Discounting Environmental Health Risks: New Evidence and Policy Implications

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When environmental policy benefits accrue over a long time horizon, the valuation problem becomes more complex. Some analysts support the use of financial market interest rates for discounting deferred health benefits, while others argue against discounting altogether. We address the discounting issue by estimating implicit discount rates for deferred health benefits exhibited by workers in their choice of job risk. Wage responses to variation in life years at risk provide direct estimates of the discount rate that workers apply to their future utility. The estimated real discount rate equals approximately 2%, consistent with financial market rates for the period. © 1990 Academic Press, Inc.

I. DISCOUNTING HEALTH EFFECTS: THE POLICY CONTROVERSY

The health benefits of environmental policies often have quite different time patterns of incidence. In the case of pesticide policies, for example, acute health outcomes such as poisoning and skin burns occur immediately after some adverse event such as product misuse. Inappropriate pesticide disposal or pesticide runoff generate health effects much more remote in nature, as there may be a lag of several decades before long-term effects such as cancer become apparent.

The health benefits of environmental policies with long time horizons do not differ from other benefits in terms of the principles that should be applied in valuing them. Society’s willingness to pay for these benefits is the appropriate benefit concept. A fundamental difference arises, however, because this willingness-to-pay amount should reflect society’s rate of time preference with respect to health risks.

The practical consequences of the discounting procedure are not innocuous. For health benefits with a two-decade lag, $1 in benefits has a discounted value of $0.67 using a 2% real rate of interest and $0.15 using a 10% real rate of interest, the rate required by the U.S. Office of Management and Budget. Members of the U.S. Congress have advocated lower rates of discount or no discounting of health effects at all. The discount rate controversy is particularly acute due to the enormity of the stakes involved. At issue is the entire thrust of EPA policy, which at present is directed primarily at long-term health risks.

Many economic policy controversies, such as the debate over benefit–cost analysis and the value of life, involve issues for which there is a reasonably broad consensus in the economics literature. The rate of time preference issue is more complex in that the underlying economic foundations are more muddled. If capital


S-51
markets are perfect then the appropriate discounting policy is straightforward, as one would simply use society's riskless rate of time preference, and the appropriate discount rate for all benefit components would be the same. In the presence of capital market imperfections matters become more complex, particularly since health is not traded explicitly in an intertemporal market. Since health status is less freely transferable across time than money or economic commodities, there is no explicit intertemporal market to observe. Some participants in the rate of time preference debate have sought refuge in this complexity in an effort to cast doubt on the legitimate economic issues at stake.

Two potentially legitimate views are the following: one might simply use the same rate of time preference that one would use for other benefit components; or one could attempt to ascertain how rates of time preference for health impacts differ, if at all. The choice between these perspectives is not simply a theoretical issue. Rather, it must be resolved empirically. In particular, do rates of time preference for health differ from the interest rates for trading financial resources and, if so, what is the nature of the discrepancy? It is especially noteworthy that many advocates of abandoning the use of uniform discount rates for all benefit components typically have no empirical basis that indicates the direction of any discrepancy that might exist.

In this article we will report on a line of research in which we have attempted to obtain estimates of individual rates of time preference with respect to health risks. The research has utilized labor market data to assess the implicit rates of time preference that workers reveal through their willingness to incur risks on the job. Section II of the paper provides a brief summary of our earlier research and presents a new econometric formulation for the estimation of worker's rates of time preference. The econometric results discussed in Sections III and IV are in line with our earlier work and lie in a far more reasonable range than the estimates of implicit discount rates that have been obtained in other economic contexts. Overall, we find no evidence of a significant discrepancy between rates of discount for health risks and financial rates of time preference. To the extent that such a difference exists, implied discount rates for health appear somewhat greater than financial market rates. Procedures for applying these discount rates to health are not entirely straightforward, and in the concluding Section V we indicate how proper recognition of the income elasticity of demand for health affects the choice of the net discount rate for health effects.

II. MODELING DISCOUNT RATES FOR JOB RISKS

The empirical approach used here to estimate implicit rates of time preference with respect to health risks utilizes data on labor market decisions. In particular, workers exposed to job risks endanger lives of different duration and accept wage compensation in return for these risks. More specifically, what is at risk is a tradeoff between current utility and a risk to one's discounted expected utility stream and, at least in theory, one could use information based on these decisions to infer the implied rate of discount involved in workers' weighting of their foregone utility stream.

Earlier studies of job risks abstracted from the discount rate issue by analyzing only compensation per unit of death risk, thus ignoring both discounting and the
differing duration of human life at risk for workers at different ages. Although two studies introduced into the wage equation an interaction between death risks and age, this approach does not capture changes in life expectancy with age or the role of discounting the utility stream associated with life expectancy.

To address the inadequacy, we have developed three different approaches to obtaining estimates of the implicit discount rate for health risks. Each of these approaches involves a different mix of compromises involving the theoretical realism of the model and the assumptions that must be imposed to achieve estimability. No single modeling technique is dominant in terms of its theoretical and empirical properties. Our intent is to explore a diversity of approaches to structuring the implicit rate of discount issue and to assess the robustness of the results with respect to these variations.

The simplest empirical approach amends the standard wage equation to recognize the duration of life and the role of discounting. The standard hedonic wage equation provides the reference point for the estimation of discount rates:

\[ \text{Wage} = \alpha + \beta'X + \gamma (\text{Death Risk}) + \epsilon, \]  

(1)

where \( \alpha \) and \( \gamma \) are coefficients, \( \beta' \) is a coefficient vector, \( X \) is a variable vector of job and worker attributes, and \( \epsilon \) is a random error term. In Moore and Viscusi [7] we amend Eq. 1, replacing the death risk variable by the discounted expected number of life years lost during the job risk, or

\[ \text{Wage} = \alpha + \beta'X + \gamma (\text{Discounted Expected Life Years Lost}) + \epsilon. \]  

(2)

Using a nonlinear least squares procedure one can obtain estimates of workers' implicit real rate of discount used in calculating the Discounted Expected Life Years Lost, which was 10–12% for the 1977 University of Michigan Quality of Employment Survey data analyzed. This approach has the advantage of simplicity and the need to make few assumptions for the estimation, but the structure of the model has no formal theoretical basis. The variable capturing the Discounted Expected Life Years Lost can be viewed as an empirical proxy for the discount rate-related issues.

A second model we have developed, which is reported in Viscusi and Moore [15], lies at the opposite extreme. That paper develops a Markov model of lifetime job choice that recognizes the stochastic structure of one's discounted expected lifetime utility maximization problem. The functional form of the estimated wage equation is derived explicitly from the theoretical model. Although the resulting nonlinear structure is quite complex, the estimation is feasible if one imposes additional structure on the model, where the most important restriction is that one can only analyze specific functional forms for the utility functions. The empirical results, which were obtained using the 1982 University of Michigan Panel Study of Income Dynamics, yield an estimated real discount rate with respect to foregone expected utility of 11.0%. This result is quite similar to that obtained using the more direct Discounted Expected Life Years approach.

In this paper we will explore a third approach based on a life-cycle model of job risks. In particular, we will estimate the implicit discount rate based on the assumption that all life years are equally valued and that job risks serve to shorten the duration of life. This assumption is no more restrictive than that employed in the more ad hoc Discounted Expected Life Years approach, and it yields an explicit
functional form for the wage equation. Although the model does not recognize the stochastic, period-by-period nature of the potential loss in longevity, as in the case of the Markov model, it also does not require as many restrictive assumptions for estimation.

Our development of a conceptual model of the life-cycle job risk problem that yields estimable discount rate parameters begins with a specification of a lottery structure that follows our earlier work. We assume state-dependent, time-separable preferences \( U^j(x_j) \), where \( x_j \) equals income in state \( j \). For simplicity, we consider only two states. In state 1, the worker is healthy and earns a wage \( w \) that increases with the job risk \( p \). Utility in the healthy state, \( U^1(w(p)) \), exceeds utility in the injured state, which we normalize with no loss of generality to equal zero. An individual’s time horizon equals his expected remaining lifetime. This variable, \( T \), also depends upon job risk, with longevity a decreasing function of the risk \( (T_p < 0) \). Future utilities are discounted at a rate of time preference \( r \), so that the expected discounted lifetime utility of a worker with \( T \) years of life remaining who chooses a job with risk \( p \) equals

\[
V = \int_0^{T(p)} U^1(w(p)) e^{-rt} dt.
\]

We assume for simplicity that the job risk is constant over time, and that the worker works until death. The individual worker chooses the risk level \( p \) to solve the problem

\[
\text{Max}_{p} V = \frac{1}{r} U(w(p))(1 - e^{-rT(p)}).
\]

(3)

The first-order condition is

\[
V_p = -\frac{1}{r} (1 - e^{-rT(p)}) U_w \frac{\partial w}{\partial p} + \frac{\partial T}{\partial p} U e^{-rT(p)} = 0,
\]

which yields, upon rearrangement of terms,

\[
\frac{\partial w}{\partial p} = -r \frac{\partial T U}{\partial p U_w} (e^{rT(p)} - 1)^{-1}.
\]

(4)

Equation 4 describes the worker’s marginal rate of substitution between current period wages and the job risk. This rate of tradeoff depends upon the expected remaining life, \( T \), the discount rate, \( r \), and the effect of the risk on longevity, \( \partial T/\partial p \). Taking logarithms of each side of Eq. 4 gives

\[
\ln \frac{\partial w}{\partial p} = \alpha - rT + \varepsilon,
\]

(5)

\footnote{See Rosen [12] for a formal analysis of the life-cycle job risk problem.}
where the error term captures errors in the approximation

$$\ln(e^{rT(p)} - 1)^{-1} = -rT.$$  

These errors will be small for typical values of $r$ and $T$. For example, if $r = 0.1$ and $T = 35$, the approximation error is less than 1% of the true value. Estimation of Eq. 5 requires information on the implicit price of risk ($\partial w/\partial p$), longevity ($T$), and individual characteristics. The coefficient on the longevity term provides a direct estimate of the individual’s rate of time preference.

III. THE SAMPLE AND THE VARIABLES

Table I defines the key variables used in the analysis, and Table II presents their means and standard deviations. As in Viscusi and Moore [15], data for this study are from the University of Michigan Panel Study of Income Dynamics for the year 1982. This particular year corresponds to the time frame covered by the risk data used. Nonfarm household heads who did not suffer long-term unemployment during the sample period are included in the sample, and cases with missing data are excluded.

The basic dependent variable used in the analysis is the worker’s hourly wage. We use the wage rate along with estimated coefficients to compute the dependent variable in Eq. 5. The auxiliary control variables include the standard wage equation variables representing human capital effects due to years of schooling, job tenure, and experience; personal characteristic dummy variables indicating the worker’s race, sex, health status, and marital status; geographical indicators of city size and
TABLE II
Descriptive Statistics: Means and Standard Deviations

<table>
<thead>
<tr>
<th>Variable</th>
<th>Mean (Standard Deviation)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Wage</td>
<td>6.92 (2.42)</td>
</tr>
<tr>
<td>Implicit price (\ln(\frac{\partial w}{\partial p}))</td>
<td>(8.7E - 3) ((4.4E - 3))</td>
</tr>
<tr>
<td>Number of dependents</td>
<td>1.01 (1.16)</td>
</tr>
<tr>
<td>Education</td>
<td>12.90 (2.51)</td>
</tr>
<tr>
<td>Black</td>
<td>0.08 (0.27)</td>
</tr>
<tr>
<td>Female</td>
<td>0.17 (0.38)</td>
</tr>
<tr>
<td>Job tenure</td>
<td>5.07 (6.27)</td>
</tr>
<tr>
<td>Experience</td>
<td>11.96 (10.56)</td>
</tr>
<tr>
<td>Union status</td>
<td>0.29 (0.46)</td>
</tr>
<tr>
<td>Job risk (per 100,000)</td>
<td>7.83 (9.66)</td>
</tr>
<tr>
<td>Mortality risk (per 100,000)</td>
<td>2271.6 (2687.6)</td>
</tr>
<tr>
<td>Northeast region</td>
<td>0.21 (0.41)</td>
</tr>
<tr>
<td>Southeast region</td>
<td>0.31 (0.46)</td>
</tr>
<tr>
<td>North Central region</td>
<td>0.29 (0.45)</td>
</tr>
<tr>
<td>West region</td>
<td>0.19 (0.39)</td>
</tr>
<tr>
<td>Sample size</td>
<td>1463</td>
</tr>
</tbody>
</table>

region (regional dummy variables); and an occupational dummy variable indicating whether the worker is employed in a blue-collar occupation.

The pivotal variable in this study is the riskiness of the worker's job. We use as our fatality risk measure a newly developed set of data on job risks generated by the National Traumatic Occupational Fatality Survey (NTOF), an on-going census of job fatalities conducted by the National Institute for Occupational Safety and Health. These data, which have been analyzed in related contexts by Moore and Viscusi [8, 9] and Viscusi and Moore [15], represent a considerable improvement over data used previously in the analysis of the labor market implications of job risks for two reasons. First, the NTOF data are a census of all industrial fatalities, and therefore do not contain any sampling error, unlike all other publicly available industry risk data. Second, since the NTOF data represent a 5-year average of fatality risks, they are not subject to any error due to transitory fluctuations in job risks that might result from a major catastrophe. The role of random fluctuations in accidents for narrowly defined industry cells will consequently be much less than if the data were available for only a single year.
IV. EMPIRICAL RESULTS

The theoretical model developed in Section II derived an expression for the individual worker's marginal rate of substitution between wages and job risks as a function of his remaining life and other control variables. A number of estimation issues arise in the empirical implementation of this model. These issues include the endogeneity of the remaining life term, T, which in our model depends on the chosen risk level, and the computation of the implicit price variable, which equals the marginal rate of substitution in equilibrium.

Our solution to both of these problems uses recently developed approaches to the estimation of structural hedonic equilibrium models.3 In particular, we construct implicit prices from a first-stage regression of wages on a vector of individual characteristics and on a nonlinear function of the job risk. We then use the estimated coefficients on the job risk variable to construct measures of the implicit price, \( \partial w / \partial p \). This implicit price variable is then used as the dependent variable to estimate the parameters of Eq. 5. To control for the potential endogeneity of the longevity term, we use instruments that are formally derived from the structural model.4 These instruments include interactions of regional dummy variables with all of the exogenous characteristics and the regional dummies themselves. We present two-stage least squares and ordinary least squares estimates of the model to illustrate the sensitivity of the estimated discount rate to the treatment of the longevity variable.

The wage equation describes the opportunities available to workers in an implicit market for job risks. Workers choose from available combinations of wages and risks according to their preferences for risk, as described in Viscusi [13] and Moore Viscusi [10]. The position of the market opportunities locus is fixed by the characteristics of the worker and the job (X), which include variables such as race and sex that reflect the effects of possible market discrimination, and the economic factors such as education, experience, and job tenure that reflect accumulated human capital. Also included in X are location variables such as city size and job characteristic variables for collective bargaining status and occupation.

In addition to the control variables, the wage equation also includes measures of the job risk, p. The risk variable is entered as a quadratic term, with both the linear and squared job risks interacted with one of the four regional dummy variables. Thus, variation in the implicit price of risk arises due to differences in region and in the risk level.

The functional form of the wage equation is

\[
\ln w_i = X_i \beta + \sum_{k=1}^{4} \left( \gamma_1 R_{ik} p_i + \gamma_2 R_{ik} p_i^2 \right) + \nu_i, \tag{6}
\]

where \( R_i \) is a dummy indicator of the region of residence of worker \( i \). The estimated

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4See Kahn and Land [6].
implicit price of the worker then equals

\[
\frac{\partial w_i}{\partial p_i} = (\gamma_1 R_{ik} + 2\gamma_2 R_{ik} p_i) w_i.
\]

Thus, the implicit price of risk for each worker depends on his wage, risk level, and region of residence.

Table III presents estimates of the market wage equation. The coefficients on the

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Coefficient (Standard Error)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Death risk × Northeast</td>
<td>0.025** (0.008)</td>
</tr>
<tr>
<td>(Death risk × Northeast)^2</td>
<td>-9.4E - 4* (4.7E - 4)</td>
</tr>
<tr>
<td>Death risk × North Central</td>
<td>0.023** (0.005)</td>
</tr>
<tr>
<td>(Death risk × North Central)^2</td>
<td>-7.6E - 4** (1.9E - 4)</td>
</tr>
<tr>
<td>Death risk × South</td>
<td>0.004* (0.002)</td>
</tr>
<tr>
<td>(Death risk × South)^2</td>
<td>-0.2E - 4 (0.4E - 4)</td>
</tr>
<tr>
<td>Death risk × West</td>
<td>0.021** (0.004)</td>
</tr>
<tr>
<td>(Death risk × West)^2</td>
<td>-5.4E - 4** (1.6E - 4)</td>
</tr>
<tr>
<td>Race</td>
<td>-0.049* (0.029)</td>
</tr>
<tr>
<td>Sex</td>
<td>-0.191** (0.023)</td>
</tr>
<tr>
<td>Education</td>
<td>0.039** (0.004)</td>
</tr>
<tr>
<td>Number of Children</td>
<td>0.033** (0.007)</td>
</tr>
<tr>
<td>Job tenure</td>
<td>0.007* (0.003)</td>
</tr>
<tr>
<td>(Job tenure)^2</td>
<td>-0.9E - 4 (0.1E - 4)</td>
</tr>
<tr>
<td>Experience</td>
<td>0.023** (0.003)</td>
</tr>
<tr>
<td>Experience^2</td>
<td>-4.7E - 4** (0.8E - 4)</td>
</tr>
<tr>
<td>Union status</td>
<td>0.172** (0.018)</td>
</tr>
<tr>
<td>Adjusted R^2</td>
<td>0.413</td>
</tr>
</tbody>
</table>

**Statistically significant at the 0.01 confidence level (one-tailed test).
*Statistically significant at the 0.05 confidence level (two-tailed test).
control variables in the estimated wage equation accord with a priori expectations. Wages increase with education, job tenure, and union status, and fall for workers who are black or female. Furthermore, the summary statistics for the estimated wage equation are consistent with those estimated elsewhere in the literature, as the model explains about 40% of the variation in individual wages.

The job risk variables perform in the expected manner, with one minor exception. In general, significant compensating wage differentials are paid to workers for increases in job risks. These differentials decrease significantly with the risk level, except in the South. However, the linear and squared risk coefficients are always jointly significant within region.

The focus of the empirical estimates is the estimation of the discount rate in the model given by Eq. 5, which is produced here for convenience:

\[
\ln \frac{\partial w_i}{\partial p_i} = \alpha - rT_i + \varepsilon_i.
\]

The term \( \alpha_i \) in Eq. 5 is a function of the discount rate, the longevity–risk tradeoff, and the utility function of the individual worker. We assume that this term can be approximated by the vector of variables \( Z \) that includes proxies for differences in tastes such as race, sex, and education. We also include occupational dummy variables in \( Z \) to capture differences in tastes reflected by occupational choice.

One problem that arises in the estimation of the implicit price equation is that observations with a negative implicit price are lost in the logarithmic transformation. To circumvent this problem, we use the wage–risk tradeoff, \( \partial w/\partial p \), as our dependent variable, and transform the coefficient on the longevity term by dividing by the implicit price after estimation to allow interpretation of the coefficient as the discount rate. We also estimated the semilogarithmic implicit price equation using the positive price observations. The estimated discount rate was not sensitive to this restriction.

The empirical model is thus

\[
\frac{\partial w_i}{\partial p_i} = Z_i \delta - r^* T_i + \varepsilon_i,
\]

and \( r \) is estimated using the expression

\[
r = \frac{\partial \ln(\partial w/\partial p)}{\partial T} = \frac{1}{\partial w/\partial p} \frac{\partial (\partial w/\partial p)}{\partial T} = \frac{1}{\partial w/\partial p} r^*.
\]

We estimate Eq. 7 using both ordinary and two-stage least squares. Table IV presents the results of this estimation.

The Table IV results provide strong support for our model of intertemporal choice. Individual rates of time preference, as measured by the scaled coefficient on the life-years lost variable, differ significantly from zero at then most stringent significance levels. The low estimated real discount rates, which equal 1.6 (OLS) and 2.0% (2SLS), also indicate that individuals are not myopic with respect to the valuation of their future health risks. In fact, the estimated real rates accord roughly with financial market interest rates for the period, once these nominal rates are adjusted for inflation.
TABLE IV
Implicit Price Equation Estimates

<table>
<thead>
<tr>
<th></th>
<th>OLS</th>
<th>2SLS</th>
</tr>
</thead>
<tbody>
<tr>
<td>Life years lost (−r)</td>
<td>−0.016**</td>
<td>−0.020**</td>
</tr>
<tr>
<td>(0.003)</td>
<td></td>
<td>(0.003)</td>
</tr>
<tr>
<td>Race</td>
<td>−0.018**</td>
<td>−0.019**</td>
</tr>
<tr>
<td>(0.007)</td>
<td></td>
<td>(0.007)</td>
</tr>
<tr>
<td>Sex</td>
<td>−0.007</td>
<td>−0.006</td>
</tr>
<tr>
<td>(0.006)</td>
<td></td>
<td>(0.006)</td>
</tr>
<tr>
<td>Education</td>
<td>0.005**</td>
<td>0.005**</td>
</tr>
<tr>
<td>(0.001)</td>
<td></td>
<td>(0.001)</td>
</tr>
<tr>
<td>Constant</td>
<td>0.060**</td>
<td>0.064**</td>
</tr>
<tr>
<td>(0.015)</td>
<td></td>
<td>(0.015)</td>
</tr>
<tr>
<td>Adjusted $R^2$</td>
<td>0.106</td>
<td>0.109</td>
</tr>
</tbody>
</table>

**Statistically significant at the 0.01 confidence level (one-tailed test).

Equation 7 is essentially a linear demand curve for longevity, where the implicit price of longevity is a decreasing function of the quantity demanded. Calculation of the value of longevity using a willingness to pay approach is accomplished by summing the area under this demand curve, as shown in Fig. 1. The willingness to pay for longevity $T_0$ equals

\[
V(T_0) = T_0 \left( \frac{\partial w}{\partial \rho} \right)(T_0) + 0.5T_0 \left( \frac{\partial w}{\partial \rho} \right)(0) - \left( \frac{\partial w}{\partial \rho} \right)(T_0)
\]

where, e.g., $(\partial w/\partial \rho X_0)$ equals the implicit price evaluated at $T = X$. Using the coefficients in Table 4 and the mean values of the explanatory variables, the implicit price when $T = 0$, $(\partial w/\partial \rho)(0)$, equals 0.127. Then, for example, a worker is willing to sacrifice $V(T_0) = 35(0.079) + (17.5)(0.127 - 0.079) = $3.61 in hourly wages for a risk exposure of 35 additional years of life. In terms of annual wage premiums, the

![Fig. 1. Demand curve for longevity.](image-url)
same worker would accept annual compensation with a present value of approximately $7210 for putting 35 years of longevity at risk. On a per year present value basis, this translates into approximately $2030. Two features of this calculation are striking. First, in terms of absolute magnitude, this willingness to pay is fairly substantial, as it constitutes about 2% of annual earnings. Second, it accords roughly with many of the value of life calculations from more traditional analyses in the literature. This last result implies a degree of robustness to our results, thereby strengthening our confidence in the estimated discount rate.

V. POLICY IMPLICATIONS

Although the 2% real rate of interest with respect to health risks is below the 10–12% range in our earlier studies, it is more in line with the real return to capital in the U.S. economy. Moreover, given the standard errors around the earlier discount rate estimate, the results are quite similar in terms of overall discount rate range implied by the results.

It should also be noted that within the context of the implicit discount rate literature, our results are very stable. In a study of consumers’ valuation of the energy efficiency of refrigerators, Hausman [5] found that consumers exhibited discount rates of 20% or more. In a sequel to this study by Gately [4], consumers were found to exhibit discount rates between 45 and 300%. An interview study of consumers’ financial rates of time preference by Fuchs [2] yielded discount rates with an average of 30%. Even corporate executives, who might be viewed as an informed respondent group, report financial discount rates of 15%. In view of the imprecision of past studies of financial discount rates, our estimated discount rates lie in a fairly narrow range.

One should also be cognizant of the ultimate objective of our study, which is to ascertain whether systematic differences exist between rates of time preference for health and financial rates of return. In each case the confidence intervals for the discount rate estimates overlap available market rates of return. Moreover, since the point estimate of the discount rate falls short of the market rate in one case and exceeds the market rate in two cases, we find no clear evidence of systematic differences between discount rates for health and financial rates of time preference.

Our results consequently provide no empirical support for utilizing a separate rate of discount for the health benefits of environmental policies. Employing such discount rates need not, however, have the sweeping effects that many observers fear. Suppose that the discount rate is $r$, the rate of real income growth is $g$, the implicit value of the health outcome is $V$, and that a lag of $t$ years precedes the health effect. If the income elasticity of the value of the health equals 1, as Viscusi and Evans [14] suggest, then the discounted value per health impact prevented equals

$$
\frac{V(1 + g)^t}{(1 + r)^t} \approx V(1 + g - r)^t.
$$

---

For rates of real income growth close to the rate of discount, discounting has little impact when earnings growth is considered simultaneously. The basic theme of our results is that proper discounting of health effects will not greatly diminish the attractiveness of environmental health policies with deferred effects. Discounting is appropriate, just as it should be for any deferred benefit component. Moreover, our empirical results do not provide any justification for utilizing a separate rate of discount for health effects. However, recognition of the positive income elasticity of the valuation of the health effects offsets much of the discounting effect.

REFERENCES


See Viscusi and Evans [14].