + g_{0Q} = 1$. Together with this normalization, the two nonzero elements of the unrestricted estimate of Π determine g_{0S}, g_{0Q}, and g_{0H}.

These unrestricted estimates were used to get an estimate of the restricted reduced form, which was then used to get estimates of the g_{if}'s and g_{ie}'s (the parameters of the distributed lags on Q_{ft} and e_t in the demand curve), as follows. As explained in the Appendix, there are $2n$ nonlinear restrictions on the $4n$ parameters contained in the matrices $\Gamma_1, \ldots, \Gamma_n$. Conditional on a given set of values of Π and of $\Phi_1, \ldots, \Phi_p$, however, these restrictions are linear. To exploit the computational convenience of this linearity, I set the restricted estimates of $\Pi, \Phi_1, \ldots, \Phi_p$, equal to the unrestricted ones and then imposed the $2n$ linear restrictions.

A test of the cross-equation restrictions, as well as confidence intervals around various parameter estimates, was constructed by bootstrapping; in Rocke's (1989) simulations, bootstrap tests of linear restrictions in a time series model yielded accurately sized tests even in a sample as small as that used here. Specifically: the residuals from the restricted regression were sampled with replacement and data were generated recursively using the restricted estimates, with the actual historical values as initial conditions. The unrestricted and restricted reduced forms were estimated using the generated data, the usual likelihood ratio statistic was computed, and parameter estimates were computed from the restricted regression. This was done 1000 times. The 95% confidence intervals reported below were constructed by dropping the lower and upper 2.5% of the empirical distribution.

The Appendix shows that the n free parameters in the matrices $\Gamma_1, \ldots, \Gamma_n$ determine $2n$ linear combinations of the $2(n + 1)$ parameters $g_{0f}, g_{0e}, \ldots, g_{nf}, g_{ne}$. Two additional bits of prior information are needed to identify the entire sets of g_{if}'s and g_{ie}'s. The information I used were figures for $\Sigma_0^n g_{if}$ and $\Sigma_0^n g_{ie}$, picked to yield implied elasticities for sample means that matched values estimated by others. Noland (1989) and some earlier studies using Japanese data summarized in Goldstein and Kahn (1985) suggest an elasticity of exports with respect to foreign output of about 1 to 5 and of exports and of imports with respect to the real exchange rate of about 0.5 to 5. From (1), the long-run elasticity of S_t with respect to Q_{ft}

evaluated at sample means is $(\partial S_t/\partial Q_{ft})$ (mean Q_{ft}/mean S_t) = $(\Sigma g_{if}/g_{0S})$ (mean Q_{ft}/mean S_t). We have $(\partial S_t/\partial Q_{ft})$ (mean Q_{ft}/mean S_t) = $(\partial S_t/\partial$exports) × [$(\partial$exports$/\partial Q_{ft})$ (mean Q_{ft}/mean exports)] × (mean exports/mean S_t). Since $\partial S_t/\partial$(exports) = 1, then [$(\partial$exports$/\partial Q_{ft})$ (mean Q_{ft}/mean exports)] = $(\Sigma g_{if}/g_{0S})$ (mean Q_{ft}/mean S_t). I thus chose Σg_{if} so that 2 = $(\Sigma g_{if}/g_{0S})$ (mean Q_{ft}/mean exports). Similarly, I chose Σg_{ie} so that 2 = $(\Sigma g_{ie}/g_{0S})$ [(mean e_t/mean exports) + (mean e_t/mean imports)].

Since mean exports and imports were each about a sixth of mean final sales, the implied long-run elasticities of sales with respect to foreign output and with respect to the real exchange rate evaluated at sample means are each about 0.3. As noted below, experimentation with alternative imposed values for these elasticities resulted in very little change in final results.

C. *Unit Roots and Lag Lengths*

The first step in the empirical work was to model the trends in the series. Univariate Dickey–Fuller tests were run on each of the four series. See Table IA. The p values in the last column come from Fuller (1976); simulations in Schwert (1987) suggest that these p values are unlikely to be misleading even in samples of the size used here.

The tests involved regressing the difference of each series on a constant, time trend, a lagged level, and either no differences or four-lagged differences. The Q statistics (not reported) suggested that either lag length sufficed to reduce the disturbance to white noise. The t statistics on the lagged levels, reported in lines (1) to (4) of Table IA indicate that the null of a unit root cannot be rejected in any of the series at conventional significance levels.

To see whether second differencing appeared to be required to induce stationarity, I then regressed the second difference of each series on a constant, a lagged first difference, and either none or four additional second differences. The t statistics on the lagged first differences, reported in lines (5) to (8) of Table IA, indicate that the null of a unit root can be rejected for a lag length of 0 but not for a lag length of 4. This latter lack of rejection, however, seemed to occur because of collinearity among the regressors: none of the regressors was significant at the 5% level in any of the equations. Since examination of the autocorrelations of each of the first differenced series (not reported) suggested that each was stationary, I conclude that first differencing is likely to induce stationarity in each of the four series.

Additional evidence consistent with this is in multivariate tests reported in Table IB. These tests were done as suggested in Johansen (1988) and Johansen and Juselius (1989), with the two tests differing only in that the regressions underlying lines (1), (3), and (5) include both constants and trend terms, those underlying lines (2), (4), and (6) only constants. The

TABLE I
TESTS FOR UNIT ROOTS

A. Univariate tests

	Variable	Lags	t statistic	p value
(1)	H_t	0	−1.49	>0.10
		4	−2.47	>0.10
(2)	Q_t	0	−2.45	>0.10
		4	−1.25	>0.10
(3)	Q_{ft}	0	−0.94	>0.95
		4	−1.53	>0.10
(4)	e_t	0	−0.82	>0.95
		4	−1.46	>0.10
(5)	ΔH_t	0	−4.88	<0.01
		4	−2.88	>0.10
(6)	ΔQ_t	0	−7.44	<0.01
		4	−2.46	>0.10
(7)	ΔQ_{ft}	0	−4.51	<0.01
		4	−2.84	>0.10
(8)	Δe_t	0	−4.59	<0.01
		4	−3.18	>0.10

B. Multivariate test, four-variable system

	Test	Statistic	p value
(1)	H_0: 4 unit roots in H_t, Q_t, Q_{ft}, and e_t, against H_A: no unit roots around deterministic trend	53.6	0.14
(2)	H_0: 4 unit roots in H_t, Q_t, Q_{ft}, and e_t, against H_A: no unit roots, no deterministic trend	44.4	0.18
(3)	H_0: 2 unit roots in H_t and Q_t, against H_A: no unit roots around deterministic trend	4.7	0.92
(4)	H_0: 2 unit roots in H_t and Q_t, against H_A: no unit roots, no deterministic trend	5.5	0.99
(5)	H_0: 2 unit roots in Q_{ft} and e_t, against H_A: no unit roots around deterministic trend	15.2	0.39
(6)	H_0: 2 unit roots in Q_{ft} and e_t, against H_A: no unit roots, no deterministic trend	5.1	0.90

bivariate tests in lines (3) to (6) were executed as a check on the four variable tests in lines (1) and (2). All regressions include five lagged levels of all four (lines (1) and (2)) or two (lines (3) to (6)) variables. The p values and computer code for these tests were kindly supplied by James H. Stock. As can be seen in Table IB, none of the tests rejects the null of as many unit roots as variables.

Of course, all these tests have low power against borderline stationary alternatives, particularly since the sample includes only about 50 data points. But I doubt that anything of importance hinges on the conclusion that there are four unit roots in the four series: in an earlier, related study

using U.S. data (West, 1990c), I found little substantive difference between difference stationary and trend stationary specifications. Thus I consider the unit root assumption a convenient but not essential simplification, and I do not consider trend stationary specifications as well only because I doubt that they will produce qualitatively different results.

I now turn to selection of the lag lengths for the ΔX_t and ΔY_t autoregressions in (7), for which I used asymptotic likelihood ratio tests with the degrees of freedom correction suggested by Sims (1980). My aim is not to select a best model in any traditional time series sense, but simply to begin my analysis with an unrestricted reduced form that is as parsimonious as possible and still consistent with the model and data.

Consider first selection of n, the number of lags of ΔX_t appearing in equation (7a), the reduced form for ΔY_t. Tests in Table IIA suggest that four lags are appropriate: the first lag of either or both variables are by themselves insignificant (lines (1) to (3)), but tend to become significant when lags 2–4 are included (lines (6), (9), and (12)). These additional lags are significant as well (lines (5), (8), and (11)) as are lags 1–4 as a whole (lines (4), (7), and (10)). I thus set $n = 4$ in (7).

Table IIB presents some tests for lags of ΔQ_t and ΔH_t in the ΔY_t vector autoregression, computed under the maintained assumption that $n = 4$. It does not appear that the ΔY_t vector autoregression, which according to the model includes only one lag of ΔH_t and none of ΔQ_t, is misspecified in the sense of including an irrelevant lag of ΔH_t (line (1)) or omitting relevant lags of ΔH_t and ΔQ_t (lines (2) to (7)). Note in particular that one cannot reject the null that an additional lag of ΔQ_t does not belong in the reduced form (lines (2) and (6)). It is this test statistic that explains why the model does not include a cost of changing production, which is standard in studies using U.S. data (e.g., West, 1990b): it can be shown that such a cost will put a lag of ΔQ_t in the reduced form.

Consider now selection of p, the lag length of the ΔX_t autoregression. It may be shown that, according to the model, one possible reason that four lags of ΔQ_{ft} and Δe_t would appear in the ΔY_t equations is that $p = 4$ as well. But univariate tests on the ΔQ_{ft} equation reported in lines (1) to (4) of Table IIIA suggest $p = 1$ for the ΔQ_{ft} equation: when one lag is included, it is highly significant (line (1)); when lags 2–4 are included as well, the null that the additional coefficients are zero is easily rejected (line (4)). Bivariate tests reported in lines (5) to (7) of Table IIIA suggest that the null that Δe_t does not Granger cause ΔQ_{ft} cannot be rejected at conventional significance ratios. I thus modeled ΔQ_{ft} as a univariate AR(1).

The parallel tests for the Δe_t equation are reported in Table IIIB. Once again, lines (1) to (4) suggest that a good univariate model is an AR(1);[6]

[6] Some mild mean reversion in the spot yen/dollar exchange rate has been noted previously (e.g., Huizanga, 1987). The sharp estimate of a nonzero autocorrelation at lag 1 in the index used here probably reflects the time averaging of the point in time spot rate.

TABLE II
LAG LENGTHS IN ΔY_t VECTOR AUTOREGRESSION

A. Lags of ΔQ_{ft} and Δe_t in ΔY_t vector autoregression

Row	No. of lags		Lags tested equal to zero		Degrees of freedom	χ^2	p value
	ΔQ_{ft}	Δe_t	ΔQ_{ft}	Δe_t			
(1)	1	1	1	1	4	1.73	0.786
(2)	1	1	1		2	0.78	0.675
(3)	1	1		1	2	1.15	0.563
(4)	4	4	1–4	1–4	16	47.43	0.000
(5)	4	4	2–4	2–4	12	45.92	0.000
(6)	4	4	1	1	4	13.12	0.011
(7)	4	4	1–4		8	25.57	0.001
(8)	4	4	2–4		6	23.95	0.001
(9)	4	4	1		2	2.30	0.317
(10)	4	4		1–4	8	31.34	0.000
(11)	4	4		2–4	6	29.18	0.000
(12)	4	4		1	2	11.43	0.003

B. Lags of ΔH_t and ΔQ_t in ΔY_t vector autoregression

Row	No. of lags		Lags tested equal to zero		Degrees of freedom	χ^2	p value
	ΔH_t	ΔQ_t	ΔH_t	ΔQ_t			
(1)	1	0	1		2	21.24	0.000
(2)	1	1		1	2	2.74	0.253
(3)	2	0	2		2	1.17	0.357
(4)	2	1	2	1	4	3.59	0.467
(5)	4	0	2–4		6	2.44	0.875
(6)	1	4		1–4	8	12.41	0.135
(7)	4	4	2–4	1–4	14	15.20	0.367

Note. The "No. of lags" in A is the variable "n" in equations (1) and (7a). In A, all equations included a constant and one lag of ΔH_t. In B, all equations include a constant and four lags of ΔQ_{ft} and Δe_t.

lines (5) to (7) suggest that a univariate forecast is likely to be good as a bivariate forecast. I thus modeled Δe_t as a univariate AR(1) as well.

In sum, in the notation of equation (7), $p = 1$ and Φ_1 is a diagonal matrix. Since the data seem to suggest that four lags of ΔQ_{ft} and e_t are appropriate in the ΔY_t equation (Table II), the implication is that $n = 4$ in the demand curve (1).

I have modeled ΔQ_{ft} and Δe_t as evolving according to an exogenously specified vector autoregression. Although I did not spell this out in the discussion above, the solution of the model implicitly involves forecasting ΔQ_{ft} and Δe_t from this vector autoregression. If ΔQ_t and/or ΔH_t Granger

TABLE III
LAG LENGTHS IN ΔX_t VECTOR AUTOREGRESSION

A. Lags of ΔQ_{ft} and e_t in ΔQ_{ft} autoregression

Row	No. of lags ΔQ_{ft}	No. of lags Δe_t	Lags tested equal to zero ΔQ_{ft}	Lags tested equal to zero Δe_t	Degrees of freedom	χ^2	p value
(1)	1	0	1		1	11.56	0.001
(2)	4	0	1		1	4.94	0.026
(3)	4	0	1–4		4	7.71	0.103
(4)	4	0	2–4		3	0.84	0.841
(5)	1	1		1	1	2.21	0.137
(6)	1	4		1–4	4	5.71	0.222
(7)	4	4		1–4	4	5.24	0.264

B. Lags of ΔQ_{ft} and e_t in Δe_t autoregression

Row	No. of lags ΔQ_{ft}	No. of lags Δe_t	Lags tested equal to zero ΔQ_{ft}	Lags tested equal to zero Δe_t	Degrees of freedom	χ^2	p value
(1)	0	1		1	1	8.72	0.003
(2)	0	4		1	1	9.91	0.002
(3)	0	4		1–4	4	11.52	0.021
(4)	0	4		2–4	3	1.68	0.642
(5)	1	1	1		1	0.96	0.327
(6)	4	1	1–4		4	2.40	0.663
(7)	4	4	1–4		4	2.62	0.623

C. Lags of ΔH_t and ΔQ_t in ΔX_t vector autoregression

Row	No. of lags ΔH_t	No. of lags ΔQ_t	Lags tested equal to zero ΔH_t	Lags tested equal to zero ΔQ_t	Degrees of freedom	χ^2	p value
(1)	4	4	1–4	1–4	16	23.41	0.103
(2)	4	0	1–4		8	18.76	0.016
(3)	0	4		1–4	8	5.97	0.650

Note. The "number of lags in the autoregression" in A and B is the variable "p" in equation (7b). In A, B, and C, all equations included constants. In C, each equation included one lag of the left-hand side variable in addition to the indicated lags of ΔH_t and ΔQ_t.

cause ΔQ_{ft} or e_t, more efficient forecasts can be produced using this information. Table IIIC tests for additions of such lags to the AR(1) process just identified. There is mixed evidence, but it appears from line (2) that ΔH_t helps predict ΔX_t. If so, it may be shown that, according to the model, there should then be lags of ΔH_t beyond the first in the ΔY_t autoregression, which is contrary to the evidence in Table IIB.

TABLE IV
UNRESTRICTED REDUCED FORM

Independent variables/ summary statistics	Dependent variables			
	ΔH_t	ΔQ_t	ΔQ_{ft}	Δe_t
ΔH_{t-1}	0.505	−0.128		
	(0.228,0.661)	(−0.855,0.571)		
ΔQ_{ft-1}	16.51	85.73	0.347	
	(−29.1,57.06)	(−188.8,125.5)	(0.079,0.533)	
ΔQ_{ft-2}	−9.42	45.11		
	(−62.6,29.48)	(−176.4,139.1)		
ΔQ_{ft-3}	0.71	−134.30		
	(−47.8,41.60)	(−165.6,143.5)		
ΔQ_{ft-4}	81.07	176.15		
	(29.0,117.5)	(−169.3,120.5)		
Δe_{t-1}	−2.12	35.31		0.429
	(−11.2,5.303)	(−30.56,18.58)		(0.121,0.611)
Δe_{t-2}	10.89	−45.92		
	(1.43,17.64)	(−29.35,23.01)		
Δe_{t-3}	−1.39	43.55		
	(−9.33,5.45)	(−27.94,23.84)		
Δe_{t-4}	8.45	−19.32		
	(−0.41,13.80)	(−27.48,23.39)		
$Q(21)$	21.30	18.32	18.67	18.95
[p value)	[0.44]	[0.63]	[0.61]	[0.59]
$\bar{R}^2$	0.57	0.38	0.61	0.18

Note. The reduced form is Eq. (7). The 95% conficence intervals are in parentheses, from bootstrap. Constant terms were included in all regressions.

One possible reconciliation is to assume that the tests in lines (5) of (7) of Table IIB and line (1) of Table IIIC incorrectly accept the null that ΔX_t evolves exogenously, the other that the test in line (2) of Table IIIC incorrectly rejects that null. Given how short the sample is, I have decided that the principle of parsimony calls for the latter resolution. So I will proceed under the assumption that ΔX_t evolves exogenously.

D. Empirical Results

I discuss in turn parameter estimates, variance decompositions, variance decompositions under some alternative sets of assumptions, and a historical decomposition of errors in forecasting GNP into cost, demand, and foreign components.

Table IV presents the unrestricted seemingly unrelated regression estimates of the four-variable system. Constant terms were also included in

all four regressions but are not reported here, to conserve space. Table V has estimates of the restricted reduced form, with 95% confidence intervals from bootstrapping in parentheses. Table VA indicates that the test of cross-equation restrictions rejects at the 0.043 level. In comparing Tables IV and VA, it appears that the numerical sense in which the data are unhappy with the restriction is that the restriction substantially changes the coefficients in the ΔQ_t equation (as noted in the tables, $\bar{R}^2$ falls from 0.38 to -0.28). It is unclear to me how the model should be modified to better match the data. In a similar model in West (1990c) tests of overidentifying restrictions also rejected, but a more loosely parameterized (and more difficult to interpret) model yielded substantively similar results. Whether the same applies in the present case I have not shown. But in light of West (1990c), I will proceed, though cautiously, despite the test statistic.

Table VB has estimates of the implied parameters of supply and demand. Three of the estimated parameters (g_{0S}, g_{0H}, and g_{0Q}) were also estimated in West (1990c) for the United States. The point estimates of the demand curve slope g_{0S} and the slope of the inventory marginal cost curve g_{0H} are quite similar to those for the United States estimated in West (1990c), as is the slope of the production marginal cost curve (though confidence intervals are so large that the estimates are consistent with practically any hypothesis of interest).[7] For the United States, however, I estimated g_{0Q} to be negative, and the positive slope to the marginal cost curve came from steep costs of adjusting production. In Japan, however, we see from Table VB and from the tests reported in lines (2) and (6) of Table IIB that costs of adjusting production appear to be zero and the positive slope to the marginal cost curve comes from steep production costs. That costs of adjustment are larger in the United States than in Japan is consistent with the results of Morrison (1989, pp. 21–22), who, using a traditional production function approach, found that over the 1960–1981 period Japan had much excess capacity relative to the United States.

The g_{if}'s (the coefficients in the distributed lag on Q_{ft} in the demand curve) begin positive and then turn negative. The point estimate of g_{0f} indicates that the instantaneous impact of a one-point increase in the index

[7] The large confidence intervals raise the question of the reliability of the asymptotic tests used in selecting the lag lengths in the reduced form, since such tests were able to reject a number of hypotheses quite sharply. That bootstrap tests would likely lead to similar conclusions is suggested by the perfect match between asymptotic and bootstrap tests of whether coefficients in the unrestricted reduced form were significantly different from zero at the 95% level (not reported in any table; these bootstrap tests used the unrestricted residuals rather than the restricted residuals used in Table IV). Thus here as in West (1990c) the data yield sharper inferences about the unrestricted reduced form than about the underlying parameters of interest.

TABLE V
RESTRICTED ESTIMATES

Independent variables/ summary statistics	A. Reduced form[a] Dependent variables			
	ΔH_t	ΔQ_t	ΔQ_{ft}	Δe_t
ΔH_{t-1}	0.505	−0.128		
	(0.228,0.661)	(−0.855,0.571)		
ΔQ_{ft-1}	13.99	−24.44	0.347	
	(−30.65,57.87)	(−103.9,55.89)	(0.079,0.533)	
ΔQ_{ft-2}	−14.95	−12.84		
	(−65.18,27.68)	(−58.58,37.82)		
ΔQ_{ft-3}	−4.18	−16.96		
	(−55.00,37.77)	(−77.71,45.32)		
ΔQ_{ft-4}	72.76	−18.39		
	(9.01,108.5)	(−96.99,56.29)		
Δe_{t-1}	−2.61	−4.951		0.429
	(−11.97,4.33)	(−20.89,11.23)		(0.121,0.611)
Δe_{t-2}	9.86	−3.600		
	(0.66,16.99)	(−16.48,9.74)		
Δe_{t-3}	−2.13	−1.130		
	(−10.62,4.81)	(−8.75,8.13)		
Δe_{t-4}	6.73	−1.702		
	(−3.16,12.59)	(−10.09,6.12)		
$Q(21)$	20.51	15.81	18.67	18.96
[asymptotic	[0.49]		[0.61]	[0.59]
p value]		[0.78]		
$\bar{R}^2$	0.56	−0.28	0.12	0.18
	B. Structural parameters[b]			
g_{0S}	g_{0H}	g_{0Q}		
0.258	0.095	0.742		
(−1.172,1.689)	(−1.713,0.365)	(−0.757,2.170)		
g_{0f}	g_{1f}	g_{2f}	g_{3f}	g_{4f}
213.5	−102.2	−9.0	−15.9	−37.2
(−516.2,872.4)	(−451.6,248.4)	(−83.1,70.8)	(−97.1,56.4)	(−175.0,106.8)
g_{0e}	g_{1e}	g_{2e}	g_{3e}	g_{4e}
113.9	−53.2	−6.1	−0.6	−3.4
(−404.5,678.0)	(−325.7,167.3)	(−28.4,19.0)	(−15.2,18.3)	(−18.3,10.8)

[a] Likelihood ratio test of restrictions = 31.99 [bootstrap p value = 0.043]. The 95% confidence intervals are in parentheses, from bootstrap.

[b] The 95% confidence intervals are in parentheses, from bootstrap. The parameters are related by the normalization $1 = g_{0S} + g_{0Q}$. The individual g_{if}'s and g_{ie}'s were identified by imposing values for Σg_{if} and Σg_{ie} consistent with the long-run elasticity of S_t with respect to both ΔQ_{ft} and e_t being approximately 0.3.

of foreign output (roughly 3 to 6 trillion yen, depending on the exchange rate) results in a 214 billion yen increase in aggregate demand; the implied elasticity at sample means is about 1.4. In the absence of further changes in the index, aggregate demand shifts back partway; as noted above, the imposed long-run elasticity is about 0.3.

The instantaneous impact of a one-point increase (depreciation) in the index of the real exchange rate is a shift upward in aggregate demand by about 114 billion yen, yielding an implied elasticity at sample means of about 0.74. In the absence of further changes, aggregate demand again shifts back partway, with the imposed long-run elasticity again being about 0.3. That the initial impact of the depreciation is to shift aggregate demand upward is inconsistent with the *J* curve literature (e.g., Noland, 1989), but consistent with a number of studies surveyed in Goldstein and Khan (1985). On the other hand, it appears that it is unusual to estimate a long-run elasticity that is smaller than a short-run one. So these point estimates perhaps are implausible.

Table VI has variance decompositions.[8] The column labeled "foreign residual" has the contribution of $V_{xt} - E(V_{xt}|U_{ct}, U_{dt})$ (the difference between the innovation in the autoregression for ΔX_t and its projection onto the cost and demand shocks). For each horizon (between 1 and 12 quarters), the percentage of variance due to each shock is given; the total may not add to 100 because of rounding.

The decompositions settle down quite quickly, requiring not much over a year. This reflects the small size of the roots in the relevant autoregressive polynomials. About one-tenth of the movement in output is due to costs, three-fifths to demand, and one-third to foreign shocks. The last figure emphasizes the degree to which Japan's economy is open, since net exports have been less than 5% of GNP (in absolute value), and exports and imports each less than 15%. The one-third figure seems consistent with the view in Komiya and Yasui (1984) that the world economy affects Japan substantially beyond the extent that might be suggested by casual examination of the level of trade flows. On the other hand, the one-third figure is perhaps higher than would be suggested by the vector autoregressions in Horiye *et al.* (1987). This estimate and, indeed, all the estimates in Table VI are very imprecise; the 95% confidence intervals are far larger than those in my U.S. study (West, 1990c), perhaps because the sample here is only a third as long.

About three-fifths of the movement in inventories is due to costs, almost

[8] Impulse response functions for the levels of the variables (not reported in any table, to conserve space): positive demand shocks cause Q_t and Q_{ft} to increase, H_t and e_t to decrease; positive cost shocks cause Q_t to decrease, all other variables to increase. I am not sure what the explanation is for the counterintuitive increase in inventories in response to cost shocks.

TABLE VI
VARIANCE DECOMPOSITIONS

	ΔQ_t			ΔH_t		
Horizon	Cost	Demand	Foreign residual	Cost	Demand	Foreign residual
1	7.4	58.9	33.8	58.6	0.9	40.5
	(0,79.4)	(0.3,91.0)	(0.2,97.4)	(14.7,99.2)	(0.0,37.5)	(0.2,65.6)
2	8.1	58.4	33.5	57.5	1.6	40.9
	(0.9,79.4)	(0.5,89.1)	(0.2,92.2)	(16.8,96.8)	(0.1,40.4)	(0.9,64.5)
3	8.7	58.0	33.3	60.4	2.3	37.3
	(1.2,79.4)	(1.1,88.7)	(0.2,90.3)	(16.1,94.7)	(0.6,41.0)	(1.9,57.6)
4	9.0	57.7	33.3	60.4	2.8	36.8
	(1.4,79.4)	(1.3,88.4)	(0.2,89.5)	(17.6,92.7)	(1.0,40.7)	(2.8,57.4)
8	9.5	57.3	33.3	61.2	3.5	35.3
	(1.5,79.4)	(1.7,88.0)	(0.2,87.9)	(15.7,91.1)	(1.4,44.8)	(3.4,48.6)
12	9.5	57.3	33.3	61.2	3.6	35.3
	(1.5,79.4)	(1.7,88.0)	(0.2,87.8)	(15.6,91.0)	(1.4,44.9)	(3.4,48.6)

Note. The 95% confidence intervals are in parentheses, from bootstrap. Additional estimates (95% confidence interval in parentheses): $h = -0.12$ $(-1.26,.45)$; $(\sigma_c/\sigma_d) = 0.84$ (0.05,1.93); breakdown for ΔQ_{ft}, for all horizons, cost = 8.8 (0.0,40.7), demand = 1.4 (0.0,26.4), foreign residual = 89.8 (50.2,99.6); breakdown for Δe_t, for all horizons, cost = 54.3 (0.7,79.4), demand = 21.6 (5.4,89.1), foreign residual = 24.3 (2.4,63.1).

all the remainder to the foreign residual. I know of no direct or indirect evidence of the plausibility of this large effect of the foreign residual, which indicates that a very important role for inventories in Japan is to buffer output from shocks to foreign output and the real exchange rate.

For both GNP and inventories, the fraction due to costs is distinctly less than what I found for the United States in West (1990c), in which the figures for output and inventories were about one-half and over 90%, respectively. A higher figure for the United States is not, however, surprising, given that U.S. inventories move procyclically and Japanese inventories do not.

As indicated in the note to Table VI, most of the variance of ΔQ_{ft} is due to foreign shocks. Even though, as noted above, Q_{ft} is largely U.S. GNP, the small figure for costs (9%) is perfectly consistent with West's (1990c) result that about half the movement in U.S. GNP is due to cost shocks: factors that shift the U.S. aggregate supply curve need not be perfectly correlated with factors that shift the Japanese aggregate supply curve. This is also consistent with Stockman (1988), who also found a substantial nation-specific component to fluctuations in industrial production.

About three-fifths the movement in the real exchange rate is due to cost shocks, one-tenth to demand shocks, and one-third to foreign shocks. The

TABLE VII
EFFECTS OF ALTERNATIVE SPECIFICATIONS ON 8-QUARTER VARIANCE DECOMPOSITIONS

(1)	(2)	(3)	(4)	(5)			(6)			(7)			(8)		
	LR elasicity of S_t with respect to			Variance decompositions											
				ΔQ			ΔH			ΔQ_f			Δe		
Line	Q_f	e_t	g_{HS}	C	D	X	C	D	X	C	D	X	C	D	X
(1)	0.30	0.30	0	10	58	33	61	4	35	9	1	90	54	22	24
(2)	0.15	0.30	0	10	57	33	61	3	36	9	3	88	54	23	23
(3)	1.50	0.30	0	10	55	36	61	5	34	9	0	91	54	18	28
(4)	0.30	0.15	0	10	72	18	61	2	37	9	0	91	54	11	34
(5)	0.30	1.50	0	10	29	62	61	6	33	9	7	84	54	37	9
(6)	0.30	0.30	0.5	10	11	82	55	18	27	5	2	93	49	25	26

Note. Columns (2) and (3) report the approximate long-run elasticities of S_t with respect to Q_{ft} and e_t that were imposed in the estimation. In columns (5) to (8), C, D, and X denote the percentage of variance due to costs, demand, and the foreign residual. Totals may not be 100 due to rounding. Line (1) repeats for convenience the entries from Table VI.

predominance of cost shocks is consistent with the evidence for the dollar/yen and mark/yen rate in Meese and Rogoff (1988).

Table VII reports the 8-quarter variance decomposition of some alternative specifications. Line (1) repeats for convenience the result in Table V. Lines (2) to (5) vary the assumed long-run elasticity of S_t with respect to Q_{ft} and e_t in accord with the range of parameter estimates reported in part IIIB. As may be seen, these have little effect on the decomposition.

Line (6) generalizes the cost function of line (2) to

$$C_t = g_{0Q}Q_t^2 + g_{0H}(H_t - g_{HS}S_{t+1})^2 + 2U_{ct}(hH_t + Q_t),$$

and considers the effect of setting g_{HS} to 0.5. The additional term, $g_{HS}S_{t+1}$, a target level for inventories, may be rationalized as necessary to adequately capture stockout costs (West, 1986). U.S. studies usually yield a value less than 1; the aggregate data used in West (1990c) in fact delivered a (insignificantly) negative point estimate.

I imposed $g_{HS} = 0$ here because this was suggested by likelihood ratio tests, which, after casual inspection, suggested that the likelihood declines as g_{HS} increases from 0; the decline becomes asymptotically significant around $g_{HS} = 0.5$. The figures in line (6) of Table VII indicate that the

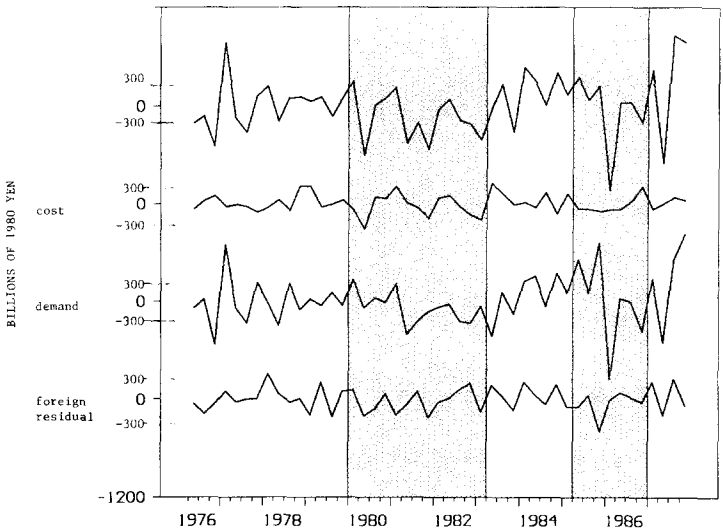

FIG. 3. Components of GNP forecast error.

results are somewhat sensitive to the value of g_{HS} (a result also found in U.S. data (West, 1990b)). I do not believe there is any particular message to be drawn from this, other than to underscore the fact that the confidence intervals around the point estimates in Table VI are large.

Figure 3 follows Blanchard and Watson (1986) in plotting $\Delta Q_t - E_{t-8}\Delta Q_t$, the error in forecasting changes in GNP 8 quarters ahead, and breaking down this forecast error into three orthogonal components. The first 13 quarters of the sample have been lost to initial conditions and to the need to forecast 8 quarters ahead. The contraction phase of the business cycle is again marked by shading. That demand shocks are the most important souce of output fluctuations is evident from the graph. It can be seen that even though the shocks are uncorrelated by construction, they tend to move together at cyclical turning points. As has been found in many studies of the U.S. economy, then, no single shock tends to dominate at business cycle turning points.

The eye is drawn to the downturn of 1986 : 1, which the model attributes mainly to demand shocks. I believe that this is consistent with conventional wisdom: An anonymous referee has informed me that the 1986 *Annual Report of the Economic Planning Agency* gives the proximate cause of this downturn as the 1985 : 4 appreciation of the yen, which in turn was a deliberate result of the coordinated demand management that followed the Plaza Accord of September 1985. In terms of the present model, this argument gives the ultimate source of the downturn as coordi-

nated shocks to foreign and Japanese demand; since the 1985 : 4 foreign and Japanese demand shocks had a large common component, the decomposition attributes most of the downturn to a movement in Japanese demand (see the discussion at the end of Part II).

IV. Conclusion

During 1975–1987, cost shocks are estimated to account for about one-tenth of the movement in Japanese GNP and one-half of the movement in inventories. Most of the remaining movement in GNP is due to demand shocks, and in inventories to shocks to the output of some of Japan's major trading partners and to the real exchange rate. While the point estimate of the importance of cost shocks for GNP is, at least superficially, inconsistent with simple real business cycle theory that holds that demand shocks are of secondary importance in explaining business cycles, the confidence intervals around this and other estimates are so large that it is difficult to discriminate sharply between simple real business cycle and other theories.

A priority for future research is to sharpen the estimates to allow better discrimination. One possibility is to extend the number of data points, to allow more business cycle variation, by lengthening the sample or by using monthly data. Another is to use economic theory to model the determinants of foreign output and the real exchange rate, which in this paper are modeled simply as an unrestricted vector autoregression. Finally, an important and challenging extension is to allow for nominal shocks and nominal rigidities.

Appendix

Let β_0 be a normalization, $\beta_0 \equiv g_{0S} + g_{0Q}$. Then the matrices in (5) that have yet to be defined are

$$A_0 = \begin{vmatrix} \beta_2 & \beta_1 \\ \beta_1 & 1 \end{vmatrix}, A_1 = \begin{vmatrix} \beta_1 & 0 \\ -\beta_1 & 0 \end{vmatrix},$$
$$D_0 = \beta_0^{-1} \begin{vmatrix} h & 1 \\ 1 & -1 \end{vmatrix}, D_1 = \beta_0^{-1} \begin{vmatrix} 0 & -1 \\ 0 & 0 \end{vmatrix}, \tag{A1}$$

where $\beta_1 \equiv -\beta_0^{-1} g_{0S}$ and $\beta_2 \equiv \beta_0^{-1}[(1 + b)g_{0S} + g_{0H}]$.

To solve for the cross-equations restrictions: Guess a solution for Y_t of the form

$$Y_t = \Pi Y_{t-1} + \tilde{\Gamma}_0 X_t + \tilde{\Gamma}_1 X_{t-1} + \cdots + \tilde{\Gamma}_n X_{t-n} + FU_t, \qquad \text{(A2)}$$

where the 2×2 matrices Π, F, $\tilde{\Gamma}_1$,and $\tilde{\Gamma}_n$ are functions of the underlying structural parameters. Define

$$\underset{2(n+1)\times 2}{\tilde{\Gamma}} \equiv (\tilde{\Gamma}_f | \tilde{\Gamma}_e) \equiv (\tilde{\Gamma}_0', \ldots, \tilde{\Gamma}_n')',$$

$$\underset{(n+1)\times 2}{G} \equiv (G_f | G_e) \equiv (G_0, \ldots, G_n)',$$

where $\tilde{\Gamma}_f$ and $\tilde{\Gamma}_e$ are $2(n + 1)$ column vectors of coefficients on lags of Q_{ft} and e_t in (A2); G_f and G_e are the coefficients on Q_f and e in the demand curve (1). Substitute (A2) and (A2) led one time period into (5), using $X_t = (\mathrm{I} + \Phi_1)X_{t-1} + \cdots + (-\Phi_n)X_{t-n-1} + V_{xt}$. Since the right-hand side of (5) is zero, the coefficients on each variable in the resulting equation must be zero. This implies

$$bA_1'\Pi^2 + A_0\Pi + A_1 = 0, \qquad \text{(A3a)}$$

$$[bA_1'(I + \Pi) + A_0]F = D_0 + bD_1, \qquad \text{(A3b)}$$

$$P_0(\tilde{\Gamma}_f | \tilde{\Gamma}_e) + P_1 = P_2(G_f | G_e) + P_3, \qquad \text{(A3c)}$$

where P_0 is $2(n + 1) \times 2(n + 1)$, with $bA_1'\Pi + A_0$ on its 2×2 diagonal blocks, bA_1' on the 2×2 band immediately above the diagonal, zeroes elsewhere; P_1 is $2(n + 1) \times 2$, $P_1 \equiv ([(bA_1'\tilde{\Gamma}_0(I + \Phi_1)]', \ldots, [bA_1'\tilde{\Gamma}_0(-\Phi_n)]')'$; P_2 is $2(n + 1) \times (n + 1)$, with $-B_0(1)$ in its $(2i - 1, i)$ entries, $-B_0(2)$ in its $(2i, i)$ entries $(i = 1, \ldots, n + 1)$, $-bB_1(1)$ in its $(2i - 3, i)$ entries, $-bB1(2)$ in its $(2i - 2, i)$ entries $(i = 2, \ldots, n + 1)$, zeroes elsewhere; $P_3 = ([-bB_1G_0'(I + \Phi_1)]', \ldots, [-bB_1G_0'(-\Phi_n)]')'$.

Equation (A3a) just identifies the two unknowns in A_0 and A_1. Equation (A3c) embodies $4(n + 1)$ equations that depend on only $2(n + 1)$ unknowns (the elements of G_f and G_e), apart from those parameters identified from (A3a); the aim is to impose the overidentifying restrictions. Now, since U_t is a random walk, and $E\Delta X_t U_t \neq 0$, the reduced form equation that was actually estimated was obtained from (A2) by differencing and substituting out for ΔX_t: $\Delta Y_t = \Pi\Delta Y_{t-1} + \Gamma_1\Delta X_{t-1} + \cdots + \Gamma_n\Delta X_{t-n} + V_{yt}$, where

$$\Gamma_1 = \tilde{\Gamma}_0\Phi_1 + \tilde{\Gamma}_1, \ldots, \Gamma_n = \tilde{\Gamma}_0\Phi_n + \tilde{\Gamma}_n, V_{yt} = FU_t + \tilde{\Gamma}_0 V_{xt}. \qquad \text{(A5)}$$

Define the $2n \times 2$ matrix $\Gamma \equiv (\Gamma_f | \Gamma_e) \equiv (\Gamma_1', \ldots, \Gamma_n')'$ where Γ_f and Γ_e are

each $n \times 1$ column vectors. Note that (A5) implies that for certain $2n \times 2(n+1)$ matrices T_f and T_e that depend on the elements of $\Phi_1, \ldots, \Phi_n$, $T_f\tilde{\Gamma}_f = \Gamma_f$, $T_e\tilde{\Gamma}_e = \Gamma_e$. This can be used to transform (A3c) to

$$\Gamma_f = P_{3f}G_f, \tag{A6a}$$

$$\Gamma_e = P_{3e}G_e, \tag{A6b}$$

where the $2n \times (n+1)$ matrices P_{3f} and P_{3e} depend on $b, A_0, A_1, \Pi, \Phi_1, \ldots, \Phi_n$. P_{3f} and P_{3e} are of rank n rather than rank $n+1$: the $4n$ equations in (A6a) and (A6b) overidentify some linear combinations of the $2(n+1)$ elements of G_f and G_e but do not identify all the individual coefficients. (Consideration of the order condition necessary to identify the entire set of elements of G_f and G_e from the Euler equation (5) suggests that the deficiency results because there are fewer lags of ΔX_t in its AR(p) than lags of X_t appearing in the demand curve (1).) As noted in the text, to identify the individual elements of G_f and G_e, values for $\Sigma_i g_{if}$ and $\Sigma_i g_{ie}$ were imposed. One can then substitute out for one of the elements of each vector—say, g_{0f} and g_{0e}—in terms of the known values $\Sigma_i g_{if}$ and $\Sigma_i g_{ie}$ and the $2n$ unknowns $g_{1f}, \ldots, g_{nf}, g_{1e}, \ldots, g_{ne}$:

$$\Gamma_f = P_{5f}G_f^* + P_{6f}, \qquad G_f^* \equiv (g_{1f}, \ldots, g_{nf})', \tag{A7a}$$

$$\Gamma_e = P_{5e}G_e^* + P_{6e}, \qquad G_e^* \equiv (g_{1e}, \ldots, g_{ne})', \tag{A7b}$$

where P_{5f} and P_{5e} are each $2n \times n$, P_{6f} and P_{6e} are each $2n \times 1$. This implies a set of n restrictions on $\Gamma \equiv (\Gamma_f | \Gamma_e)$, where the restrictions are linear: with the restricted values of $\Pi, \Phi_1, \ldots, \Phi_n$ held fixed at their unrestricted estimates, P_{5f}, P_{6f}, P_{5e}, and P_{6e} are matrices with known values.

Once the restricted estimates are available a restricted $\tilde{\Gamma}_0$ can be computed from (A7) and (A5). Since $F(EU_tU_t')F' = E(V_{yt} - \tilde{\Gamma}_0 V_{xt})(V_{yt} - \tilde{\Gamma}_0 V_{xt})'$ (see Eq. (A5)), F and then h, σ_c^2 and σ_d^2 can then be recovered from (A3b) as described in West (1990c).

To compute impulse response functions and variance decompositions: Write the projection of V_{xt} on U_t as PU_t, where an estimate of P is obtained by OLS. So $V_{xt} = PU_t + V_{xt}^*$, where $EV_{xt}^*U_t' = 0$ but the two elements of V_{xt}^* are in general correlated with one another. Write the 4×1 vector of reduced form residuals as $V_t = (V_{yt}' V_{xt}')'$ as

$$V_t = \begin{vmatrix} \tilde{\Gamma}_0 P + F & \tilde{\Gamma}_0 \\ P & I_2 \end{vmatrix} \begin{vmatrix} U_t \\ V_{xt}^* \end{vmatrix},$$

from which impulse response functions and variances decompositions can be calculated in the usual fashion.

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