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Health and Residential Mobility in Later Life:

A New Analytical Technique to Address an Old Problem

Lynda M. Hayward

SEDAP Research Paper No. 34

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Health and Residential Mobility in Later Life: *

A New Analytical Technique to Address an Old Problem

By

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Abstract

For some time researchers have known that the relationship between health and the residential mobility of the elderly is not straight forward and changes with age. Attempts to examine this relationship in multi-variate models using cross-sectional data have resulted in contradictory or ambiguous findings. One solution has been to create separate models for different age groups. However, the onset of poor health differs considerably by individual, particularly for the "young-old". Multi-variate proportional hazards models using longitudinal data offer a new approach to address this problem. As an example, data from the Ontario Longitudinal Study of Aging have been analyzed using proportional hazards models as compared with logistic regressions. The logistic regressions yield typically ambiguous results. The proportional hazards models indicate a reversal with time in the relationship between one of the two mid-life health measures and residential mobility, and the results for both measures are consistent with the theoretical literature.

Health, in terms of functional ability, plays an important role in developmental life cycle models of residential mobility in later life. Litwak and Longino (1987) have argued that the elderly make different types of moves depending on their stage in the life cycle. Briefly, these are: an amenity or life style move made upon retirement; a support or kinship seeking move that accompanies the development of chronic disabilities, and an institutional move when seniors can no longer live independently. However, the relationship between health and residential mobility is not straightforward. While on the one hand, good health can facilitate amenity migration upon retirement, poor health is thought to be the main motivation for support seeking moves in later life (Litwak and Longino, 1987; Patrick, 1980). Moreover, a lack of health resources may prevent older people from making moves, particularly long distance migrations (Carter, 1988; Patrick, 1980; Wiseman, 1980). As a result, the canceling effect of these trends can create the impression that health has little or no effect.

Patrick (1980) has suggested that at any age the relationship of health with the likelihood of moving may be U-shaped as illustrated in Figure 1. Coupled with the negative relationship between health and age, he argues that it becomes difficult to analyze the impact of health changes on migration for the elderly population overall. This could possibly explain some of the contradictory or ambiguous findings concerning this relationship in the literature, especially in cases where age and health are both included in the models (*e.g.* Meyer and Speare, 1985; Sommers and Rowell, 1992; Speare *et al.*, 1991). These problems are further confounded by the use of retrospective cross-sectional data in which the causal direction of the relationship is unclear and differences attributed to age may reflect cohort differences such in the age of the onset of poor health.

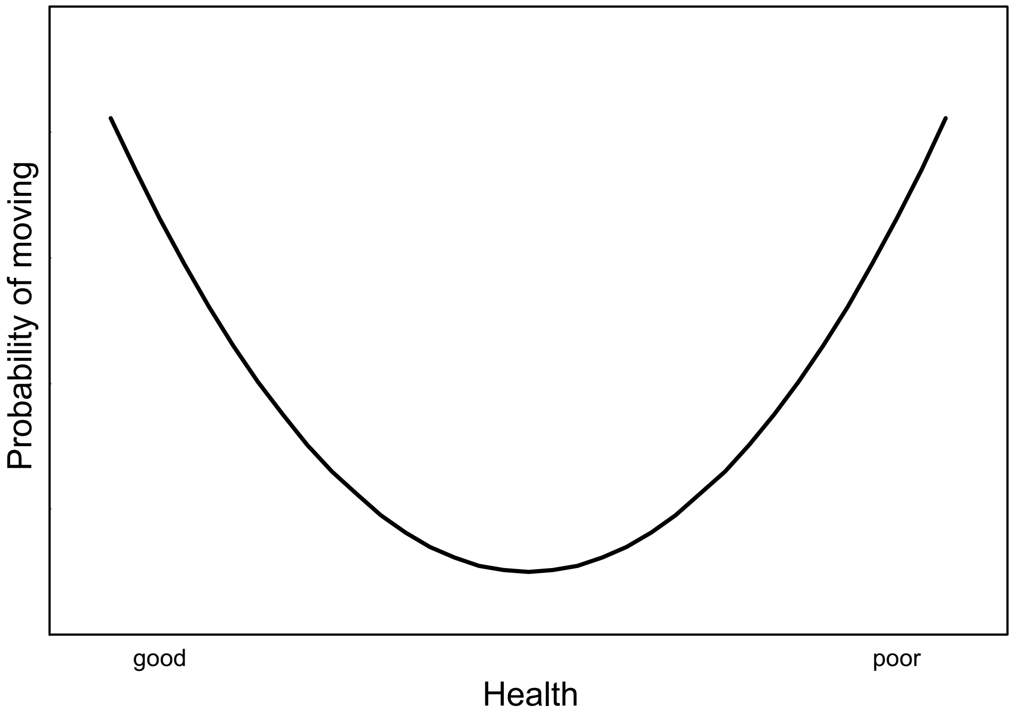

Figure 1 Probability of moving by health for the elderly

One approach commonly taken to address the analytical problems associated with different types of moves being made at different points in the life cycle is to develop separate migration or residential mobility models for different age groups of older people (*e.g.* Moore *et al.*,1997), or to focus on the "old-old" age group (*e.g.* Longino *et al.*, 1991; Speare *et al.*, 1991). Since disability rates increase exponentially with age (Forbes *et al.*, 1993; Moore *et al.*, 1997), it is reasonable to expect that in the older age groups an increasing proportion of moves are associated with declining health or functional ability.

The logistic regression models of elderly migration developed by Moore and his associates (1997), present a good example of the results obtained from the development of separate models for different age groups of older people. In one of the two models developed for respondents aged 55 to 64, those with severe disability were found to be less likely to move than those with no disabilities. This relationship disappeared in the second model when geographic region was entered into the equation. The level of disability was not found to have a statistically significant relationship with migration in either of the models developed for the 65 to 74 age group ($p>.05$). On the other hand, for the oldest group, aged 75 and over, those with mild or moderate disabilities were found to be half again as likely to migrate as compared with those with no disabilities, while those with severe disabilities did not differ significantly from those with no disability. As can be seen in this example, the creation of separate models for the analysis of cross-sectional data can be helpful when examining the relationship between health and residential mobility for the "old-old". However, because the age of onset of disability or poor health varies considerably, results concerning this relationship for the "young-old" remain ambiguous.

With the increasing availability of longitudinal data, alternative analytical approaches are now possible. Comparisons of cross-sectional and longitudinal analyses of residential mobility and migration have shown clear advantages for the use of a longitudinal approach (Clark, 1992; Davies and Pickles, 1985). This paper examines how an event history approach, specifically survival analysis, can be used to address many of the problems associated with the analysis of the relationship between health and the residential mobility of the elderly, particularly for the younger group. Using data from the Ontario Longitudinal Study of Aging, the results obtained from the logistic regression modeling techniques commonly used with cross-sectional data, are compared with those obtained from survival analyses of longitudinal data, specifically Cox proportional hazards models.

Formulation of the Models

To facilitate comparison of the results of these two different analytical techniques, it is useful to have some understanding of the formulation of the models, beginning with logistic regression.

Logistic Regression

Logistic regression has been the preferred analytical technique for multi-variate analyses of the residential mobility of the elderly using cross-sectional or short-term longitudinal data (*e.g.* Bradsher *et al.* 1992; Findley, 1988; Jackson *et al.*, 1991; Longino *et al.*, 1991; Moore *et al.*, 1997; Silverstein and Zablotsky, 1996; Sommers and Rowell, 1992; Speare *et al.*, 1991). This procedure is particularly well suited for estimating the association between independent variables and binary dependent variables, such as moving/not moving, while adjusting for possibly confounding factors. As described by Henkins and Buring (1987:317), this procedure is a variant of multiple regression in which the log odds of the occurrence of a dichotomous dependent variable (the logit), " can be expressed as a simple linear function of the independent predictor variables", as below:

$$\ln [Y/1\text{-}Y] = a + \beta_1 X_1 + \beta_2 X_2 + ...$$

where, in the present analysis: Y would be the probability of moving, a is a constant; β_1 and β_2 would be the parameter estimates; and X_1 and X_2 the independent health variables. This procedure also has the advantage that the "coefficients can be directly converted to an odds ratio that provides an estimate of the relative risk that is adjusted for confounding" variables (Henkins and Buring,1987:317).

Proportional Hazards

Cox proportional hazards modeling is also a log-linear procedure which describes the distribution of survival time (T) as a hazard function h(t), or the conditional failure (moving) rate. It is "defined as the probability of failure [moving] during a very small time interval, assuming that the individual has survived [aged-in-place] to the beginning of the interval" (Lee, 1980:12). This can be expressed as:

$$h(t) = \lim_{\Delta t \to 0} \frac{p\{\text{ an individual aging-in-place to t, moving in the interval } (t, t+\Delta)\}}{\Delta t}$$

In practice, the hazard function for this example could be estimated as the proportion of individuals moving per year in a time interval (*e.g.* year 2 to year 3), given that they have aged-in-place to the beginning of the interval (*e.g.* year 2). It can also be interpreted as the instantaneous probability of moving at a specific time.

When explanatory variables are introduced into the model, the hazard function could depend on time and the independent variables. As explained by Afifi and Clark (1990:358-359), the Cox model expresses the associated hazard function as the product of two parts: one that depends on time only, called the baseline hazard, $h_0(t)$; and one that depends on the explanatory variables (X_i), the regression part of the model. Mathematically, represented by them as:

$$h(t,X) = h_0(t) \exp(\beta_1 X_1 + \beta_2 X_2 + ...)$$

It is easier to understand the regression portion of the model, if one looks at the example where there is one explanatory variable β_1 such as poor health with the value of 1 if the respondent had poor health or 0 if not. The hazard function for the respondent with poor health would be:

$$h(t, \text{poor health}) = h_0(t) \exp(\beta_1)$$

while the hazard function for those with good health would be:

$$h(t, \text{good health}) = h_0(t) \exp(0) = h_0(t)$$

If one wished to obtain the ratio of the hazard function of those with poor health to those with good health, this would be:

$$h(t, \text{poor health})/h(t,\text{good health}) = \exp(\beta_1)$$

a constant which does not vary with time. "In other words, the hazard function for group 1 [poor health] is proportional to the hazard function for group 2 [good health]" (Afifi and Clark, 1990:359), hence the name "proportional hazards regression model". In a multi-variate model, the $\exp(\beta_1)$ in the above example would be an estimate of the relative risk of moving for those with poor health as compared with those with good health when controlling for all other explanatory variables (Hirdes and Brown, 1994). This is estimated by maximizing a partial likelihood function. Proportional hazards models are considered to be semi-parametric, since it is not necessary to make any assumptions concerning the form of the baseline hazard, to obtain estimates of the relative risk of moving.

The advantages of using proportional hazards models for analyses such as this one are that: the moving rate can be modelled directly; attrition is accommodated; relative risks can be estimated without specifying the baseline hazard; and explanatory variables that change over the study period can be treated as time-dependent covariates and updated as necessary (Hirdes and Brown, 1994). One disadvantage is that by ignoring the baseline hazard, it is not possible to estimate the magnitude of the hazard experienced by each group, and both may be very small.

A major assumption of the proportional hazard model is that the estimate of the relative risk is constant over time. This assumption can be tested for each explanatory variable by entering the interaction of this variable with time into the model as a time-dependent covariate (Hirdes and Brown, 1994). A statistically significant interaction term would indicate that the explanatory variable has a different association with the outcome variable, such as moving, at an earlier or later stage. The nature of this change can be examined graphically by plotting the combined direct and interaction effects against time. This can be a particularly useful part of the analysis if you have a theoretical reason to believe that the factors that are associated with a move by the elderly may change over time as suggested by the developmental model of Litwak and Longino (1987). The relative advantages of the inclusion of a time interaction when modeling the relationship between health and residential mobility in later life can be seen in the example below.

Example Data

The Ontario Longitudinal Study of Aging was a government initiative, which began in 1959 with a stratified quota sample of 2000 employed, 45 year old men (Forbes et al., 1989; Ontario, 1962). The sample was stratified by the Department of Public Welfare district, type of community (metropolitan, small urban and rural) and to some extent occupation (after 75% of quota was reached). The sampling ratios for the districts varied from five to eight per cent. The subjects were interviewed in person, once a year (except in 1977) until 1978, at which time they were 64 years of age. During this first phase of the study, questions were asked repeatedly about a number of things including changes in living conditions, social activities, marital status, employment, and health. In 1990, a follow-up telephone survey of 545 remaining respondents (aged 76), 49 proxies, and 276 survivors was conducted. Because the study ended when the respondents were 76 years of age, analyses of these data focus on the residential mobility of the "young-old".

To replicate the data usually available for cross-sectional analyses, only the 594 subjects who were alive at the end of the study in 1990 were used for the logistic regression models. For these models, information from earlier interviews was treated as if it had been obtained from retrospective questions asked in 1990.

On the other hand, survival analyses are able to make use of information up to the point that a respondent is lost from the study. Hence, the sub-sample of the Longitudinal Study of Aging data used for the proportional hazards models was composed of the 1063 respondents who were still in the study in the year of retirement or upon turning age 65, which ever came first. This increased sample size is one of the advantages of proportional hazards modeling.

Measurement of the Variables

The analytical examples developed for this paper were part of a larger study examining the relationship of mid-life patterns with residential mobility in later life (Hayward, 1998). The dependent variable used for this comparison, measures whether or not a move was made after retirement or upon turning age 65, whichever came first. At age 65 there are major changes in legal rights and obligations, such as mandatory retirement and access to government benefits (McPherson, 1990), that mark a beginning of a new stage in adult social development. However, early retirement (prior to age 65) also represents a major change in social roles that is known to be associated with residential mobility. For this reason, chronological age is not the only criteria used to define the beginning of the "young-old" period of the life cycle in these analyses. For the logistic models residential mobility was coded as a dichotomy with a value of 0 for those who had not made a move during the period after retirement/age 65, and 1 for those who had made a move. For the proportional hazards models this variable took the form of the time in years to the first move after retirement/age 65.

Two measures of perceived health during mid-life were constructed for the present comparison. The first focused on the period of the onset of poor health during mid-life. The subjects of the Longitudinal Study of Aging were frequently asked how they would rate their health. To create a summary measure of the onset of poor health, the data for each of the years in which the question was asked were dichotomized into two groups, those with good to excellent health, and those with fair, poor, or very poor health. These were then compared and a dummy variable series (health rating) was created, composed of: the early poor health group who reported fair, poor, or very poor health in the period from age 45 to 50; the group who reported poor health at a later time in mid-life; and the reference category of those who consistently reported good to excellent health throughout the mid-life period before retirement or age 65, whichever came first.

The second measure of perceived health examined the stability of health throughout the mid-life period. Most years the respondents were asked if there had been any changes related to health in the past year - coded as better, the same, or worse than the year before. Since elderly residential mobility choices are often thought to be related to declining health, a summary measure was calculated using the number of years in which the subject reported that his health was worse, divided by the number of years for which he answered the health change questions, and multiplied by 100. The resulting proportion of years with declining health was then trichotomized into approximate thirds and coded as three dummy variables: low (reference category), moderate (17 to 22% of mid-life with declining health), and high (over 22% of mid-life with declining health).

For the purposes of this study, the interactions with time were entered as time-dependent covariates in the form of the product of each variable and the natural logarithm of time. The logarithmic transformation of time was particularly appropriate in this analysis since the sample size decreased rapidly after 12 years due to age at retirement.

Results

During the period after retirement or age 65 (whichever came first), approximately a third of the subjects made at least one move before the age of 76 (35.0% of the logistic regression sample; 32.3% of the proportional hazards sample). It is not surprising that those who lived to the age of 76 (the logistic regression sample) were slightly more likely to have moved, since they had more years in which to make a move. Moreover, the proportional hazards sample included those who had died suddenly, who may have made a move if they had lived longer. Both are probably conservative estimates, since it is reasonable to expect that a sizable proportion of the subjects who could not be located for the follow-up survey at age 76 (5.7%) made a move.

The distribution of mid-life health characteristics of the respondents in the two samples can be seen in Table 1. Once again, it is not surprising that the sample used for the logistic regression example were more likely to have good health throughout mid-life, since they survived to age 76. On the other hand, the proportional hazards models included those subjects who had died, many of whom had a relatively early onset of poor health. There was little difference between the two samples in the proportion of mid-life years with declining health.

Logistic Regression Model Results

The results of the logistic regression analyses can be seen in Tables 2 and 3. In both the bi-variate and multi-variate analyses, those respondents with an early onset of poor health during mid-life were more likely to move after retirement/age65 than were those with good health throughout mid-life ($p<.05$). In the multi-variate analysis, those with early poor health were

Table 1 Distribution of mid-life health characteristics for Longitudinal Study of Aging respondents

Mid-life Variable	Logistic Regression Example Percentage[1] (n)	Proportional Hazards Example Percentage (n)
Health rating		
good	50.4 (298)	45.2 (481)
later poor	30.5 (180)	34.1 (363)
early poor	19.1 (113)	20.6 (219)
Years with declining health		
low	20.0 (118)	20.5 (218)
moderate	43.5 (257)	42.1 (447)
high	36.6 (216)	37.4 (398)

[1] Percentages may not sum to 100 due to rounding

Table 2 Bi-variate logistic regression models for first move after retirement/age 65 by mid-life health variables for living respondents (n=591)

Mid-life Variable	Parameter Estimate	Standard Error	Odds Ratio	95% Confidence Limits
Health rating				
good (ref)			1.00	
later poor	.220	.199	1.25	0.84 - 1.84
early poor	.523*	.227	1.69	1.08 - 2.64
Years with declining health				
low (ref)			1.00	
moderate	- .223	.229	0.80	0.51 - 1.25
high	- .302	.237	0.74	0.47 - 1.18

~ p<.10; * p<.05; ** p<.01; *** p<.001

Table 3 Multi-variate logistic regression models for first move after retirement/age 65 by mid-life health variables for living respondents (n=591)

Mid-life Variable	Parameter Estimate	Standard Error	Odds Ratio	95% Confidence Limits
Health rating				
good (ref)			1.00	
later poor	.308	.206	1.36	0.91 - 2.04
early poor	.637**	.241	1.89	1.18 - 3.03
Years with declining health				
low (ref)			1.00	
moderate	- .194	.231	0.82	0.52 - 1.30
high	- .451~	.247	0.64	0.39 - 1.03

~ p<.10; * p<.05; ** p<.01; *** p<.001

almost twice as likely to make a move after retirement or age 65 than were those with good health. These results are not surprising, since one would expect that those with an early onset of poor health would be making earlier support seeking moves than those with consistent good health in mid-life.

On the other hand, no statistically significant relationship was found between the proportion of years with declining health during mid-life and the likelihood of moving after retirement /age 65, in either the bi-variate or multi-variate logistic regressions (p>.05). In the multi-variate logistic regression, there was some evidence that respondents with a relatively high proportion of years with declining health were less likely to move in later life than those with few years of declining health during mid-life, at a marginal level of significance (p<.10). However, the direction of this relationship was opposite to that found for the onset of poor health. These findings appear to be contradictory and difficult to interpret. The lower level of statistical significance and different sign on the regression coefficients for declining health during mid-life could indicate a canceling effect due to the relationship reversing with time. The proportional hazards models allow this possibility to be examined.

Proportional Hazards Model Results

The results from the bi-variate proportional hazards models are summarized in Table 4. The relationship between mid-life health rating and residential mobility in the bi-variate proportional hazards models is similar to that found for the logistic regression models - the direction of the relationship is the same and the risk ratio is slightly smaller. Again, those with an early onset of poor health were more likely to move after retirement/age65 in later life than those with consistent good health throughout mid-life. The interactions with time did not obtain even a marginal level of statistical significance ($p>.10$), which suggests that this relationship did not change with time. These findings are consistent with the interpretation that those who have an early onset of poor health make earlier support seeking moves.

The bi-variate proportional hazards models of a move after retirement/age 65 with the proportion of years with declining health, are interesting. As can be seen in Table 4, the model without the time interaction does not obtain statistical significance ($p>.05$) and is similar to the logistic regression model. Once the time interactions are entered into the model there is strong evidence of a relationship between the number of years of declining health during mid-life and the likelihood of moving, which changes with time. Graphically presented in Figure 2, it can be seen that just after retirement or turning age 65, the men who had a relatively high proportion of years with declining health were least likely to move, followed by those with a moderate proportion of years with declining health. Those with a low proportion of mid-life years with declining health were most likely to move. This finding is consistent with Litwak and Longino's

Table 4 Bi-variate proportional hazards models for a move after retirement/age 65 by mid-life health variables (without and with time interaction, n= 1063)

Mid-life Variable	Parameter Estimate	Standard Error	Risk Ratio	95% Confidence Limits
Health rating				
good (ref)			1.00	
later poor	.170	.126	1.19	0.93 - 1.52
early poor	.380**	.137	1.46	1.12 - 1.91
Health rating				
good (ref)			1.00	
later poor	.024	.207	1.02	0.68 - 1.54
(later poor x ln time)	.122	.138	1.13	0.86 - 1.48
early poor	.193	.228	1.21	0.78 - 1.90
(early poor x ln time)	.155	.150	1.17	0.87 - 1.57
Years with declining health				
low (ref)			1.00	
moderate	-.204	.142	0.82	0.62 - 1.08
high	-.153	.144	0.86	0.65 - 1.14
Years with declining health				
low (ref)			1.00	
moderate	-.645**	.217	0.53	0.34 - 0.80
(moderate x ln time)	.408**	.157	1.50	1.11 - 2.04
high	-.786***	.228	0.46	0.29 - 0.71
(high x ln time)	.557***	.159	1.75	1.28 - 2.38

~ p<.10; * p<.05; ** p<.01; *** p<.001

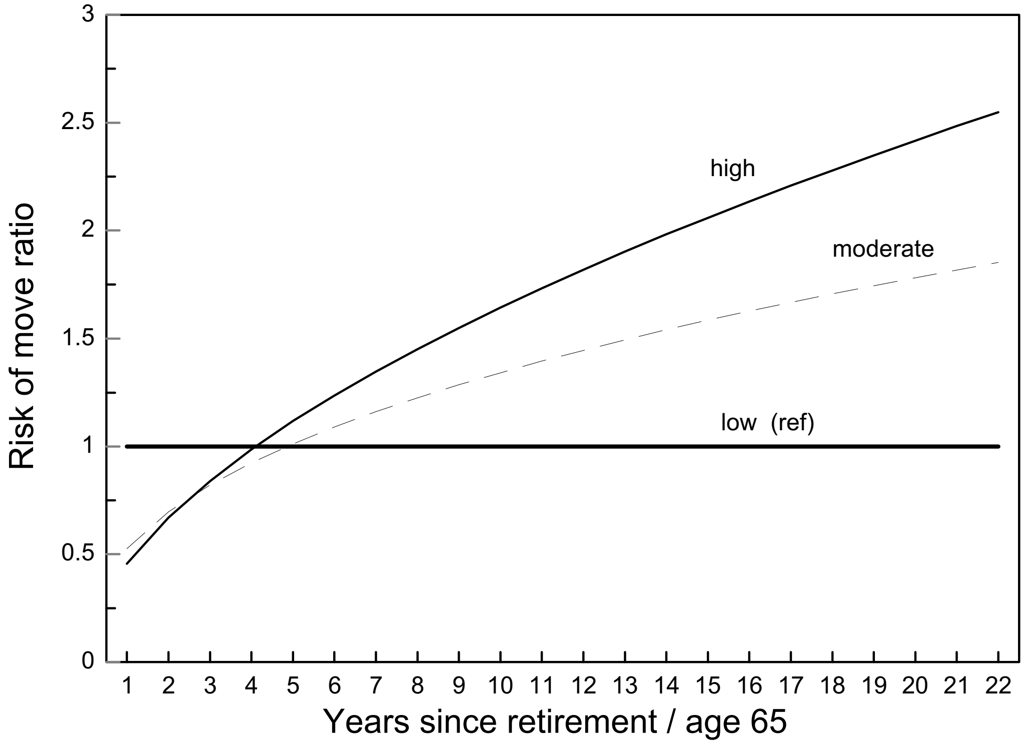

Figure 2 Bi-variate risk of first move ratios after retirement / age 65 by proportion of mid-life years with declining health

(1987) life cycle model of the residential mobility of the elderly which suggests that those making amenity retirement moves are more likely to be in good health. In the present analysis, those with a low proportion of mid-life years with declining health are twice as likely to make a retirement move compared with those with either a relatively moderate or a high proportion of years with declining health. However with time, this relationship changes. Those with a relatively high proportion of mid-life years with declining health are increasingly more likely to make a move as time goes by, and rapidly become the most likely group to move. This is consistent with the argument that those with poor health are increasingly more likely to make a support seeking move.

The same results can be found in the multi-variate proportional hazards model in Table 5, graphically presented in Figure 3. It is interesting to observe how the two measures of mid-life health patterns work together in the multi-variate model. Both have independent associations with the relative risk of moving in later life in the expected direction. Respondents with an early onset of poor health are more likely to move in later life. However, those with a relatively high proportion of mid-life years with declining health are least likely to move shortly after retirement. A possible explanation of this pattern is that those with less stable health histories are more conservative in their mobility choices in later life because of an inability to predict future health, or an anticipation of rapid decline, which makes them less willing or able to cope with the added stress of a move until the need for assistance makes a move unavoidable. On the other hand, those with more stable health, even if it is poor, may be better able to assess the relative costs and benefits of an earlier move.

Summary and Conclusions

This analysis has shown that proportional hazards models which include time interactions can be more informative than cross-sectional logistic regression models, when there is reason to believe that the relationship being studied changes with time. In the multi-variate logistic regression model, the relationship between the proportion of years with declining health during mid-life and residential mobility after retirement/age 65 was negative, contrary to what Litwak and Longino's theory (1987) would lead one to expect, and only obtained a marginal level of statistical significance ($p<.10$). On the other hand, a highly significant relationship ($p<.001$), which reversed with time as predicted by their theory, was found in the multi-variate proportional hazards model with time interactions. Those with a relatively high proportion of mid-life years

Table 5 Multi-variate proportional hazards models for a move after retirement/age 65 by mid-life health variables (n=1063)

Mid-life Variable	Parameter Estimate	Standard Error	Risk Ratio	95% Confidence Limits
Health rating				
good (ref)			1.00	
later poor	.195	.128	1.22	0.95 - 1.56
early poor	.405**	.142	1.50	1.14 - 1.98
Years with declining health				
low (ref)			1.00	
moderate	- .624**	.217	0.54	0.35 - 0.82
(moderate x ln time)	.408**	.157	1.50	1.11 - 2.04
high	- .861***	.230	0.42	0.27 - 0.66
(high x ln time)	.556***	.159	1.74	1.28 - 2.38

~ p<.10; * p<.05; ** p<.01; *** p<.001

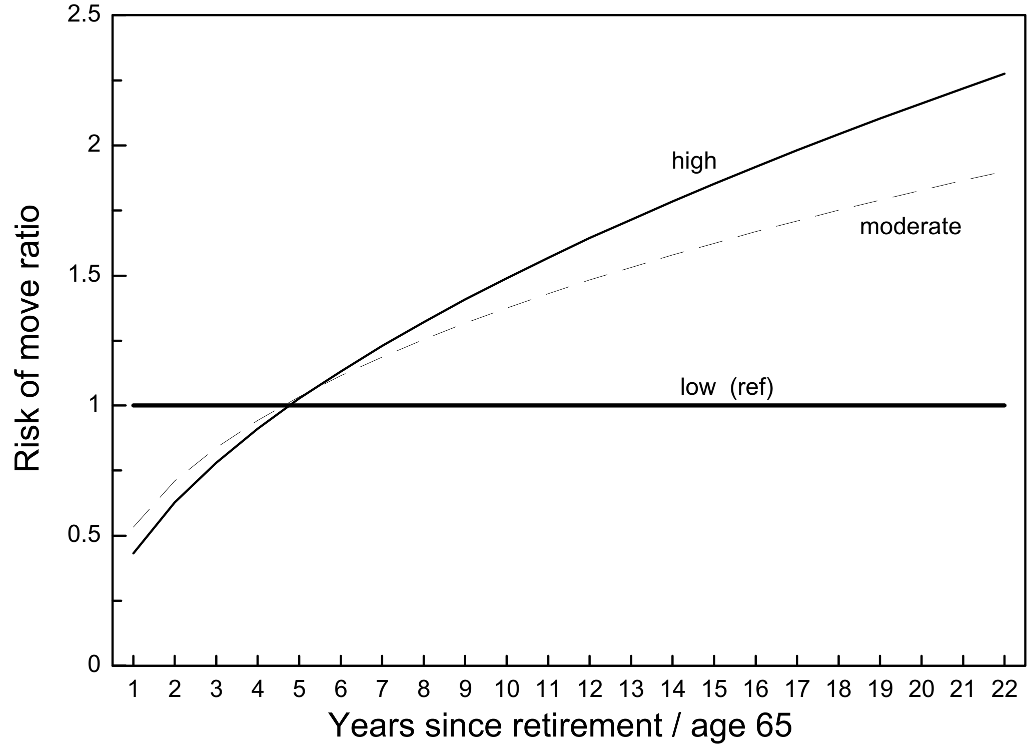

Figure 3 Multi-variate risk of first move ratios after retirement / age 65 by proportion of mid-life years with declining health

with declining health were least likely to make a move just after retirement, which is consistent with the literature which suggests that those with poor health are less likely to make an amenity move. However, with time this group was increasingly more likely to make a move, presumably to obtain assistance. In combination, these findings suggest that for this "young-old" group of men, the canceling effect observed by Patrick (1980) was masking the relationship between the proportion of years with declining health during mid-life and residential mobility in later life in the logistic regression model.

In sum, proportional hazards models can be used to look more closely at the relationship between health and the residential mobility of the elderly and how it changes with time. In future research, as more complete longitudinal data become available which span a larger portion of the life course, it will also be possible to enter health changes in later life into proportional hazards models of residential mobility as time-dependent covariates, to examine both the relationship of health patterns throughout the life course with a move in later life, and the role of health changes in later life as events triggering a move. This is a proportional hazards modeling option which could not be pursued in the present example using the Ontario Longitudinal Study of Aging data, due to 12 year gap in information concerning health in later life.

In conclusion, this paper has illustrated the potential utility of longitudinal proportional hazards models which include health measures and time interactions with health, in a single model of elderly residential mobility. Using this analytical procedure, it will no longer be necessary to develop separate models for different age groups of older people to minimize the canceling effect noted by Patrick (1980).

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