ADJUSTING FOR ATTRITION IN EVENT-HISTORY ANALYSIS

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I. Introduction

One of the major strengths of panel data is that change measures are derived from current reports rather than from recollected statuses. This removes recall error as a source of bias affecting parameter estimates of event-history and other dynamic models. Unfortunately, another characteristic of panel data is that respondents tend to drop out or attrite as the panel ages. This introduces a different potential source of bias in event-history models. Which of these biases is most important (and implicitly, whether panel data is superior to retrospective) is a complicated question -- the answer to which probably varies from one substantive application and set of surveys to another.

The seriousness of the biases introduced to event-history models by panel attrition, for instance, will depend on the amount of attrition and on the relationship between the propensity to leave the sample and the propensity to undergo the substantive change being analyzed. If these propensities are related, then the seriousness of the bias will depend on whether the relationship is confined to the covariates included in the event-history model specification or whether there are excluded, or even unmeasured, factors which affect both propensities. If the covariates do fully account for the relationship between the propensity to change and the propensity to attrite, then the parameter estimates of the substantive model will be unaffected by attrition. If, on the other hand, there is a residual relationship, then explicit corrections for attrition will be required to obtain unbiased estimates.

Previous work (Hill, 1994) suggested that attrition adjusted weighting was not capable of correcting for attrition bias in event history models. The reason is that all the weighting schemes investigated assumed that attrition was a form of independent censoring. The alternative investigated in this paper is modeling attrition and the substantive change of interest as correlated competing hazards (i.e., competing hazards with correlated unmeasured heterogeneity). The estimated net survival function for the substantive change is interpretable as the one we would obtain if attrition were eliminated.

The context of this investigation is divorce (or separation) in the 1986 Panel of the Survey of Income and Program Participation (SIPP). Although the SIPP appears to remain roughly representative of marital statuses once adjustments for attrition are made (see e.g., Singh, 1988), there is concern that it seriously under represents change in marital status, particularly divorces, over the panel period (see e.g., M. Hill, 1987 and D. Hill, 1993). Since divorce is rare, both absolutely and relative to attrition, the parameter estimates from event-history models of it are especially vulnerable to attrition bias. Divorce is also important substantively. Not only is it of interest to behavioral scientist in its own right, but it is also an important determinant of family income, program participation and economic well-being. Furthering our understanding of these conditions in the population is the fundamental reason for conducting the survey.

2. Background

The Survey of Income and Program Participation is a large panel survey of individuals in the United States which has been in operation since 1984 (see Jabine, et al. (1990) for a detailed description of the SIPP). The survey is comprised of a set of panels which are fresh cross-sections introduced annually. The members of each panel are interviewed every four months for roughly two and one-half years and retrospective information on income, employment, program participation and family composition is obtained for each month of the four month reference period.

Although SIPP study procedures call for following and interviewing all panel members when they leave the original sample households, this is not always possible. In the 1986 Panel just under one-fourth of the individuals originally in interviewed households became non-response at some point in the panel period and two-thirds of these individuals were never successfully recontacted (i.e., they attrited). Early on in the survey the concern was raised that attrition was particularly problematic when individuals experienced a marital status change. This is clearly the case in Table 1 which presents the final attrition patterns for husbands and wives in the 1986 SIPP panel. The sample consists of all couples who were married at some point in the panel period. In the vast majority (6,333+130) of these couples neither spouse attrited. The 24-month divorce/separation rate among these panel members was just over two percent. This rate is in sharp contrast to the 60 percent (=100*92/(92+64)) divorce/separation rate among those couples in which one or the other (but not both) partners attrited. The more common situation in which both partners attrite is even more problematic when it comes to estimating marital status change. The reason is that we do not know how many of the 1,004 couples who were still married at the time of their last interview, divorced or separated subsequent to (or concurrently with) their attriting.

Table 1
Ultimate Attrition Pattern
by Known Divorce/Separation Status
(Initially Married Couples 1986 SIPP)

	Status @ Exit	
#Spouses Attriting	**Married**	**Divorced**
0	6,333	130
1	64	92
2	1,004	9

It is also clearly obvious from Table 1 that divorce and attrition can not be independent and adjustment procedures which assume they are are not likely to be effective.

3. Attrition and Divorce as Correlated Competing Hazards

Hill, 1994, found the estimated parameters and implicit cumulative divorce rates from an event history model of divorce to be virtually unaffected by which of four radically different weighting schemes was used. This robustness would be encouraging were it not for the fact that, judging from Vital Statistics data, all of the estimates obtained implied divorce hazards which are far too low. Apparently weighting, at least of the sort used in his study, is not the solution. An alternative is to model attrition and divorce as potentially correlated competing hazards. This way, the effects of attrition on the estimated divorce hazards can be removed by examining the net hazard function for divorce. The major deficiency of the weighting approaches was that they all implicitly assumed independent censoring--i.e., that attritors behaved the same way after their last interview as before. It is quite likely, however, that divorce and attrition are both symptoms of what might be called marital distress--a shared unmeasured risk factor. Thus, individuals in distressed marriages are more likely both to divorce and to attrite than are people in happier marriages. The precise timing of the SIPP interview relative to the timing of a marital disruption is certainly unimportant to these people. It is, of course, crucial to the divorce analyst using SIPP data.

4. The SURF Model

Most competing hazards models also assume independent censoring and, as a result, are not likely to be any more successful in removing bias than the weighting approaches. An exception is the Shared Unmeasured Risk Factor (SURF) model of Hill, Axinn and Thornton (1993). It is most useful to formulate this model in terms of the propensities to divorce (D^*_{ti}) and to leave the sample via attrition (A^*_{ti}). These propensities can be represented according to:

$$\begin{aligned} D^*_{ti} &= \alpha_D + \beta_D' X_{Dti} + \epsilon_{Dti} \\ A^*_{ti} &= \alpha_A + \beta_A' X_{Ati} + \epsilon_{Ati} \end{aligned} \qquad (1)$$

The covariate vectors (X_{Dt} and X_{At}) may or may not have common elements and there may or may not be constraints imposed across the coefficient vectors. The dynamic mechanism assumed is that couples remain in the base state (married and responding) until such time that <u>either</u> D^*_{ti} or A^*_{ti} exceeds some threshold τ. At this time, the couple moves to whichever competing state has the highest propensity score.

Unlike most competing hazards models, the SURF model assumes that the random components of the competing propensities are related via:

$$F(\epsilon_{Dt},\epsilon_{At}) = \exp(-[\exp(-\epsilon_{Dt}/\rho)+\exp(-\epsilon_{At}/\rho)]^{\rho}) \qquad (2)$$

where ρ, known as the index of dissimilarity, is confined to the half-open interval (0,1]. This distribution is known as Gumbel's Type B bivariate extreme-value distribution. The correlation of the ϵ's can be shown to be:

$$r_{\epsilon_{Dt}\epsilon_{At}} = 1-\rho^2 \qquad (3)$$

In the special case where $\rho = 1$, the correlation is zero and the SURF model reduces to the ordinary discrete-time competing hazards model with independent censoring discussed by Allison (1982). In this case it can be shown that the competing hazards model is mathematically equivalent to a single hazard model which treats exits to alternatives as an independent form of right-censoring (see Petersen, 1991). In the discrete choice literature this independence assumption is known as Independence of Irrelevant Alternatives (IIA) and appears in many guises.

Columns 3 and 4 of Table 2 present the results obtained when the independence assumption is relaxed. The estimated index of dissimilarity (the bottom right entry of the table) of .60 is significantly less than 1.0

and implies a correlation between the random portions of the divorce and attrition propensities of roughly .64. Relaxing the IIA assumption also has the effect of reducing the estimated effects of age, foodstamp recipiency, and home ownership. Evidently, while they remain important and significant, part of the apparent effect of these factors in the independent specification was due to their effects on attrition propensities. The most dramatic impact of allowing for non-independent censoring in the form of attrition is to double the estimated effect of whether the month of transition was a "seam" month--from .48 to .96. The meaning of this is that many of the exits recorded at the seam month were attributed to attrition under the independence assumption when they were actually due to divorce. This, of course, is an almost unavoidable consequence of the survey design in which interviews are attempted only periodically--unless at least one of the ex-spouses of a new divorce remains reachable until the next SIPP interview, the case will be recorded as attrition.

It is important to note that the inclusion of the seam month as a predictor in both the attrition and divorce portions of the model is crucial to the stability of the estimates. With this variable included, all specifications of the model examined yielded dissimilarity index estimates in the .5 to .75 range. When the seam month is excluded, on the other hand, the estimated ρ ranged as high as 1.5--a value for which the Gumbel distribution is undefined. The reason this predictor is so important to the estimation of the dissimilarity index is that more than any other factor, knowing its value allows us to distinguish between attrition and divorce. Interestingly, the point estimate for ρ remains at about .6 even when the seam is included as the only covariate in the model. After years of struggling with the "seam" problem it is gratifying that, in this instance at least, it proves to be useful.

Unlike the results of the alternative weighting procedures, the effects of correcting for attrition using the SURF model are quite striking. The twenty-four month cumulative hazard of divorce net of attrition under the independence assumption is 3.6, which is some 20 percent higher than the crude. This is roughly what we would expect from the independence of irrelevant alternatives assumption. The corresponding net divorce hazard rate obtained when the independent censoring assumption is relaxed is 5.6%, which is 56% percent higher than the crude rate. This brings SIPP divorce rate estimates almost exactly up to those implied by Vital Statistics data.

5. Conclusions and Recommendations for Future Research

This paper has examined the effectiveness of two methods of adjusting for attrition in event- history models. Because there was some evidence that attrition in the SIPP had a significant impact on the observed divorce rates, divorce was chosen as the substantive example. The first method of adjusting for attrition consisted of using attrition-adjusted sampling weights in the likelihood function of the event-history model. This method was found to have virtually no effect on the model estimates. The second method involved modeling attrition as a competing alternative means of exiting the base state (married and responding). When the stochastic portions of the propensity to attrite was allowed to be correlated with the corresponding random component of the propensity to divorce, the estimated cumulative hazards function was found to increase significantly for a 24-month rate of roughly 4% to one of over 5.5%. This increase in implied divorce rates brings the SIPP estimates almost in line with those from outside sources.

The results suggest that attrition and divorce are intimately related in that there are shared, or at least correlated, unmeasured risk factors affecting each. This results in a significant stochastic dependence between them which violates the underlying assumption of independent censoring upon which the weighting adjustments are based. Only when this dependency is explicitly recognized and corrected do the estimates change appreciably.

While the results of the Shared Unmeasured Risk Factors competing hazards model are encouraging as a means of correcting event-history model estimates for attrition, more work needs to be done. First, the method needs to be applied to a wide variety of substantive events. Exits from poverty spells and from spells of participation in means tested programs in the SIPP should be investigated. Also, however, the technique should be tested on data from other panel surveys such as the PSID.

Additionally, while none of the weighting schemes investigated in this paper had any discernable effect on the parameter estimates from the divorce event-history model, there are a wide variety of weighting schemes which were not analyzed. Future research should concentrate on those weighting schemes which would allow for non-independent censoring.

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TABLE 2
Structural Model of Attrition and Divorce

	IIA		SURF	
	ATTRITION	DIVORCE	ATTRITION	DIVORCE
CONSTANT	-4.90** (.20)	-4.39** (.26)	-4.70** (.18)	-4.65** (.20)
AGE OF COUPLE	.39** (.12)	-2.57** (.36)	.26* (.12)	-1.82** (.28)
RELATED	- .42** (.15)	-- --	- .39** (.14)	-- --
WHITE	- .27* (.10)	-- --	- .23* (.09)	-- --
NORTH EAST	.21* (.09)	-- --	.20* (.08)	-- --
HOME OWNER	- .17* (.08)	- .44** (.14)	- .21** (.07)	- .37** (.11)
IMPUTATIONS	.65** (.10)	-- --	.60** (.09)	-- --
EMPLOYED	- .22** (.07)	-- --	- .23** (.06)	-- --
FOODSTAMPS	- .48* (.22)	1.29** (.19)	- .26 (.19)	1.00** (.18)
INCOME	- .46** (.18)	-- --	- .40* (.16)	-- --
TIME	.17** (.06)	-- --	.14* (.06)	-- --
SEAM	2.83** (.11)	.48** (.14)	2.70** (.11)	.97** (.16)
ENTERED MARRIED	-1.59** (.12)	-- --	-1.49** (.11)	-- --
CATHOLIC	.46** (.09)	- .16 (.21)	.43** (.08)	- .03 (.16)
LN(L),ρ	-6667.56	1.0 (--)	-6659.74	.60** (.07)

*Significant at the .95 level.
**Significant at the .99 level.