Marine Resource Economics, Volume 13, 197–213
0738-1360/98 $3.00 + .00

Fishing Behavior and the Length of the Fishing Season

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Abstract *The basic hypothesis of this paper is that the amount of fishing that a fish harvester undertakes during a year is not determined entirely by circumstances which are exogenous to the fisher, such as weather conditions and resource availability, but is also partially a matter of individual choice. The paper develops a behavioral model of fishing from the perspective that the decision to modify the period of time over which fishing takes place is governed by a comparison of the marginal benefits and costs of doing so. The model is tested econometrically as an error-components model using a 10% longitudinal sample of recipients of seasonal fishermen's unemployment insurance benefits in Newfoundland over the period 1971–93. The results suggest that the Canadian unemployment insurance program has reduced the length of the fishing season in Newfoundland by about 8–10 weeks.*

Key words Error-components model, fishing season, unemployment insurance.

Introduction

The basic hypothesis of this paper is that the amount of fishing that a fish harvester undertakes during a year is not determined entirely by circumstances which are exogenous to the fisher, such as weather conditions and resource availability, but is also partially a matter of individual choice. As such, the decision can be analyzed with the usual apparatus that economists apply to choice decisions at the margin, from the perspective that the decision to modify the period of time over which fishing takes place is governed by a comparison of the marginal benefits and costs of doing so. Changes in these marginal benefits and costs, through institutional modifications or otherwise, can alter the balance between these two, and so lead to a change in the decision as to how long during the year to pursue the fishing activity.

The marginal benefits and costs of fishing an additional week are an individual matter that depends on both the productivity and the preferences of the individual fisher. Therefore, testing the basic hypothesis should also be done at the individual level. The existence of a social insurance program in Canada, to which fishers have access, provides us with the opportunity to engage in such testing.

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The analysis underlying this paper was supported by a research grant from the Institute of Social and Economic Research, Memorial University; a Donner research fellowship from the Canadian-American Center, University of Maine; a Science Subvention from the Department of Fisheries and Oceans and the Natural Sciences and Engineering Research Council of Canada; and a grant from the Social Sciences and Humanities Research Council of Canada. We wish to thank the Department of Employment and Immigration of the Government of Canada for providing us with the data used in this paper, and two anonymous referees for useful comments.

The Canadian unemployment insurance program was established in order to enable the government to insure employees against the consequences of job loss. It has been a fixture of the Canadian social welfare system since 1940. The program was governed by an Unemployment Insurance Act[1] which was enacted in 1971 and only recently (May 1996) supplanted by new legislation (which has been renamed the Employment Insurance Act).

Normally, an unemployment insurance claimant must have been engaged in an employment contract in order to qualify for benefits. Self-employed persons normally do not qualify for coverage. The one exception to this general prohibition is provided in section 130 of the Act, which enables the Canada Employment and Immigration Commission to operate a scheme of unemployment insurance for "self-employed persons engaged in fishing." Such a scheme has been in existence since 1956. The program provides fishers with benefits during their off-season, in amounts dependent upon earnings during the fishing season. The Special Seasonal Fishermen's Benefits Program has become a fundamental support mechanism for inshore fisheries on the Atlantic coast of Canada.[2]

The Unemployment Insurance Act that governed the program was approved by the Parliament of Canada in 1971, replacing less generous legislation. After an initial period of stability, perceived difficulties with the Act led to a series of amendments between 1976 and 1980. Other than amendments passed in 1983 in response to recommendations by the Kirby Task Force on the Atlantic Fisheries, the Act remained essentially unchanged until 1996 insofar as inshore fishermen in Newfoundland, the focus of our analysis, were concerned.

The existence of the program enables us to test hypotheses on fishers' decisions about the duration of their fishing activity during the year. First, the program provides us with a longitudinal panel of data covering the individual earnings of fishers, the number of weeks worked during the fishing season, and the unemployment insurance benefits received as a result. Second, the program structure creates considerable variation in the incentives provided to fishermen, and does so both longitudinally and across individuals. The reason for this variation is that the benefit entitlement of fishers varies considerably from case to case, depending on individual circumstances. It also varies from year to year as the rules governing the determination of benefits change. This contrast in incentives enables us to infer matching contrasts in behavior.

The Unemployment Insurance Act could potentially have influenced the behavior of inshore fishers in three ways: (*i*) by altering the attractiveness of fishing relative to other forms of economic activity, unemployment insurance can affect the number of people engaging in fishing as a full-time or part-time occupation; (*ii*) by modifying the returns to fishing, it can change the length of time that fishers engage in this activity during the year; and (*iii*) it can also change the intensity with which fishers are prepared to fish at any particular time.

Changes of the first kind have been analyzed by Ferris and Plourde (1980, 1982), who concluded, on the basis of aggregate data over the period 1956–68, that the presence of unemployment insurance in this period accounted for one-half of the inshore fishing boats in Newfoundland (Ferris and Plourde 1980, p. 116). Since the database used in this paper contains no information on alternative employment opportunities, we can provide no additional insight into this

[1] *An Act Respecting Unemployment Insurance in Canada,* R.S.C. 1985, C. U-1.

[2] The historical background underlying this anomaly is discussed in considerable detail in Schrank (1996).

particular question. The paper instead focuses on changes of the second kind, which are changes in the length of time spent fishing during the year as a result of the unemployment insurance program. Such changes have been the subject of early speculations by Copes (1972, p. 69) and formal theoretical modeling by Ferris and Plourde (1980, 1982). It has, however, appeared to have escaped detailed empirical analysis.

The basic structure of the Canadian unemployment insurance program as it affects inshore fishermen is outlined in the next section of the paper. Then, the paper briefly outlines a behavioral model of the decision to fish in a particular week within the season.[3] The next section presents econometric estimation of the relationship between fishing earnings and fishing weeks, while the subsequent section extends the model to incorporate the decision by individual fishermen as to how many weeks in the year to engage in fishing. Some limitations and possible extensions of the analysis are discussed in the final section of the paper.

The Canadian Unemployment Insurance Program

In the period under analysis, the Canadian Unemployment Insurance program operated in the following manner. Employees earning income in excess of a predefined minimum in a week are deemed to have insurable earnings in that week. Both the employee and the employer then contribute premiums at a given rate to an Unemployment Insurance Account. If the employee works a sufficient number of insured weeks, then upon an interruption of earnings, he may, after a two-week waiting period, obtain weekly unemployment insurance benefits equal to a percentage of the average weekly insured earnings received during the qualifying period. The level of weekly earnings that is insurable is subject to a ceiling, which limits the level of both the premiums that must be contributed and the benefits that can be received.

The length of the period over which benefits can be received depends on the number of insured weeks in the qualifying period, and on the national and regional rates of unemployment. Earnings received during the benefit period may be kept if they are less than 25% of the weekly benefit; earnings in excess of this amount result in a dollar-for-dollar reduction in benefits.

Most inshore fisheries in Atlantic Canada are organized on a coadventurer rather than on an employer-employee basis. In the coadventurer system, the boat owner receives a predefined share of the value of the catch net of operating costs. The remainder is shared evenly among the crew of the vessel. The structure of the unemployment insurance program is not well suited to this kind of arrangement. It is not obvious who should be considered the employer, how insured earnings should be defined, and when an interruption of earnings is deemed to have taken place.

Usually, the fish buyer is deemed to be the employer (Regulations, s.76).[4] The insured earnings of a crewman consist of the crewman's share. For the boat owner (or lessee), insured earnings are deemed to be the net value of the catch after deducting (*a*) crewmen's shares and (*b*) 25% of the value of the catch to account for operating expenses. If the boat owner's earnings fall short of the minimum level of earnings required for a week's earnings to be insurable under the Act, these earnings

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[3] The theoretical model is discussed only briefly here, since it has been presented in greater detail in Roy *et al.* (1994).

[4] *Regulations Respecting Unemployment Insurance*, C.R.C. 1978, C. 1576.

are deemed to be at that minimum level (Regulations, s.78). Thus, even a minimal level of fishing activity qualifies a boat owner (but not his crewmen) for unemployment insurance.[5]

Fishermen are categorized into year-round and seasonal for unemployment insurance purposes. The requirements for classification as a year-round fisherman are extremely stringent (Regulations, s. 84), and almost all inshore fishermen in Newfoundland are classified as seasonal. A seasonal fisherman can receive benefits only during the off season, which for most fishermen is the period between November and May [Regulations, s. 85(7)]. For this reason, potential claimants regard fishing benefits as inferior to benefits from regular employment, which can be taken at any time in the year, and usually for a longer period of time. However, to qualify for regular benefits, regular employment must be maintained for a minimum number of weeks during the qualifying period. For the most part, fishermen in Newfoundland have not been able to maintain sufficient regular employment to qualify for regular benefits.

A Model of the Length of the Fishing Season

Fishing in Newfoundland is a seasonal occupation, and boat owners, if they are rational, will decide whether to fish in a given week on the basis of a comparison of the marginal benefits and costs of doing so. We model this decision-making process on the basis of a theoretical analysis similar to that formulated by Ferris and Plourde (1980, 1982).[6]

The model is based on the supposition that fishing income varies from week to week during the year, primarily because of changes in resource availability (although ice and weather conditions are also factors).[7] Fishing income can also vary over the season because of changes in prices and costs. However, these changes are not a significant source of intraseason income variation in Newfoundland. Ex-vessel fish prices are set at the beginning of the season through a process of collective bargaining between the processors' association and the fishermen's union. While these prices are technically minimum prices, in practice market prices rarely rise above the negotiated levels (Department of Fisheries and Oceans 1989, 65 ff.). There is no evidence of significant intraseasonal variation in harvesting costs. As a result, fishing income is subject to diminishing returns as the fishing season is extended. This

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[5] Fishermen are permitted to arrange their affairs with buyers in such a way as to accumulate their catches over more than one week, and to average the accrued value over that number of weeks [*Regulations*, 79(5)]. As a result, earnings in weeks during which catches are high can be applied to weeks in which earnings are lower. This procedure enables fishermen to obtain increased benefits from weeks during which earnings exceed the ceiling, and to include as an insured week one in which earnings would otherwise be considered below the minimum level.

[6] Ferris and Plourde build upon substantial literature examining the effect of tax and social security considerations on labor supply. See Rea (1977), Hannoch and Hoenig (1978), and Wales and Woodland (1979). This literature depends on an empirical separation of income and substitution effects on the supply of labor, which seems to have had its earliest expression in Mincer (1962), although the idea goes back to Robbins (1930) and Jevons (1888). This literature is surveyed in Hausman (1985) and Pencavel (1986).

[7] The relative importance of these factors is species-specific. Some species (*e.g.*, cod, caplin) show seasonal migration patterns which leave them vulnerable to inshore gear for only a part of the year. Other species (*e.g.*, lobster, crab) are more in the nature of pure depletion fisheries.

Fishing income can also vary over the season because of changes in prices and costs. However, these changes are not a significant source of intraseason income variation in Newfoundland. Ex-vessel fish prices are set at the beginning of the season through a process of collective bargaining between the processors' association and the fishermen's union. While these prices are technically minimun prices, in practice, market prices rarely rise above the negotiated levels (Department of Fisheries and Oceans 1989, 65 ff.). There is no evidence of significant intraseasonal variation in harvesting costs.

relationship between fishing income and fishing weeks can be represented as a concave function[8] $f = f(L)$, where $f' > 0$, $f'' < 0$. It is further assumed that fishers select the level of fishing activity which places them on the highest possible indifference curve between work and income. That is to say, they seek to maximize the value of a utility function $U(f, L)$, where $U_f > 0$ and $U_L < 0$. This occurs at the tangency between the $f(L)$ function and the highest possible indifference curve,[9] where $f'(L) = MRS(f, L)$.

The seasonal fishermen's benefits program alters this pattern of incentives. The benefits received supplement earnings from fishing, and, in so doing, alter the incentives to fish, through both income and substitution effects. The income effects are straightforward and consistent with the standard literature referenced above. If "leisure" (understood to mean time spent in activities other than fishing)[10] is a normal good, then higher income from unemployment insurance leads to an increase in the demand for leisure, and thus a reduction in the number of weeks spent fishing, in order to enjoy this additional leisure.

The substitution effects are more complex and usefully represented through formal modeling. The amount of unemployment insurance income, S, earned in a benefit period is the product of three factors: (*i*) the benefit-earnings, or replacement ratio, r, which is the proportion of average weekly insured earnings during the qualifying period that is returned to the claimant as benefits during a week of unemployment; (*ii*) the average level of weekly insured earnings, E, during the claimant's qualifying weeks, that is the basis on which the level of weekly benefits is calculated; and (*iii*) the number of weeks, B, over which the claimant is entitled to draw benefits. Since the variables E and B depend on the number of insured weeks, L, in the qualifying period, this relationship can be written as

$$S(L) = r\, E(L)\, B(L). \tag{1}$$

The benefit-earnings ratio, r, is a constant, which was equal to 66% from 1972–78, then 60% until 1990, and 57% thereafter.

Until 1983, the average value of insured earnings, $E(L)$, was calculated over qualifying weeks in either the entire qualifying period (which usually begins in April), or in the last 20 weeks of this period, whichever was to the fisherman's advantage. We assume that fishing income increases at a diminishing rate as the number of fishing weeks increases. This assumption implies that if weekly earnings are below the insurable ceiling, average weekly insured earnings will decline as the number of fishing weeks increases. This decline would have a negative effect on the level of unemployment benefits, acting as a disincentive for the fisherman to extend the number of fishing weeks. Thus, we can specify that $E'(L) \leq 0$, with the strict inequality holding, where $f'(L)$ is below the maximum level of weekly insurable earnings.

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[8] The fishing weeks variable, L, should not be necessarily taken as chronologically ordered. The basic idea is that since fishers prefer weeks when fishing income is high to weeks when it is lower, then as the number of weeks spent fishing is increased, the income earned in the marginal week must be less than income earned in intra-marginal weeks. The relationship between fishing income and fishing weeks is, therefore, concave. However, marginal weeks need not be weeks at the end (or beginning) of the fishing season.

[9] Roy *et al.* (1994) also consider the case in which fishers switch between fishing and wage employment as the marginal returns to fishing fall below the wage which could be earned in shore employment. We do not consider this case here, partly because of data limitations, and partly because we doubt that this case is representative of the employment options available in rural Newfoundland, where unemployment rates are typically in excess of 20%, and shore employment is generally sporadic and unreliable. Because of low skill levels, moving costs, and a general lack of liquid cash resources, mobility into other areas of Canada is not generally a feasible short-term option.

[10] We use the term "leisure" here because it has become institutionalized in the labor economics literature. However, the term "nonmarket household production" (from which true leisure is one possible output) would probably be a more accurate description of the alternative use of time by inshore fishermen.

In 1983, the definition of the average value of insured earnings was modified so that fishermen with at least 15 qualifying weeks of fishing would receive benefits based on earnings in their best 10 weeks of fishing. This change would render $E(L) = E(10)$ for $L \geq 15$, and so remove the disincentive to extend the number of fishing weeks for those fishing at least 15 weeks.

The relationship, $B(L)$, between the number of benefit weeks and the number of insured weeks during which income was earned can be separated into four sequential stages.

Stage 1: Here the number of insured weeks, L, is less than the minimum number of qualifying weeks required to entitle a claimant to benefits, denoted by q. In Newfoundland, this minimum level of insured weeks was 8 until 1978, when it was raised to 10. In this stage, obviously, there are no benefits; *i.e.*,

$$B(L) = 0 \text{ if } L < q.$$

Stage 2: Once a claimant qualifies for benefits, the number of weeks that benefits can be claimed increases with the number of insured weeks, up to some maximum which is governed by the length of the off-season during which benefits can be claimed. In Stage 2, this maximum has not yet been reached, so there is a positive relationship between the number of benefit weeks and the number of insured weeks. This positive relationship creates an incentive to extend the number of fishing weeks in order to qualify for a longer period of unemployment benefits. Specifically, claimants are entitled to 5 weeks of benefits for every 6 qualifying weeks, so $B'(L) = 5/6$. Beginning in 1976, fishermen were entitled to a certain number of weeks called "extended benefits," B_{ext}, which is independent of the number of qualifying weeks they have worked. Thus, in Stage 2, the number of benefit weeks can be written as the linear relationship

$$B(L) = (5/6)L + B_{ext}.$$

Stage 3: Here the fisherman works a sufficient number of insured weeks to qualify for the maximum number of benefit weeks, B_{max}. Consequently, additional fishing does not increase the length of the period that a fisherman is entitled to benefits. Thus, the positive incentive to extend the number of fishing weeks which exists in Stage 2 is removed. In this stage, we have

$$B'(L) = 0, \text{ and } B(L) = B_{\max}.$$

Stage 4: Ultimately, as the number of fishing weeks is extended, fishing takes place during that part of the year when seasonal benefits can be claimed. In this stage, every additional week spent fishing is one in which unemployment benefits could have been received. Therefore, there is an incentive to reduce the number of fishing weeks. Here we have $B'(L) = -1$, and so (allowing for the two-week waiting period), $B(L) = 50 - L$.

Overall, the relationship $B(L)$ is piecewise linear, with four distinct segments, and can be expressed as

$$\begin{aligned} B(L) &= \min\left[\frac{5}{6}L + B_{ext}, B_{max}, 50 - L\right], L \geq q \qquad (2) \\ &= 0, L < q. \end{aligned}$$

Table 1
Unemployment Insurance Benefit Week Parameters
(Newfoundland, 1971–93)

	q	B_{ext}	B_{max}
1971–75	8	0	22
1976	8	18	27
1977–93	10	20	27

This relationship was modified by amendments in 1976 and in 1977, which altered the q, B_{ext}, and B_{max} parameters. These changes are summarized in table 1. It should be noted that, as a result of the 1977 changes, the levels of q, B_{ext}, and B_{max} increased sufficiently so that $(5/6)q + B_{ext} > B_{max}$, and Stage 2 was eliminated.

When income from seasonal fishermen's benefits is added to earned fishing income, total fishing income, F, becomes

$$\begin{aligned} F(L) &= f(L) + S(L) \\ &= f(L) + r\,E(L)\,B(L). \end{aligned} \tag{3}$$

The tangency condition can then be written as

$$F'(L) = f'(L) + S'(L) = MRS(F, L) \tag{4}$$

thus, the increased income (inclusive of unemployment insurance) from fishing an additional week equals the marginal rate of substitution between income and leisure.

In the next two sections, we shall estimate an econometric model that consists of the earnings-weeks relationship, $f(L)$, and the labor supply relationship implied by the tangency condition (4). The unemployment insurance benefits equation, $B(L)$, is part of the model as well, but as an identity, depending on known institutional parameters and without a stochastic component. The model is estimated on a 10% sample of seasonal benefit recipients in Newfoundland over the period 1971–93 provided by the Department of Employment and Immigration of the Government of Canada. The sample contains panel observations on 21,447 benefit spells involving 5,999 recipients. The data consists of information on weeks of insured employment resulting in the benefits claim, the sum of insured earnings during this employment spell, weekly benefit rates when unemployment occurred, and a total number of benefit weeks during the unemployment spell.[11] A limited set of demographic data (such as age and sex) and economic data (such as occupational and industrial classifications) is also included. Since the data covers a fairly long period (twenty-three years), it is typically the case that observations on a particular individual span only a small portion of this period, and need not be contiguous observations.

[11] An anonymous referee expressed concerns about data corruption arising from a possible incentive to over-declare earnings in order to increase benefit. However, the deemed employer is required to maintain detailed records on fishermen's earnings and premiums payable (*Regulations*, s. 77). As well, since the employer must pay premiums which depend on the fisherman's insured earnings, the employer has an incentive to minimize this amount. Outright fraud does not appear widespread as a result. There are, however, reliability concerns arising from the fact that the effect of the minimum and maximum insurable earnings levels is to censor such earnings both from above and from below. As a result, insured earnings may not be the same as actual earnings. This issue is discussed in the final section of the paper.

Econometric Estimation of the Earnings-Weeks Relationship

Let us begin with the specification of the relationship between earnings and insured weeks. For estimation purposes, we use a log-linear approximation as follows:

$$f_i = \theta_i L_i^{\beta} \tag{5}$$

where earnings of fisherman i, in a particular year are represented by f_i, and weeks worked by fisherman i, in that year by L_i. The elasticity of earnings with respect to the number of weeks is measured by the β parameter, which is assumed to be the same for all fishermen and which should be between 0 and 1. The greater the diminishing returns to an extended fishing season, the lower is the value of β. The θ_i parameter reflects the productivity of the fisherman for a particular value of L. This level of productivity is known to vary considerably from person to person, depending on factors such as experience, location, luck, and innate skills.

From the beginning, we were confronted with a serious identification problem, the nature of which is captured in figure 1, representing the earnings-weeks relationship of two fishermen with different values of the parameter θ_i. Ideally, we would like to be able to trace this relationship by controlling for the value of θ_i. Unfortunately, we cannot observe θ_i directly, and so even if its value is partly captured through the use of various correlates, much of its effect will be reflected in the equation disturbance term. However, the number of weeks spent fishing is an endogenous variable and is unlikely to be independent of θ_i. For example, high values of θ_i may be associated with a lengthy fishing season. Therefore, variations in θ_i which are not captured in the regression could be correlated with the independent variable, L_i, causing the estimate of the β parameter to be biased. In terms of figure 1, instead of tracing an earnings-weeks relationship, such as OF, we may instead trace a locus of tangencies such as AB.

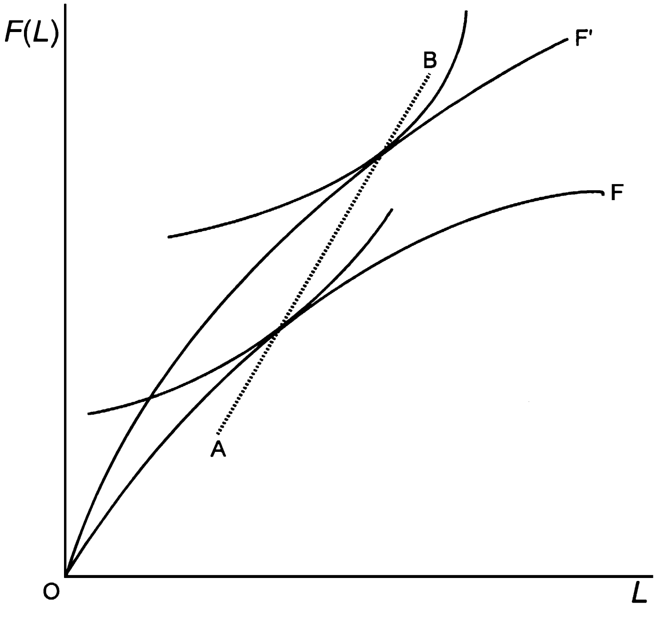

Figure 1. Biased Estimation of the Earnings-Weeks Relationship

The standard solution to this problem is to use an instrumental variable for L_i which is related to L_i, but is independent of variations in the earnings-weeks relationship as captured in θ_i. The latter requirement precludes any determinants of fishing productivity as instrumental variables, since these affect the earnings-weeks relationship. Variables which are related to the income-leisure preferences of fishermen, but not to productivity differences between fishermen, would be appropriate candidates for consideration as instrumental variables. Unfortunately, our database does not include any variables which clearly satisfy this criterion.

Notwithstanding this dilemma, we consider that we have been able to obtain a reasonable (although not perfect) solution to the problem in specifying the earnings-weeks relationship [equation (5)] as an error-components relationship. Specifically, we decompose the θ_i term into an individual-specific component, α_i, and a time-specific component, η_t. The individual-specific component captures those aspects of productivity which are specific to the individual over all time periods, such as innate skill. The time-specific component, on the other hand, captures those effects which are specific to the time period for all individuals, such as changes in resource availability, weather conditions, and product price.[12] We have also incorporated an age variable into the regression specification in order to capture individual-specific differences that are correlated with age. The intent is to factor in differences in productivity that are due to accumulated experience.[13] To the extent that older fishermen can afford larger boats, differences in physical capital accumulation may be reflected here as well.

The econometric literature provides two alternative specifications for the error-components model. The fixed-effects model specifies the individual-specific and/or the time-specific effect as a fixed parameter to be estimated through the use of dummy variables. To the extent that these fixed effects capture that part of θ_i which is correlated with the regressors (and particularly with L_i), the estimate of the β parameter is unbiased.

The random-effects model, by contrast, specifies the individual-specific effects, and/or the time-specific effects, as a random variable possessing specific characteristics. Usually, but not always, the random variable is identically and independently distributed, and is estimated through Generalized Least Squares (GLS) (see Judge, *et al*. 1985, ch. 13; Greene 1997, ch. 14.4). The random-effects model has some methodological attractions and can be shown to lead to efficient estimation when the model specification is valid. The main disadvantage of this model is that when the random effects are correlated with the explanatory variables, least-squares bias results. Fortunately, the latter situation is testable by using the Hausman specification test (Greene 1997, pp. 632–34).

We specify the time-effect as a fixed (dummy variable) effect in both models, because the number of time periods in our analysis is small (twenty-three years), and the time interval over which observations on particular individuals spans does not overlap much. We have modeled the individual effect, however, as both a fixed effect and a random effect. The fixed-effects model, after logarithmic transformation, is as follows:

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[12] In principle, some of these factors could be estimated directly. However, both prices and resource availability are species-specific, and we have no information on the species makeup of the landings of individual fishermen in our sample. Similarly, both weather conditions and resource availability vary spatially, and the location information which is included in our sample is rather coarse and of uncertain reliability.

[13] Measuring this effect is complicated by the fact that fishing is conducted in teams; therefore, the income earned by a fisherman depends on the accumulated experience of the entire crew, particularly the skipper, rather than that of only the individual fisherman. I am grateful to an anonymous referee for reminding me of this.

$$\ln f_{it} = \alpha_i + \eta_t + \gamma \ln a_{it} + \beta \ln L_{it} + \varepsilon_{it}, \quad \varepsilon_{it} \sim \text{IID}(0, \sigma^2_\varepsilon) \tag{6}$$

where f_{it} is the earnings of individual i at time t; a_{it} is the age of individual i at time t; L_{it} is weeks worked by individual i at time t; α_i is the fixed effect specific to individual i; η_t is the fixed effect specific to time period t; and ε_{it} is the equation disturbance for individual i at time t, assumed to be identically and independently distributed with zero mean and constant variance.

The random effects model, by contrast, is specified as

$$\ln f_{it} = \alpha + \eta_t + \gamma \ln a_{it} + \beta \ln L_{it} + u_i + \varepsilon_{it}, \qquad u_i \sim \text{IID}(0, \sigma^2_u), \varepsilon_{it} \sim \text{IID}(0, \sigma^2_\varepsilon) \tag{7}$$

where the individual effect, u_i, is now a random variable which is identically and independently distributed with constant variance. The model is estimated by Feasible GLS, with the variance components estimated using the technique outlined in Greene (1997, pp. 626–28).

For comparison purposes, we also estimate a constrained Ordinary Least Squares (OLS) model in which the individual effects, α_i, are equal for all individuals.

$$\ln f_{it} = \alpha + \eta_t + \gamma \ln a_{it} + \beta \ln L_{it} + \varepsilon_{it}, \quad \varepsilon_{it} \sim \text{IID}(0, \sigma^2_\varepsilon) \tag{8}$$

The regression estimates and the estimated standard errors of the constant parameters for all three models are presented in table 2. In all three models, the estimated value of β is fairly high, but significantly below unity, ranging from 0.85 in the fixed-effects model, to 0.94 in the constrained OLS model. The hypothesis of diminishing returns is, therefore, confirmed. The estimate of the age effect is more variable, but in all cases, it is significantly positive. This is consistent with the maintained hypothesis that productivity rises with age. In the fixed-effect model, however, the value ascribed to this effect seems to be implausibly high.

The fixed-effects model can be used to test the hypothesis that the individual effects take the same value; in this case, the constrained OLS model would be valid. Under the null hypothesis, when the ε_{it} are normally distributed, the F-statistic has a value of 5.80 with 5.998 and 15.424 degrees of freedom. At these values, the null hypothesis is decisively rejected at any reasonable level of significance.[14] The Hausman statistic for orthogonality of the random effects with the independent variables, which under the null hypothesis is asymptotically distributed as chi-squared with 24 degrees of freedom, has a value of 137.03, also leading to rejection of the null hypothesis.

We conclude from these results that the fixed-effects model is the most satisfactory for our purposes, because the alternative models possess significant least-squares bias which is absent from the fixed-effects model. We do note that the size of the age effect estimated in this model is substantially (and implausibly) higher than it is in the other two models. There may remain some least-squares bias in the fixed-effects model to the extent that the model errors ε_{it}, which, by design, are orthogonal to both the individual effects and the time effects, are nonetheless corre-

[14] Similarly, in the random-effects model, the Breusch-Pagan (1980) Lagrange multiplier test can be utilized to test the hypothesis that $\sigma^2_u = 0$, implying that random effects are absent. The value of the Lagrange multiplier, which under the null hypothesis is distributed as chi-squared with one degree of freedom, is 37.105, so the null hypothesis is once more decisively rejected.

Table 2
Regression Coefficients and Standard Errors,
Earnings-Weeks Relationship

	Constrained OLS Model	Fixed-Effects Model	Random-Effects Model
α	4.088	—	4.077
	(0.033)		(0.050)
β	0.945	0.856	0.866
	(0.007)	(0.008)	(0.007)
γ	0.063	0.873	0.133
	(0.006)	(0.070)	(0.012)
R^2	0.750	0.923	—

lated with weeks worked. However, since most of the variance in ln L_{it} is between-group variance rather than within-group variance, and since this between-group variance is, by design, independent of the model error, we think that most of the least-squares bias has been removed from the fixed-effects model. We also note that the estimates of β generated by the fixed-effects and random-effects models are quite similar, despite the fact that there is some bias in the random-effects model which is absent from the fixed-effects model.[15]

Econometric Estimation of the Tangency Condition

We now construct a model of the tangency condition (4), which equates the marginal return to an additional week of work $F'(L)$ to the marginal rate of substitution between income and leisure. The former concept, which we refer to as the real marginal return to work (RMRW), was derived as follows. By differentiating equation (1), substituting the result into the left-hand portion of equation (4), and adding appropriate subscripts, we obtain

$$F'_{it}(L_{it}) = f'_{it}(L_{it}) + r_t\left[E_{it}(L_{it})B'_t(L_{it}) + E'_{it}(L_{it})B_t(L_{it})\right]. \tag{9}$$

We derive $f'(L)$ by differentiating the log-linear specification in equation (5) for $f(L)$ to obtain

$$f'_{it}(L_{it}) = \beta\alpha_{it}L_{it}^{\beta-1} = \beta\,\frac{f_{it}}{L_{it}}. \tag{10}$$

The expression $B(L)$ was derived from equation (2), and $B'(L)$ was obtained by differentiating equation (2). The expression $E'(L)$ was set equal to 0 if a fisherman was at the maximum level of insurable earnings, or working 20 weeks or more in years before 1983 and 15 weeks or more afterward. Otherwise, $E'(L)$ was derived by differentiating $E(L) = f(L)/L$ to obtain

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[15] In the random-effects model, the standard error of the u_i term (which is absent from the fixed-effects model) is estimated as 0.420, while that of the ε_{it} term is only 0.199.

$$E'_{it}(L_{it}) = \frac{f'_{it}(L_{it})}{L_{it}} - \frac{f_{it}(L_{it})}{L^2_{it}} \qquad (11)$$
$$= (\beta - 1)\frac{f_{it}(L_{it})}{L^2_{it}}$$
$$= (\beta - 1)\frac{E_{it}(L_{it})}{L_{it}}.$$

Values for E_{it}, f_{it}, F_{it}, and L_{it} were obtained from the database. The value of β was set equal to the fixed-effects estimate obtained in the previous section. Finally, the entire expression was divided by the Consumer's Price Index for St. John's, Newfoundland, in order to convert the expression to a real return to work.

Because of the presence of kinks and discontinuities in both the $B(L)$ and $E(L)$ equations resulting from the design of the Unemployment Insurance Program, the term $F'(L)$ is not defined at several values of L, and so the tangency condition (4) is not, in general, satisfied at these points.[16] We have dealt with this problem by excluding such observations from the sample used for estimation of the parameters underlying the tangency condition. This is not entirely satisfactory, since a considerable amount of information (8,140 observations) is removed as a result. This includes those cases in which the claimant qualifies for unemployment insurance with a minimum number of weeks—a group which is of considerable interest for policy purposes. We shall consider this issue more fully in a subsequent section.

The expression used for the marginal rate of substitution is based on the assumption that the underlying preference functions of fishermen between real income and leisure can be approximated by the constant-elasticity-of-substitution form

$$U_{it} = \left[(F_{it}/P_t)^{\rho} + \delta_i(52 - L_{it})^{\rho}\right]^{1/\rho} \qquad (12)$$

where F_{it}/P_t is real income of fisherman i in period t; $52 - L_{it}$ is leisure enjoyed by fisherman i at time t; δ_i is an individual-specific parameter reflecting the relative preference of fisherman i for income versus leisure; and ρ is a parameter, assumed to be the same for all fishermen, which is related to the elasticity of substitution σ between income and leisure by the relationship $\rho = 1 - (1/\sigma)$. Since

$$MRS = -\frac{\partial U}{\partial L}\Big/\frac{\partial U}{\partial(F/P)}$$

through differentiation we obtain

$$MRS_{it} = \delta_i\left[\frac{52 - L_{it}}{F_{it}/P_t}\right]^{\rho-1}. \qquad (13)$$

[16] Specifically, $B'(L)$ fails to exist at the boundaries between the separate stages of the $B(L)$ function, and $E'(L)$ fails to exist (after 1982) at $L = 15$, where average insured earnings are deemed to equal the average earnings in the best ten weeks of fishing.

Setting MRS_{it} equal to $RMRW_{it}$, taking logarithms of both sides, and rearranging, we obtain the regression equation

$$\ln\left[\frac{52 - L_{it}}{F_{it}/P_t}\right] = \sigma \ln \delta_i - \sigma \ln\left[\frac{F'_{it}}{P_t}\right] + \zeta_{it}, \ \zeta_{it} \sim \text{IID}(0, \sigma_\zeta^2) \tag{14}$$

when the individual preference term δ_i is treated as a fixed-effect parameter. On the other hand, when the preference term is treated as a random effect, the optimization equation becomes

$$\ln\left[\frac{52 - L_{it}}{F_{it}/P_t}\right] = \sigma \ln \delta - \sigma \ln\left[\frac{F'_{it}}{P_t}\right] + v_i + \zeta_{it}, \tag{15}$$
$$v_i \sim \text{IID}(0, \sigma_v^2), \quad \zeta_{it} \sim \text{IID}(0, \sigma_\zeta^2).$$

In both cases the ζ_{it} term may be considered to be the effect of optimization errors occurring due to mechanical breakdown, incorrect anticipations, and so on. It should be noted that F'_{it} is a function of L_{it} [see equation (9) above], so there is some possibility that this variable is correlated with the equation disturbances v_i and ζ_{it}, creating least-squares bias.

We estimate both models, along with a constrained OLS model that imposes the restriction that all fishermen have the same preferences, in which case the regression equation can be written as

$$\ln\left[\frac{52 - L_{it}}{F_{it}/P_t}\right] = \sigma \ln \delta - \sigma \ln\left[\frac{F'_{it}}{P_t}\right] + \zeta_{it}, \quad \zeta_{it} \sim \text{IID}(0, \sigma_\zeta^2). \tag{16}$$

The regression estimates and estimated standard errors of the constant parameters for the three models are presented in table 3. The estimate of the elasticity of substitution parameter, σ, is, in all cases, statistically significant but fairly low, ranging from 0.14 in the fixed-effects model, to 0.29 in the constrained OLS model. We can conclude that real income and leisure are not regarded as close substitutes by this population of fishermen.

The fixed-effects model can be used to test the hypothesis that individual preferences (as reflected in the δ_i parameter) are identical, so that the Constrained OLS model is valid. Under the null hypothesis, when the ζ_{it} are normally distributed, the F-statistic has a value of 3.68 with 4.981 and 8.324 degrees of freedom. The null hypothesis is clearly (and not unexpectedly) rejected.[17] Finally, the Hausman statistic for orthogonality of the random preference effects with the independent variable, which under the null hypothesis is asymptotically distributed as chi-squared with 1 degree of freedom, has a value of 242.25, once more leading to rejection of the null hypothesis.

Again, we conclude from these results that the fixed-effects model is the most satisfactory one, since it appears that the value of RMRW is correlated with indi-

[17] Similarly, in the random-effects model, the Breusch-Pagan (1980) Lagrange multiplier test is utilized to test the hypothesis that $\sigma_v^2 = 0$, so that random effects are absent. The value of the Lagrange multiplier, which under the null hypothesis is distributed as chi-squared with one degree of freedom, is 14.836, so the null hypothesis once more is decisively rejected.

vidual income-leisure preferences, and, therefore, with the individual effect in the random-error model. There may remain some least-squares bias in the fixed-effects model to the extent that the optimization errors, ζ_{it}, are correlated with RMRW. Some suggestions as to how to deal with this possibility are discussed in the next section.

While it would be premature to base a full-model simulation on these results, it is nonetheless of interest to derive some estimates of the impact that the introduction of the Unemployment Insurance Program had on the length of the fishing season. Table 4 presents three examples of individuals with preference functions and earnings-weeks functions based on the fixed-effects models presented in this and the previous section. The Unemployment Insurance parameters utilized are representative of the situation in 1983–89.

While the details differ from case to case, the results suggest that the program reduced the fishing season by 8–10 weeks, and a concomitant reduction in earned income which was more than compensated for by unemployment insurance receipts of approximately CDN$3,000 or more per year. These are fairly significant effects. There could, however, be counteracting effects from marginal fishermen who increase their fishing to the 10-week minimum in order to qualify for unemployment insurance. Since we have excluded such individuals from our sample in the estima-

Table 3
Regression Coefficients and Standard Errors, Tangency Condition

	Constrained OLS Model	Fixed-Effects Model	Random-Effects Model
$\sigma \ln \delta$	–3.986	—	–4.343
	(0.043)		(0.043)
σ	0.297	0.142	0.222
	(0.008)	(0.009)	(0.008)
R^2	0.099	0.923	—

Table 4
Simulated Effects of Unemployment Insurance Program on Fishing

	Weeks Worked L_i	Earned Income $f_i(L_i)$	UI Income $S_i(L_i)$	Total Income $F_i(L_i)$
Case I				
With UI	12	$2,933	$3,933	$6,892
Without UI	23.3	$5,156	—	$5,156
Case II				
With UI	20	$4,528	$4,069	$8,597
Without UI	28.8	$6,178	—	$6,178
Case III				
With UI	30	$6,391	$3,014	$9,405
Without UI	37.8	$7,773	—	$7,773

Note: Income values are in Canadian dollars.

Table 5
Simulated Effects of Change in Replacement Parameter, *r*, on Fishing

	Weeks Worked L_i	Earned Income $f_i(L_i)$	UI Income $S_i(L_i)$	Total Income $F_i(L_i)$
Case I				
$r = 0.6$	12	\$2,933	\$3,933	\$6,892
$r = 0.5$	15	\$3,546	\$3,391	\$6,937
Case II				
$r = 0.6$	20	\$4,528	\$4,069	\$8,597
$r = 0.5$	21.4	\$4,805	\$3,391	\$8,196
Case III				
$r = 0.6$	30	\$6,391	\$3,014	\$9,405
$r = 0.5$	32.5	\$6,835	\$2,202	\$9,037

Note: Income values are in Canadian dollars.

tion of the tangency condition, it is arguably illegitimate to apply our model to such individuals, and so we have not done so.

The analysis presented in table 4 is open to the objection that the span of the simulated change lies far outside the sample space, so that the results may be critically dependent on the assumed functional form of the earnings and utility functions. We may also be interested in the effects of a change in the parameters underlying the unemployment insurance system which falls short of scrapping the system entirely. Table 5 simulates the effect of lowering the replacement parameter, *r*, from 0.6 to 0.5. Such a change is probably at the outer limits of political feasibility. The model predicts that, in the three cases that were analyzed in the previous example, fishermen would be induced to work an additional 1.5–3 weeks as a result of this change. Unemployment insurance benefits would fall by an amount in the range of CDN\$500 to \$800, some of which is offset by higher earned income from the longer fishing season. The response in the short-season Case I is particularly interesting, in that the reduction in *r* would induce the fisherman to increase fishing to 15 weeks in order to take advantage of the "best 10 weeks" rule which determines average insured earnings when the number of insured weeks is at this level. In this case, total income would actually rise as a result of the reduction in benefits.[18]

Reducing the maximum number of benefit weeks, B_{max}, from 27 to 22 would have effects similar to those of the reduction in *r* in Case I and Case II, but would have no effect on Case III, since, in this case, the season is sufficiently long that the fisher would be taking less than 22 weeks in benefits at her preferred position. An increase in the minimum number of qualifying weeks, *q*, or a reduction in the number of weeks of extended benefits, B_{ext}, would only affect the decisions of those fishing a short season, such as Case I.

[18] This counterintuitive result is due to the discontinuous nonconvexity in the benefits function $B(L)$ at $L = 15$ which results from the operation of the "best ten weeks" rule at this point.

Limitations and Extensions

The analysis presented above has two limitations which particularly concern us. First, as discussed above, the level of earnings which qualify for unemployment insurance is subject to a weekly maximum. If earnings exceed this maximum, they are deemed to be at that maximum for unemployment insurance purposes. Similarly, a week's earnings must exceed a particular level in order to qualify for insurance. If earnings fall short of this minimum, they are deemed equal to that minimum for a skipper, but not for a crewman. Therefore, our data on insured earnings are indirectly censored upward—indirectly, because the censoring is imposed on a week-by-week basis. Therefore, the total is censored to the degree that weekly earnings have been censored, a matter which cannot generally be determined from our database. Similarly, the data are indirectly censored downward for a skipper, and indirectly truncated downward for a crewman.

Truncation and censoring can lead to bias in the regression estimates, and the bias can sometimes be severe (see the Monte Carlo results presented in Davidson and MacKinnon 1993, p. 538). Maximum-likelihood estimation is normally used to estimate models with censored and truncated data. This method necessitates the specification of an exact functional form (typically, although not necessarily, normality) for distribution of the disturbances. The methods used in the previous sections, by contrast, require only the assumption of identical and independent distributions with finite variance, a considerably weaker assumption.

While there is rich and extensive literature on the handling of censored and truncated data, with our data, the standard approach must be modified to incorporate the fact that the data consist of sums of a series of censored and truncated data, rather than directly truncated or censored data. The required modification is not a trivial one.

The second area of concern arises from the piecewise nature of the benefits function (3), which led us to remove observations which occurred at the kinks and discontinuities in the estimation of the tangency function. There are two problems with this procedure. First, the procedure is inefficient because information is discarded. Second, optimization errors on the part of fishermen (as a result of uncertainty, for example) can result in behavior which would introduce bias into the parameter estimates.

The source of this bias is twofold (Pudney 1989, pp. 198–201). First, there is the simultaneity bias already discussed, in that model error affects weeks worked, which, in turn, affects RMRW. Therefore, the RMRW variable is correlated with the model error. The second source of bias that results from optimization errors arises when the observed and optimum positions of a given data point lie on different segments of the benefits function. For example, a fisherman who would best locate in Stage 2 may mistakenly fish into Stage 3 as a result of overly optimistic expectations of the return to fishing in that stage. When such points are grouped into the "wrong" segment, the value of RMRW at the optimum is calculated incorrectly. Therefore, the resultant parameter estimates are subject to errors-in-variables bias.

A related problem is that data points whose optimum position is at the corner vertex formed by two segments of the earnings-weeks relationship will generally not satisfy a tangency condition defined over either segment. Inclusion of these boundary points in the estimation will, therefore, bias the results. When fishermen make optimization errors, however, we cannot readily identify those cases that possess corner optima.

These are not new problems. There exists an extensive literature (Wales and Woodland 1979; Zabalza 1983; Phipps 1990) on the impact of piecewise-linear constructs, such as progressive income taxes and unemployment insurance on the length of work spells. Maximum likelihood methods have been successfully utilized to resolve these difficulties (Pudney 1989, pp. 201–205). The solution of such problems, however, is simplified considerably by the assumption usually made in this literature that work is available throughout the year at a fixed wage. Unfortunately, this assumption is unsupportable in the present case. Our objective in future work is to adopt these techniques to contexts, such as the present one, in which the "wage" varies systematically through the year.

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