LONG-RUN EQUILIBRIUM IN THE EMPIRICAL STUDY OF MONOPOLY AND COMPETITION

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A long-run tendency of industry profit rates to converge to a single competitive level has been a fundamental tenet of the industrial organization approach to the study of competitiveness in a market economy. This paper shows that for the post World War II period a weak equalization can be econometrically identified with different reaction speeds by industry. However, persistent profit rate differences endure. Finally, a portfolio theory of risk is considered as an explanation of these differentials.

I. INTRODUCTION

One of the central propositions of economic theory, formulated by (among others) Hirshleifer [1976, 319] and Malinvaud [1972, 239], is that industry rates of return should equalize in the long run. Although the actual observed rates of profit at any point in time may differ between industries, according to this proposition, such differences should not persist.[1] According to Stigler [1963, 55],

> Economic analysis tells us that the rates of return in competitive industries will be strictly equal...in long-run equilibrium.... But this very concept of long-run equilibrium reminds us that, in a world where all events are not perfectly anticipated, there will be a stream of unexpected disturbances that call for a stream of changes in the allocation of resources: unanticipated shifts in consumers' desires; the impact upon international markets of wars and political events; the irregular march of major advances in technology, and others.

However, rates of return in long-run equilibrium are not directly observable; and it is further unclear whether they can be proxied by averages, and, if so, over what time periods. Bain [1951, 309] recognized the problem.

> True profits...will seldom in individual cases exactly average through time to a long-run equilibrium level, but will be influenced by lagging response to, and faulty anticipation of, a never-ending succession of dynamic changes.

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1. This paper restricts its analysis to the examination of industry rates of return. It can be argued that these averages should be equal even if there are differences between establishments in each industry. Intraindustry differences in profitability can arise from a variety of reasons and can coexist with equal average industry rates of return when entry and exist is unimpeded (see Mueller [1986, 9]). This study considers only the effect of market forces on the interindustry equalization process and does not consider the structure of individual markets.

Economic Inquiry
Vol. XXVIII, January 1990, 151–162

Because of this ambiguity, there has been debate regarding the proper interpretation of empirical studies of rates of return and market structure. Numerous studies, such as those by Bain [1951], Mann [1966], Stigler [1963], Sherman [1968], Weiss [1963], and Fuchs [1961], have found a positive correlation of rates of return with concentration ratios and barriers to entry. However, because these studies were cross sectional, the concept of long-run equilibrium was either ignored or proxied by short-term industry averages. Bain, for example, considered a selection of four-digit industries for the years 1936–40. The four-year average was a proxy for long-run equilibrium, in the sense that "persistence" was defined as above average levels sustained over four years. He found a weak positive relationship between concentration and these proxies. Stigler used three-digit industries and chose ten-year profit averages rather than four-year proxies. He also reported a weak relationship between profit rate and concentration. Likewise, Mann related ten-year average industry rates of profit to concentration and barriers to entry and discovered a weak positive relationship.

Brozen [1970; 1971] criticized these studies, arguing that if they truly revealed that monopolistic structures caused above average profit rates to persist, then these industries should be able to sustain high profit rates for longer than four to ten years. When he examined the same industries in a later period he found that the high rates of profit tended to decline while the low profit rates tended to increase as time passed. He concluded that Bain and others had modeled long-run equilibrium inadequately, because the time periods considered in those studies were too short. Brozen also scrutinized the selection methods by which industries were included in the sample, and claimed that the samples were biased toward overstating the effects of concentration.

While Brozen argued that the positive relationships observed in prior studies were due to either sample bias or to disequilibrium, he never undertook a study which might have helped settle the issue of whether profit rate differentials were persistent. Even after Brozen, studies published in the industrial organization literature were still cross sectional, and the observed relationship between industry profit rates and market structure varied depending on the particular year chosen.[2]

It is an interesting anomaly in the history of economic thought that Brozen's criticism has received "surprisingly little attention" (according to Mueller [1986]) in the decade following the publication of his work. Mueller attributes this in part to the continued focus on the profits-concentration relationship, which can be best addressed by the analysis of highly disaggregated data sets and sophisticated cross sectional analysis (e.g., the line of

2. Compare Gambles [1970, 68]. A notable exception is Mueller's [1986] study of firm rates of profit. Mueller developed a model which in spirit was similar to this one (published while the present paper was under review) but which relied on a different statistical specification.

business studies at the Federal Trade Commission). The explanation in this paper is more simple. Prior to the publication of the industrial capital stock series by the Department of Commerce in the early 1980s there were simply no reliable data on long-run profit rates.[3]

The present study is devoted to the question of whether long-run industrial rates of return equalize. Knowledge of the long-run behavior of profits rates, as many economists have pointed out, is analytically prior to the issue of the relationship between industrial profit rate differentials and market structure or other explanatory variables—especially if the short-run behavior of these relations is studied. Knowledge of the dynamics of this equalization process (and a clear characterization of the rate of profit driving it) can help build a well-specified explanatory model in the future.

The procedure for the examination of the issue of persistence incorporates the following principles.

1. Using panel data, the cross-sectional interdependence and the time series development of industry rates of return are estimated simultaneously.
2. Since the effects of market structures are not specified as explanatory variables, they are incorporated into the statistical disturbances. Therefore, a robust estimation method must be employed, one which allows the residuals to be autoregressive or heteroscedastic.
3. The equilibrium level is not proxied by an average, but is taken as the apparent target of a short-run dynamic adjustment process.

II. THE ECONOMETRIC MODEL

We model the rate of profit, r_{tj}, in industry j in year t, as the sum of three components:

$$r_{tj} = \alpha_t + \beta_j + u_{tj}\,; \quad t = 1, \ldots, T; \quad j = 1, \ldots, J. \tag{1}$$

In this expression,

α_t is the general rate of profit prevailing in year t. Its inclusion takes common cyclical and trend components out of the rates of return.

β_j is an industry-specific component which differs from industry to industry but remains stable over time.[4]

3. Prior to the 1980s, industry data were constructed from samples of firms classified by major product line. It was dangerous, therefore, to compare anything but a few years close in time, due to changing distributions of resources between product lines and changing industry definitions. The data used here avoid this problem because they are constructed on an establishment basis. Mueller has likewise avoided this issue by considering only the long-run behavior of firm-level rates of return.

4. In equation (1), α_t and β_j are identified only up to a joint additive term: adding a constant to all α_t and subtracting it from all β_j gives an equally good fit. But only one such choice admits the given interpretation of α_t and β_j. By normalizing the industry specific components β_j in such a way that the weighted mean (weighted by the capital stock) of all industry specific profit rate components is zero, the α_t become the weighted averages of industrial target rates of return.

u_{tj} is a disturbance term that reflects short-run influences on the rates of return. Their temporary nature is reflected in the assumption $E(u_{tj}) = 0$ and the assumption that their effects are, due to short-run market mechanisms, reduced geometrically over time; i.e.,

$$u_{tj} = \rho_j u_{t-1,j} + v_{tj}. \tag{2}$$

A small "reaction coefficient" ρ_j means that short-term variations in the rates of return of industry j decay quickly.[5]

It is usually assumed in this kind of model that the v_{tj}, the parts of the disturbances representing new random shocks, are homoscedastic and uncorrelated over time. Such an assumption is not appropriate here. Noncompetitive conduct and many other disturbances of the equalization of profits last for more than one year. Using as estimates of α_t, β_j, and ρ_j those values which allow one to explain the actual time path of the profit rates with the smallest random shocks means that any assumption concerning the variances and covariances of the v_{tj}, can be dispensed with.[6] "Smallest" is defined to mean the minimum of the sum of the squared estimated shocks, weighted by the capital stock of the industry affected. Due to this weighting, disturbances which affect larger industries are considered to be "larger."[7] Choosing the common treatment of the initial shock $v_{1j} := u_{1j}\sqrt{1-p_j^2}$, the estimates $\hat{\alpha}_t$, $\hat{\beta}_j$, and $\hat{\rho}_j$ minimize the following weighted sum of squares:

$$\frac{1}{T}\sum_{j=1}^{J} w_j \left(\sum_{t=1}^{T} \hat{v}_{tj}^2\right) =$$

$$\frac{1}{T}\sum_{j=1}^{J} w_j \left\{(1-\hat{\rho}_j^2)(r_{1j} - \hat{\alpha}_1 - \hat{\beta}_j)^2 + \sum_{t=2}^{T} [r_{tj} - \hat{\alpha}_t - \hat{\beta}_j - \hat{\rho}_j(r_{t-1,j} - \hat{\alpha}_{t-1} - \hat{\beta}_j)]^2\right\}, \tag{3}$$

where w_j is a weight proportional to the capital stock in industry j.

5. Besides the AR(1) process of equation (2), AR(2) and MA(1,1) processes were also tried. There was no improvement in fit, and the shocks still appeared to be autocorrelated. There are too few observations for a more precise estimation of the error structure.

6. The choice of the smallest shocks can be justified by the principle of parsimony: if there is a simple and a complicated explanation of the same phenomenon, give more credence to the simple one. This principle is commonly used in econometric model building, and implicitly also in estimation whenever a least squares principle is applied.

7. This weighting procedure also ensures that the estimation results remain unchanged if one combines two industry groupings with identical rates of return.

III. DATA

The data examined below are derived from "GNP and Components (14) by Industry, 1948–79," a National Income and Product Accounts data base available from the Bureau of Economic Analysis (BEA), and from the Bureau of Labor Statistics' Bulletin 2034, "Capital Stock Estimates for Input-Output Industries: Methods and Data," revised by Ken Rogers of the BEA staff.

The National Income and Product Account (NIPA) industry-level data, only recently available, offer a number of advantages over the industry data used in earlier studies. First, the data are compiled on an establishment basis for entire industries rather than extrapolated from a small sample of large firms, which avoids the problem of sample bias. Second, the unincorporated sector of each industry is included (see Bosworth [1982, 36–37]).[8] Finally, the data permit tangible assets to be valued at replacement cost rather than at book value (see Stigler [1963, 61]).[9]

The measure of the rate of return used in this paper is:

$$\frac{\textit{profit}}{\textit{gross fixed capital} + \textit{inventories}},$$

where profit is defined as GNP by industry less (wages and salaries + supplements + noncorporate wage equivalent + indirect taxes).[10] This concept of profit measures net income gross of corporate taxes. Gross fixed capital and inventories are measured at replacement cost.[11]

8. A wage equivalent for the noncorporate sector was obtained as follows. Total wages were divided by total employment, to get the average compensation per worker. The average compensation per worker was then multiplied by the number of self employed to obtain the estimated noncorporate wage. This procedure is widely used and was recommended by John Gorman of the BEA. It overstates the wage bill since unincorporated businesses typically pay lower wages; but in manufacturing, the unincorporated sector is small, and does not significantly affect our results.

9. In studies of profitability and market structure it is desirable to attempt to approximate an economically meaningful definition of "industry" as closely as possible. Most studies are careful to use three-digit and four-digit levels of aggregation. Unfortunately, only a two-digit definition is available from the BEA for the large number of years necessary to measure persistence. NIPA industry data are available at the two-digit level back to 1948.

10. The definition of the rate of profit examined in this paper is widely employed in previous literature. In addition, in two earlier works, Glick [1985] and Glick and Ehrbar [1988] considered the sensitivity of profit rate dispersion to the definition of the rate of profit; see also Gale and Branch [1982]. The results reported below are not affected (within a broad range) by the specific definition of the profit rate.

11. Fisher and McGowan [1983] argue that accounting data cannot be used to reflect accurately economic variables. However, they only consider the case of an individual firm. This paper contends that important insights can be gained from analysis of accounting data. It is doubtful whether the Fisher and McGowan analysis of economic and accounting rates of return can be extended to industry aggregates considered over thirty years. In a recent paper, Dumenil and Levy [1986] have shown that the economic rate of return and the average accounting rate of return for the aggregate U.S. economy display an identical trend. Their result adds support to the assumption made in this paper.

IV. RESULTS

Model (1) was estimated for every interval between 1949 and 1979 whose length T is at least five years. Every estimation generated the average squared estimated random shocks (i.e., the value of the objective function (3)), the sample variances of the industry specific target rates β_j, and the average of the persistence coefficients ρ_j. As predicted by Brozen [1971], the sample variances of the target profit rates decline steadily as the estimation period is lengthened. However, contrary to an implicit claim by Brozen, this decline has a floor. Fitting a constant and a term declining with $1/T$ to these variances[12] one obtains

$$\frac{1}{J-1}\sum(\beta_j-\bar{\beta})^2 \approx 0.0023+0.0030/T.$$

In other words, in an estimation period of ten years, the sample variance of the estimated industry-specific components is around 0.0026 (corresponding to a standard deviation of approximately 5 percent), and increasing the length of the estimation period will reduce this variance only by about 10 percent.

The time pattern of the profit rate variances also reveals that intervals straddling the year 1974 generally had a lower variance than intervals which did not, the difference being approximately 0.0004. Apparently, the oil crisis disturbed the old noncompetitive structures, resulting in either greater competition or the formation of new, randomly different noncompetitive structures.

Concerning the estimated random shocks, there is a "calm" period from 1956–73, in which the average shocks were very low. If the estimation period is extended into this calm period from either side, one obtains increasingly better fits (which lead to shorter confidence intervals for the parameters estimated). This result is not due to the increasing length of the estimation interval as such, but rather to the fact that a larger and larger portion of this estimation interval is located in a period with few stochastic shocks. The issue of the optimal length of the estimation period is therefore complicated by the succession of periods displaying varying intensities of the disturbances.

The averages of the adjustment coefficients, $\bar{\rho}$, *increased* as the sample periods were lengthened. They increased rapidly until a length of approximately ten years, where they reached a plateau with only small further increase. One possible explanation for this phenomenon is that the target profit rates of the short-term adjustment process characterized by the ρ_j are not the

12. This reflects the assumption that part of the industry-specific components is constant, while a portion of it changes randomly from year to year.

equilibrium rates themselves but fluctuate around them with a period of approximately ten years.[13]

Another empirical feature of the adjustment processes is revealed by plotting the short-term adjustment coefficients ρ_j versus the industry-specific components β_j. Apparently there is no systematic relation between the two. If an industry shows signs of a speedy short-run adjustment, this does not mean that its profit rate remains close to the general average.

How big are the differences in industry-specific target rates? Choosing the longest estimation period with low average random shocks and a low $\bar{\rho}$ (indicating that it avoids most of the fluctuations in target rates of return), 1954–73, the industry-specific rates of return were calculated. Table I presents the results of this estimation. The stars indicate that the parameters were significantly different from zero at the 0.05* and 0.01** levels, assuming serially uncorrelated random shocks.[14] This assumption is not actually satisfied, since the estimated random shocks are clearly autocorrelated; however, they still seem random enough that the significance levels give an indication of the relative precision of the estimates. According to these tests, six industries displayed profit rates which were on average 5.3 percent above the general profit rate, five had rates of profit which were on average 5.7 percent below the general profit rate, while nine industries produced profit rates which were not significantly different from the general profit rate.

V. RATES OF RETURN AND RISK

Although the focus has not been on explaining profit rate differentials, but instead on examining their long-run behavior, economic theory only postulates an equalization of returns in a competitive market once an appropriate adjustment for risk has been made. The issue of risk and its relationship to rates of return in manufacturing is a very complex one. One simple measure of risk is the industry standard deviation. Figure 1 plots for every industry, marked by its Standard Industrial Classification (SIC) code, the average target profit rate $\bar{r}_j = \bar{\alpha} + \beta_j$ (where $\bar{\alpha}$ is the average of all α_t for 1949–79, which was $\bar{\alpha} = 0.166$) against the estimated standard deviations σ_j of the industry returns.[15] Obviously, part of the industry-specific differences can be ex-

13. Assume the target profit rate is first low and then high. If one fits to this an adjustment to an unchanging target, the residuals will first tend to be negative, and then positive, which is misinterpreted by the model as strong serial correlation. If in a third period the target profit rate is low again, extending the estimation into this third period will not further increase this apparent autocorrelation.

14. Since our robust estimation does not easily yield confidence intervals, the significance levels of the maximum likelihood estimation of the same model were used, which gives results very similar to those in Table I.

15. The full period 1949–79 had to be used, since for the shorter period 1954–73 an estimate of the mean-variance frontier would not have been possible.

TABLE I
Estimation Period 1954–73

SIC	Industry	Industry-Specific Return β_j	Standard Deviation σ_j	Adjustment Coefficient ρ_j	Weight w_j
20	Food and Kindred Products	–0018*	0.016	0.64	0.102
21	Tobacco Manufactures	0091*	0.064	0.85	0.011
22	Textile Mill Products	–0050**	0.018	0.79	0.039
23	Apparel and Related Products	0006	0.018	–0.01	0.016
24	Lumber and Products	0038	0.057	0.97	0.028
25	Furniture and Fixtures	0001	0.018	0.10	0.010
26	Paper and Allied Products	0023	0.040	0.95	0.054
27	Printing and Publishing Industries	0000	0.023	0.85	0.039
28	Chemicals and Allied Products	0038*	0.029	0.95	0.107
29	Petroleum and Coal Products	–0152**	0.029	0.61	0.059
30	Rubber Products	–0016	0.029	0.78	0.025
31	Leather and Leather Products	–0004	0.029	–0.13	0.005
32	Stone, Clay, and Glass Products	0017	0.044	0.97	0.041
33	Primary Metal Industries	–0043**	0.022	0.88	0.133
34	Fabricated Metal Products	–0020**	0.014	0.64	0.060
35	Machinery (Except Electrical)	0027**	0.013	–0.69	0.084
36	Electrical Equipment	0030**	0.022	0.21	0.059
37	Transportation Equipment	0067**	0.051	0.36	0.101
38	Instruments and Related Products	0069**	0.031	0.40	0.018
39	Miscellaneous Manufactures	0013	0.037	0.71	0.012

plained by these standard deviations. If industry 29 is ignored as an obvious outlier, a regression yields

$$\bar{r}_j = \underset{(0.020)}{0.083} + \underset{(0.434)}{2.045}\,\sigma_j \qquad R^2 = 0.57$$

i.e., 57 percent of the variation in industry-specific levels can be explained by differences in standard deviation, and the "risk premium" is 2 percent for each 1 percent of standard deviation.[16]

Unfortunately, modern portfolio theory would argue that the standard deviation is not a good indicator of risk. Variability of the asset returns which can be diversified away by investing in more than one industry cannot be considered a disadvantage for the investor. Instead of comparing rates of return and standard deviations of single industries, the Capital Asset Pricing Model (CAPM) explains the link between risk and return by a portfolio approach. The basic assumption is that every investor owns an "efficient" portfolio that has highest expected rate of return for its given variance, or (what is the same) lowest variance for its expected return. It is one of the important results of mean-variance theory that the combination of efficient portfolios is efficient as well. Therefore, the sum total of all portfolios, the "market portfolio," must be efficient if everyone holds an efficient portfolio. Using the efficiency of the market portfolio, one can derive a linear relation between the expected return of every asset and its contribution to the volatility of the market portfolio (the "beta" of that asset).

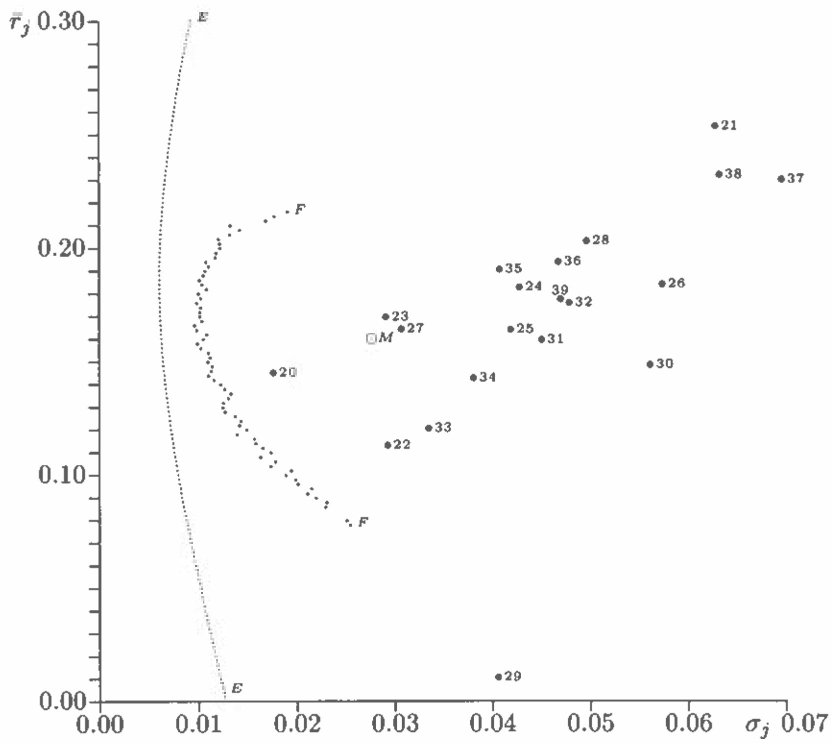
The dotted line *EE* in Figure 1 is the hyperbola representing the best combinations of means and standard deviations that can be obtained forming portfolios, allowing short positions, of the individual industries. *M* represents mean and standard deviation of the whole manufacturing sector combined. If the CAPM were valid, *M* would have to lie on this frontier. Various tests have been designed to test for this fact, but they require a far larger number of observations than our annual data can provide.

Data limitations are not the only obstacles to applying mean-variance theory to the present data set. It can be argued that on theoretical grounds the CAPM is simply not appropriate here. Richard Roll [1977] pointed out that the usual tests of the efficiency of the market portfolio were of questionable validity, since they usually worked with very poor proxies of the true market portfolio, which should contain all the assets any investor holds. This criticism is certainly valid in this case, since the portfolios considered are only those which consist of direct investments in manufacturing.

A second criticism raised by Haim Levy [1983] is also pertinent to an application of the CAPM to the present situation. Efficient portfolios usually

16. Estimates by Fisher and Hall [1969] obtain a similar R^2 but only half the coefficient of σ_j.

FIGURE 1
Average Target Profit Rates $\overline{r}_j$ plotted against standard deviations σ_j



contain about 50 percent "short positions" (negative holdings), while the mechanism of investment flows and diversification in the background of the present model excludes negative holdings. For the given data this is a relevant restriction. To the right of *EE*, Figure 1 shows a second frontier *FF*, which is composed of portfolios with only nonnegative holdings. (It is only approximate since it was constructed by randomly generating several million portfolios with nonnegative holdings and selecting those with the lowest variances.) This second frontier is far inside the first one, presumably due to business cycle effects which cannot be diversified away with nonnegative holdings. Clearly, the larger the difference between frontiers *EE* and *FF*, the less applicable are the usual tests of efficiency of the market portfolio. In addition, if only nonnegative holdings are allowed, the theoretical result that the market portfolio is efficient no longer holds. Even if every investor holds a portfolio that lies on *FF*, the sum of all portfolios need no longer lie on *FF*.

Due to the inapplicability of the capital asset pricing model, only a verbal caution can be given regarding the regression result of standard deviation versus return. The standard deviation is a deficient measure of risk due to the possibility of diversifying the portfolio (if not for other reasons as well). Furthermore, a risk premium may not be the only explanation of the relation between standard deviation and rate of return. For example, it can be argued that an industry with stronger fluctuations requires higher amounts of working capital. If this is the case, the denominator in the profit rate calculation underestimates the actual capital applied in these industries, and therefore the rate of return itself is overestimated.

VI. CONCLUSION

Although economic theory predicts that, under competitive conditions, industry rates of return will be equalized only in the long run, applied research on this topic has generally adopted a short-run methodology. Brozen [1970], among others, has raised doubts concerning the persistence of profit rate differentials and has suggested that the time period necessary to adjust these differentials in actual competition is longer than the time periods usually examined in cross-section studies. Although this paper lends support to Brozen's suggestion that previous short-run studies have failed to capture long-run equilibrium, the increase in equalization gained from lengthening the estimation period seems, at first glance, disappointingly small, and a persistent differential still endures in the long run. Further, the estimates indicate that, beyond the short-run reactions to profit rate disparities, there may be medium-term fluctuations of the target profit rates, lasting around ten years. A third result is that the short-term reaction speeds of the industries are very different, and a faster adjustment in the short run does not necessarily lead to more profit rate equality in the long run. Brozen's hypothesis of equalization in the long run may still hold, but the interaction of many markets, each of which reacts with a different speed, may make this adjustment process a very complex one.

The main focus of this study was an assessment of the size of industry-specific differences, without the attempt to explain where these differences come from. However, the inclusion of additional explanatory variables was not necessary for exploring whether the industry differences could be explained by differences in industry volatility. Indeed there is a strong and significant relation between industry standard deviations and rates of return.

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