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Note on Research

Transitions from and Returns to Out-of-Home Care

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This longitudinal study employs bivariate probit regression analysis to examine factors associated with returns home from, and reentry to, out-of-home care for 21,484 children placed by child welfare authorities in California. A central focus of the study is to determine the degree to which phenomena unaccounted for in analyses of the returns home from care are correlated with factors unaccounted for in analyses of the process of reentry to care. Although a previous analysis implies direct effects of race and age on foster care reentry, current results suggest that these effects are mediated by other factors. A second concern of the study led to the finding that Aid to Families with Dependent Children eligibility is associated with delays in the return home of children initially placed in kinship foster care.

During the past 2 decades, a substantial portion of research on out-of-home care has focused on the various types of transitions of children into and out of care. The transitions include, among others, initial entries into care, returns to family, adoptions, and reentries to care. The research has not been without serious shortcomings.[1] Among the most significant of these has been investigators' reliance on small, nonprobability samples and their use of designs based on point-in-

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time surveys of children who have made the transitions of interest. Taken together, these practices are likely to give rise to findings that are biased and that fail to capture the dynamics of these transitions.

In recent years, several studies have attempted to overcome these problems.[2] The investigations are multivariate and longitudinal, often covering several years, and employ very large samples derived from statewide administrative foster care databases. Although limited by the relatively narrow range of case information available from these sources, the studies have generated some potentially important findings on attributes and experiences of children, as well as service conditions, that affect the probability that these young people will make transitions to or from out-of-home care. However, data limitations aside, these more recent studies are themselves not necessarily free of problems.

Despite probabilistic sampling of data files on children returned home from care, analyses may nevertheless be subject to selection bias. In the absence of correction, parameter estimates may be in error, thus leading to misinterpretations of conditions affecting the likelihood of children being returned to care. Using data from a probability sample drawn from the population of foster children in California, we show how "uncorrected" and "corrected" findings lead to different interpretations. Because the data and analytic objectives we employ overlap those of two investigations recently published by one of the authors, Mark Courtney, we contrast the results of these earlier investigations to our own.

In the course of undertaking the above analysis, another issue arose, this one dealing with the possibility that children in foster care who are placed with relatives might experience different placement outcomes depending on whether their kin caregivers receive a foster care maintenance payment or an Aid to Families with Dependent Children (AFDC) payment. The exploration of this possibility is a second concern of our study.

Problems of Transitional Research

In 1994, Courtney published the results of a study on family reunification among foster children.[3] The investigation was based on a random sample of 8,741 children, or approximately 10 percent of the children who entered foster care in California for the first time between January 1988 and May 1991, whose domicile status was observed through May 1991. In 1995, Courtney published a second study dealing with reentry into foster care among a sample of 6,831 children who had been reunified with their families between January 1 and June 30, 1988.[4] The foster care status of these children was monitored through June 1991. The samples in the two studies were not nested, although some individuals may have been present in both. Approximately 35 percent of the children initially observed in foster care returned to their families by May 1991, and about 19 percent of the children who returned home from an initial spell in out-of-home care were found to have subsequently reentered care by June 1991. The two studies were based on statewide administrative data on out-of-home care in California.

Largely employing Cox proportional hazard regression models, Courtney examined the impact of children's various attributes and experiences on the rates of family reunification and foster care reentry. In the 1994 study, family reunification rates were significantly affected by several factors, including children's age, race and ethnicity, AFDC eligibility, and type of out-of-home placement.[5] Specifically, infants, AFDC eligibles, African Americans, and children initially placed with relatives returned home at slower rates than did, respectively, older children, non-AFDC eligibles, non–African Americans, and those placed in nonkin homes. Findings in the foster care reentry study indicated, among other findings, that African-American and AFDC-eligible children were more likely to reenter care and that children last placed with relatives were less likely to reenter care.[6] No clear effect of age was observed.

For our current study, two implications of Courtney's investigations are of particular importance: first, controlling for other factors, African-American and AFDC-eligible children appear less likely to experience family reunification than other children; second, and again controlling for other factors, while children in kinship foster placements return to their families at a slower rate than those in nonkin placements, this may reflect greater preparation for family reunification, as evidenced by their relatively lower rate of reentry into care.

There is, however, potential sample selection bias in earlier studies of foster care reentry such as Courtney's. That is, in the foster care reentry study, the sample of children returned home from foster care was not a random sample of children in care. Thus, its members may have been selected in part by unmeasured attributes differentiating children returned home from those remaining in care (e.g., parental substance abuse). If these unmeasured factors in the error terms in predicting return home were related to factors in the error terms in predicting reentry, then the possibility arises that Courtney's estimates of the effects of the independent variables in his analysis are biased. Moreover, the bias could apply to both the unconditional and conditional questions normally raised regarding the factors affecting reentry to foster care.[7]

The selection problem arises in many substantive analyses, usually dealing with criteria that are continuous (e.g., income, achievement, school grades, etc.). For such cases, the traditional approach for correcting bias has been to estimate simultaneously a selection equation (i.e., selected or not for observation) and a substantive equation, including a term that accounts for the correlation between the errors in the substantive equation and the errors in the selection equation.[8]

Bivariate Probit Analysis

We apply here a version of this approach, bivariate probit analysis, in which both the selection and substantive equations involve dichotomous dependent variables.[9] That is, the selection equation predicts y_1, an indicator of whether or not an individual has been returned home (i.e., $y_1 = 1$ or $y_1 = 0$), and the substantive equation predicts y_2, an indicator of whether the individual has been returned to foster care (i.e., $y_2 = 1$ or $y_2 = 0$). The standard approach in dealing with these situations is to assume that the occurrence of each of these events is a reflection of two index variables, y_1^* and y_2^*, where for each individual

$$y_{i1}^* = b_{01} + b_{11}x_{i11} + b_{12}x_{i12} + \ldots + \epsilon_{i1,}$$

and

$$y_{i2}^* = b_{02} + b_{21}x_{i21} + b_{22}x_{i22} + \ldots + \epsilon_{i2,}$$

where $y_{i1} = 1$ if $y_{i1}^* > 0$, 0 otherwise, and where $y_{i2} = 1$ if $y_{i2}^* > 0$, 0 otherwise, $E(\epsilon_{i1}) = E(\epsilon_{i2}) = 0$, and $\mathrm{cov}(\epsilon_1, \epsilon_2) = \rho$. The y^*'s are assumed to be normally distributed, and ρ signifies the correlation between the errors (ϵ_{ij}) in the two equations. In a simple probit model, the probability that, say, $y_{i1} = 1$ is given by

$$\begin{aligned} \mathrm{prob}(y_{i1} = 1) &= \mathrm{prob}(y_{i1}^* > 0) \\ &= \mathrm{prob}(\beta_{01} + \beta_{i1}x_{i11} + \beta_{12}x_{i12} + \ldots + \epsilon_{i1} > 0) \\ &= F(\beta_{i01} + \beta_{i11}x_{i11} + \beta_{12}x_{i12} + \ldots + \epsilon_{i1}), \end{aligned}$$

where $F(\beta_{01} + \ldots)$ represents the cumulative probability distribution of y_1^*. For purposes of exposition, $F(\beta_{01} + \ldots)$ can be written formally as

$$F(\beta_{01} + \ldots) = \mathrm{prob}(-\infty \leq z \leq y^*) = \int_{-\infty}^{y^*} \phi(z)dz,$$

where $\phi(z)$ represents the probability density function of z, and the integral represents the cumulative probability of $\phi(z)$ from $-\infty$ to y^*. In the case of bivariate probit estimation, the concern is with the joint probability of y_1^* and y_2^* and is given by

$$\mathrm{prob}(z_1 < y_1^*, z_2 < y_2^*) = \int_{-\infty}^{y_2^*} \int_{-\infty}^{y_1^*} \phi(z_1, z_2, \rho)dz_1dz_2.$$

Under well-defined conditions, if the correlation between the errors of selection and substance probit equations (ρ) is not significantly different from zero, then the problem of bias in assessing the impact of independent variables may be considered negligible. If the correlation does differ significantly from zero, then the problem of bias is potentially present but is corrected by employing the bivariate probit analysis.[10] Although we employ probit models here rather than hazard models, as in the previous research, the results from the two forms of analyses are quite similar.

Sample and Specification of Equations

The sample was drawn from California child welfare administrative data kept at the Children's Services Archive of the Child Welfare Research Center at the University of California, Berkeley.[11] The sample (N = 21,484) is composed of all abused or neglected children placed for the first time in out-of-home care by California county child welfare departments during 1988 who were 12 years or younger at the time of placement. The age restriction, not present in either of the previous studies, was imposed to ensure that all children in the sample had sufficient time to make the transitions of interest (e.g., allowing 4 years for return home and 2 years for return to care following return home). Of the children in the sample, 11,534 (54%) returned to the care of their families within 4 years of the initial placement in out-of-home care. Of those returned to their families, 2,169 (19%) reentered out-of-home care within 2 years of returning home. The sample we employ is larger and includes some but not all of the members of the two samples analyzed by Courtney. In particular, Courtney's 1995 reentry analysis included children who entered care more than 4 years prior to returning home, whereas the current sample does not.

The models we estimate, however, are similar to those previously used by Courtney.[12] In both the current reunification and reentry models, the independent variables include the following child attributes: gender, race and ethnicity, the presence of chronic health problems at time of entering care, AFDC eligibility at time of entering care, and type of county (e.g., rural or urban) from which the child was placed. Variables unique to the reunification model include the age of the child at the time of foster care entry, the type of home (e.g., two parent or mother only) from which the child was removed at the time of entry, the reason for removal, and the type of initial placement of the child (i.e., group, nonkin, or kin home). Variables entered only in the foster care reentry model were type of placement at the time of reunification, the total number of placements experienced by the child in the course of the initial foster care spell, and the total length of time in care during the spell.

Although correcting for possible selection bias is the main interest in the present investigation, we also pursue an additional line of inquiry, noted earlier, having to do with the possibility that children placed in the homes of relatives might have different experiences in care depending on the reimbursement received by their caregivers. Thus, we are interested in examining a potential interaction effect, one that was not specifically addressed in Courtney's earlier analyses.[13] Courtney used AFDC eligibility as a proxy for family socioeconomic status.[14] Although this is not unreasonable and is a fairly common practice when other measures of family income are unavailable, it may be that AFDC eligibility is more important as an indicator of the level of reimbursement provided to foster caregivers than of the financial resources of families. As it happens, AFDC eligibility can significantly affect the reimbursement paid to the child's foster caregiver *if* the caregiver is related to the child.[15] In California, as in some other states, relatives who provide foster care for AFDC-eligible children are paid more than kin caregivers of ineligible children.[16] This difference in payments may contribute to differing efforts to reunify children with their families and preserve the families after reunification. Specifically, for children placed with relatives, one might expect to see some effect of AFDC eligibility on out-of-home care transitions. This potential interaction effect is represented in the models described below.

Findings

Table 1 presents the frequency distributions of the variables used in the probit and bivariate probit models. Table 2 contains estimates of a probit equation predicting reentry to out-of-home care.[17] This equation does not take into account sample selection effects and is therefore similar in form to Courtney's reentry hazard equation. The results in table 2 for the most part mirror those reported by Courtney in his reentry study.[18] Compared with Caucasians, African Americans are more likely, and Hispanics less likely, to reenter out-of-home care. Children with health problems are more likely to reenter care. Children whose last placement was in kinship care are less likely to return to care than are children from other placements. Placement instability while in care is associated with an increased probability of reentry. Short stays in care are associated with an increased risk of reentry. Finally, children from urban counties other than Los Angeles are less likely to reenter care than are children from other counties. Neither child age nor gender was found to be associated with reentry. In addition, neither AFDC eligibility nor the interaction terms reflecting

Table 1

FREQUENCIES FOR BIVARIATE PROBIT ANALYSIS

	Frequency	%
Return-to-family variable:		
Remained in care	9,950	46.3
Went home	11,534	53.7
Reentry-to-care variable:		
Did not reenter	9,365	81.2
Reentered	2,169	18.8
Age at entry to care (years):		
0–1	6,336	29.5
1–3	5,456	25.4
4–7	5,446	25.3
8–12	4,246	19.8
Initial placement:		
Foster home	13,431	62.5
Kin home	6,588	30.7
Other placement	1,465	6.8
AFDC eligibility:		
Not eligible	10,886	50.7
Eligible	10,598	49.3
Reason for placement in out-of-home care:		
Neglect	15,395	71.7
Physical abuse	3,017	14.0
Sexual abuse	1,769	8.2
Other reason	1,303	6.1
Gender:		
Male	10,682	49.7
Female	10,802	50.3
Health problems:		
No health problem	20,472	95.3
Health problem	1,012	4.7
Race/ethnicity:		
Caucasian	8,576	39.9
Hispanic	4,911	22.9
African American	7,450	34.7
Other race/ethnicity	547	2.5
County type:		
Urban	8,517	39.6
Urban/rural	3,935	18.3
Los Angeles	7,827	36.4
Rural	1,205	5.6
Home from which the child was removed:		
Two parents	2,682	12.5
Mother	17,378	80.9
Father	926	4.3
Other relative	498	2.3
Length of time in care prior to exit:		
Under 1 month	7,239	33.7
1–3 months	1,669	7.8
3–6 months	1,027	4.8
6–12 months	2,049	9.5
1–2 years	3,766	17.5
Over 2 years	5,734	26.7
Number of placements in first episode of care:		
One	10,844	50.5
Two	5,420	25.2
Three	2,413	11.2
Four or more	2,807	13.0

Table 1 (*Continued*)

	Frequency	%
Final placement of first episode:		
Foster home	11,098	51.7
Kin home	8,868	41.3
Other placement	1,518	7.1
Age at exit from first episode (years):		
0–1	1,315	11.4
1–3	3,532	30.6
4–7	3,444	29.9
8–12	2,706	23.5
Over 12	537	4.7

AFDC-placement type interactions were significant predictors of reentry, though the main effect of AFDC eligibility approached significance at $p < .05$.

These findings agree with Courtney's with two minor exceptions.[19] First, in this analysis, Hispanics were slightly less likely to reenter care than Caucasians, while in Courtney's analysis, there was no significant difference between the two groups. Second, we found no age effect. In the earlier research, children between ages 7 and 12 at the time of exit from care had a lower probability of reentry than children under age 1.

In tables 3 and 4, we present results of a bivariate probit analysis predicting family reunification and foster care reentry. The entries in table 3 are parameter estimates from an equation predicting returns home from foster care. Similar to the findings from Courtney's analyses, the table indicates the following.[20] Compared with Caucasians, African Americans are less likely to be returned home. Children under age 1 at the time of placement are much less likely than older children to be returned. Children with health problems are less likely to be returned. Children whose initial placement is due to neglect are less likely to be returned than are children whose placement was based on physical or sexual abuse. Last, children from two-parent families are more likely than children from either mother-only or other-relative families to be returned home.

As noted above, we included in our analyses the possibility of interaction effects involving type of out-of-home care (e.g., foster, relative, other) and children's AFDC eligibility. As seen at the bottom of table 3, while there is no effect of the interaction between regular foster home care and children's AFDC eligibility, there is a significant effect involving kinship care and children's AFDC eligibility. Children's AFDC eligibility does not affect the likelihood of return home among those placed in regular foster homes but *lessens* the likelihood of return among children placed in kinship families. This result was implicit but

Table 2

Probit Model Predicting Reentry to Out-of-Home Care, Maximum Likelihood Estimates

Variable	Coefficient (1)	Standard Error (2)	Approximate p-Value (3)
Constant	−.842	.088	.001
Gender:			
Male*			
Female	−.007	.028	.811
Race/ethnicity:			
Caucasian*			
African American	.146	.034	.001
Hispanic	−.081	.035	.022
Other race/ethnicity	−.077	.082	.350
Age at return to family (years):			
0–1*			
1–4	.035	.048	.470
4–8	−.034	.048	.477
8–12	.014	.050	.784
Over 12	.055	.080	.494
Health problem:			
No problem*			
Problem	.386	.070	.001
AFDC eligibility:			
Not eligible*			
Eligible	.195	.100	.051
Last placement prior to return to family:			
Group care*			
Foster home	−.088	.076	.247
Kinship home	−.430	.080	.001
Number of placements prior to return home	.066	.014	.001
Length of time in care prior to return home:			
Under 1 month*			
1–3 months	−.052	.046	.257
3–6 months	−.089	.055	.106
6–12 months	−.196	.047	.001
1–2 years	−.316	.046	.001
Over 2 years	−.291	.056	.001
County type:			
Urban*			
Urban/rural	.188	.038	.001
Los Angeles	.071	.034	.037
Rural	.174	.057	.002
Placement type by AFDC-eligibility interactions:			
Foster home and AFDC	−.089	.106	.403
Kinship home and AFDC	.168	.111	.131
Log-likelihood		−5,382.4	
χ^2 (df = 23)		385.9	
Significance level		.001	

* Designates omitted category.

not readily observable in Courtney's earlier study because he analyzed family reunification for children in kinship care separately from reunification for children in nonkin foster care.[21]

The most obvious explanation for this finding is that the relatively high foster care payment for kin caregivers of AFDC-eligible children makes return home of these children less financially attractive to the

Table 3

Bivariate Probit Model of Reunification

Variable	Coefficient	Standard Error	Approximate p-Value
Constant	.019	.055	.726
Gender:			
Male*			
Female	−.026	.018	.149
Race/ethnicity:			
Caucasian*			
African American	−.422	.021	.001
Hispanic	−.014	.024	.544
Other race/ethnicity	.050	.057	.376
Age at entry to out-of-home care (years):			
0–1*			
1–4	.280	.024	.001
4–8	.348	.024	.001
8–12	.344	.026	.001
Health problem:			
No problem*			
Problem	−.423	.041	.001
Reason for removal from home:			
Neglect*			
Physical abuse	.346	.026	.001
Sexual abuse	.368	.034	.001
Other reason	−.174	.037	.001
Child's household prior to placement:			
Two parents*			
Mother	−.110	.028	.001
Father	−.012	.050	.806
Other relative	−.371	.061	.001
AFDC eligibility:			
Not eligible*			
Eligible	.002	.066	.980
Child's initial placement:			
Group care*			
Foster home	.008	.048	.863
Kinship home	.053	.050	.295
County type:			
Urban*			
Urban/rural	.009	.025	.708
Los Angeles	.095	.022	.001
Rural	.154	.040	.001
Placement type by AFDC-eligibility interactions:			
Foster home and AFDC	.105	.069	.130
Kinship home and AFDC	−.245	.073	.001
Measure of association between error terms in return and reentry equation (ρ)	.514	.123	.001

* Designates omitted category.

child's extended family than would be the case if the child's kin caregiver was receiving AFDC. Given this explanation, the absence of an effect of AFDC eligibility on the likelihood of return for children placed in nonkin foster care is not surprising: AFDC eligibility has no impact on foster care reimbursement rates for nonkin, and the caregiver is not a member of the child's extended family. This potential

Table 4

Bivariate Probit Model of Reentry to Care

Variable	Coefficient	Standard Error	Approximate p-Value
Constant	−1.180	.102	.001
Gender:			
Male*			
Female	−.007	.026	.793
Race/ethnicity:			
Caucasian*			
African American	.002	.047	.959
Hispanic	−.074	.034	.027
Other race/ethnicity	−.057	.079	.470
Age at return to family (years):			
0–1*			
1–4	.105	.047	.025
4–8	.090	.052	.086
8–12	.142	.054	.008
Over 12	.185	.079	.019
Health problem:			
No problem*			
Problem	.207	.080	.009
AFDC eligibility:			
Not eligible*			
Eligible	.183	.094	.051
Last placement prior to return to family:			
Group care*			
Foster home	−.086	.071	.229
Kinship home	−.395	.076	.001
Number of placements prior to return home	.066	.013	.001
Length of time in care prior to return home:			
Under 1 month*			
1–3 months	−.053	.043	.216
3–6 months	−.093	.052	.073
6–12 months	−.197	.044	.001
1–2 years	−.332	.043	.001
Over 2 years	−.319	.052	.001
County type:			
Urban*			
Urban/rural	.172	.037	.001
Los Angeles	.079	.032	.014
Rural	.189	.055	.001
Placement type by AFDC-eligibility interactions:			
Foster home and AFDC	−.062	.099	.536
Kinship home and AFDC	.121	.104	.245

* Designates omitted category.

financial disincentive for family reunification from kinship foster care has sparked considerable recent discussion by policy analysts and prompted some states to rethink their policies regarding kinship care.[22]

Although our findings appear to imply that relatives caring for AFDC-eligible children are responding to a financial incentive by prolonging provision of foster care, other explanations warrant investigation. For example, it is quite possible that the circumstances associated with AFDC-eligible children in kinship foster care are more difficult to resolve than those of children in other forms of placement. Alternatively, the higher reimbursement rate for kinship care of AFDC-eligible children may affect the likelihood of reunification through its effect on caseworkers as much as through its effect on families. That is, caseworkers may consider where children are "better-off" and compare the financial resources of birth and caregiver families in making this assessment. When children are placed with nonkin, the caseworker may be more inclined to return the child to a parent living in poverty because of the belief that children should live with their biological families. However, when children are living with kin, the caseworker may believe that the child is already with family, thus making financial resources more relevant in the return decision.

Although the finding of an interaction effect of placement type and AFDC eligibility is provocative and should be of interest to policy makers, its explanation requires much more detailed study than is possible here. Administrative data often provide the longitudinal perspective and requisite sample size to describe important outcomes that individuals experience in social service systems and identify some correlates of these outcomes. However, administrative data are often limited in their ability to shed light on *why* a particular outcome is observed. Nevertheless, sound policy analysis and development requires an understanding of the underlying reasons for the functioning of social service programs and their clients. For example, if children remain in kinship care for long periods because their caregivers are "in it for the money," then a reduction in reimbursement for kinship care might not only reduce length of stay in care, it might also significantly reduce the number of kin willing to provide foster care. On the one hand, this could create problems for an already placement-poor child welfare system. On the other hand, if caseworker behavior is driving length of stay in kinship care and the caregivers themselves are not significantly influenced by differences in reimbursement rates, then rate reductions might not seriously reduce the availability of kinship care. Of course, if this is the case, then the proper focus of reform may be in the area of casework practice rather than reimbursement policies. It is unfortunate that child welfare administrative data are not rich enough to help test such competing hypotheses.

Finally, and most relevant to the initial motivation for this analysis, there is a substantial positive association between the error terms in the reunification and reentry equations (see table 3). This suggests that unmeasured factors that influence reunification affect reentry to foster care in the same way. The association also implies that some of the parameter estimates from the single probit analysis presented in table 2 may be biased. The parameter estimates for foster care reentry in table 4 take into account the influence of these unmeasured factors.

The major departures in the findings from the uncorrected equation shown in table 2 occur in the effects of race and age of child at time of placement. Specifically, contrary to the findings shown in table 2, among those children who have been returned home, those who are African American are no longer more likely than others to reenter foster care, and following their return home, children who were infants at the time of initial placement are *less likely* than older individuals to reenter foster care.

How might these findings be interpreted? That is, what unmeasured variables, possibly linked in the same manner to reunification and to reentry, might account for the change in the results from the single probit equation? A number of factors that might be expected to influence both family reunification and reentry to care (e.g., homelessness, domestic violence, social support from extended family) might also be associated with child age and race. For example, one plausible hypothesis is that these findings are due to the unmeasured influence of factors related to parental substance abuse. Prior research has indicated that African-American parents are more likely to be substance abusers than Caucasian parents and that most infants who enter the care of the child welfare system experience in utero drug or alcohol exposure.[23] It may be that African-American children returning home from foster care are more vulnerable to foster care reentry because of possible continued parental substance abuse. In a similar manner, infants may be less likely than older children to reenter care when their own health problems associated with substance exposure, and the ongoing substance abuse by their parents, are factored into the equation. Some indirect support for this hypothesis is found in our data. Tables 2 and 4 show that the impact of child health problems is substantially weaker in the reentry equation that corrects for selection. Thus, it may be that something related to child health is also important among the unmeasured variables affecting reentry. Substance abuse is a clear possibility.

In all likelihood, however, the unmeasured contributors to reunification and reentry that play a role in our findings comprise a constellation of factors. It is unfortunate that the administrative data that are readily available to simultaneously study family reunification and reentry capture very little of the heterogeneity of families and children. In any event, by not taking these considerations into account, the single probit reentry equation (table 2) erroneously implies that there is a race effect in reentry to foster care and that infants are no less likely to reenter care than other children.

This research documents the importance of carefully taking into account the problem of sample selectivity and the need to investigate interactions implied by theory or policy. The failure to control for selectivity can lead to biased estimations. The failure to investigate potentially important interactions may lead to oversimplified findings. In either case, the results of analyses will be misleading.

Substantively, our analyses suggest that returns home from foster care are affected by an interaction such that, for reasons as yet unknown, AFDC eligibility is associated with delayed return home for children initially placed with relatives. The finding clearly needs further exploration and explanation. Our analyses also suggest that the effect of race and age on reentry to out-of-home care is complex. That is, in contrast to the implications of earlier studies suggesting that there are direct effects of race and age on foster care reentry, our results suggest that the effects are mediated by other, unidentified factors. This also needs further study and explication. Finally, although our study indicates that administrative data can be a relatively inexpensive source of important information, administrative data must continue to be improved and, in many cases, supplemented with other forms of research (e.g., surveys, panel studies) designed to answer specific questions.

Notes

We would like to thank an anonymous reviewer for offering alternative explanations for some of the study's findings and the staff of the Child Welfare Research Center at the University of California, Berkeley, for access to the data and for their helpful comments on this study.

1. For methodological critiques of this research see, e.g., Richard P. Barth, Mark Courtney, Jill Duerr Berrick, and Vicki Albert, *From Child Abuse to Permanency Planning: Child Welfare Services Pathways and Placements* (New York: Aldine de Gruyter, 1994).

2. See, e.g., Mary I. Benedict and Roger B. White, "Factors Associated with Foster Care Length of Stay," *Child Welfare* 70 (1991): 45–58; Mark E. Courtney, "Factors Associated with the Reunification of Foster Children with Their Families," *Social Service Review* 68 (1994): 81–107, and "Reentry to Foster Care of Children Returned to Their Families," *Social Service Review* 69 (1995): 226–41; Mark E. Courtney and Yin-Ling Wong, "Comparing the Timing of Exits from Substitute Care," *Children and Youth Services Review* 18 (1996): 307–34; Robert M. Goerge, "The Reunification Process in Substitute Care," *Social Service Review* 64 (1990): 422–57; and Robert M. Goerge, Fred H. Wulczyn, and Allen W. Harden, *Foster Care Dynamics, 1983–1993, California, Illinois, Michigan, New York and Texas: An Update from the Multistate Foster Care Data Archive* (Chicago: University of Chicago, Chapin Hall Center for Children, 1995), and *Foster Care Dynamics, 1983–1992, California, Illinois, Michigan, New York and Texas: A Report from the Multistate Foster Care Data Archive* (Chicago: University of Chicago, Chapin Hall Center for Children, 1994).

3. Courtney, "Factors Associated with the Reunification of Foster Children" (n. 2 above).

4. Courtney, "Reentry to Foster Care" (n. 2 above).

5. Courtney, "Factors Associated with the Reunification of Foster Children" (n. 2 above).

6. Courtney, "Reentry to Foster Care" (n. 2 above).

7. The unconditional question takes the form: "Given a sample of children in foster care, what are the phenomena affecting the likelihood of returning home and then reentering care?" The conditional question takes the form: "Given a sample of children who have returned home from foster care, what are the factors affecting the likelihood of reentering care?" For a description of the ways in which sample selection effects can

bias the answers to both of these kinds of questions, see, e.g., Richard A. Berk and Subhash C. Ray, "Sample Selection Bias in Sociological Data," *Social Science Research* 11 (1982): 352–98.

8. James Heckman, "Sample Selection Bias as a Specification Error," *Econometrica* 47 (1979): 153–61.

9. For a description of bivariate probit analysis, see, e.g., William H. Greene, *Econometric Analysis,* 2d ed. (New York: Macmillan, 1993).

10. Wynand Van de Ven and Bernard Van Praag, "The Demand for Deductibles in Private Health Insurance," *Journal of Econometrics* 17 (1981): 229–52.

11. For a more detailed description of the Children's Services Archive, see Barbara Needell, Daniel Webster, Richard P. Barth, Kristen Monks, and Michael Armijo, *Performance Indicators for Child Welfare Services in California, 1994* (Berkeley: University of California at Berkeley, School of Social Welfare, Child Welfare Research Center, 1996).

12. See Courtney, "Factors Associated with the Reunification of Foster Children," and "Reentry to Foster Care" (both in n. 2 above), to compare our models with those of Courtney. The exceptions include the following: First, Courtney employed models of family reunification separately for children placed in nonkin and kin homes, whereas we employed a single model including a variable identifying type of initial placement (i.e., kin/nonkin). Second, although Courtney's model of family reunification included a dichotomous variable identifying families that had received preplacement services, this indicator, found nonsignificant in Courtney's analyses, was omitted from our model. We omitted this variable as a result of information, not available at the time of Courtney's analyses, indicating that this variable is not recorded consistently throughout California. Finally, our model of reentry included a variable, namely, the type of county from which the child was initially placed, not included in Courtney's analysis of this transition.

13. Interaction effects refer to situations in which the effects of a given variable of interest depend on the values of some other variable or variables. For example, in this study, we consider the possibility that the observed effect of kinship care placement in lowering the likelihood of family reunification is restricted to children in kinship care who come from AFDC-eligible families. While interaction effects are easily estimated in principle, their total number can be exceedingly large and their interpretation ambiguous. Thus, the study of interaction effects is highly selective and usually guided by important theoretical or policy considerations.

14. See Courtney, "Factors Associated with the Reunification of Foster Children," and "Reentry to Foster Care" (both in n. 2 above).

15. For a description of states' policies regarding reimbursement for kinship care, see James P. Gleeson and Lynn C. Craig, "Kinship Care in Child Welfare: An Analysis of States' Policies," *Children and Youth Services Review* 16 (1994): 7–31.

16. Reimbursement rates can differ between kin caregivers who receive foster care maintenance payments and those who receive AFDC for two reasons. First, foster care maintenance rates are higher than AFDC payments. For example, the monthly AFDC rate in California for children 0–4 years old was $293 in 1996, whereas the foster care payment was $345 per month. Second, per capita AFDC payments decrease with family size, whereas foster care payments increase proportionally to the number of children placed with a given family. Thus, when sibling groups are placed, the difference between the applicable AFDC and foster care payments can be quite large. For example, the AFDC reimbursement rate in California for three foster children under 4 years old in 1996 would have been $594 per month, whereas the monthly foster care rate for the same three children was $1,035. The differences are even larger when older children are involved since foster care maintenance rates are higher for older children but AFDC rates are not.

17. The entries in col. 1 of table 2 are estimates of the effects of the variables on the value of y^*, which in turn determines the probability that $y = 1$ (foster care reentry). A value greater than zero indicates an increase in the probability of reentry, whereas a value less than zero indicates a decrease. The entries in col. 3 indicate the probability that these parameter estimates could have arisen by chance alone.

18. Courtney, "Reentry to Foster Care" (n. 2 above).

19. There are several possible explanations for the dissimilar findings. In general, these explanations involve differences in either the sample or independent variables employed in our study versus those used in Courtney's reentry study. For example,

unlike the present study, the earlier investigation failed to include controls for the type of county in which placement occurred. If the type of county is associated with the age that children return home from care (e.g., children in urban counties are returned home at an earlier age), this could account for the absence of an age effect in this analysis. In a similar manner, if county type is associated both with the race of children in care and the likelihood that children will reenter care, then the addition of county type to an analysis of reentry could lead to different parameter estimates for the effect of race on reentry. The absence of an AFDC-eligibility effect in our model resulted from inclusion of an AFDC-placement interaction term. In preliminary models that did not include the interaction term, we in fact found a significant main effect of AFDC eligibility on reentry. Finally, our use of an entry cohort of foster children, while essential for testing for selection effects across returns home and reentries to care, results in a sample that differs from the exit cohort sample Courtney used for his reentry analysis. Children in long-term care were included in Courtney's sample but excluded from ours, and this may contribute to the differences in the findings. Similarly, nearly 15 percent of Courtney's reentry sample was over 13 years old at the time of return home from care, whereas less than 5 percent of our sample was over age 12.

20. Courtney, "Factors Associated with the Reunification of Foster Children" (n. 2 above).

21. Ibid.

22. For discussions of trends in state kinship care policy and reimbursement practices see, e.g., Jill Duerr Berrick and Barbara Needell, "Recent Trends in Kinship Care: Public Policy, Payments, and Outcomes for Children," in *The Future of Children* (in press); and Mark F. Testa, "Kinship Foster Care in Illinois," in *Child Welfare Research Review,* ed. Jill Duerr Berrick, Richard P. Barth, and Neil Gilbert (New York: Columbia University Press, 1997), 2:101–29.

23. For data pertaining to the relationship between race and perinatal substance exposure, see, e.g., William A. Vega, Bohdan Kolody, and Jimmy Hwang, "Prevalence and Magnitude of Perinatal Drug Exposure in California," *New England Journal of Medicine* 12 (1993): 850–54; and for estimates of the impact of the problem on the child welfare system, see, e.g., U.S. General Accounting Office, *Foster Care: Prenatal Drug Abuse Has an Alarming Impact on Young Children* (Washington, D.C.: U.S. General Accounting Office, 1994).