An Economic Theory of Suicide

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> . . . as soon as the terrors of life reach the point at which they outweigh the terrors of death, a man will put an end to his life. [A. SCHOPENHAUER, *On Suicide*]

Although sociologists have developed and tested numerous theories about suicide, economists have not analyzed this phenomenon. We derive an economic theory of suicide and test its implications using: (1) data by age in many developed countries; (2) a time series, 1947–67, by age group in the United States; (3) a cross section by state and age group in 1960. Most of our predictions are verified. Particularly striking are the response to unemployment, as strong in the postwar period as in the Depression and stronger among older individuals, and the negative effect of increased permanent incomes among all but the youngest age group, a result found in both the time series and the cross section.

Since the appearance of Durkheim's *Le suicide* in 1897, sociologists have constructed numerous theories to explain patterns in suicide rates both within and across societies. Economists have not considered the problem of suicide, although it is surely one involving individual decision making

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and often also economic factors.[1] Just as economics has served to improve our understanding of fertility (cf. Becker 1960; Schultz 1969) and marriage (Silver 1965), so it may provide a testable theory of suicide. The empirical analysis should enable us to measure another effect on society of both long-run economic growth and cyclical fluctuations in income. In this paper we outline several sociological theories and then develop a new, economic theory of suicide. Some implications of this theory are then tested against various sets of data.

I. Sociological Theories and an Economic Theory

Most subsequent work by sociologists in the study of suicide has been motivated by Durkheim. In addition to providing a model for empirical research by sociologists, he summarized existing knowledge about suicide patterns by age, literacy, and other demographic and sociological variables and constructed a typology for classifying suicides. In a departure from the Durkheimian model, Henry and Short (1954) developed the frustration-aggression theory. According to them, suicide results from individuals' frustration in their attempts to achieve their social goals and the ensuing aggressive feelings, directed either at themselves (suicide) or at others (homicide). They postulate that economic improvement leads to a decrease in frustration and thus aggression, and they concentrate most of their empirical work on the relation of suicide rates to the business cycle. They also argue that frustration and aggression occur differentially depending upon the degree of external restraint imposed upon the individual. Since they assume that external restraint is inversely related to status, their theory implies a positive correlation between suicide rates and status. Their empirical work demonstrates conclusively only the positive correlation of suicides and unemployment in the United States before 1942.

Undoubtedly, much of the variation in suicide rates among ethnic and demographic groups can be explained by these and other sociological theories. Several aspects of this problem, particularly variations in suicides by age and income, that cannot simultaneously be reconciled by any one of these theories may well be rationalized by an economic theory. Admittedly, some suicidal behavior can in no way be attributed to economic factors. We contend, though, that some is due to economic

[1] In the earliest available records in Great Britain, covering 1770 through 1830, one-fourth of the recorded suicides were attributed to poverty (Fedden 1938, p. 203). Lendrum (1933) analyzes 419 male suicide attempts and finds that over one-third of those giving a motive cite economic problems. Shneidman and Farberow (1965, p. 36) analyze a sample of suicide notes and find 8 percent of those listing a specific cause mention poverty or business difficulties.

decision making, so that some of the variation in suicide rates should be explicable using hypotheses derived from an economic theory.[2]

Let us postulate the following utility function for the average individual in a group of people with permanent income YP:

$$U_m = U[C(m, YP) - K(m)] > 0, \tag{1}$$

where m is his age and K a technological relation describing the cost each period of maintaining himself alive at some minimum level of subsistence.[3] (Presumably $K' > 0$.) If this is the utility of the average individual age m with permanent income YP, then the present value of his expected lifetime utility at age a is:

$$Z(a, YP) = \int_a^{\omega} e^{-r(m-a)} U_m P(m)\, dm, \tag{2}$$

where r is the private discount rate, ω is the highest attainable age, and $P(m)$ is the probability of survival to age m given survival to age a. The Z, then, is a decreasing function of a and an increasing function of YP.

Let $b_i \sim N(0, \sigma^2)$ be the ith individual's taste for living, or conversely, his distaste for suicide. The variable b_i is in units commensurate with the utility function and is subject to any transformations performed on U_m.[4] It is assumed to be defined for the cohort at birth.

Following the spirit of the epigraph to this paper, we assume that an individual kills himself when the total discounted lifetime utility remaining to him reaches zero. This point must be a maximum, for the remaining lifetime utility is decreasing with age. Our hypothesis, therefore, is that the ith individual commits suicide if and when $Z_i(a, YP) + b_i = 0$. The fraction of individuals in the cohort born at time $(t - a)$ who commit suicide at age a is:

$$S(a) = f[-Z(a, YP)], \tag{3}$$

where $f(\cdot)$ is the density function for b_i.[5] In our model, the instantaneous suicide rate is simply the fraction of individuals in the cohort for whom

[2] This is not the first analysis of the economics of mortality. However, previous work neither formulates theories nor tests hypotheses. Hansen (1957) presents a calculation of the loss in GNP resulting from child mortality; Holtmann and Ridker (1965) compute the welfare loss arising from the early incurrence of burial costs of individuals who die prematurely. We calculate that the loss in human capital resulting from suicides among males ages 25–64 in 1959 was $740 million, based on a discount rate of 6 percent (*Census of Population* [1960], PC(2)-7B, and U.S. Office of Vital Statistics, *Vital Statistics* [1959]).

[3] Grossman (1972) has argued that this function can be affected by the individual's earlier choice of health expenditures.

[4] Our results are valid for any distribution which is positive over the entire real number line and has positive first and second derivatives at the left tail.

[5] We are explicitly assuming that no members of the birth cohort die from other causes before age a. Relaxing this assumption by dividing equation (3) by the fraction of the cohort actually remaining at age a changes none of our theoretical results on the relation of suicide to age and income.

$Z(a, YP)$ reaches $-b$ at age a. Differentiating equation (3) totally:

$$dS = -f'\left(\frac{\partial Z}{\partial a}\,da + \frac{\partial Z}{\partial YP}\,dYP\right). \tag{4}$$

Now $\partial Z/\partial a < 0$, both because $K' > 0$ and because there is less gross consumption left to enjoy as a person ages. Also, $\partial Z/\partial YP > 0$, so we can use equation (4) to derive the predictions that the suicide rate, S, will rise with age and decrease with permanent income.

Taking the total differential of equation (4), we obtain:

$$\begin{aligned} d^2S = {} & \left[-f'\,\frac{\partial^2 Z}{\partial a^2} + f''\left(\frac{\partial Z}{\partial a}\right)^2\right] da^2 \\ & + 2\left[-f'\,\frac{\partial^2 Z}{\partial YP\,\partial a} + f''\,\frac{\partial Z}{\partial YP}\cdot\frac{\partial Z}{\partial a}\right] dYP\,da \\ & + \left[-f'\,\frac{\partial^2 Z}{\partial YP^2} + f''\left(\frac{\partial Z}{\partial YP}\right)^2\right] dYP^2. \end{aligned} \tag{5}$$

Of the three terms, only the sign of the third can be determined unambiguously. It is positive, given the reasonable assumptions that we are dealing with the left tail of the normal distribution (so that f' and $f'' > 0$) and that the marginal utility of lifetime income decreases with increases in YP so that $\partial^2 Z/\partial YP^2 < 0$. The signs of the other terms are ambiguous. The first term in $\partial^2 S/\partial a^2$ is negative if Z is decreasing in a, and the second is positive. Similarly, the first term in $\partial^2 S/(\partial YP\,\partial a)$ is positive and the second negative.[6]

The age-specific suicide rates, then, will depend upon the distribution of tastes against suicide and the distribution of permanent incomes. Several specific hypotheses arise from the economic theory we have outlined. We expect that the rate of suicide increases with age ($\partial S/\partial a > 0$); that the suicide rate is inversely related to permanent income ($\partial S/\partial YP < 0$); but that the marginal absolute effect on suicide declines as permanent income increases ($\partial^2 S/\partial YP^2 > 0$).[7]

[6] The analysis could be expanded to include expected utility based on the probability of employment at each age. Unless severe restrictions are placed on differences by ages in the elasticities of expectations about future unemployment in response to changes in current unemployment, the only result is that suicide rates will rise as unemployment increases. In Section IV we examine the role of unemployment and, by inference, differences by age in these elasticities.

[7] In addition to the economic aspects of suicide discussed here, some demographic differences in suicide rates are well established. The rate is lower for females than males in all countries compiling reliable data, and in the United States the nonwhite suicide rate is slightly less than half that of whites (see U.S. Office of Vital Statistics, *Vital Statistics* [1967], II.a: 1.20). These differences are subsumed under the taste variable, b, in our theory, an assumption that allows us to concentrate on economic factors. A sociological theory that performs well in explaining these demographic differences is that of Gibbs and Martin (1964).

II. Empirical Evidence—Introductory Considerations

The deficiencies in the data on suicide rates are well known. (Douglas [1967] presents a detailed discussion of this problem.) Underreporting of suicide rates varies in different jurisdictions depending upon the persistence of the coroner and the degree to which society stigmatizes suicide as a form of deviant behavior. This difficulty implies that any cross-country comparisons of suicide rates are likely to be worthless; the underreporting problem is undoubtedly one reason for the low suicide rates observed in predominantly Catholic countries.[8] Within the United States, the use of suicide rates by states in a cross-section regression model could be dangerous because errors in the dependent variable are correlated with the true (observed) values of the income variables. Errors may also be introduced into any correlation of income and suicide by the requirement that an individual be covered for 2 years before an insurance company will pay off on a death by suicide (Gregg 1959, pp. 139–40).

While we admit that there are problems in using official statistics on suicide rates, a consideration of them suggests that the data should be useful if handled judiciously. Data on suicide rates by age within any country should embody very few problems; there is little reason to expect systematic differences in underreporting of suicides by age. The degree of underreporting may change over time as the moral stigma attached to suicide decreases. We can, however, account for this difficulty in Section IV and still make use of time-series data on suicide rates. Data from a narrow jurisdiction in which one official has charge of all decisions on causes of death should be fairly reliable. Finally, the use of the appropriate demographic measures should enable us in Section V to remove most of the errors arising from comparisons across jurisdictions within a fairly homogeneous society.

III. Suicide by Age—International Evidence

Table 1 presents the average male suicide rate by age in each of the 21 developed countries for which all the data are available from 1965 through 1967. In nine of these, the rate rises monotonically with age; in two it reaches a peak at the oldest age group although it does not rise steadily, and in eight others it rises monotonically to a peak in the age group 55–64. Only in two countries, Sweden and Poland, does the peak occur before this point, and in the former the rate is essentially constant after the peak. While the evidence presented in this table is by no means conclusive, it does suggest that within a country there is a general tendency

[8] This problem is vividly illustrated by the probably apocryphal quotation, cited by Alvarez (1972, p. 87), from an Irish coroner who certified an accidental death from gunshot wounds: "Sure, he was only cleaning the muzzle of the gun with his tongue."

TABLE 1
MALE SUICIDE RATES BY AGE (PER 100,000 POPULATION), 21 COUNTRIES, 1965–67 AVERAGE

	AGE					
COUNTRY	15–24	25–34	35–44	45–54	55–64	65–74
Australia	10.1	22.4	29.7	33.9	34.9	41.8
Austria	19.7	31.6	42.3	50.0	65.0	59.5
Belgium	7.8	13.0	17.8	29.5	42.1	54.9
Bulgaria	8.3	7.9	10.9	19.1	25.2	40.2
Canada	10.3	15.9	19.4	25.5	30.0	25.9
Czechoslovakia	26.0	35.9	43.6	50.2	52.7	67.3
Denmark	9.7	22.5	32.6	45.9	49.7	40.2
England and Wales	6.4	10.8	14.5	17.7	23.9	26.8
Finland	15.5	35.2	53.2	62.2	68.6	68.9
France	7.0	17.0	26.0	40.4	54.3	54.0
W. Germany	18.3	26.9	33.0	46.1	55.9	48.4
Hungary	30.5	44.1	56.0	65.2	70.7	76.4
Israel	4.8	10.2	11.4	16.4	18.2	23.0
Japan	14.7	21.0	15.5	20.9	35.2	52.1
Netherlands	4.1	7.5	8.4	13.9	18.8	24.5
New Zealand	6.7	11.0	19.4	29.0	29.9	28.3
Norway	6.2	10.5	15.6	19.5	20.6	19.8
Poland	13.0	22.6	29.2	30.4	28.5	24.4
Sweden	13.8	28.3	40.8	51.0	50.5	47.4
Switzerland	19.0	25.1	34.5	46.2	49.3	53.0
United States	9.9	17.3	22.4	28.4	36.0	35.7

SOURCE.—World Health Organization, *World Health Statistics Annual*, 1965, 1966, 1967 (Geneva: WHO, 1968–70).

for suicide rates to rise with age. Insofar as people born earlier have both lower incomes at any age and less income remaining, this evidence is consistent with our theory of the effect of income on suicide.

IV. Time Series—Males, United States, 1947–67

Data are available on the suicide rate by 5-year age groups for 1947–67, and the data on real incomes of persons exist for some closed 5-year and 10-year age intervals through age 64 in the United States.[9] These data form the basis for the empirical analysis in this section. The theoretical model we presented in Section I provides us with all but two of the

[9] The suicide data, numbers of suicides by 5-year age groups, are published in U.S. Office of Vital Statistics, *Vital Statistics* (1967) and earlier issues. These were transformed into rates per 100,000 population through the use of population data kindly made available to us by the National Center for Health Statistics. Data on money incomes of male persons by age for 1947–64 are available in Bureau of the Census, *Trends in the Income of Families and Persons in the United States, 1947–64* (Washington: Government Printing Office, 1967). For 1965–67 the data were provided by Charles Beach of Queen's University and are based on his calculations using published data. The money income series are deflated by the Consumer Price Index (CPI). The CPI series and data on unemployment rates by age group are from Bureau of Labor Statistics, *Handbook of Labor Statistics* (1971), Bulletin 1705, pp. 120 and 253. All the income data used are deflated to constant 1967 dollars.

variables to be used in the regression equations. In particular, for a cohort age A at time t, the suicide rate $S(A, t)$ will be a function of the discounted real permanent income stream remaining at time t and the age of the cohort. We also include a variable to measure transitory deviations around expected income. When unemployment rises, individuals' expectations of future incomes (and utilities) are revised downward. Holding real income of the employed constant, an increased number of people will believe future prospects to have diminished and will commit suicide.

We construct discounted permanent income remaining, $YPL(A, t)$ by assuming that individuals project the average growth rate of incomes by age group that occurred during 1947–67. We divide each 10-year age interval in two and assign each 5-year interval the value $Y(A, t)$, the mean real income at age A at time t, for the 10-year interval to which it belongs. Through a regression of this variable against time we construct $\bar{Y}(A, t)$, expected real income at age A at time t, which we project out to the end of the economic lifetime of each group.

We then compute $YPL(A, t)$ as:

$$YPL(A, t) = \bar{Y}(A, t) + \sum_{i=1}^{L} \frac{\bar{Y}(A + 5i, t + 5i)}{(1 + r)^{5i}} + \frac{\bar{Y}(N, t + N - A)}{(1 + r)^{N-A}},$$

where N is the mean age of males in the open interval 65+, and $L = (62.5 - A)/5$.[10] Age A is taken as the midpoint of the appropriate 5-year age interval. The series is computed for $r = .02, .06, .12, .20$, and ∞.

For transitory income we use the proxy $UN(A, t)$, the fraction of males age A who are unemployed at time t. This term is included alone and in an interaction with the age variable. There is evidence that periods of unemployment are longer among older workers who are laid off, and further evidence that older workers' expectations of finding new employment are affected by this phenomenon.[11] We thus expect unemployment to have an increasing effect on suicide as workers age.[12]

[10] The actual value of N in 1970 was 73.6. (Computed from Bureau of the Census, *Census of Population* [1970], General Population Characteristics, PC(1)-B1, table 50.) This method of computing permanent income remaining implicitly assumes that individuals are uncertain about their incomes in the distant future and weight them accordingly. The implication of this uncertainty is an optimal consumption path which is shaped like a monotonic function of $\bar{Y}$ as we have computed it (see Nagatani 1972 for a derivation of this optimization rule).

[11] Wilcock and Franke (1963, p. 61) find that only 40 percent of workers below age 45 failed to find jobs within one year after their layoffs; for workers 45 and over the figure was 60 percent. One older worker claimed, "It's little better than shooting me to throw me out of a job at 55" (p. 85).

[12] An additional transitory term, $Y(A, t) - \bar{Y}(A, t)$, was included in preliminary work. It was dropped because the t-statistic of its coefficient never exceeded .2 in absolute value.

Our basic estimating equation is:

$$S(A, t) = \alpha_0 + \alpha_1 YPL(A, t) + \alpha_2 YPL^2(A, t) + \alpha_3 A + \alpha_4 A^2 + \alpha_5 A^3 + \alpha_6 UN(A, t) + \alpha_7 UN(A, t) \cdot A + \nu_{A,t}, \tag{6}$$

where $\nu_{A,t}$ is a disturbance term, and income is measured in thousand-dollar units. The higher-order terms on A are included as a result of some experiments made on the international data in table 1.[13] We estimate equation (6) over a pooled cross section of the time series 1947–67 for each of the nine 5-year age intervals, 20–64. The open interval 65+ is excluded because most males in the group are not economically active. A modified version of this equation, denoted (6′), is also estimated and differs from equation (6) only by the addition of the term $\alpha_8 t$. This modification enables us to test the hypothesis that the apparent effect of increasing income over time on suicide rates merely reflects a trend in the latter. Finally, we search over the *YPL* series computed for each of the five discount rates and report results only for the value that maximizes the likelihood function in this parameter.

Examination of the residuals based on ordinary least-squares estimates of equations (6) and (6′) showed that $E(\nu_{A,t} \cdot \nu_{A,t-1}) \neq 0$ for each A. To remove this source of inefficiency in the estimation, we devised a two-round procedure that transformed the original variables by an estimate of the serial correlation of the residuals within each age group.[14] The results of this procedure are presented in table 2 for $r = .20$, the value of the discount rate that maximized the likelihood function for both equations (6) and (6′). This fairly high discount rate suggests that incomes in the near future are most important in affecting suicide. The R^2 presented is based on the second-round estimates.[15] Undoubtedly the most striking

[13] In 10 of the 20 countries in table 1 for which suicide rates generally increase with age, the suicide-age profile has a definite logistic shape. To reflect this phenomenon in equation (6), and thus to distinguish better between K and YP in testing our theory, we need the higher-order terms.

[14] No standard technique for removing serial correlation in pooled cross-section and time-series data is available. To circumvent this difficulty we employed the following method. We postulate a variance-covariance matrix of the residuals composed of nine 21×21 blocks on the prime diagonal:

$$\Omega = \sigma^2 \begin{pmatrix} 1 \rho \rho^2 & \cdots & \rho^{20} \\ \rho & \cdots & \\ \vdots & & \\ \rho^{20} & \cdots & 1 \end{pmatrix},$$

and blocks consisting entirely of zeros off the prime diagonal. The first-round estimate $\hat{\rho}$ is obtained, and each variable is transformed to produce $\Delta X_{A,t} = X_{A,t} - \hat{\rho} X_{A,t-1}$, $t = 1948, \ldots, 1967$. The second-round estimates are based on the ΔX variables for the sample 1948–67.

[15] There is no problem of nondiagonality of the variance-covariance matrix of the residuals resulting from correlation among the residuals within each age group. Only 2 percent of the total unexplained variance in equations (6) and (6′) is accounted for by this correlation.

TABLE 2
SECOND-ROUND ESTIMATES OF EQUATIONS (6) AND (6′), r = .20

Variable	(6)	(6′)
Constant	26.54	26.24
	(5.29)	(3.89)
YPL	−7.73	−7.78
	(−5.82)	(−4.10)
YPL^2	.56	.57
	(6.20)	(5.54)
A	−2.08	−2.09
	(−2.86)	(−2.69)
A^2	.065	.065
	(3.73)	(3.56)
A^3	−.0005	−.0005
	(−3.62)	(−3.47)
UN	−70.16	−69.68
	(−3.15)	(−3.12)
$UN \cdot A$	3.40	3.38
	(5.68)	(5.64)
t	...	.008
	...	(.06)
R^2	.93	.93
ρ	.517	.526

NOTE.—t-values in parentheses here and in table 3.

TABLE 3
MEAN VALUES OF DERIVATIVES AT DIFFERENT AGES, BASED ON EQUATION (6)

	GROUP		
DERIVATIVES	20–24	40–44	60–64
$\partial S/\partial UN$	6.23	74.14	142.04
	(.57)	(7.79)	(7.63)
$\partial S/\partial \bar{Y}$	2.17	−.29	−1.11
	(10.62)	(−1.62)	(−4.02)
$\partial^2 S/\partial \bar{Y}^2$	3.16	3.12	1.45
	(6.20)	(6.20)	(6.20)

feature of these results is the insignificance of the trend variable, t, in equation (6′); apparently the effects of income on the suicide rate depicted in equation (6) are not the result of the correlation between income and time. In both equations all the other variables are significant, and those for which we postulated an expected sign on the coefficient confirm our expectation. The estimates for the age terms suggest that in the United States there is a logistic shape curve relating suicide rates to age.

In table 3 we present various derivatives based on equation (6) and computed at the means for each age group listed. (We assume the means have no stochastic element and treat them as constant in computing variances of these derivatives.) Raising the fraction unemployed by .03, roughly equal to the average increase for males in postwar recessions, increases the suicide rate by amounts ranging from 0.19 per 100,000 for the youngest age interval to 4.26 per 100,000 for the 60–64 age group. In absolute numbers, we find that total male suicides per year in the age groups 20–64 rise by approximately 950 in an average recession, given the population distribution of the United States in 1960. Durkheim (1951) and others have demonstrated the relation between suicide and unemployment in periods characterized by violent swings in business activity. Our results indicate that this relation also holds for the relatively mild recessions of the postwar period and that it is stronger among higher age groups.[16]

Table 3 also presents the derivatives involving income that we discussed in Section I. The effect of a thousand-dollar increase in all real expected incomes on the suicide rate is negative for the median and oldest age groups, but positive for the 20–24-year-old group. We find that $\partial S/\partial \bar{Y} = 0$ at $A = 35$, so that we observe the predicted sign of this derivative in all but the three youngest age groups.[17] The predictions of our model on $\partial^2 S/\partial \bar{Y}^2$ are confirmed by the signs of this derivative.

As incomes have risen in the postwar period, the male suicide rate has fallen for most age groups, especially for the older groups nearing retirement age. Only for the youngest three age groups (20–34 years) has the suicide rate risen with rising incomes, a result that is contrary to our predictions. There is no simple economic explanation for this phenomenon, but the expansion in the number of people pursuing education into their twenties and the consequent postponement of consumption by an increased fraction of the people in these groups could be one cause. The decline in suicide rates for older people as their incomes have increased may be especially strong because of the decrease in variability of income resulting from the expansion of Social Security benefits. Indeed, there was a sharp drop in the relative suicide rates of groups over age 55 in the late 1930s coincident with the introduction of Social Security.

V. Cross-Section Evidence

In this section we present first a summary of studies on the relationship of suicide rates to occupation and income based on individual data and then the results of estimating the cross-section analogue to equation (6). Official published reports contain no information relating the suicide

[16] Our results imply an average peak-to-trough change in the male suicide rate for ages 20–64 of 2.07. Interestingly enough, the weighted suicide rate for this group rose in the Depression by 10.08, from a low of 28.59 in 1928 to a peak of 38.67 in 1932, while the fraction unemployed rose by .192. Our estimates of equation (6) imply a rise of 13.27 in this suicide rate in response to this change in the fraction unemployed, a result fairly close to that which occurred.

[17] Equations (6) and (6′) were reestimated for each age group separately, without the three age terms and for $r = .20$. In all nine groups the net effect of UN was positive; furthermore, $\partial S/\partial \bar{Y}$, computed at the mean for each age group and based on the coefficients for that group alone, was again positive for ages 20–34 and negative thereafter.

TABLE 4
RELATIVE SUICIDE RATES (100 = PROFESSIONAL RATE) AND FAMILY INCOMES IN 1959, BY OCCUPATION*

	Cook County (1959–63)†		New Orleans (1954–59)‡		Tulsa (1937–56)§	
Occupation	Males	Total Y	White Males	Whites' Y	White Males	Whites' Y
Professionals...	100	9,445	100	8,456		
					{100	8,522
Managers.....	107	10,487	174	8,669		
Clericals......	88	7,086	100	6,020		
					{33	6,782
Sales.........	134	8,843	102	6,861		
Craftsmen	141	8,142	137	6,262	41	6,201
Operatives	157	6,887	224	5,527	58	5,359
Service	314	6,045	502	4,689	68	4,108
Laborers......	342	6,026	481	4,276	110	4,316

* Computed from *Census of Population* (1960), PC(1)–15D, 20D, 38D, table 145, for the SMSA containing the city.
† Computed from Maris (1967, p. 249).
‡ Computed from Breed (1963, p. 181).
§ Computed from Powell (1958, p. 135).

rate to income or occupation. Several sociologists have, however, constructed detailed data by occupation for individuals within a narrow jurisdiction, and the results of these studies are summarized in table 4 along with census data on family incomes by occupation of the family head. In each study the suicide rate for professional workers is used as a base of 100, and the index for each occupation group is related to it. In all three studies presented, the suicide rate is generally lower among the occupation groups in which income is greater. As table 5 shows, the rank correlations of the suicide index with income are negative, although the small sample size makes it difficult to claim a high degree of significance for them individually. The correlations are especially strong when professionals and managers are deleted from the sample. The data presented in table 4 do not hold constant for religious practices; if there is a negative correlation between the percentage Catholic and occupational status within each narrow jurisdiction, the results presented in table 4 will be biased against observing the expected negative relation between suicide and occupational status. While there are relatively few Catholics in Tulsa, the fairly high percentage in New Orleans and Cook County suggests that the true pattern of suicide rates declines even more sharply with income. Age differences by occupation would impart a negative bias to the correlations. The correlations of the suicide index and median age by occupation in each city are all found to be nearly zero, suggesting that differences in age by occupation are not producing spurious correlations in table 5.

Dublin (1963), using data from England and Wales, presents results similar to those that Powell (1958) has produced for Tulsa. Sainsbury

TABLE 5

Rank Correlations between Suicide and Income Based on Table 4

	Cook County	New Orleans	Tulsa
All occupations	−.69*	−.63*	−.31
Excluding professionals and managers	−.83*	−.77*	−.90*

* Significantly different from zero at the 10% level, one-sided test.

(1955) finds constant suicide rates except in the highest and lowest social strata, where the rates are greater, but he, like Dublin (1963), fails to adjust for age differences. Summarizing these studies, we can conclude that, with the exception of professional workers and managers, a strong negative relation between income and suicide rates seems to be present in the data. The more general conclusion, that the suicide rate declines monotonically as permanent income increases, is supported somewhat less strongly by the available studies.

We turn now to the last test of our theory of suicide. Whereas the units of observation in Section IV are various age groups at different points in time, the units of observation here are age groups in different states of the United States at the same point in time, 1960. We restrict our analysis to the white male population and to all 38 states in the contiguous United States for which data on income by age and race are available.[18] The age and income variables are constructed similarly to those used in the time-series tests in Section IV. Five age groups—20–24 and the four 10-year groups between 25 and 64—form the basis for analysis. To avoid heteroskedasticity in the residuals, each observation is weighted by the square root of the white male population for that age and state.

Special problems in the cross-section test require the introduction of two new variables.[19] The first difficulty is that, *ceteris paribus*, suicide rates tend to be substantially higher in the Far West. The restless life often associated with this area may be a cause of this phenomenon. This consideration suggests the use of a dummy variable, assuming the value of unity for California, Oregon, and Washington. The second is that the distribution of Roman Catholics in the population is not random across states. Rather, Catholics tend to be concentrated in the heavily urban, industrial states where incomes are relatively high. Given the lower incidence of suicide among Catholics, a failure to control for their proportion in a state's population would bias the coefficient of income negatively. We might then spuriously confirm the implications of our

[18] The states excluded are Idaho, Maine, Montana, Nevada, New Hampshire, North Dakota, Rhode Island, Utah, Vermont, and Wyoming.

[19] These problems can be interpreted in the context of our theory as resulting from differences in the distributions of b across different cultures or societies.

theory for the effects of income on the suicide rate. To correct for this possibility, we experimented with several possible indices of the proportion of a state's population who are Catholics, including the fraction of foreign stock and the fraction with Spanish surname. The variable best justified theoretically and used here is *CATH*, the percentage of total elementary and secondary enrollment in Catholic parochial schools in each state. This proxy may be better than a measure of the percentage Catholic population, if, as in Ryder and Westoff's study of fertility (1971), Catholic education is the best indicator of the degree to which behavior conforms to Church dogma.[20]

Table 6 presents the estimates of

$$\begin{aligned} S(A,j) = \beta_0 + \beta_1 YPL(A,j) + \beta_2 YPL^2(A,j) + \beta_3 A + \beta_4 A^2 \\ + \beta_5 A^3 + \beta_6 CATH + \beta_7 WEST + \mu_{A,j}, \end{aligned} \tag{7}$$

where $S(A,j)$ is the suicide rate by age and state, and the *YPL* and A terms are as used in equations (6) and (6′).[21] Equation (7′) differs only in that the observations for the far western states and the dummy variable *WEST* have been deleted. Table 7 presents estimates of the derivatives $\partial S/\partial \bar{Y}$, based on equation (7), computed as in table 3.

The first thing to note about these tables is that, as in the time-series tests, the predictions of our theory on the age profile of suicide rates and the effect of income on suicide rates are borne out by the data. The sign patterns of $\partial S/\partial \bar{Y}$ by age in the cross-section and time-series tests are remarkably alike—negative for all but the youngest age groups. The only difference is for the 25–34 group, where the derivative is positive in the time series but negative here. The positive value of $\partial^2 S/\partial \bar{Y}^2$ here is contrary to our prediction in Section I. Given that the search procedure maximizes at $r = \infty$ and given the age-*YPL* profile resulting from this, the signs on *YPL* and YPL^2 are the only ones consistent with the pattern

[20] Numbers of white male suicides by state for 1960 are published in U.S. Office of Vital Statistics, *Vital Statistics* (1960), table 9-6. These were transformed into rates per 100,000 population using *Census of Population* (1960), data published in PC(1)-1B, table 59. Income data by age group for each state are published in PC(1)-D, table 134. Data on Catholic elementary enrollment in 1961–62 are from Office of Education, *Circular* no. 753, p. 34; those on secondary enrollment, available only for 1960–61, are from *Circular* no. 707, p. 27; and public school enrollment for 1961–62 is from *Circular* no. 757, p. 47.

[21] Initial estimates, using the same five values of r as in Section IV, showed that the likelihood function reaches a maximum at $r = \infty$. We then searched over a finer grid of $r = .35, .50, .75$, and 1.00 and found $r = \infty$ still to be the maximizing value, although the likelihood function is very flat for $r \geq .20$. No unemployment term is included in the estimates of equation (7) that we present because of our lack of confidence in the state unemployment data. Moreover, it is not clear that differences in unemployment rates across states at a point in time represent transitory components of income. When the measured unemployed rate, prorated by the national unemployment rate for the age group relative to the total national rate, was added to equation (7), its coefficient had a t-value of $-.6$ in equation (7); its addition produces only minor changes in the other coefficients.

TABLE 6
ESTIMATES OF EQUATIONS (7) AND (7′), $r = \infty$

Variable	(7)—38 States	(7′)—35 States
Constant	46.45	43.98
	(2.01)	(1.88)
YPL	10.97	12.84
	(3.35)	(3.69)
YPL^2	−1.33	−1.61
	(−3.86)	(−4.31)
A	−4.80	−4.88
	(−2.42)	(−2.43)
A^2	.14	.14
	(3.02)	(3.06)
A^3	−.001	−.001
	(−3.16)	(3.26)
CATH	−.43	−.39
	(−5.75)	(−5.24)
WEST	7.96	...
	(5.18)	...
R^2	.893	.867

NOTE.—*t*-values in parentheses here and in table 7.

TABLE 7
MEAN VALUES OF $\partial S/\partial \bar{Y}$ IN THE FIVE AGE GROUPS

Group	20–24	25–34	35–44	45–54	55–64
(7)	3.89	−2.02	−3.61	−2.51	−.33
	(2.50)	(−2.47)	(−3.64)	(−2.95)	(−.40)
(7′)	4.31	−2.68	−4.53	−3.21	−.57
	(2.65)	(−3.15)	(−4.33)	(−3.61)	(−.66)

of $\partial S/\partial \bar{Y}$ by age observed both here and in the time series. That the likelihood function is maximized at $r = \infty$ in the cross section is consistent with the observation that income differences across states are permanent. Therefore, current incomes are a good proxy for lifetime incomes here and may well perform better.[22]

The *WEST* dummy is positive and significant; its coefficient indicates that, once we adjust for income, age, and Catholic school enrollment, the suicide rate for males 20–64 averages eight more per 100,000 in the Far West. The proxy variable for the extent of Catholicism, *CATH*, has the expected negative coefficient. The percentage Catholic enrollment varies from 1 in North Carolina to 26 in Wisconsin; it implies a difference in the adult male suicide rate of 10 to 11 per 100,000 population, if other factors are held constant. Since the average rate in our sample is only 25 per 100,000, the effect of Catholicism is quite substantial.

[22] We are indebted to Jacob Mincer for this observation. The flatness of the likelihood function in the cross section for $r \geq .20$ suggests that there is little difference across states among these measures.

VI. Conclusion

In no sense do we claim that the individual agony resulting in suicide stems from what is solely an economic calculation; the majority of suicides can perhaps be explained on noneconomic grounds. Nonetheless, as Fedden (1938, p. 54) pointed out, poverty can be an important cause of suicide in developed societies where values are to a large extent based on the possession of material goods. Given the materialistic basis of modern society, it is reasonable to expect that variations in the suicide rate will be related to economic variables in ways predictable by economic theory.

Although our main purpose has been to demonstrate the value of applying economic theory to the problem of suicide, we have also discovered some interesting and apparently novel empirical results. Within the United States we have found that the mild cyclical decreases in economic activity since 1945 have produced the same proportionate increases in suicide rates that resulted from the stronger cyclical fluctuations before 1945. The suicide behavior of older people is significantly more sensitive to variations in unemployment than is that of younger people. Suicide rates decrease with increased incomes, both in the cross section and the time series, for all but the youngest age groups. Contrary to widely held beliefs (Shneidman and Farberow 1957, pp. 60–67) based on studies using inappropriate data or inadequate analysis and on what we call the "Richard Cory myth," our results imply that suicide rates are generally lower among higher-income groups.

Our results suggest the value of using economic theory to analyze social phenomena such as suicide. Like other theories, ours has successfully predicted the relation of suicide by age; however, ours also entails predictions about both the response of suicide to increased income and the change in that response as income increases that seem to be confirmed, with some exceptions, by the data. It is more convincing when one remembers the trend toward decreased underreporting of suicides that should bias our time-series results against confirmation of our theory. We believe the study demonstrates the need for work using data on individuals to separate out economic influences on patterns of suicide from noneconomic ones.

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