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EVIDENCE FROM SEVEN COUNTRIES ON WHETHER INVENTORIES SMOOTH AGGREGATE OUTPUT

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ABSTRACT

Casual examination of annual postwar data on inventories and aggregate output for seven developed countries – Canada, France, West Germany, Italy, Japan, United Kingdom, United States – suggests that in these countries the primary function of aggregate inventories is not to smooth aggregate output in the face of aggregate demand shocks. Japan is a possible exception to this generalization.

1. INTRODUCTION

A number of recent papers have considered whether inventories smooth output fluctuations in the United States (e.g., Blinder [1]). There does not, however, seem to have been much research on this question, for other countries.* This paper is a preliminary attempt to help fill this gap.

It uses annual postwar data from the seven 'G7' countries – Canada, France, West Germany, Italy, Japan, the United Kingdom and the United States – to see whether aggregate inventories serve mainly to buffer aggregate output from demand shocks. For all these countries, except possibly Japan, the answer seems to be no, in two senses. First, aggregate output (measured by either GDP or GNP) is more variable than aggregate final sales. Second, positive sales shocks tend to make inventories increase, with output rising more than one to one with such shocks. As is well known for the U.S., then, aggregate inventory behavior does not seem to be consistent with the production smoothing model of Holt et al. [3].

It should be emphasized that this is a first, preliminary effort in what is likely to be a somewhat larger research project. I have yet to obtain quarterly or monthly data, for example, which probably are more appropriate for studying production smoothing (although annual data will still reflect the low frequency fluctuations implied by the model given below). I also have yet to obtain a figure for the level (as opposed to change) in inventories; this precludes computation of such elementary statistics as a mean inventory-sales ratio. In

*In a paper the author became aware of only after drafting the present paper, Wilkinson [2] touches on whether inventories smooth production, for exactly the countries considered in this paper. Wilkinson's main focus, however, is estimation of a general model of inventory demand for a subset of these countries.

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addition, my approach is casual, and no standard errors have been calculated.

2. MODEL AND TESTS

Let Q_t be real aggregate output, S_t real aggregate final sales, H_t real aggregate inventories. The variables are linked by the identity $Q_t = S_t + \Delta H_t$.

Much recent U.S. research on inventories has assumed a variant of the Holt et al. [3] production smoothing model. The representative firm minimizes the expected present discounted value of costs over an infinite horizon, with a constant discount rate. In a general version of this model, per period costs are

$$a_0(\Delta Q_t + u_{1t})^2 + a_1(Q_t + u_{2t})^2 + a_2(H_t - a_3 E_t S_{t+1} + u_{3t})^2 \tag{1}$$

where the a_i are positive parameters, the u_{it} are zero mean i.i.d. cost shocks and E_t denotes mathematical expectations conditional on period t information. The three terms in (1) capture costs of changing production, costs of production and costs of having inventories deviate from a target level. See West [4] for further discussion. Constant and linear terms are allowed in the empirical work but are omitted from (1) for simplicity.

I will consider two implications that follow when the model is specialized, as in Blinder [5] or Belsley [6], so that inventories serve mainly to buffer output from demand fluctuations. This requires that $a_2 a_3$ be small relative to a_0 and a_1, and that the cost shocks u_{it} have minor effects (e.g., because the standard deviation of cost shocks is small relative to that of demand shocks). The first of my two tests looks at some sample moments. The specialized model suggests that production Q_t should be smoother than demand S_t; inventories will be adjusted to avoid the costs that result when the level or change of production is varied. Let 'var' denote variance, 'corr' correlation. The model then suggests

$$\mathrm{var}(Q_t)/\mathrm{var}(S_t) < 1, \tag{2a}$$

$$\mathrm{var}(\Delta Q_t)/\mathrm{var}(\Delta S_t) < 1, \tag{2b}$$

$$\mathrm{corr}(S_t, \Delta H_t) < 0. \tag{2c}$$

See West [3] for a formal argument for the first two inequalities. The last inequality follows from the first, since

$$\mathrm{var}(Q_t) = \mathrm{var}(S_t) + \mathrm{var}(\Delta H_t) + 2\mathrm{cov}(S_t, \Delta H_t).$$

We look at (2c) separately because it focuses on the elementary production smoothing notion that inventory investment ΔH_t should be countercyclical.

The first inequality does not make sense if variables have unit roots. One can, however, calculate an analogue to $\mathrm{var}(Q_t) - \mathrm{var}(S_t)$ that has a meaningful population counterpart, even in the presence of unit roots. See below.

A second test of the model looks at how inventories respond to sales shocks. If the cost of having inventories deviate from a target level is small ($a_2 a_3$ is small relative to a_0 and a_1), inventories should be drawn down when there is a positive sales shock. (This does not hold under all circumstances. See Blinder [7].) One admittedly crude way to check this is to suppose that only lagged sales are used to forecast future sales. Suppose that the sales process follows an autoregression,

$$S_t = f_1 S_{t-1} + f_2 S_{t-2} + \cdots + f_q S_{t-q} + v_t, \tag{3}$$

where the f_i are parameters and v_t is the zero mean i.i.d. sales shock. Constant and trend terms, included in the empirical work, are omitted for simplicity. The lag polynomial $(1 - f_1 L - \cdots - f_q L^q)$ has roots on or outside the unit circle, with a root on the unit circle implying that differencing is required to induce stationarity.

By algebra such as in Blanchard [8], the decision rule for inventories is

$$H_t = r_1 H_{t-1} + r_2 H_{t-2} + d_1 S_t + \cdots + d_1 S_{t-q+1} + u_t, \tag{4}$$

where constant and trend terms have again been suppressed. The r_i and d_i are functions of the cost parameters a_i, the sales parameters f_i and the rate for discounting future costs. The disturbance u_t is a linear combination of the u_{it}.

Suppose, finally, that cost and demand shocks (u_t and v_t) are uncorrelated. If S_t and H_t are stationary (the lag polynomial in (3) does not have a root on the unit circle) one can estimate (3) and (4) by OLS. If S_t and H_t have unit roots, it is more efficient to impose the unit root in (3). In either case, one can then use the estimates to trace out the impact of a sales shock on inventories: $\partial H_t/\partial v_t = d_1$, etc.

3. RESULTS

3.1 Data

Annual data on nominal and real GNP or GDP and on nominal change in inventories was taken from the *International Financial Statistics* (IFS) tape of the International Monetary Fund. Annual rather than quarterly data were used in part because they seem likely to be more reliable; figures on inventory investment in Germany, for example, are benchmarked against data on inventory levels only annually, with preliminary quarterly figures simply computed as a residual (OECD [9,p.13]). The definition of inventory investment, incidentally, does not appear to be identical in all countries, since there seems to be some variation in the treatment of certain stocks held by the government (OECD [10,11,9]).

Data were available 1957–1986 for six of the countries, 1961–1986 for Canada. Aggregate output Q was measured by GNP when this was available (Germany, Japan, United Kingdom, United States), GDP otherwise (Canada, France, Italy). For all countries, the base year for the real data is 1980 and all data are expressed in billions of units of home currency.

A deflator was calculated by dividing nominal by real output. Real inventory investment ΔH_t was calculated by dividing the nominal IFS figure by the deflator. Real final sales S_t was then computed as $S_t \equiv Q_t - \Delta H_t$. A real inventory series H_t was created by accumulating the changes in real inventories: $H_1 = \Delta H_1$, $H_2 = H_1 + \Delta H_2$, etc. (The IFS tape does not seem to supply a figure for the level of inventories.) All such manufactured values of H_t are of course too low by a constant value of H_0, the presample value of the inventory stock. Note that the series being off by a constant will affect only the constant term in regressions, and will leave estimates of, for example, variances and correlations unchanged.

The procedure for computing a real series for H_t and ΔH_t is nonetheless unsatisfactory in that it uses the output deflator to convert the inventory data. In the U.S., at least, a more subtle and complicated procedure is employed by the Department of Commerce in constructing constant dollar inventory series (Hinrichs and Eckman [12]). To get an idea of how substantial the biases induced by the deflation procedure are, we compared the deflated IFS data for the United States to the constant dollar Department of Commerce data, with the latter obtained from Citibase.

The results are in Table 1, with notes at the foot of the table describing the procedure used. Since the Department of Commerce is the source for the IFS data, the correlation between the two real GNP series is virtually perfect (Table 1, panel A). (See below for the qualifier 'virtually'.) The differing deflation procedures led to only slight discrepancies between the two sets of inventory and sales figures, with correlations of about 0.99, in levels or differences (panel A). In addition, the correlation of moments within each data set are very close. Compare panels B and C. (Note that the figures for Q and ΔQ are not identical, for the two data sets. We believe that the minor discrepancies resulted because of errors introduced when we converted the Department of Commerce data from its 1982 base year to the 1980 base year that IFS uses.)

It seems from Table 1, then, that the use of

TABLE 1

	Q	S	H		ΔQ	ΔS	ΔH
(A) *Correlations between deflated IFS and department of commerce data*							
	1.000	0.9993	0.9878		1.0000	0.9985	0.9931
(B) *Correlations within deflated IFS data*							
Q	2758.5	0.96838	0.81880	ΔQ	2942.1	0.93789	0.77325
S	2340.1	2116.9	0.82621	ΔS	2134.1	1759.8	0.65576
H	649.68	574.29	228.23	ΔH	619.64	406.41	218.26
(C) *Correlations within deflated department of commerce data*							
Q	2759.2	0.97086	0.78475	ΔQ	2943.0	0.94018	0.74336
S	2345.1	2114.7	0.79109	ΔS	2156.4	1787.6	0.61899
H	710.49	627.02	297.08	ΔH	581.63	377.46	208.02

Notes:
1. Annual data, 1957–1986.
2. Moments for Q, S and H calculated around a constant and time trend, for ΔQ, ΔS and ΔH around a constant. For each, a shift in these deterministic terms was allowed in 1974.
3.In panels B and C, variances and covariances are on and below the diagonal, correlations are above the diagonal.

an output deflator to deflate nominal data on inventory investment introduces only very slight errors. We will therefore proceed on the tentative assumption that the use of data deflated in this way is unlikely to introduce serious biases.

3.2 Empirical results

Columns (2) to (4) of Table 2 report inequalities (2a) to (2c). Column (5) reports essentially a measure of $\text{var}(Q) - \text{var}(S)$ that is legitimate in the presence of unit roots; inequality (2a) indicates that this difference should be positive. Column (5) was calculated as described in West [13], using five lags of ΔS_t. Column (6) is presented to scale the column (5) figure. With the possible exception of Japan, the well-known U.S. experience is typical – aggregate output is about 15 to 100% more variable than final sales (columns (1) and (2)). In Japan, however, output is not even 10% more variable. Column (5) indicates that the column (2) result is not a spurious result

TABLE 2

Relative variability of output and final sales

(1) Country	(2) $\text{var}(Q)/\text{var}(S)$	(3) $\text{var}(\Delta Q)/\text{var}(\Delta S)$	(4) $\text{corr}(S,\Delta H)$	(5) $E(Q^2-S^2)$	(6) $\text{var}(\Delta Q)$
Canada	1.55	2.20	0.18	−24.2	29.4
France	1.15	1.48	0.07	−1751.7	870.2
W. Germany	1.39	1.76	0.22	−151.8	664.0
Italy	1.64	2.53	0.15	-13.3×10^6	4.5×10^6
Japan	1.02	1.07	−0.01	-12.2×10^6	16.7×10^6
U.K.	1.75	1.94	0.37	-6.9×10^6	17.0
U.S.	1.30	1.67	0.33	−713.6	3047.0

Notes:
1. See notes to Table 1.
2. For columns (5) and (6), units are billions of real 1980 units of home currency, squared.
3. As explained in the text, column (5) essentially calculates $\text{var}(Q) - \text{var}(S)$ in a fashion that is robust to the presence of unit roots. Column (6) is presented solely for comparison to column (5).

of inappropriate treatment of unit roots – output is more variable than sales even when unit roots are explicitly allowed. In all countries but Japan, inventory investment is procyclical (column (3)).

Table 3 contains the impulse response functions of inventories to a positive sales shock, of magnitude one 1980 unit of home currency (e.g., one 1980 French franc for France). Panel A presents the results when (3) and (4) were estimated in levels, panel B when a unit root was imposed in (3). The lag length q was set to 2; the Q statistic in all of the regressions suggested that this sufficed to whiten the residuals. Deterministic terms were included as described in note to Table 1. Detailed regression results are in an appendix available from the author on request.

To read the table, consider the entry for Canada in panel A. If sales unexpectedly rise by one Canadian dollar, inventories initially rise by 42 Canadian cents. The next year they rise by an additional 9 cents ($9=52-41$), before beginning to fall back toward their trend line. Although for panel B equation (3) was estimated in differences, the figures in panel B apply to the level and not the difference of inventories.

A positive sales shock initially causes inventories to rise; with the exception of Japan, in differences, all entries in year 0 are positive. In differenced specifications, the year 5 figure suggests that a positive sales shock also causes a rise in the steady state level of inventories, again with the exception of Japan.

TABLE 3

Inventory response to one unit sales shock

Country	Year					
	0	1	2	3	4	5
(A) *Regression estimates in levels*						
Canada	0.42	0.51	0.37	0.18	0.05	−0.01
France	0.16	0.56	0.61	0.60	0.60	0.62
W. Germany	0.31	0.26	0.03	−0.13	−0.14	−0.06
Italy	0.28	0.22	0.12	−0.04	−0.08	−0.06
Japan	0.08	0.17	0.16	0.11	0.08	0.06
U.K.	0.20	0.26	0.16	0.05	−0.01	−0.02
U.S.	0.25	0.27	0.15	−0.02	−0.12	−0.10
(B) *Regression estimates in differences*						
Canada	0.39	0.41	0.34	0.27	0.21	0.17
France	0.26	0.42	0.49	0.52	0.54	0.54
W. Germany	0.29	0.33	0.32	0.31	0.29	0.27
Italy	0.28	0.24	0.31	0.31	0.33	0.34
Japan	−0.09	−0.12	−0.12	−0.11	−0.10	−0.08
U.K.	0.17	0.20	0.21	0.21	0.20	0.20
U.S.	0.21	0.31	0.35	0.37	0.38	0.38

Notes:
1. See notes to Table 1.

4. CONCLUSION

Casual examination of annual postwar data suggests that in the 'G7' group of countries, aggregate inventories do not serve mainly to smooth output fluctuations in the face of aggregate demand shocks. Japan provides a possible exception, although even in Japan production smoothing behavior, if present, is not particularly marked. That inventory behavior is qualitatively similar in these countries is consistent with the work of Moore [14], which gives the level and change in inventories the same position in the NBER reference cycle in each of the seven countries.

A simple extension of this work is to consider quarterly data as well, at least in those countries where the quarterly data are reasonably reliable. The work of Wilkinson [2] suggests that quarterly results are likely to be broadly similar, although it also suggests that at quarterly frequencies Japanese inventory behavior is not qualitatively different from that of the other countries. More generally, desireable areas for future research include considering the role of inventories in business cycles in the light of international differences in tax systems, in the degree to which various economies are open, and in the sources of business cycle shocks.

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