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TAX POLICY AND INVESTMENT BEHAVIOR: REPLY AND FURTHER RESULTS

By ROBERT E. HALL AND DALE W. JORGENSON*

In our earlier paper in this *Review*,[1] we presented the results of an empirical study of the impact of changes in American tax policy on investment expenditures. The basic principles of the investment function underlying our work are the familiar ones: starting from a hypothesis about the form of the production function, we derived the profit-maximizing demand for capital input as a function of output and the relative price of capital services. We then estimated the parameters of the lagged adjustment of actual capital to this desired level. One property of the demand function, namely its elasticity with respect to the price of capital services, obviously has a crucial role in the application of such an investment function to the measurement of the effects of tax policy. We chose the relatively simple Cobb-Douglas parametrization for the production function. This choice implied the particular numerical value of unity for the own-price elasticity of demand for capital (or elasticity of substitution). At the time we made the choice, we believed that it was a reasonable one in the light of evidence from studies of production functions and factor demand. Evidence accumulated since then has strongly confirmed our hypothesis, as we will show here.

Robert Coen [4] has correctly pointed out that in our original exposition we did not provide a detailed justification for the choice of this hypothesis. Our first objective is to review the extensive evidence on the elasticity of substitution now available. The elasticity of substitution is difficult to estimate from time series since it is essentially a second order parameter in the relationship between products and factors, as Kmenta [18] and Nelson [21] have pointed out. As we observe in Section I, relatively minor errors in specification or seemingly unimportant differences in methods of measurement may have a substantial impact on the estimated elasticity of substitution. The elasticity can be estimated more reliably from individual cross sections and from successive cross sections, as Griliches [9] has demonstrated. New evidence made available by Zarembka [26] reinforces Griliches' conclusions and removes important sources of ambiguity in the interpretation of earlier studies.

Our second purpose is to present a new set of estimates of the time form of the response of investment to changes in its determinants. These estimates, based on a refinement of our previous econometric method, suggest that the impact of tax policy is considerably more rapid than we originally estimated. Our earlier estimates of the timing of changes in investment following a change in tax policy need revision in light of this finding. Our new results are in much better agreement with the results of studies of the time

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[1] See [10] in this *Review*, June 1967.

structure of investment behavior at the level of industry groups or for individual firms.

Our third objective is to extend our previous estimates to measure the impact of policy changes since 1963 on investment expenditures. Since 1963 several important changes in tax policy have been carried out; these changes provide a further opportunity for evaluation of the responsiveness of investment expenditures to tax policy. We first consider the effects of reduction of the rate of taxation on corporate profits from 52 percent to 48 percent in 1964. This reduction was combined with elimination of the requirement that the base for depreciation charges be reduced by the amount of the investment tax credit, making the tax credit adopted in 1962 considerably more effective. We find that the main stimulus to investment resulting from the 1964 tax cut came from changes in the investment tax credit rather than changes in the tax rate. If the tax rate alone had been reduced with no alteration in the investment tax credit, the direct impact of the tax cut of 1964 on the level of investment would have been small and negative. A second change in tax policy we analyze is the suspension of the investment tax credit and certain types of accelerated depreciation in late 1966 and early 1967. These changes had a substantial impact on investment and would have had a much greater impact if the suspension had been kept in force for the period stipulated at the time of the suspension.

I. *The Elasticity of Substitution*

In our econometric model of investment behavior we have maintained the hypothesis that the elasticity of substitution is equal to unity. Although relatively minor errors in specification or small differences in methods of measurement have substantial impact on estimates of the elasticity of substitution from time series, this parameter can be estimated reliably from cross sections, either through the study of factor demand or from direct estimates of the production function. We first review estimates of the elasticity of substitution from empirical studies of factor demand. Our maintained hypothesis is strongly supported by this evidence. Second, we consider recent attempts by Eisner and Nadiri [6] and by Eisner [7] to estimate the elasticity of substitution from data on investment expenditures. The resulting estimates are extremely sensitive to errors in specification and prove to be highly unreliable. We conclude that these estimates of the elasticity of substitution from data on investment do not provide a useful alternative to our hypothesis that the elasticity of substitution is unity.

In the original study of the elasticity of substitution by Arrow, Chenery, Minhas, and Solow [1, henceforward ACMS] the elasticity of substitution was significantly different from zero for every industry group included in the study. For nine of the twenty-four industry groups the elasticity of substitution was significantly below unity. These findings are shown in the first column of Table 1. Fuchs [8] has re-examined these conclusions, inserting a dummy variable representing the level of development of countries included in the sample. For only two industry groups is the elasticity of substitution significantly different from unity—for one group the elasticity is above unity and for the other below. The results of ACMS and Fuchs are compared in Table 1. The first conclusion supported by these results is that the elasticity of substitution is significantly different from zero. However, Fuchs has thrown into question the finding of ACMS that the elasticity of substitution may be less than unity.

The results of ACMS and Fuchs are based on international cross sections. For the United States a number of studies have been made of the elasticity of substitution for two-digit industries within manufacturing. Dhrymes [5], Minasian [20], and Solow [23] have fitted relationships involving value added per man and the wage rate. More recently, this relationship has been fitted to data for successive cross sections by Griliches [9] and by Zarembka [26]. Second, Bell [2] and Dhrymes [5] have fitted analogous relationships involving capital and its rental price. All of these estimates provide support

TABLE 1—COMPARISON OF ESTIMATES OF THE ELASTICITY OF SUBSTITUTION FOR MANUFACTURING INDUSTRIES BY ACMS AND BY FUCHS FROM INTERNATIONAL CROSS SECTIONS, 1950–1955*

Industry Group	ACMS	Fuchs
Food and kindred products		
Dairy products	0.72	0.90
	(.07)	(.08)
Fruit and vegetable canning	0.86	1.09
	(.08)	(.10)
Grain and mill products	0.91	1.32
	(.10)	(.17)
Bakery products	0.90	1.07
	(.07)	(.11)
Sugar	0.78	0.90
	(.12)	(.18)
Tobacco	0.75	1.22
	(.15)	(.21)
Textile mill products		
Spinning and weaving	0.81	0.98
	(.07)	(.10)
Knitting mills	0.79	0.95
	(.06)	(.08)
Lumber and wood products	0.86	1.08
	(.07)	(.14)
Furniture and fixtures	0.89	1.04
	(.04)	(.09)
Pulp, paper, and products	0.97	0.91
	(.10)	(.18)
Printing and publishing	0.87	1.02
	(.06)	(.09)
Chemicals and products		
Basic chemicals	0.83	1.11
	(.07)	(.10)
Misc. chemicals	0.90	1.06
	(.06)	(.09)
Fats and Oils	0.84	1.06
	(.09)	(.18)
Leather and leather goods	0.86	0.98
	(.06)	(.10)
Stone, clay, and glass products		
Clay products	0.92	0.66
	(.10)	(.20)
Glass	1.00	1.27
	(.08)	(.10)
Ceramics	0.90	1.08
	(.04)	(.13)
Cement	0.92	1.31
	(.15)	(.22)
Primary metal products		
Iron and steel	0.81	0.76
	(.05)	(.11)
Nonferrous metals	1.01	0.94
	(.12)	(.20)

TABLE (*Continued*)

Industry Group	ACMS	Fuchs
Fabricated metal products	0.90	1.01
	(.09)	(.17)
Electrical Machinery	0.87	1.03
	(.12)	(.21)

* Source: Nerlove [22], Table 1, Columns 1 and 2, pp. 60–63.

for the hypothesis that the elasticity of substitution is unity.

Minasian's estimates of the elasticity of substitution are based on data for individual states for 1957; he finds two estimates of the elasticity of substitution significantly above unity, three significantly below, and thirteen not significantly different from unity among a total of eighteen industry groups.[2] Solow's results are based on data for Census regions for 1956; he finds no estimates significantly below unity, only one significantly above, and seventeen not significantly different from unity.[3] Dhrymes employs data for individual states in 1957, essentially the same data analyzed by Minasian; however, he finds very low estimates of the elasticity of substitution for regressions in which value added per man is the dependent variable and the wage rate is an independent variable.[4] Zarembka has recently analyzed these same data and has discovered that Dhrymes inadvertently regressed the wage rate on value added per man rather than vice versa. With this error removed, estimates based on data for individual states in 1957 are similar to those of Minasian.[5]

In fitting relationships involving capital and its rental price to data for individual states for 1958, Bell finds no estimates significantly below unity, eleven significantly above, and seven not significantly different

[2] See [9], Table 1, p. 287.

[3] See [9], Table 1, p. 287.

[4] Dhrymes [5], Table 1, p. 358, column labeled σ_1.

[5] Zarembka [26], p. 9, fn. 2. Dhrymes' estimates are reproduced by Nerlove [22], Table 1, pp. 60–63, column labeled "Regression I." Corrected estimates are given by Zarembka, Table 1, p. 10, column labeled $\hat{\sigma}$ for 1957.

from unity.[6] Analyzing data on capital for 1957, Dhrymes finds one estimate significantly below unity, two significantly above, and fourteen not significantly different from unity.[7] The overall conclusion from these four studies of cross sections of individual states or Census regions for the years 1956, 1957, and 1958 is that the elasticity of substitution is not significantly different from unity. Significant deviations above unity are about as frequent as significant deviations below unity. We conclude that our maintained hypothesis that the production function is Cobb-Douglas in form is supported by evidence from studies of factor demand, including demand for labor and demand for capital.

Estimates of the elasticity of substitution from time series are much more erratic than those from cross sections. Estimates substantially above unity and substantially below unity have been obtained for alternative specifications of factor demand equations. The estimates appear to be extremely sensitive to changes in the stochastic specification of factor demand relationships; minor differences in measurement produce substantially different estimates.[8] The primary explanation for the lack of reliability of time series estimates of the elasticity of substitution is that the elasticity is essentially a second-order parameter in the relationship between the growth of output and the growth of factor inputs, as Nelson [21] has demonstrated.

A partial reconciliation of time series and cross section results has been undertaken by Griliches [9] and by Zarembka [26], taking into account variations in labor quality over the cross section of individual states and allowing for serial correlation in successive cross sections. Griliches estimates the elasticity of substitution for two-digit industries from cross sections of individual state observations for 1957. The results are given in the first column of Table 2. Griliches summarizes these results as follows:

TABLE 2—ESTIMATES OF THE ELASTICITY OF SUBSTITUTION FOR TWO-DIGIT MANUFACTURING INDUSTRIES IN THE UNITED STATES BY GRILICHES*

Industry Group	a. Cross Section	b. Lagged Adjustment	c. Serial Correlation
Food	0.91	1.01	**
	(.10)		
Textile	0.94	1.11	1.09
	(.17)		(.54)
Apparel	1.06	0.84	0.63
	(.19)		(.21)
Lumber	1.07	1.11	1.18
	(.06)		(.20)
Furniture	1.04	0.99	**
	(.07)		
Paper	1.67	1.30	**
	(.30)		
Printing	0.83	0.45	0.68
	(.18)		(.28)
Chemicals	0.71	0.59	0.70
	(.22)		(.40)
Petroleum	***	***	***
Rubber and Plastics	1.28	2.21	0.90
	(.42)		(.43)
Leather	0.84	1.60	1.16
	(.26)		(.40)
Stone, clay, glass	0.91	0.77	1.88
	(.19)		(.49)
Primary metals	1.41	3.49	2.37
	(.42)		(.47)
Fabricated metals	0.85	1.17	1.20
	(.14)		(.28)
Machinery, exc. electrical	1.24	2.40	2.00
	(.38)		(.29)
Electrical machinery	0.66	0.40	1.53
	(.31)		(.53)
Transportation equipment	0.96	1.09	**
	(.55)		
Instruments	0.75	0.82	**
	(.43)		

* Source: Griliches [9], Table 3, p. 293.
** Contradicts serial correlation model; see Griliches [9], note c, Table 3, p. 294.
*** No significant relationships found.

> The first set of σ (elasticity of substitution) estimates is comparable to, though substantially better than (in terms of fit and t ratios) the Minasian and Solow estimates and is generally of the same order of magnitude. Only one of these σ's (out of 17) is significantly different from unity, and that one is above unity.[9]

[6] See [9], Table 1, p. 287.
[7] Dhrymes [5], Table 1, p. 358, column labeled σ_2.
[8] Time series studies are reviewed by Nerlove [22], pp. 82–100, and by Griliches [9], pp. 285–290.
[9] [9], p. 292.

Griliches next fits two alternative dynamic models to cross sections of individual states for 1957 and 1958—a lagged adjustment model and a model that takes serial correlation into account. The results are given in the second and third columns of Table 2. Griliches summarizes these results as follows:

> The second set of σ (elasticity of substitution) estimates is based on the partial adjustment equation, while the third is based on the serial correlation model. In 12 out of 17 cases the latter model is the one consistent with the data. In general, all of the estimated σ's are not very (statistically) different from unity, the significant deviations if anything occurring above unity rather than below it.[10]

In Griliches' serial correlation model equality between elasticities of substitution for cross sections in 1957 and 1958 is taken as a maintained hypothesis. This hypothesis can be tested by estimating elasticities of substitution separately for the two cross sections and testing their equality, given serial correlation in the residuals. A test of this type has been carried out by Zarembka. Except for one industry, the hypothesis is accepted. Zarembka also presents combined estimates of the elasticity of substitution, using a statistical model similar to that of Griliches. For only two industries of the thirteen studied by Zarembka is the elasticity of substitution significantly different from unity; for both the elasticity is significantly less than unity.[11] Our overall conclusion for estimates of the elasticity of substitition from successive cross sections is that this parameter is not significantly different from unity. For the individual cross sections analyzed by Bell, Dhrymes, Griliches, Minasian, Solow, and Zarembka, the elasticity of substitution is significantly different from zero. This conclusion also holds for the estimates from successive cross sections by Griliches and Zarembka.

We have already indicated that estimates of the elasticity of substitution from time series data on capital input are unreliable. This implication is strongly supported by the estimates of the elasticity of substitution from quarterly data on investment expenditures reported by Eisner [7]. Eisner's empirical results extend those of Eisner and Nadiri [6], incorporating a variable that represents changes in the difference between two measures of the rental price of capital. Eisner demonstrates that the effect of this variable is not significantly different from zero in any of the regressions he presents so that Eisner's model effectively reduces to that presented by Eisner and Nadiri.

The Eisner-Nadiri model is alleged to provide a direct test of a model presented earlier by Jorgenson and Stephenson [14] in which the elasticity of substitution is taken equal to unity. Closer examination of the two models reveals that they are mutually contradictory and that evidence predicated on the validity of one is not relevant to the validity of the other.[12] In addition, Bischoff [3] has drawn attention to an important error of specification in the econometric model used by Eisner and Nadiri. Since the model presented by Eisner effectively reduces to that of Eisner and Nadiri, Bischoff's analysis carries over directly to Eisner's model. The basic difficulty in the Eisner-Nadiri model arises from the fact that errors in the distributed lag function employed in the model are autocorrelated. Eisner and Nadiri have ignored this autocorrelation so that their estimates of parameters of the distributed lag function are inconsistent. The estimated elasticity of substitution is highly sensitive to this error in specification. Bischoff demonstrates that elimination of the error in specification from the model used by Eisner and Nadiri and by Eisner results in the following conclusions:

1. The elasticity of capital with respect to the ratio of the price of output to the rental price of capital is not significantly different from unity.[13]

2. The elasticities of capital with respect

[10] [9], p. 292.

[11] Zarembka [26], Table 2, p. 16.

[12] Jorgenson and Stephenson [17], Section 2, pp. 3–5.

[13] Bischoff [3], Table 2, tests for the maintained hypothesis that the errors are a first order autoregressive process. This conclusion directly contradicts the first conclusion of Eisner and Nadiri [6], p. 380.

to output and prices taken together are not significantly different from unity.[14]

3. The best fitting elasticity of capital with respect to output is negative with a high standard error, hardly consistent with flexible accelerator models of investment as Eisner and Nadiri and Eisner have claimed.[15]

4. Results do not contradict our maintained hypothesis that the production function is Cobb-Douglas in form.[16]

Our conclusion is that the results of Eisner [7] and Eisner and Nadiri [6] rest on an error in the stochastic specification of the Eisner-Nadiri model and that all of the conclusions Eisner has drawn concerning our analysis of tax policy are directly reversed for a correctly specified model of the Eisner-Nadiri variety. The sharp conflict between estimates of the elasticity of substitution by Eisner and Nadiri and estimates from individual and successive cross sections we have reviewed above is removed when the error in specification of the Eisner-Nadiri model is eliminated. Estimates of parameters such as the elasticity of substitution and the degree of homogeneity of the production function from the Eisner-Nadiri model are highly unreliable. This fact is demonstrated both by the large standard errors associated with estimates from a correctly specified model and by the high degree of sensitivity of the estimates to errors in specification. These estimates do not provide a useful alternative to the hypothesis that the production function is Cobb-Douglas in form.

We turn now to Coen's estimates of the elasticity of substitution. His model is a straightforward generalization of ours, obtained by replacing our demand function for capital by one derived from a *CES* production function. Thus he is able to avoid the specification error of Eisner and Nadiri, but at the cost of estimating a nonlinear equation. His estimates of σ are clearly inconsistent with the empirical results on the elasticity of substitution that we have summarized above. Unfortunately, Coen does not test the hypothesis that σ is different from unity, so the reader has no information about the statistical significance of his results. But whether or not his estimates are significantly different from one, it is not appropriate to proceed on the basis that σ is in the neighborhood of 0.2 in investment research in the face of the large body of evidence to the contrary from research on production functions and from Bischoff's work on investment functions. Until a reconciliation is available for the low elasticities of substitution in Coen's investment equations and the roughly unitary elasticities in production equations, the most reasonable course is to impose the value of unity in investment equations.

II. *The Conceptual Basis for the Measurement of the Effects of Tax Policy.*

Coen expresses concern about the "legitimacy" of the appearance of output in the demand function for capital; he concludes that investigators who have studied investment equations in which an output variable appears in the right-hand side have made a serious conceptual error. Thus Coen rejects our demand function for capital,

$$K_t^* = \alpha \frac{p_t Q_t}{c_t}$$

on the grounds that Q_t is set by the firm and cannot "legitimately" appear on the right hand side. What he presumably means is that this equation could not be used for decision-making within the firm; until the firm decides on Q_t, it cannot calculate K_t^*.[17]

Our interpretation of this equation is the following: It is the demand for capital conditional on the level of output. The question

[14] Bischoff [3], Table 2. This conclusion directly contradicts the second conclusion of Eisner and Nadiri [6], p. 380.

[15] The numerical value given by Bischoff is $-.434$ under the hypothesis that the autoregressive parameter is zero. This contradicts the third conclusion of Eisner and Nadiri [6], pp. 380–81. This conclusion is repeated by Eisner [7], pp. 5–6.

[16] This contradicts the fourth conclusion of Eisner and Nadiri [6], p. 381. The remaining conclusions of Eisner and Nadiri are refuted by Bischoff [3] and by Jorgenson and Stephenson [17], esp. Section 4, pp. 11–12.

[17] Coen's argument must be distinguished from the entirely different econometric point that if a jointly dependent variable appears on the right-hand side the choice of estimation method should take account of that fact.

of how firms determine output is an interesting one, but, as we shall see, it is irrelevant to our particular use of the investment equation. We should note a very serious difficulty which Coen's approach might encounter: if firms produce under constant returns to scale, their choice of output is indeterminate. Our conditional demand function still holds, but Coen's equation breaks down. The reader is invited to substitute parameter values with $\alpha+\beta$ equal to one into Coen's equation 17 to see how this difficulty arises. Since empirical evidence generally supports constant returns to scale,[18] Coen's method has serious shortcomings.

The point of Coen's argument is that tax policy affects not only the conditional demand for capital given output, but also the level of output itself. Since we measured only the first effect and assumed that output remained unchanged, Coen suggests that we understated the true effect of tax policy. Coen's charge that we thereby committed a conceptual error cannot be sustained. At worst, we failed to make clear the question we were answering with our calculations. To be precise, the question was the following: Suppose that instead of increasing investment incentives, the government had pursued an alternative fiscal policy which resulted in exactly the same dollar value of output, interest rate, and capital goods price. Then by how much would investment have been less under the alternative policy?

Another question which could be asked is the general equilibrium one: What were the overall effects of tax incentive policies, assuming no alternative compensatory policy? The answer waits for the development of a satisfactory model of general equilibrium in the United States economy.

Coen prefers to ask a third question: Suppose that there had been no investment incentive policy and that the government had pursued an alternative fiscal policy that reduced output to the point that every firm was in equilibrium at the output price, wage, interest rate, and capital goods price that existed under the incentive policy. Then by how much would investment have been reduced? The difficulty with this question is that unless decreasing returns to scale are strong, a very large decrease in output is necessary to compensate for the higher rental price of capital goods in the absence of the investment incentives. In the case of constant returns to scale, output and capital would always fall to zero in this circumstance. This suggests that Coen's question is rather less interesting than he supposes. By contrast, our calculations are not affected by the assumption of constant returns to scale. Even if a general equilibrium calculation could be made, we would want to carry out our calculations in order to separate the total effect of tax policy into an incentive effect and an output effect. We have attempted to provide a method for measuring the incentive effect which will be useful in further research in this area.

III. *New Estimates of the Parameters of Investment Functions.*

Our earlier evaluation of the effects of tax policy on investment behavior was based on policy changes through 1963. Since 1963, a number of important changes in tax policy have taken place, providing a further opportunity for evaluation of the responsiveness of investment expenditures to changes in tax policy. For this purpose we have re-estimated our model, using data that have become available since our earlier study. We have also revised our econometric technique to take account of recently developed methods for analysis of distributed lag functions. With these changes we obtain a new set of investment functions for the non-farm sector of the United States. In the following section we employ these investment functions to evaluate the effects of changes in tax policy that have taken place since 1963.

Our basic econometric model is a dis-

[18] Zarembka has examined this question very carefully in the study previously cited. For three of his thirteen industry groups the degree of homogeneity differs significantly from unity; for one of these groups, the degree of homogeneity is below unity as assumed by Coen. For the other two, the degree of homogeneity is above unity. Overall, there is little evidence that the degree of homogeneity differs significantly from unity; deviations above unity appear, if anything, to be slightly more important than deviations below unity.

tributed lag function in net investment and changes in desired capital:

$$N_t = \sum_{\tau=0}^{4} \beta_\tau (K^*_{t-\tau} - K^*_{t-\tau-1}) + \epsilon_t$$

where N_t is net investment in period t, K^*_t is desired capital, and ϵ_t is a random error.[19] Under the assumption that the production function is Cobb-Douglas in form, the desired level of capital is given by:

$$K^*_t = \alpha \frac{p_t Q_t}{c_t},$$

where α is the elasticity of output with respect to capital, p_t is the price of output, Q_t the quantity, and c_t the rental price of capital.

The rental price of capital services depends on the tax rate u, the after tax rate of of return r, the investment goods price q, the rate of replacement δ:

$$c = \frac{1 - k - uz}{1 - u} q(r + \delta),$$

where z is the discounted value of depreciation allowances allowed for tax purposes and k is the investment tax credit. This formula is appropriate for the period after 1964 when the tax credit is not deducted from allowable depreciation. For 1962 and 1963 the appropriate formula is:

$$c = \frac{(1 - k)(1 - uz)}{1 - u} q(r + \delta);$$

during these years the depreciation base was reduced by the amount of the investment tax credit.

Through 1953 the rental price is that appropriate to straight line depreciation. Since 1954 the rental price is that appropriate to the sum of the years' digits depreciation.[20] The investment tax credit was introduced in 1962 at a rate nominally equal to 7 percent of the value of investment in equipment. In practice, certain limitations on the applicability of the investment tax credit reduce its effective rate to 6 percent for manufacturing equipment and 5.8 percent for non-farm, non-manufacturing equipment.[21] From October 1966 to March 1967 the investment tax credit was suspended.

We took the tax rate to be the statutory rate prevailing during most of each year. We did not allow for excess profits taxes during the middle thirties or the Korean War. For all years we took the rate of return before taxes ρ to be constant at 20 percent. This value is higher than the value of 14 percent used in our previous studies. The higher value is consistent with the results of Jorgenson and Griliches [15]. Under our assumption of a constant before-tax rate of return the after-tax rate $r = (1-u)\rho$ varies with the tax rate. The investment goods price is the same as that used to deflate investment expenditures in current prices and the rate of replacement is the same as that used to calculate capital stock. Estimates of lifetimes of assets allowable for tax purposes were obtained from a special Treasury study [24]. These estimates are the same as those employed in our previous studies:

Period	Equipment	Structures
1929–54	17.5	27.8
1955	16.3	25.3
1956–61	15.1	22.8
1962–65	13.1	22.8

New estimates of these lifetimes for recent years would require that the special Treasury study be updated.

The statistical technique for fitting our econometric model to data on investment expenditures is described in detail in our paper, [12]. Briefly, we include a lagged value of net investment together with current and five lagged values of the change in the desired level of capital. The coefficient of lagged net investment in this first stage regression is interpreted as an estimate of autocorrelation. It is used to perform an

[19] For further details, our paper [10], esp. pp. 392–98, may be consulted.

[20] Depreciation under the sum of the years' digits formula has a higher present value for the range of lifetimes and rates of return of interest for this study. See [10], Table 1, p. 395.

[21] These estimates of the effective rate of the tax credit are based on data from tax returns for 1963 [25].

TABLE 3—FITTED INVESTMENT FUNCTIONS, 1935–40 AND 1954–65, SECOND STAGE RESULTS

Asset Class	$\hat{\alpha}\hat{\beta}_0$	$\hat{\alpha}\hat{\beta}_1$	$\hat{\alpha}\hat{\beta}_2$	$\hat{\alpha}\hat{\beta}_3$	$\hat{\alpha}\hat{\beta}_4$	R^2	s	d
Mfg. Equipment	.0130 (.0047)	.0200 (.0034)	.0208 (.0040)	.0153 (.0040)	.0036 (.0053)	.602	.620	2.099
Mfg. Structures	.0041 (.0033)	.0082 (.0030)	.0093 (.0035)	.0073 (.0032)	.0024 (.0034)	.186	.513	1.304
Non-farm, Non-mfg. Equipment	.0374 (.0083)	.0282 (.0038)	.0211 (.0063)	.0160 (.0047)	.0129 (.0090)	.800	1.190	1.724
Non-farm, Non-mfg. Structures	.0059 (.0034)	.0105 (.0032)	.0118 (.0035)	.0098 (.0031)	.0046 (.0029)	.169	.506	1.825

autoregressive transformation of net investment and change in desired capital, which are then used in a second stage regression. In the second stage we impose the constraint that the lag coefficients lie on a parabola. The estimated lag functions and other regression results are given in Table 3. It should be noted that R^2 for these regressions is a measure of the degree of explanation of the autoregressively transformed values of net investment, and is therefore not comparable to the R^2 in our earlier results. The overall goodness of fit as measured by the standard errors is superior to that of our previous investment functions for 1931–41 and 1950–63 except for manufacturing structures. This improvement is mainly due to the change in time period and to revisions of the basic investment data; however, it is also partly due to the change in our specification of the distributed lag function. We have added three lagged changes in desired capital, which improves the results to some extent.

The parameters of the distributed lag function $\{\mu_t\}$ may be estimated by employing the constraint that the sum of the coefficients of this function must be unity to estimate the parameter α.[22] The resulting estimates are given in Table 4. The mean lag for each function is also given in Table 4. Comparing these mean lags with estimates from our earlier studies, we find that the new estimates were very similar for investment in equipment. The mean lag is now estimated to be slightly lower for manufacturing equipment and slightly higher for non-farm, non-manufacturing equipment. For structures, the new estimates differ substantially from the old. The old estimate of the mean lag for manufacturing structures was 3.84 years whereas the new estimate is 1.86; the old estimate of the mean lag for non-farm, non-manufacturing structures was 7.49 years whereas the new estimate is 1.92. For both sets of results the lags for structures are estimated to be longer than for equipment.

A disturbing feature of our earlier results is that the lag pattern fails to agree with the substantial body of evidence from studies at

TABLE 4—FITTED INVESTMENT FUNCTIONS, 1935–40 AND 1954–65, DERIVED RESULTS

Asset Class	Mean Lag (years)	$\hat{\alpha}$
Manufacturing Equipment	1.67	.0727
Manufacturing Structures	1.86	.0312
Non-farm, Non-manufacturing Equipment	1.47	.1160
Non-farm, Non-manufacturing Structures	1.92	.0426

$\hat{\alpha}$: Estimate of the elasticity of output with respect to the capital input.

[22] For detailed discussion of this restriction and its use in estimation of the parameter α, see Jorgenson [13], pp. 135, 147–48.

the level of two-digit industries by Jorgenson and Stephenson [14] and studies at the level of the individual firm by Jorgenson and Siebert [16]. For manufacturing, Jorgenson and Stephenson estimate the average lag at about two years, while results from individual industries range from six to eleven quarters with results clustering in the neighborhood of the overall average. The results for individual firms are characterized by more variability than the results for industries, as would be expected. The average lags estimated by Jorgenson and Siebert range from less than a year to over three years with values between one and two years predominating. "Mayer estimated a seven quarter average lag from the decision to undertake investment to the completion of the project for manufacturing and electric power combined on the basis of surveys [19]. We conclude that our new estimates agree closely with Mayer's survey results and with estimates derived from investment functions for industry groups and for individual firms. Our previous estimates of the average lags for structures are evidently biased by specification errors in the underlying distributed lag functions and should be replaced by our new estimates.

IV. *The Impact of Tax Policy on Investment Behavior.*

Two major changes in tax policy affecting investment behavior have been made since 1963. First, in the tax cut of 1964, the corporate income tax rate was reduced from 52 percent to 48 percent. At the same time, business firms were no longer required to reduce the base for depreciation by the amount of the investment tax credit. This alteration in the tax law substantially enhanced the effectiveness of the investment tax credit adopted in 1962. A second change in tax policy is that from October 1966 to March 1967 the investment tax credit and certain types of accelerated depreciation were suspended. In the legislation implementing suspension of the tax credit, the period of suspension was to run from October 1966 to December 1967. We analyze the effects of the suspension that actually took place and the hypothetical effects of a longer suspension, as stipulated at the time the suspension began.

The qualitative features of the response of investment to a change in tax policy are essentially the same for all changes. To evaluate the effects of particular tax measures it is useful to assess the response of investment quantitatively. Our calculations are based on a partial equilibrium analysis of investment behavior. We hold all determinants of investment expenditures except for tax policy equal to their actual values. We then measure the impact of tax policy by substituting into our investment functions parameters of the tax structure—tax rate, depreciation formulas, and tax credit—appropriate to alternative tax policies. The difference between investment resulting from actual tax policy and investment that would have resulted from alternative tax policies is our measure of the impact of tax policy. In our new calculations both investment and capital stock are measured in prices of 1965. We estimate the impact of all changes in tax policy through 1970. In order to make these estimates, we employed a rough set of projections of the determinants of investment. No great precision was required in these projections, since the estimates of the differential impacts of alternative policies are not at all sensitive to the assumed level of investment.[23]

In analyzing the effect of the tax cut, we assume that the before tax rate of return was left unchanged. Under this condition, the effect of a change in the tax rate on the rental price of capital services is neutral, provided that depreciation for tax purposes is equal to economic depreciation.[24] Under the conditions actually prevailing in 1964, depreciation for tax purposes was in excess of economic depreciation for both plant and equipment in manufacturing and non-farm, non-manufacturing sectors. Accordingly, the rental price of capital services resulting from the tax cut was actually greater than the rental price before the cut. Following are the

[23] Further details are given in [12], pp. 54–55.
[24] See Section 2 of our paper [12].

TABLE 5. CHANGES IN GROSS INVESTMENT (I), NET INVESTMENT (N), AND CAPITAL STOCK (K) RESULTING FROM THE TAX CUT OF 1964 (Billions of 1965 Dollars)

Year	Manufacturing						Non-Farm, Non-Manufacturing					
	Equipment			Structures			Equipment			Structures		
	I	N	K	I	N	K	I	N	K	I	N	K
1964	−.049	−.049	0	−.020	−.020	0	−.135	−.135	0	−.044	−.044	0
1965	−.136	−.129	−.049	−.062	−.061	−.020	−.267	−.241	−.135	−.125	−.123	−.044
1966	−.181	−.155	−.178	−.088	−.083	−.081	−.217	−.145	−.376	−.167	−.155	−.167
1967	−.186	−.137	−.333	−.087	−.077	−.164	−.223	−.123	−.521	−.168	−.146	−.322
1968	−.155	−.086	−.470	−.066	−.051	−.241	−.274	−.150	−.644	−.142	−.110	−.468
1969	−.120	−.038	−.556	−.043	−.025	−.292	−.245	−.092	−.794	−.099	−.059	−.578
1970	−.119	−.032	−.594	−.043	−.023	−.317	−.213	−.043	−.886	−.088	−.044	−.637

rental prices for 1965, the first full year of the tax cut:

	Without Tax Cut	With Tax Cut
Manufacturing Equipment	.296	.299
Manufacturing Structures	.237	.240
Non-Farm, Non-Manufacturing Equipment	.352	.355
Non-Farm, Non-Manufacturing Structures	.247	.250

Our estimates of the decrease in net investment, gross investment, and capital stock resulting from this change are given in Table 5. In general, the effects of the rate reduction are small and negative. It should be emphasized that these estimates are conditional on the level of output actually resulting from the tax cut; quite clearly the overall effect of the tax cut was to stimulate investment by increasing output. A second, little noticed change in tax policy in 1964 was the repeal of the Long Amendment; after repeal, the tax credit was no longer deducted from the depreciation base for tax purposes. This change raises the effective rate of the tax credit to almost 10 percent as compared with approximately 6 percent under the Long Amendment. The rental price of capital services for equipment in 1964 with and without repeal of the Long Amendment are:

	With Long Amendment	Without Long Amendment
Manufacturing Equipment	.302	.293
Non-Farm, Non-Manufacturing Equipment	.363	.352

Estimates of the increase in net investment, gross investment, and capital stock resulting from this change are given in Table 6.

These increases are quite substantial. The peak effect for manufacturing equipment took place in 1965 at which time the net investment in equipment attributable to the repeal was 10.4 percent of total net investment. In the non-farm, non-manufacturing sector, the peak effect in 1964 was over a billion dollars and accounted for 16.3 percent of net investment in equipment in that sector. Once again, a dip in the effect of this policy change can be seen in 1966–67 and one or two years after, resulting from the suspension of the investment credit. The lag structure in the non-manufacturing sector makes the dip much more noticeable there than in the manufacturing sector.

TABLE 6—CHANGES IN GROSS INVESTMENT (I), NET INVESTMENT (N), and CAPITAL STOCK (K) RESULTING FROM THE REPEAL OF THE LONG AMENDMENT (Billions of 1965 Dollars)

Year	Manufacturing Equipment			Non-Farm, Non-Manufacturing Equipment		
	I	N	K	I	N	K
1964	.238	.238	0	1.042	1.042	0
1965	.400	.365	.238	.958	.758	1.042
1966	.412	.329	.603	.706	.360	1.800
1967	.349	.217	.932	.750	.335	2.160
1968	.229	.067	1.149	1.021	.541	2.495
1969	.236	.064	1.216	.761	.177	3.036
1970	.297	.115	1.280	.792	.174	3.213

Table 7—Change in Gross Investment (I), Net Investment (N), and Capital Stock (K) Resulting from Suspension of the Investment Tax Credit for Equipment and Accelerated Depreciation for Structures from October 10, 1966 to March 8, 1967 (Billions of 1965 Dollars)

Year	Manufacturing						Non-Farm, Non-Manufacturing					
	Equipment			Structures			Equipment			Structures		
	I	N	K	I	N	K	I	N	K	I	N	K
1966	−.177	−.177	0	−.046	−.046	0	−.762	−.762	0	−.119	−.119	0
1967	−.271	−.245	−.177	−.089	−.086	−.046	−.599	−.452	−.762	−.200	−.192	−.119
1968	−.153	−.091	−.422	−.060	−.052	−.132	.069	.303	−1.214	−.126	−.104	−.311
1969	−.009	.066	−.513	.011	.012	−.184	.051	.226	−.911	−.011	.018	−.415
1970	.157	.223	−.447	.074	.075	−.172	.018	.150	−.685	.111	.139	−.397

Table 8—Change in Gross Investment (I), Net Investment (N), and Capital Stock (K) Resulting from Hypothetical Suspension of the Investment Tax Credit for Equipment and Accelerated Depreciation for Structures from October 10, 1966 to December 31, 1967 (Billions of 1965 Dollars)

Year	Manufacturing						Non-Farm, Non-Manufacturing					
	Equipment			Structures			Equipment			Structures		
	I	N	K	I	N	K	I	N	K	I	N	K
1966	−.177	−.177	0	−.046	−.046	0	−.762	−.762	0	−.119	−.119	0
1967	−.872	−.846	−.177	−.250	−.247	−.046	−3.190	−3.043	−.762	−.614	−.606	−.119
1968	−.567	−.416	−1.023	−.234	−.216	−.293	.208	.940	−3.805	−.475	−.425	−.725
1969	−.181	.031	−1.439	−.063	−.032	−.509	.171	.722	−2.865	−.154	−.074	−1.150
1970	.270	.477	−1.408	.117	.152	−.541	.093	.505	−2.143	.193	.278	−1.224

In 1966, an important objective of economic policy was to restrain investment. After a number of alternative changes in tax policy were considered and rejected,[25] the investment tax credit for equipment was suspended beginning October 10, 1966; at the same time, accelerated depreciation for structures was replaced by 150 percent declining balance depreciation. In the original legislation implementing these changes in tax policy, the suspension was to remain in effect until the end of 1967, a total period of almost fifteen months. The suspension was lifted on March 9, 1967, so that the total period of suspension was a little less than five months. The effect of the suspension on the annual rental price of capital in 1967 was the following:

	Without Suspension	With Suspension
Manufacturing Equipment	.320	.351
Manufacturing Structures	.259	.276
Non-Farm, Non-Manufacturing Equipment	.379	.414
Non-Farm, Non-Manufacturing Structures	.270	.287

Our estimates of the effects of the suspension on net investment, gross investment, and capital stock are given in Table 7.

For all categories of assets, the suspension had a restraining effect on the level of investment in 1967. We estimate that this effect continued into 1968 for all assets except non-farm, non-manufacturing equipment. For all classes of assets, the restoration of the original tax credit for equipment and accelerated depreciation for structures will result in a stimulus to investment in 1969 and 1970. For no class of assets is the level of capital stock as high at the end of 1970 as it would

[25] Policies under consideration during early 1966 and their potential impact on investment expenditures are discussed in our earlier study [11].

have been in the absence of the suspension. The total gross investment for the five year period 1966–70 is considerably lower than it would have been in the absence of the five month suspension.

If the suspension of the investment tax credit for equipment and accelerated depreciation for structures had continued for fifteen months, the impact on the level of investment would have been much more substantial. Our estimates are given in Table 8. For investment in structures the restraining effect of the suspension would have continued into 1969, although the impact would have been very slight in that year. For investment in equipment as well as for structures the magnitude of the impact would have been much greater. As a result the stimulus from restoration of the tax credit and accelerated depreciation would have been correspondingly increased.

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ALTERNATIVE INTEREST RATES AND THE DEMAND FOR MONEY: COMMENT

By HARVEY GALPER*

In a recent issue of this *Review*, T. H. Lee [3] examined the influence on the demand for money of the interest rate paid on savings and loan (*S* and *L*) shares. His regressions revealed a strong effect of the *S* and *L* rate on holdings of money. Furthermore, the *S* and *L* rate had a stronger influence than all other rates examined, and, on an annual basis at least, complete adjustment of the actual to the desired money stock was found when the *S* and *L* rate was used. These findings imply strong substitution effects between holdings of money and savings and loan shares and call into question Friedman's [5] notion of the money stock which assumes that time deposits at commercial banks are the closest substitute to money proper.

Lee's conclusions, however, are based on only fifteen annual observations, and, in several instances, he mentions that the lack of quarterly data prevents the development of a more thorough analysis.[1] By interpolating semiannual data from the Federal Home Loan Bank Board (*FHLBB*), I have been able to develop a quarterly series on savings and loan rates for the period 1956–66. This series has then been applied in regressions similar to Lee's to see if his results would be changed by the use of quarterly data. By and large his conclusions have been confirmed, although quarterly data allow a more precise investigation of the lag structure in adjusting actual to desired money holdings.

In the following sections of this note, I shall first present the technique for deriving the quarterly savings and loan rate, along with the resulting series, and then shall present the regression results based on the interpolated quarterly series.

I. *Derivation of a Quarterly Series on Interest Rates Paid on Savings and Loan Shares*[2]

For the six month periods ending in June and December of each year, 1956–65, the Federal Home Loan Bank Board has prepared data on the average interest rates paid by all insured savings and loan associations. These rates are the advertised rates of each institution weighted by the savings capital held by the institution at the end of

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[1] [3, for example, p. 1177 and fn. 11, p. 1178].

[2] This section, in particular, owes much to discussions with Frank de Leeuw.