Gender and the Nexus of Parental and Community Control:

The Social Context of Sexual Transition Behavior During Early Adolescence[*]

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Gender and the Nexus of Parental and Community Control:

Contextual Effects on the Timing of First Intercourse

We integrate social capital and collective efficacy theory with emerging research on the intersection of parental supervision, child gender, and neighborhood to develop a contextual model of the timing of first intercourse during early adolescence. We hypothesize that the effects of family and neighborhood controls on the timing of first intercourse will exhibit opposing patterns by gender. Family controls are likely to be more consequential for girls while neighborhood controls will target the more visible behavior of boys. Findings from multilevel discrete-time models of the timing of first intercourse indicate that family supervision exerts significantly greater influence on the timing of first intercourse for girls than boys. Neighborhood supervision capacity—captured by the concept of collective efficacy—delays sexual onset for boys. For girls, the effect of collective efficacy is conditional, emerging only for those who experience lower levels of parental supervision.

Recent decades have witnessed marked shifts in the occurrence of sexual activity during early adolescence (Allen Guttmacher Institute 1994). Though the overall prevalence of adolescents who were sexually active declined during the 1990s, the proportion of younger adolescents experiencing sexual onset increased (Terry and Manlove 2000). Adolescent sexual activity, particularly at earlier ages, has implications for the subsequent occurrence of sexually transmitted disease, unwanted fertility, diminished mental health, and other risk behavior (Meier 2002; Resnick, et al. 1997; Tubman, Windle, and Windle 1996). Consequently, understanding the origins of early adolescent sexual behavior has become a focus of increasing concern (Albert, Brown, and Flanigan 2003).

An emerging emphasis in research on the etiology of adolescent sexual behavior and problem behavior more generally is the role of parent-child relations. In particular, family management and supervision practices have been an important focus of recent attention, yielding significant findings on the role of parental and caregiver control in reducing the likelihood of social and behavioral problems, poor academic achievement, and school adjustment difficulties (Vandivere, Tout, Capizzano, and Zaslow 2003). Research on youth sexual behavior also offers evidence that children who are more closely supervised are less likely to be sexually active (Miller 1998). To date, however, careful investigation of parental supervision and monitoring effects on adolescent sexual transitions has been limited. The relative dearth of research on parental control has persisted despite its implications for youth's exposure to another important determinant of adolescent sexual behavior—the neighborhood environment.

With the emergence of theory and research on the role of geographic context in organizing adolescents' behavior, researchers have increasingly considered the potential for neighborhood characteristics to influence the prevalence of early sexual activity above and beyond the contribution of family and individual level factors (Billy, Brewster, and Grady 1994; Brewster 1994b). Recent contributions to neighborhood theory on adolescent behavior have highlighted the role of neighborhood structural disadvantage in diminishing the community-level capacity to mobilize on behalf of shared goals (Leventhal and Brooks-Gunn 2000; Sampson 1997). Sampson and colleagues have described this aspect of community competence as *collective efficacy* and emphasized its implications for the effective socialization and supervision of children and adolescents (Sampson, Morenoff, and Earls 1999).

We extend the collective efficacy approach to include insights from the expanding literature on the intersection of parental supervision patterns, child gender, and neighborhood. Although mounting empirical evidence points to more powerful effects of neighborhood characteristics on well-being for boys (Leventhal and Brooks-Gunn in press), research on neighborhood influence has yet to explicitly address the potential for gender differences in the effects of social environments on behavior. We argue that early adolescent boys experience lower levels and weaker influence of parental controls by comparison with girls, contributing to differences in the potential for neighborhood influence by gender. To the extent that girls' exposure to neighborhood environments is relatively limited by comparison with boys, features of community context may exert less influence on their behavioral outcomes. Similarly, the more visible behavior of boys in

neighborhood contexts may elicit collective supervision efforts directed specifically toward control of their behavior. Parental supervision practices, in turn, may condition the influence of neighborhood characteristics, with youth who are insulated from external environments exhibiting little response to community variation (Furstenberg, Cook, Eccles, Elder, and Sameroff 1998; Kupersmidt, Griesler, Rosier, Patterson, and Davis 1995).

In what follows, we develop and test this gendered approach to the influence of familial and community controls on the timing of sexual onset. Our analyses employ multiple data resources drawn from the Project on Human Development in Chicago Neighborhoods (PHDCN), a study designed to elucidate the role of neighborhood context in the lives of children and youth (Earls and Buka 1997). Our goals are threefold: First, we examine the micro level etiology of early adolescent sexual activity for boys and girls, highlighting the potential for differences in the effects of family supervision and monitoring by gender. Second, we explore whether neighborhood characteristics exert independent and differential influence on the timing of first intercourse for boys and girls, emphasizing the extent to which collective efficacy directly affects sexual transition behavior, above and beyond features of neighborhood structure. Third, we consider the potential for interaction effects between collective and family supervision.

THEORETICAL BACKGROUND

PARENTAL CONTROL AND ADOLESCENT SEXUAL TRANSITION BEHAVIOR

A number of theoretical models linking family process and adolescent behavior have emphasized the role of parental control and family management (Crouter and Head 2002; Gottfredson and Hirschi 1990; Patterson 1982; Sampson and Laub 1993). In this view, "direct" parental controls—or the extent to which parents and caregivers supervise and monitor children's behavior in the context of clearly conveyed behavioral rules—are important in limiting opportunities and incentives for engaging in problem behavior. Indeed, parental supervision has been demonstrated to reduce the likelihood of a range of problem behaviors among youth (Loeber and Stouthammer-Loeber 1986) and is likely to lead to later onset of sexual activity (Miller, McCoy, Olson, and Wallace 1992; Miller 1998; Small and Luster 1994).

Parental controls, however, may have distinct implications for the behavioral trajectories of boys and girls. Developments in control theoretical orientations (Hagan 1988; Hagan, Simpson, and Gillis 1987) and recent empirical research (Galambos and Maggs 1991) suggest the possibility of differential application and impact of parental controls by child gender. In a power-control theory of common delinquency, Hagan (1988) emphasizes the relative freedom from family-based social controls experienced by boys. Although Hagan seeks to explain gender differences in the occurrence of common delinquency, we suggest that gender-specific application of family based social controls is likely to be *at least as* pronounced with respect to limitations on sexual behavior.

Moreover, differential application of parental control may be manifest both in the total *quantity* of supervision as well as the *impact* of regulatory practices. Although boys are likely to be subject to parental regulation, they may simultaneously be encouraged to cultivate a preference for riskier behavior by comparison with girls. For girls, socialization practices that encourage dependence on, and obedience to, parents and caregivers more heavily than boys are likely to reinforce the effect of parental controls on girls' behavior. Girls may also experience more strict enforcement of supervisory rules due to prevalent gendered cultural conceptions of appropriate adolescent conduct. The perceived social consequences of early sexual involvement for girls, including pregnancy and childbirth, disrupted education, and diminished marital prospects are also likely to focus parental regulatory practices on the *sexual* activity of girls (as opposed to, e.g., the violent or substance abuse behavior of boys), strengthening the impact of parental controls on this aspect of girls' behavior. Consequently, girls may experience both more extensive and etiologically significant effects of parental control on the timing of sexual transitions.[1]

We also address the role of youth emotional attachments toward caregivers. Emotional attachments between parents and children facilitate socialization to value orientations that support delaying sexual activity and raise the psychological costs of engaging in risk behaviors such as early sex (Small and Luster 1994). Research on parenting more generally has shown that parent-child relationships characterized as warm

[1] An alternative gendered hypothesis would view the quantity and impact of parental control as greater for boys, perhaps due to their increased likelihood of participating in problem behavior by comparison with girls. Although some evidence suggests that boys may be more vulnerable to the effects of familial stressors such as divorce and parental conflict (Moffitt, Caspi, Rutter, and Silva 2001), this research has focused on nonsexual problem behavior and has not directly considered the impact of parental supervision.

and supportive are associated with lower risk of adolescent problem behavior (Jessor and Jessor 1977; Miller 1998; Steinberg and Sheffield-Morris 2001; Steinberg and Silk 2002). Again, with respect to sexual behavior, girls may experience the deterrent effects of attachments to parents more powerfully than boys.

An exclusive focus on links between micro-level contexts and adolescent sexual risk behavior, however, ignores the larger social environment in which youth are embedded. Accordingly, we turn to theoretical orientations emphasizing macro-level influences, developing a collective efficacy perspective on the community context of youth sexual behavior.

NEIGHBORHOOD STRUCTURE AND SEXUAL ONSET

Research on the neighborhood context of adolescent sexual behavior has focused largely on the structural characteristics of communities thought to concentrate risky sexual behavior. In this tradition, collective efficacy theory links key aspects of urban neighborhood structure with variability in dimensions of social organization relevant to adolescent outcomes. Of central theoretical interest is the concentration of economic disadvantage within contemporary urban neighborhoods. The emergence of stark socioeconomic inequalities at the neighborhood level (Wilson 1996) and associated patterns of racial segregation (Massey and Denton 1993; Krivo and Peterson 1996) during the past few decades have called attention to the consequences of resource poor environments for the development of children and adolescents (Jencks and Mayer 1990). Above and beyond family and individual factors, widespread poverty and unemployment

substantially limit the availability of economic and social resources with which to sustain key institutions including schools, local voluntary organizations, and more informal neighbor networks (Wilson 1996). Indeed, growing evidence suggests that neighborhood socioeconomic disadvantage is associated with early sexual debut and related outcomes, such as teenage pregnancy, after accounting for individual and family background characteristics (Baumer and South 2001; Billy, Brewster, and Grady 1994; Brewster 1994a; Brewster 1994b; Brewster, Billy, and Grady 1993; Brooks-Gunn, Duncan, Klebanov, and Sealand 1993; Ku, Sonenstein, and Pleck 1993; Ramirez-Valles, Zimmerman, and Newcomb 1998).

A second structural emphasis—and a potential outcome of economic disadvantage—is the level of instability in residential tenure. Rapid turnover in local population and diminished rates of homeownership inhibit the emergence of viable social networks and weaken sentiments of attachment to local communities. Third, ethnic and racial heterogeneity may complicate informal communication within neighborhoods, impeding the development of cross-group ties.

The consequences of structural disadvantage—weak local institutions, sparse or fragmented informal networks and attenuated neighborhood attachments—diminish the capacity of communities to come together to achieve common goals, including the informal social control of neighborhood youth (Sampson 1997). Thus structural characteristics of neighborhoods may influence *community-level* mechanisms that are consequential for adolescent behavior; specifically, neighborhood social cohesion and informal social control capacity, or *collective efficacy*.

NEIGHBORHOOD COLLECTIVE EFFICACY AND SEXUAL ONSET

The concept of collective efficacy,[2] rooted in the work of Bandura (1986), refers to the level of mutual trust, solidarity, and shared values among community residents (i.e., social cohesion) combined with shared expectations for pro-social action (i.e., informal social control) (Portes 1998; Sampson, Raudenbush, and Earls 1997). Cohesion among residents serves as social capital that may facilitate joint action on behalf of the community. Attachments to the local community and trust that neighbors are invested in a core set of recognized values makes collective efforts to achieve shared objectives possible. Collective efficacy emerges with the crystallization of expectations that neighborhood residents will act on behalf of shared values and collective goals, including those oriented toward beneficial outcomes for local youth.

Like its individual level counterpart (self-efficacy), collective efficacy is goal-specific. Extant research on collective efficacy has established it's regulatory impact on crime and delinquency within urban neighborhoods (Sampson, Raudenbush, and Earls 1997). However, collective efficacy relevant to the control of neighborhood adolescent delinquency is not necessarily equivalent to the community level ability to regulate risky sexual behavior. Though evidence suggests that communities maintaining high levels of collective efficacy with respect to the informal social control of crime and criminogenic activity tend to be capable of influencing a wide range of outcomes (Browning and

[2] Collective efficacy *theory* encompasses the set of causal links between neighborhood structural disadvantage, intervening social organization, and the resulting broad range of deficits experienced among residents of vulnerable neighborhoods. The *concept* of collective efficacy, however, captures the key dimensions of social organization hypothesized to channel the impact of structural disadvantage on neighborhood outcomes.

Olinger-Wilbon 2003), the conceptual links between neighborhood social organization and the typically more private and consensual aspects of adolescent sexual behavior require clarification.

How might communities with high levels of collective efficacy contribute to the regulation of behavior that, unlike delinquency, often occurs "behind closed doors"? Recent evidence suggests that a substantial proportion of adolescent sexual activity is initiated in private spaces, potentially beyond the reach of community-oriented regulatory strategies. In fact, a majority of adolescents (56%) report that their first sexual experience occurred at home or in their partner's parents' house (Child Trends 2002). The situational characteristics of adolescent sexual behavior suggest that the community level capacity to influence its prevalence may rest, in part, on the existence of social ties between parents of (potentially) sexually involved youth. That is, what Coleman (1990) has described as *intergenerational closure*—social ties linking parents with the parents of their children's friends—may provide an important source of community based social capital relevant to the management of adolescent sexual behavior. Coleman has argued that intergenerational closure enhances social control capacity significantly beyond what would be possible for individual parents without such ties. The capacity to efficiently disseminate information, share supervision responsibilities, and reinforce norms regarding acceptable behavior across extensive neighborhood based networks of parents and caregivers is likely to reduce the prevalence of early sexual onset among neighborhood youth.

Collective efficacy theory, however, points to the combined role of community cohesion and expectations for beneficial action as the crucial element of social organization relevant to youth outcomes. In this view, social capital as represented by the structural linkage among parents of interconnected children represents a stock of social resources that *may* be mobilized to achieve the shared goal of limiting adolescent exposure to sexual risk. The prevalence of social ties alone, however, is not sufficient to yield collective benefits for youth (Sampson, Morenoff, and Earls 1999). The effective regulation of adolescent behavior is facilitated when extra-familial intergenerational social relationships combine with expectations for adult action on behalf of local youth. These emergent normative orientations, in turn, may lead not only to more effective joint supervision of local youth, but to the development of local institutions and opportunities (e.g., after school programs, parks and recreation centers) that discourage early sexual behavior as a byproduct of adolescents' social and psychic investment in conventional activities.

Accordingly, we expect that the combination of social cohesion, intergenerational closure, and expectations for active support and supervision of youth by local adults will exert regulatory effects on the prevalence of early adolescent sexual activity. Empirical investigation of the determinants of youth sexual behavior has yet to examine the unique effects of neighborhood level intergenerational closure and collective efficacy on adolescent sexual behavior.

THE NEXUS OF GENDER, PARENTAL SUPERVISION, AND COLLECTIVE EFFICACY

Collective efficacy theory does not explicitly hypothesize distinct effects of community characteristics by gender. Nevertheless, the potential for gender specific effects of family social controls points to the possibility that extra-familial environments may have unique effects for boys and girls. In the extreme case of relative freedom from family based social control for boys compared with highly restrictive environments for girls, the degree of exposure to neighborhood context for the latter will be limited (Ensminger, Lamkin, and Jacobson 1996). Assuming that collective social control efforts are gender neutral, girls will experience less significant total influence due to lower levels of extra-familial exposure. Yet, the assumption of equal application of neighborhood-based social control may also be questionable to the extent that boys are perceived to be a greater threat to the collectivity and in need of more vigilant regulation. In this gendered collective control hypothesis, boys are more exposed to, and controlled by, neighborhood environments and will experience greater effects of collective supervision regardless of differential familial control.[3]

The relatively rare analyses of neighborhood effects estimated separately for boys and girls offer some evidence of gender-specific contextual influence (Leventhal and Brooks-Gunn 2000). Ramirez-Valles, Zimmerman, and Juarez (2002), for instance, identified

[3] An alternative gendered collective control hypothesis would argue that collective supervision capacity is more oriented toward the behavior of girls than boys, particularly with respect to sexual behavior. The social control of female sexuality may operate at multiple levels (the family, community, and state). Hagan, (1988), however, argues that while family controls are of greater significance for girls, state controls focus on the behavior of boys and men, suggesting that macro level *informal* social controls may also be male-focused (see also Black [1976] for a discussion of the inverse relationship between the quantity of controls applied by family and kinship groups and the state).

distinct predictors of the timing of first intercourse for African American boys and girls. In particular, living in a neighborhood with low housing values accelerated sexual onset for boys but was not a significant predictor for girls.[4] Neighborhood effects research on nonsexual outcomes has also found evidence of more powerful associations of neighborhood characteristics for boys with respect to school dropout (Crane 1991; Ensminger, Lamkin, and Jacobson 1996) and math achievement (Entwisle, Alexander, and Olson 1994).

Finally, the simultaneous emphasis on familial and community control of early adolescent sexual activity calls for consideration of their intersecting effects. The effect of collective efficacy may be more pronounced for youth who experience lower levels of familial supervision and greater exposure to neighborhood environments (Coley and Hoffman 1996; Kupersmidt et al. 1995). We consider this hypothesis separately for boys and girls, allowing a test of the gendered collective supervision hypothesis. Given comparably low levels of parental supervision, collective control focused more heavily on boys would yield larger effects of this neighborhood level process on boys' sexual transitions. No studies to date have specifically examined the potential for opposing effects of familial and collective control by gender, nor have contextual models of adolescent problem behavior considered the *nexus* of these controls—that is, the potential for interactive effects of parental and collective supervision.

[4] Effects of neighborhood social processes on boys' sexual onset might be expected to limit girls' sexual behavior as well, given that girls are involved in the vast majority of boys' sexual debuts. Yet, comparable neighborhood effects by gender would be required only if boys' and girls' sexual onset experiences occurred with opposite sex partners from their neighborhoods who were also experiencing sexual onset. In fact, a substantial proportion of adolescents' first sexual intercourse experiences may occur with older and/or sexually experienced partners. Early adolescent dating or romantic relationships with older partners, for instance, are significantly more likely to involve sexual intercourse (Albert, Brown, Flanigan 2003).

ADDITIONAL FAMILY, PEER, AND INDIVIDUAL LEVEL DETERMINANTS OF SEXUAL ONSET

Although we focus on the role of familial and collective supervision from the standpoint of our theoretical model, we also address a range of additional potential determinants of early adolescent sexual activity. First, we consider the role of family socioeconomic status (Haveman and Wolfe 1994), parental absence (Flewelling and Baumann 1990; Thornton and Camburn 1987), and household size (Dornbusch, Carlsmith, and Bushwall 1985) as key structural factors relevant to adolescent sexual activity. Family socioeconomic factors are thought to regulate the quantity, quality, and diversity of resources to which children and youth theoretically have access (Brooks-Gunn, Klebanov, Liaw, and Duncan 1995; Haveman and Wolfe 1994). Limited family economic resources are likely to place constraints on the capacity of parents to provide opportunities for children and otherwise promote positive outcomes, such as the delay of sexual activity during adolescence (Duncan and Brooks-Gunn 1997; Lauritsen 1994; McLanahan and Sandefur 1994). Family low-income is also associated with less optimal parenting—including low warmth/supportiveness, harsh discipline, and parent-child conflict—and, in turn, child and adolescent outcomes (Conger, Rueter, and Conger 2000; Conger, Ge, Elder, Lorenz, and Simons 1994; McLoyd and Wilson 1990). Single and stepparent households and larger households (independent of income) also may face constraints on the capacity to provide social support, nurturing, and supervision to children (Dornbusch, Carlsmith, and Bushwall 1985; Wu and Thomson 2001).

Second, variation in family level factors may also contribute to risk-enhancing (or reducing) peer characteristics (Thornberry 1987) with which early sexual onset is likely to be associated. Evidence suggests that peers who are sexually active or are perceived to be sexually active (Kinsman, Romer, F. F. Furstenberg, and Schwarz 1998), as well as peers who engage or are perceived to engage in nonsexual problem behavior, such as alcohol and drug use (Blum, Beurhing, and Rinehart 2000; Costa, Jessor, Donovan, and Fortenberry 1995), encourage sexual onset by positively reinforcing such activity and providing opportunities for its occurrence. Rodgers and Rowe (1993), for instance, have used a social contagion framework to explain the spread of adolescent sexual behavior among peer groups. Thus, deviant peer group affiliation is anticipated to be associated with an increased likelihood of early sexual onset. In contrast, supportive and emotionally rewarding relationships with peers may reduce the likelihood of early sexual onset (Bearman and Bruckner 1999).

Third, at the individual level, a host of biological, behavioral, temperament, and academic risk factors for adolescent sexual activity has been identified. A key biological risk factor is early pubertal maturation (Zabin, Hirsch, Smith, and Hardy 1984). This link may result from hormonal effects on behavior or through contextual responses to premature physical development by parents and peers (Brooks-Gunn and F. F. Furstenberg 1989; Graber, Brooks-Gunn, and Galen 1998; Halpern, Udry, and Suchindran 1997; Stattin and Magnusson 1990).

Continuity in deviant behavior across childhood and adolescence is well established (Robins 1966), suggesting that prior problem behavior trajectories may also be associated

with early and heightened levels of sexual activity (Elliott and Morse 1987; Jessor 1983; Weiher, Huizinga, Lizotte, and Kammen 1991). A number of studies have demonstrated the intercorrelation between sexual risk behavior and other forms of problem behavior (Osgood 1988; Paternoster and Brame 1998). Independent of behavioral orientations, temperamental characteristics such as social inhibition and sociability may hinder or promote social, and potentially sexual, contact among youth (Capaldi, Crosby, and Stoolmiller 1996). Finally, among adolescents, evidence suggests that academic ability and attachment to school are negatively associated with sexual onset (Halpern, Joyner, Udry, and Suchindran 2000; Hayes 1987; Hofferth 1987; Resnick et al. 1997). Below, we consider the impact of familial and collective supervision in the context of extensive multivariate models including family structural, peer, and individual level factors as well as neighborhood structural context.

DATA AND METHODS

We use multiple independent data sources gathered by the Project on Human Development in Chicago Neighborhoods (PHDCN) to examine individual, family, and neighborhood correlates of adolescent sexual risk behavior. All individual and family measures are drawn from the PHDCN Longitudinal Cohort Study. Neighborhood measures are constructed from 1990 Census data, vital statistics data on Chicago communities, and the PHDCN Community Survey.

DESIGN

For the Longitudinal Cohort Study, Chicago's 847 census tracts were combined into 343 neighborhood clusters (NCs) that maintained relative population homogeneity with respect to racial/ethnic, socioeconomic, housing, and family structure characteristics (NCs average roughly 8,000 people).[5] NCs were also defined on the basis of ecologically meaningful boundaries such as railroad tracks and freeways. Next, a two-stage sampling procedure was employed that included selecting a random sample of 80 of the 343 Chicago Neighborhood Clusters (NCs) stratified by racial/ethnic composition (7 categories) and SES (high, medium, and low). The aim was to have an equal number of NCs in each of the 21 strata that varied by racial/ethnic composition and SES. This objective was well approximated with only 3 exceptions; low-income primarily European American, high-income primarily Latino, and high-income Latino/African American neighborhoods did not exist.Within these 80 NCs, children falling within 7 age cohorts (birth and ages: 3, 6, 9, 12, 15, & 18) were sampled from randomly selected households. Extensive in-home interviews and assessments were conducted with these children and their primary caregivers at two points in time over a 4-year period, at roughly 2-year intervals (Wave 1 in 1995-1996 and Wave 2 in 1998-1999).

The Community Survey is a probability sample of 8,782 residents of Chicago focusing on respondent assessments of the communities in which they live. The Community Survey was conducted in 1994-1995 in conjunction with the first wave of the

[5] Respondents were given the following definition of "neighborhood:" "By neighborhood...we mean the area around where you live and around your house. It *may* include places you shop, religious or public institutions, or a local business district. It is the general area around your house where you might perform routine tasks, such as shopping, going to the park, or visiting with neighbors."

Longitudinal Cohort Study; however, these samples are independent. The Community Survey used a three-stage sampling strategy: First, city blocks were randomly selected within each of the identified 343 NCs; second, dwelling units within blocks were randomly selected; and third, respondents (one adult, age 18 or over, per household) within dwelling units were randomly selected to complete surveys. The Community Survey sampling strategy ensured that the number of cases collected per NC would be sufficient to estimate neighborhood characteristics based on aggregated individual level data.[6] The final response rate was 75%.

SAMPLE

This study uses Wave 1 data on subject and primary caregiver characteristics to predict the timing of first intercourse among respondents between the ages of 11 and 16. However, information on age at first intercourse is drawn from Wave 2. By focusing on earlier adolescent sexual onset, we consider a population for whom the capacity to manage sexual activity is less developed and the costs of sexual behavior may be significant (Koyle, Jensen, Olsen, and Cundick 1989; Seidman, Mosher, and Aral 1994). The sample is roughly evenly split by sex of the adolescent and a range of socioeconomic groups is represented, by design. For the total sample of youth, 52% are girls and 48% are boys. The racial/ethnic composition of the sample reflects the diversity of Chicago's population in the 1990s: 48% are Latino, 33% are African American, 16% are white, and

[6] The overall within-neighborhood *n* averaged about 25. The 80 target NCs for the longitudinal study included oversamples resulting in an average of about 50 interviews per NC.

3% are from other racial groups (we combine the white and other racial categories in the analyses below). Sample retention across waves was 82%.

To maximize the internal validity of the analysis, we include only respondents who report an age at first intercourse that is equal to or greater than their Wave 1 age. By dropping respondents who report sexual onset prior to their Wave 1 age, the analysis avoids predicting an outcome with characteristics clearly measured after the event has occurred.[7] Of the 526 boys with data available on the dependent variable at Wave 2, 62 had sex prior to their Wave 1 age and an additional 29 cases were missing data on one or more of the independent variables (analysis $N = 435$). Of the 541 girls with Wave 2 information on the timing of first intercourse, 30 had sex prior to their Wave 1 age and an additional 31 were missing data on one or more independent variables (analysis $N = 480$).[8]

DEPENDENT VARIABLE

The dependent measure used in the analysis is the respondent's age at first sexual intercourse in years. During the second wave of the Longitudinal Cohort Study,

[7] For respondents who report sexual onset during their age at Wave 1, we cannot determine temporal ordering between Wave 1 explanatory variables and first intercourse. However, less than 8% of the sample (and 22% of those who reported an intercourse experience) report that their age at first intercourse was the same as their wave 1 age. Assuming a random distribution of ages among the sample and intercourse experiences that are likely skewed toward the latter half of any given age year (i.e., those who report an age at first intercourse of 14 are more likely to have had sex during the latter half of their 14th year), less than half of those who report equivalent Wave 1 and first intercourse ages will have been interviewed after their first sexual experience.

[8] A drawback to dropping respondents who reported sexual onset prior to their Wave 1 age is the potential for sample selection bias. To address this possibility, we ran our analyses including respondents with ages of first intercourse prior to Wave 1. These expanded analyses yielded essentially the same results as our restricted sample, enhancing our confidence in the findings reported below.

adolescents were asked, "How old were you when you first had sexual intercourse?"[9] An effort was made to protect confidentiality during gathering of sensitive information such as ensuring privacy and making youth aware of the fact that a Federal Certificate of Confidentiality was obtained for the study, which made it illegal to disclose adolescents' responses to anyone, such as parents or school officials (the exception being if youth posed a potential threat to themselves or others). Among respondents in their 16th year, 58% of boys and 49% of girls reported having had sexual intercourse.

INDEPENDENT VARIABLES

We include a variety of independent variables designed to measure key family, peer, and child characteristics as well as features of neighborhood environment.

Child and family demographic characteristics. Key individual level demographic characteristics include gender, age, and race/ethnicity (two dummy codes indicating African-American and Latino, with White/other serving as omitted referent). Structural background measures at the family level include socioeconomic status as measured by the first principal component of annual household income, education (highest education level achieved by primary caregiver in the household[10]), and occupation for the adolescent's primary caregiver.[11] Family structure is captured by a dummy variable marking the presence of a biological mother and father in the household versus all other

[9] Questions related to first sexual intercourse were prefaced with the following statement: "Sometimes people refer to sexual intercourse as 'doing sex', 'having sex', 'making love', or 'going all the way'."
[10] If the subject had two primary caregivers, the caregiver who had the highest education level was used.
[11] Occupational prestige was based on a coding strategy developed by Nakao and Treas (1994) using the updated 1990 Census Occupational Classification System.

family arrangements. A measure of household size enumerates the number of individuals in the household.

Family processes. We employ two measures of parental supervision capturing both overall extent of supervision and monitoring as well as parental practices with specific implications for youths' level of exposure to neighborhood environments. First, a continuous global measure of parental supervision and monitoring is a 16-item scale based on an expanded version of the adolescent supervision subscale of the Home Observation for Measurement of the Environment (HOME; see the Appendix for item descriptions) (Bradley, et al. 2000). Items were based on primary caregiver reports and include dichotomous (yes/no) responses to questions asking about supervision and monitoring practices. Items cover caregiver regulation of leisure time and daily schedules, curfews, behavior with peers, television watching, knowledge and communication regarding the hazards of drug and alcohol use, and consistency of rule application across children. The final measure is the empirical bayes residual from a two-level Rasch model (reliability = .63). We employ item response theory in the context of multilevel models to construct individual level scales when the number of items used to create the scale is large (increasingly the likelihood of missing data on component items). An advantage of the multilevel item response model is the capacity to adjust for missing data, obviating the need to drop entire cases if data are missing on any one item in the scale. Item response models also take into account item "severity" and, if necessary, frequency in generating person-level "ability" scores (Wright and Masters 1982). Scale scores from multilevel item response models are person level empirical

bayes (EB) residuals. EB residuals regress scale scores toward the grand mean by a factor proportional to the unreliability with which they have been estimated (Raudenbush and Bryk 2002).[12]

An additional measure of parental supervision was constructed in order to capture parental control practices related to youth exposure to neighborhood environments ("place monitoring"). We performed a latent class analysis of three dichotomous items capturing whether (1) After school subject goes somewhere that adult supervision is provided, (2) Subject is not allowed to wander in public places without adult supervision for more than two hours, and (3) Primary caregiver has had contact with two of the subject's friends in the last two weeks.[13] Latent class analysis offers a useful method of data reduction when considering a smaller number of categorical variables. A two class solution produced the best fitting model, with a high supervision category (all PCs responded affirmatively to the three items) and a low supervision category (25% reported that the subject had no supervised place to go after school, 57% reported that subject was allowed to wander unsupervised in public places for more than two hours, and 46%

[12] Although the component items of the HOME supervision subscale were not designed to be analyzed using test construction theory, we view the benefits of IRT methodology and the reasonable reliability of the resulting scale as warranting application of this scaling approach to the HOME inventory items. Multivariate analyses of both global supervision and place monitoring measures (and the consistency of findings) also allay concerns over the specific choice of scaling methodology used for the global measure.

[13] Contact with friends may be viewed as an additional indicator of parental control over youth exposure to external environments. Unsupervised time after school and in public places outside of the home may be considered less hazardous if youth are accompanied by known and trusted peers. We considered more than one operationalization of place monitoring with comparable results in multivariate models (e.g., a latent measure of place monitoring constructed with an additional item—"PC establishes rules for subject's behavior with peers and asks questions to determine whether they are being followed"—yielded nearly identical findings as those reported below.

reported no contact with subject's friends in the last two weeks; all PCs in the low supervision category responded negatively to at least one of the items).[14]

Second, *family attachment and support* is based on a five item scale tapping the extent to which the adolescent feels the following statements are true: (1) "No matter what happens, I know that my family will always be there for me should I need them," (2) "My family lets me know they think I'm a worthwhile (valuable) person," (3) "People in my family have confidence in me," 4) "people in my family help me find solutions to my problems," and (4) "I know my family will always stand by me." Responses were given on a three-point scale ("not true," "somewhat true," or "very true"). The scale constitutes the sum of the item responses divided by the number of items (reliability = .75) (Turner, Frankel, and Levin 1987).

Peer influences. We include two measures of peer influence. The first scale includes a three-item measure of *positive peer attachment* (adolescent's level of agreement with the following statements: (1) "I have at least one friend I could tell anything to," (2) "I feel very close to some of my friends," and (3) "My friends would take the time to talk about my problems, should I ever want to" Responses were given on a three-point scale ("not true," "somewhat true," or "very true"). The scale constitutes the sum of the item responses divided by the number of items (reliability = .63) (Turner, Frankel, and Levin 1987). The second measure assesses *peer deviance* and was constructed from seventeen items asking adolescents to report on the behavior of people they "spend time with."

[14] Final likelihood ratio chi-square was 4.76 (versus the saturated model) on 3 degrees of freedom (the conditional probabilities of experiencing each supervision item were constrained to be 1 within the high supervision category).

Questions asked about how many of these people engage in activities such as alcohol and drug use, property and violent crime, and "sexual intercourse."[15] Responses were given on a three-point scale ("none of them," "some of them," or "all of them"). The measure used in the analyses is the empirical bayes residual from a multilevel ordinal logit (rating scale) analysis (reliability = .86).

Individual risk factors. We focus developmental, behavioral, temperament, and academic risk factors for early sexual onset. Measures of *pubertal development* were constructed separately for boys and girls based on their reports of physical status; higher scores indicate greater physical maturity. Boys were asked whether they had experienced a height spurt, growth of body hair, any skin changes (such as pimples), voice deepening, and growth of facial hair (Responses were given on a four-point scale: "no," "yes (barely)," "yes (definitely)," or "development completed"). The scale constitutes the sum of the item responses divided by the number of items (reliability = .69). Girls were asked whether they had experienced a height spurt, growth of body hair, any skin changes (such as pimples), breast growth, and menstruation (reliability = .72) (Petersen, Crockett, Richards, and Boxer 1988).

Prior problem behavior was assessed by adolescents' reported participation (yes/no) in nineteen activities involving violent behavior, property crime, and use of illegal drugs; items were combined using a multilevel Rasch model (Cheong and Raudenbush 2000; Raudenbush, Johnson, and Sampson forthcoming). The final scale used in analyses is the empirical bayes residual from the level-two model (reliability = .75).

[15] Including only the latter item in multivariate models resulted in essentially the same results.

In addition, a subscale of the Emotionality and Sociability Inventory (EASI, Buss and Plomin 1984) was employed to assess sociability. This widely used inventory was administered to primary caregivers, who reported how characteristic each behavior was of their children (from "1" uncharacteristic to "5" characteristic). Sample items include "makes friends easily" and "likes to be with people." The scale constitutes the sum of the item responses divided by the number of items (reliability = .65). Academic achievement/competence was assessed by adolescents' performance on the reading component of the *Wide Range Achievement Test* (WRAT-3; (Wilkinson 1993)), which is a standardized test (mean = 100; standard deviation = 15) that evaluates reading achievement.

Neighborhood characteristics. All neighborhood characteristics were derived for each of the 80 NCs (or neighborhoods) used in the analyses. Measures of neighborhood level structural characteristics were constructed using data from the 1990 Decennial census. Based on theory and extensive prior investigation (Land, McCall, and Cohen 1990; Sampson, Raudenbush, and Earls 1997), principal components analyses were performed on selected census-based measures. First, *concentrated disadvantage* is the first principal component of the percent below the poverty line, percent receiving public assistance, percent unemployed, and the percent of households headed by a female. For analytic purposes, the natural log of this principal components scale score was used.[16]

[16] Ideally, the racial composition of the neighborhood clusters should also be considered, independent of concentrated disadvantage. Unfortunately, the relatively high correlation between these conceptually distinct dimensions renders investigation of their unique effects statistically problematic. To assess whether the key neighborhood effects we observe for are actually capturing the racial composition of the neighborhood, we included a measure of percent African-American in Model 1 of Table 3 and Model 6 of Table 4. The coefficients for concentrated disadvantage and collective efficacy changed negligibly with the

Second, a *residential stability* component combines measures of continuity of residence (percent living in the same house as 1985) and the percent of housing occupied by owners. Third, an *immigrant concentration* component combines the percent Latino and percent foreign born.

Our measure of *collective efficacy* combines information from two scales administered as part of the Community Survey: First, a social cohesion scale was constructed from a cluster of conceptually related items measuring the respondent's level of agreement (on a five-point scale) with the following statements: (1) "People around here are willing to help their neighbors," (2) "This is a close-knit neighborhood," (3) "People in this neighborhood can be trusted," (4) "People in this neighborhood generally don't get along with each other," and (5) "People in this neighborhood do not share the same values." The latter two items were reverse coded.

Second, an intergenerational closure and informal social control scale was constructed from items tapping the respondent's level of agreement that (1) "Parent's in this neighborhood know their children's friends," (2) "Adults in this neighborhood know who the local children are," (3) "There are adults in this neighborhood that children can look up to," (4) "Parents in this neighborhood generally know each other," and (5) "You can count on adults in this neighborhood to watch out that children are safe and don't get in trouble." By design, this scale captures aspects of both adult-child social ties as well as expectations for active support and informal social control of local youth by neighborhood adults (Sampson, Morenoff, and Earls 1999).

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inclusion of the percent African American in the neighborhood, which did not achieve significance at the conventional level in either model.

The two scales were highly correlated ($r > .65$) and were combined into a single measure of youth-related intergenerationally-oriented collective efficacy (hereafter "collective efficacy"). The measure of collective efficacy used in the analysis was constructed using a three-level linear item response model (Raudenbush and Bryk 2002; Sampson, Raudenbush, and Earls 1997). At level one, an item-response model adjusted individual level latent collective efficacy scores for missing data, taking into account the difficulty level of items for which a response was provided. At level two, neighborhood collective efficacy scores (intercepts in between-individual models) were adjusted for the social composition of Chicago neighborhoods through inclusion of controls for gender, age, race/ethnicity (Black, Hispanic vs. White), education, employment status (employed vs. not employed), marital status (never married, separated or divorced vs. married), home ownership, years resident in the neighborhood, and number of moves in the last five years. At level three, adjusted neighborhood intercepts varied randomly around the neighborhood grand mean. The standardized neighborhood level empirical bayes residual constitutes the collective efficacy score employed as an independent variable in subsequent analyses of the timing of first intercourse. The three-level reliability of the combined scale was .81.[17]

We also include a control for the prevalence of *adolescent childbirth* in order to capture the salience of sexual activity at the neighborhood level. Collective supervision capacity may be confounded with, or contribute to, the concentration of early sexual activity among youth. In turn, aggregate levels of sexual activity among neighborhood

[17] See Raudenbush and Sampson (1999) for a discussion of reliability in three level models.

youth may facilitate the proto-cultural transmission of sexual behavior (Wilson 1996). We consider the effect of collective efficacy both with and without a control for the prevalence of adolescent childbearing to assess the robustness of any observed effects. The measure was constructed using vital statistics data from the City of Chicago on births to adolescent women, ages 15 to 19 for 1992-94. The measure used in the analysis is the standardized empirical bayes log birth rate per 1,000 population.[18]

Table 1 reports descriptive statistics on variables included in the analysis, disaggregated by gender. Of note are comparisons of parental supervision levels for the two measures employed. Although there is modest evidence of overall supervision differences in the expected direction (with girls receiving higher levels [$p < .10$]), the proportions of boys and girls receiving place monitoring are comparable. Nevertheless, the etiological significance of both measures of supervision may vary across gender.

ANALYTIC STRATEGY

Our analysis of sexual onset employs recently developed multilevel discrete-time event history techniques (Barber, Murphy, Axinn, and Maples 2000; Reardon, Brennan, and Buka 2002). The two level discrete time model, estimating effects separately for boys and girls, takes the following form:

[18] See Raudenbush and Bryk (2002) for a discussion of a comparable two level poisson model as applied to official data on homicide. Data on the number of adolescent women ages 15 to 19 with each neighborhood were taken from the 1990 Census as a proxy for the 1992-94 denominator.

$$h_{ijt} = \ln\left(\frac{h_{ijt}}{1-h_{ijt}}\right) = D_{Mij}\left(p_{Mj} + \sum_{p=1}^{P} b_{Mp} X_{Mpij}\right) + D_{Fij}\left(p_{Fj} + \sum_{p=1}^{P} b_{Fp} X_{Fpij}\right)$$

$$\pi_{Mj} = \gamma_{M0} + \sum_{q=1}^{Q} \gamma_{Mq} Z_{qj} + \delta_{Mj}$$

$$\pi_{Fj} = \gamma_{F0} + \sum_{q=1}^{Q} \gamma_{Fq} Z_{qj} + \delta_{Fj}$$

where, at level 1, h_{ijt} is the hazard of sexual onset for subject i in neighborhood j at age t, D_{Mij} is an indicator variable (male = 1), p_{Mij} is a male-specific intercept, and X_{Mpij} is covariate p for (male) subject i in neighborhood j (with associated male-specific coefficients b_{Mp}). A second set of terms estimates a female-specific intercept p_{Fij} and a set of coefficients b_{Fp} for the equivalent set of covariates at the subject level. At level two, adjusted male- and female-specific intercepts are regressed on a series of neighborhood level covariates where γ_{M0} and γ_{F0} are level two intercepts, Z_{qj} is covariate q in neighborhood j, γ_{Mq} and γ_{Fq} are associated coefficients describing the effect of neighborhood level covariates for boys and girls, respectively, and the d's are gender-specific level two random effects. Estimating gender-specific models simultaneously allows for statistical tests of the difference in coefficient magnitudes by gender.

Our first set of multivariate analyses (Table 2) focuses on individual, family, and peer influences on adolescent sexual behavior, estimating the effects of covariates separately for boys and girls. In a baseline model, we examine the effects of age, race/ethnicity, and

immigrant generation on the timing of first intercourse (Model 1). We then add family structural background (SES, composition, and size) and our theoretically central measures of family supervision (along with parent-child attachment; Model 2). Finally, we add peer influences and individual risk factors (peer deviance, positive peer attachment, pubertal development, prior problem behavior, sociability, and reading ability; Model 3) to assess the extent to which parental supervision effects remain after child level controls are included.

The next set of models (Table 3) considers neighborhood measures. First, we include our key neighborhood structural characteristics—concentrated disadvantage, residential stability, and immigrant concentration—to test whether these characteristics aid in the explanation of timing of sexual onset over and above family, peer, and individual characteristics. Concentrated disadvantaged is examined first (Model 1), followed by inclusion of residential stability and immigrant concentration (Model 2). We then enter our measures of collective efficacy (Model 3) and the prevalence of adolescent childbearing (Model 4), to consider their main effects on the timing of first intercourse.

Finally, in Table 4, we consider the cross-level interaction between neighborhood and family level supervision in order to assess the extent to which collective efficacy varies in its effects by level of parental control. These models employ our measure of parental place monitoring. We first consider the main effects of this alternative measure of supervision (Model 1) and then consider the cross-level interaction between place monitoring and collective efficacy estimated in separate models for boys and girls

(Models 2 and 3). Models 4 and 5 add adolescent childbearing to the model for place monitoring slopes. Finally, Model 6 assesses neighborhood effects in reduced individual level models (controlling only demographic background) to gauge the impact of intervening family process, peer, and individual characteristics as potential mediators of neighborhood characteristics.

RESULTS

We begin the discussion of multivariate results by considering multilevel discrete time logit models[20] conditional only on the linear and quadratic effects of age (not shown). As expected, age is highly significant and powerfully associated with sexual onset for both boys and girls. The quadratic effects of age are also significant for both genders, indicating that the proportionate increase in the log odds of first intercourse declines in magnitude with age.[21] Initial models conditional on age also allow for a decomposition of variance in the dependent variable across levels of analysis. For boys, the estimate of the intercept variance is .68 ($p < .05$). For girls, the estimate of the intercept variance is .43 but does not achieve significance at the conventional level. Variance components estimates suggest that the timing of boys' sexual onset is characterized by greater variation across neighborhood than is girls. Nevertheless, two points should be noted:

[20] We estimated both random effects and population-average models with robust standard errors. Both models yielded essentially the same pattern of results. We present results from the population-average model with robust standard errors below.

[21] Dummy variables for age categories did not significantly improve model goodness-of-fit over the linear and quadratic specification. We also evaluated the extent to which the proportionality assumptions of the discrete-time logit model held at both levels of the analysis. That is, interactions between individual level covariates and age were tested in order to determine whether the effect of age could be considered constant across key groups. We found no significant interactions with age at level 1. We allowed the linear and quadratic effects of age to vary randomly across neighborhoods but found no evidence of significant variability in the effect of this covariate across contexts.

First, relatively small or insignificant variance components do not rule out significant neighborhood effects in multivariate models. Second, the magnitude and significance of variance components should be interpreted with caution if a cross-level interaction is hypothesized (Raudenbush and Bryk 2002).

We turn next to multivariate models of the timing of first intercourse including additional demographic, family, peer, and individual level covariates (Table 2). Model 1 includes demographic background characteristics including age, age squared, race/ethnicity, and immigrant generation. African American boys and girls report substantially earlier ages of first intercourse by comparison with their same gender white peers, as do Latino girls. First generation immigrants of both genders also report earlier ages of first intercourse. Second generation girls also report later ages of sexual onset by comparison with 3rd or higher generation girls.

Model 2 considers family structure and process variables, including the effects of parental supervision and monitoring. Family socioeconomic status does not achieve significance for either boys or girls, though the presence of two biological parents is negatively associated with early sexual onset for both genders. Increasing family size is associated with earlier sexual transitions for boys but not girls. Consistent with the gendered familial control hypothesis, the effect of parental supervision and monitoring is insignificant for boys and highly significant and negative for girls. The odds of sexual onset, for girls, decrease by 24% with a one standard deviation increase in parental supervision. Likelihood ratio chi-square tests of the hypothesis of coefficient equality reveal that the effect of parental supervision for girls is significantly larger than boys ($p <$

.05).[22] Family attachment and support is negatively associated with early sexual onset for both boys and girls and does not significantly differ in magnitude by gender.

Model 3 adds a host of peer and individual level covariates. For boys, but not girls, prior problem behavior is powerfully associated with the timing of first intercourse. In contrast, pubertal development and sociability predict early sexual onset for girls but are not significant predictors for boys. Positive peer attachment, peer deviance, and verbal aptitude are not significantly associated with sexual onset for either gender. Family structure effects, supervision (for girls), and family attachment/support remain significant in Model 3, suggesting that these factors do not operate principally through their impact on individual vulnerabilities and peer characteristics. Of note in Model 3 is the emergence of a significant positive effect of family socioeconomic status on sexual onset for boys. Apparent differences in the etiology of sexual onset by gender based on comparisons of *p*-values, however, are not supported in separate tests of coefficient equality. Other than parental supervision, only the effect of Latino ethnicity varied significantly ($p < .05$) by gender.

Table 3 reports the results of models including additional neighborhood level covariates to fully specified micro-level models (Model 3 of Table 2). We report coefficients for neighborhood effects only. Model 1 includes the logged concentrated poverty scale. Poverty is positively associated with sexual onset, powerfully so for boys,

[22] The models were estimated using penalized quasi-likelihood in HLM5 (Raudenbush and Bryk 2002). Full maximum likelihood estimates could not be computed in HLM with the model as specified, precluding tests of coefficient equality across gender. In order to perform hypothesis tests on the fixed effects, we reran key models using Generalized Linear Latent and Mixed Models (GLLAMM), producing alternative maximum likelihood estimates (Rabe-Hesketh 2001). Coefficient estimates produced by GLLAMM were negligibly different from the HLM results. Likelihood ratio chi-square tests of coefficient equality across gender are based on the GLLAMM results.

above and beyond micro level factors. Model 2 adds residential stability and immigrant concentration, which do not achieve significance. Model 3 adds our measure of collective efficacy and reveals a significant negative effect for boys but not for girls. The introduction of neighborhood adolescent childbearing in Model 4 does not achieve significance, but reduces the magnitude of the effects of both poverty and collective efficacy for boys (Models including adolescent childbearing without collective efficacy reveal a significant positive effect on the timing of first intercourse for boys).

These findings are potentially consistent with a gendered theory of the effects of collective efficacy on adolescent sexual behavior. Though the effects of collective supervision on sexual onset are not statistically different for boys and girls (in contrast to the effects of familial supervision), they nevertheless point to the potential for distinct experiences of neighborhood characteristics by gender. In order to test whether the differential effect of collective efficacy is due to the unique etiological significance of parental supervision for girls (i.e., that they experience less exposure to neighborhood environments) and/or gendered application of collective control, we examined the cross-level interaction between collective efficacy and a measure of parental place monitoring that more directly taps parenting practices that relate to neighborhood exposure. A significant effect (comparable to boys) of collective efficacy on the sexual behavior of girls who are subject to lower levels of monitoring would be consistent with a gender-neutral collective control hypothesis, whereas a consistently larger effect of collective efficacy for boys by comparison with unmonitored girls would be evidence in support of the gendered collective control hypothesis.

The results of these analyses are reported in Table 4. First, Model 1 reproduces Model 3 of Table 2 in order to evaluate the unique effect of our alternative supervision measure. Consistent with the results reported in Table 2, parental place monitoring is associated with delayed sexual onset for girls but is not significantly associated with sexual onset for boys. The difference in the magnitudes of the effects by gender is statistically significant ($p < .05$). The positive effect of adolescent childbearing, for boys, becomes statistically significant in Model 2. Models 2 and 3 of Table 4 add the cross-level interaction between neighborhood collective efficacy and place monitoring, running this interaction model separately for boys and girls.[23] Although the interaction is not significant in Model 2 (boys), Model 3 (girls) reveals a significant positive coefficient for the interaction between collective efficacy and place monitoring. Fitting the cross-level interaction for girls also results in a significant negative effect of collective efficacy for unmonitored girls. Thus, girls who experience lower levels of place monitoring (as measured by our dichotomous indicator designed to tap supervision relevant to neighborhood exposure) also experience the effects of neighborhood collective efficacy on their sexual transition behavior. In contrast, the effects of collective efficacy are completely wiped out for girls who experience high place monitoring. This finding holds when we include the measure of adolescent childbearing in the model for place monitoring slopes. Moreover, the effects of collective efficacy hold for both boys and girls when we fit a reduced individual level model, including only demographic background characteristics (Model 1 of Table 2) and parental supervision, suggesting that

[23] We ran the cross-level interaction models both separately by gender and simultaneously, producing essentially the same results.

additional family, peer, and individual level factors do not mediate the effects of collective efficacy on sexual transition behavior.

Figures 1 and 2 plot the cumulative predicted probability of sexual onset by age.[24] For boys, we plot the predicted probabilities for those in high (+1 standard deviation) and low (-1 standard deviation) collective efficacy neighborhoods (holding other model covariates at their sample means). Clear differences emerge in the cumulative probability of sexual onset over the course of early adolescence. For instance, by age 16, 48% of boys in high collective efficacy neighborhoods are predicted to have experienced first intercourse by comparison with 68% of boys in low collective efficacy neighborhoods. For girls, we plot the cross-level interaction between collective efficacy and parental place monitoring. For girls who experience higher levels of place monitoring, no statistically significant effect of collective efficacy emerges. Among high place monitored 16 year-old girls in low collective efficacy neighborhoods, 31% are predicted to have experienced sexual onset by comparison with 37% in high collective efficacy neighborhoods. Among girls who experience lower levels of place monitoring, in contrast, the effect of collective efficacy is significant and nontrivial. Less monitored girls who live in high collective efficacy neighborhoods benefit from their environments such that their probability of having experienced first intercourse by age 16 (34%) is comparable to their monitored peers. The combination of low parental place monitoring and low collective efficacy, however, dramatically increases the probability of sexual

[24] Based on Model 1 (boys) and Model 5 (girls) of Table 4.

transition: 70% of girls in this group are predicted to have experienced first intercourse by age 16—comparable to boys in low collective efficacy neighborhoods.[25]

DISCUSSION

Increasingly, research on the etiology of early sexual behavior is extending a longstanding interest in family, peer, and individual factors to include the neighborhood context of adolescents' lives (Brooks-Gunn, et al. 1993). We build on this research to develop and test a theory of the link between neighborhood context and adolescent sexual behavior, emphasizing the role of structural disadvantage and collective efficacy. We extend a model focused on the influence of neighborhood level disadvantage and collective supervision capacity to include the potential for gender differences in the impact of neighborhood environments, emphasizing the role of parenting practices in differentially exposing boys and girls to neighborhood conditions.

Most research on the emergence of adolescent sexual risk behavior has focused attention primarily on family and individual level factors. Indeed, the findings of micro level analyses reported here were consistent with a host of prior studies stressing, in particular, the significance of key demographic background and family-level characteristics. African American adolescents initiate sexual activity earlier than white youth as do Latino girls (though not boys). First generation immigrant youth also delay intercourse when compared with youth whose families established US residence at an

[25] Note that the simulation assumes that youth maintain residence in high or low collective efficacy neighborhoods across the age range considered.

earlier generation. Family structure effects were also robust, particularly the impact of living with both biological parents, reinforcing the findings of prior studies.

Of theoretical interest was the effect of parental supervision and monitoring and the potential for differential significance of this factor by gender. Consistent with a gendered familial control hypothesis, the effect of a global measure of supervision was quite powerful and negative for girls but insignificant for boys. The effect of parental supervision and monitoring was statistically different by gender in separate hypothesis tests, conforming to the expectations of theoretical approaches stressing heightened parental concern with the early adolescent behavior of girls by comparison with boys. Although we found some evidence in support of the hypothesis that girls experience a significantly greater quantity of supervision and monitoring when considering a global measure of parental control, a dichotomous indicator designed to capture supervision directly relevant to neighborhood exposure (place monitoring) revealed comparable levels by gender. Thus, early adolescent girls and boys may be subject to broadly similar parental oversight and behavioral rules with respect to neighborhood exposure but with differential *consequences*.

A number of explanations for this phenomenon are possible. First, socialization patterns that cultivate more assertive and riskier behavioral orientations among boys and greater dependence on and obedience to parents among girls may result in greater impact of supervisory rules on girls (Hagan, et al. 1987). Second, rules may be less stringently enforced for boys than girls, due to the influence of traditional orientations toward the appropriate conduct of early adolescent girls. Third, parents may apply supervision rules

with reference specifically to behaviors that may place girls at risk of early sexual activity. Thus, parents may scrutinize girls' friendships, recreational activities, and other routine activity patterns with the specific goal of limiting exposure to *sexual* opportunities. Parents may be motivated by concern over the perceived severe negative consequences of sexual risk behavior for girls by comparison with boys.

Additional factors at the individual level that emerged as significant determinants of early adolescent sexual activity included family attachment and support; pubertal development and sociability (for girls); and prior problem behavior (for boys). Although the pattern of statistical significance for individual level factors was inconsistent by gender, statistical tests of the difference in magnitude of these effects revealed remarkably little evidence of gender-specific etiologies of early sexual onset beyond the effects of parental supervision (see also Moffitt, Caspi, Rutter, and Silva 2001).

Our principal concern was the additional and interactive impact of neighborhood level factors—particularly collective supervision capacity—on the emergence of sexual behavior during adolescence. Neighborhood factors were important determinants of sexual onset for both boys and girls. The introduction of the neighborhood concentrated poverty measure to models including family, peer, and individual effects resulted in significant coefficients for both genders. This finding indicates that socioeconomic features of neighborhood context play a consequential role in the unfolding sexual trajectories of urban adolescents. No evidence of unique effects of immigrant concentration or residential stability, however, emerged.

A key contribution of our analysis was the investigation of community level supervision capacity as captured by a measure of collective efficacy (Brooks-Gunn, Duncan, Klebanov, and Sealand 1993; Sampson 1997; Sampson, Raudenbush, and Earls 1997; Wilson 1996). We argued that the collective capacity to manage adolescent sexual behavior emanates from the social integration of local adults and children, a larger sense of cohesion among community residents and shared expectations for pro-social, intergenerationally-oriented action of adults on behalf of local youth. Norms regarding the responsibility of adults for children may help foster inter-family communication and joint supervision necessary to regulate the behavior of adolescents.

Our theoretical model, however, calls attention to the potentially gendered nature of neighborhood influence. Drawing on the logic of gender-based theoretical orientations in delinquency research and recent findings regarding the relative impact of neighborhoods on boys and girls, we hypothesized that boys experience greater independence from parental controls, particularly during the early stages of adolescence. In turn, this relative freedom results in more exposure to neighborhood environments and enhanced opportunity for extrafamilial influence on behavioral trajectories. We also proposed that the application of collective supervision may be gendered due to a heightened collective concern with the consequences of boys behavior.

Separate models of the timing of first intercourse for boys and girls offered evidence of distinct neighborhood-level etiologies. Gender differences in the effects of neighborhood collective efficacy were evident, providing support for the hypothesis that boys experience greater influence from neighborhood environments. For boys, collective

efficacy emerged as a significant negative predictor of sexual onset—an effect not observed for girls. The significance of collective efficacy in delaying sexual activity among boys held in the context of models including neighborhood structural characteristics as well as a measure of the prevalence of adolescent childbearing.

The effect of collective supervision capacity on the behavior of boys, by comparison with girls, is consistent with the hypothesis that the relative freedom from the effects of parental controls experienced by boys results in greater exposure to the influence of neighborhood environments. A gendered collective control hypothesis, however, implies that girls who are subject to lower levels of place monitoring (and greater exposure to neighborhood environments) would still experience weaker effects of collective supervision capacity by comparison with boys. Tests of the cross-level interaction between collective efficacy and a measure of parental place monitoring revealed a significant coefficient for girls. Collective efficacy delays sexual onset for girls who experience lower levels of parental place monitoring but has no effect for girls who experience higher levels of familial control. The consequences of low levels of familial and collective control combine to place girls at particularly high risk of early sexual activity—comparable to boys in low collective efficacy neighborhoods. These findings suggest that the effects of collective control on the timing of first intercourse are gender-neutral for youth who are exposed to neighborhood environments.

Our analyses point to the importance of considering the *community-level social processes* that may be relevant in organizing individual adolescent behavior. Multilevel approaches to youth sexual activity that consider family, peer, and individual level factors

supplemented only by a consideration of neighborhood *structural* context will fail to capture the additional impact of collective social organization. In particular, intergenerational closure, social cohesion, and expectations for informal social support of youth are likely to play a significant role in influencing the course of adolescent behavioral trajectories. Critically, the nexus of familial and collective supervision capacity may be central, particularly for girls.

Our analysis is not without limitations. First, although the focus of the PHDCN on a single city provides an opportunity to gather extensive data both on individual youth and their caregivers as well as broader features of neighborhood environments, the study is nevertheless limited to Chicago, complicating inferences to other populations. Second, neighborhood residence is not random, as families have some choice as to where they live. Some unmeasured characteristics may account for any observed neighborhood influences (Tienda 1991). Nevertheless, we observe neighborhood effects even after including a rich battery of micro level controls made available in the Project on Human Development in Chicago Neighborhoods data.

Future research on the etiology of early adolescent sexual behavior will benefit from more careful attention to the community contexts in which youth transitions are enacted and to the interactive effects of family-based processes and neighborhood environments. Parents are actively involved in organizing children's exposure to extra-familial environments (Furstenberg 1993) and may deliberately insulate younger adolescents, particularly girls, from the effects of deleterious community contexts. As data resources linking information on adolescent behavior to detailed features of multiple social

contexts—including families, peers, neighborhoods, and schools—researchers will be able to further elucidate the complex interactive process by which context shapes lives.

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Table 1. Descriptive Statistics for Variables in the Analysis (Wave 1)[*]

Independent Variables	Boys Mean	Boys Std. Dev.	Girls Mean	Girls Std. Dev.
Demographic and family background				
Race/ethnicity				
African-American	.306	-	.350	-
Latino	.494	-	.467	-
White/other	.200	-	.183	-
Age	14.694	1.536	14.810	1.522
Immigrant generation				
First[a]	.182	-	.115	-
Second	.306	-	.333	-
Third plus	.513	-	.552	-
Family socioeconomic status	-.054	1.473	-.188	1.431
Two biological parents[a]	.526	-	.454	-
Family size	5.361	1.967	5.281	1.922
Family attachment and support	2.801	.297	2.775	.313
Supervision (continuous)	.029	.634	.082	.602
Supervision (high place monitoring)	.609	-	.619	-
Peers				
Peer support[a]	2.442	.492	2.665	.429
Peer deviance[a]	-.303	1.247	-.481	1.303
Individual characteristics				
Pubertal development[a]	2.135	.512	2.510	.587
Prior problem behavior[a]	-.197	.974	-.343	.901
Sociability	3.613	.781	3.714	.778
Verbal aptitude	97.763	19.456	98.781	17.858

* Boys: N = 435; Girls: N = 480.

[a] Gender difference is significant ($p < .05$)

Table 2. Multilevel Discrete-Time Logit Models of The Timing of First Intercourse[a]

Independent Variables	Model 1 Boys	Model 1 Girls	Model 2 Boys	Model 2 Girls	Model 3 Boys	Model 3 Girls
Race/ethnicity						
African american	.973 **	.840 **	.806 **	.662 *	.669 *	.769 *
	(.316)	(.347)	(.353)	(.383)	(.337)	(.417)
Latino	-.373	.851 *	-.416	1.009 *	-.364	1.009 *
	(.352)	(.510)	(.403)	(.596)	(.379)	(.599)
Age	.615 ***	.718 ***	.643 ***	.695 ***	.524 ***	.579 ***
	(.083)	(.106)	(.088)	(.120)	(.106)	(.135)
Age squared	-.079	-.142 *	-.089	-.100	-.076	-.048
	(.066)	(.065)	(.065)	(.070)	(.066)	(.066)
Immigrant generation (vs. third)						
First	-.609 *	-1.470 *	-.439	-1.561 *	-.090	-1.160
	(.369)	(.765)	(.397)	(.807)	(.417)	(.859)
Second	-.265	-.616 *	-.163	-.698 *	-.102	-.489
	(.339)	(.373)	(.323)	(.412)	(.315)	(.428)
Family socioeconomic status	-	-	.125	-.047	.190 *	-.034
			(.108)	(.097)	(.107)	(.118)
Two biological parents	-	-	-.655 **	-.700 **	-.680 **	-.663 *
			(.272)	(.268)	(.269)	(.331)
Family size	-	-	.139 **	.038	.119 *	.044
			(.056)	(.070)	(.062)	(.076)
Supervision (continuous)	-	-	-.038	-.448 **	.101	-.526 **
			(.176)	(.156)	(.181)	(.180)
Family attachment and support	-	-	-.814 **	-.930 *	.658 *	.860 *
			(.319)	(.464)	(.323)	(.522)
Positive peer attachment	-	-	-	-	-.054	-.516
					(.241)	(.378)
Peer deviance	-	-	-	-	.155	.163
					(.102)	(.119)
Pubertal development	-	-	-	-	.358	.487 *
					(.231)	(.282)
Prior problem behavior	-	-	-	-	.372 ***	.080
					(.106)	(.205)
Sociability	-	-	-	-	.183	.476 *
					(.181)	(.212)
Verbal aptitude	-	-	-	-	-.009	-.003
					(.007)	(.010)
Intercept	-1.644 ***	-2.499 ***	-3.038 ***	-3.729 ***	-3.329 **	-6.036 ***
	(.297)	(.369)	(.602)	(.757)	(1.448)	(1.698)

[a] Neighborhood level N=77; Person-period level N=1699

* $p < .05$ ** $p < .01$ *** $p < .001$ (one-tailed tests). Standard errors in parantheses.

Table 3. Multilevel Discrete-Time Logit Models of The Timing of First Intercourse. Models Including Neighborhood Effects[a]

Independent Variables	Model 1		Model 2		Model 3		Model 4	
	Boys	Girls	Boys	Girls	Boys	Girls	Boys	Girls
Log concentrated disadvantage	.615 ***	.350 *	.628 ***	.345 *	.599 ***	.331	.404 *	.260
	(.163)	(.202)	(.172)	(.203)	(.175)	(.211)	(.239)	(.254)
Residential stability	-	-	.074	-.048	.132	-.008	.147	.008
			(.141)	(.175)	(.149)	(.194)	(.147)	(.199)
Immigrant concentration	-	-	-.229	-.195	-.322	-.229	-.295	-.217
			(.221)	(.222)	(.235)	(.226)	(.237)	(.225)
Collective efficacy	-	-	-	-	-.279 *	-.111	-.226 *	-.094
					(.129)	(.187)	(.130)	(.181)
Prevalence of adolescent childbearing	-	-	-	-	-	-	.276	.120
							(.190)	(.368)
Intercept	-2.807 *	-6.027 ***	-2.742 *	-5.919 ***	-2.474 *	-5.905 ***	-2.442 *	-5.936 ***
	(1.417)	(1.650)	(1.428)	(1.680)	(1.446)	(1.686)	(1.440)	(1.669)

Note: Models include all family, peer, and individual level covariates (see model 3 from Table 2).

[a] Neighborhood level N=77; Person-period level N=1699.

* $p < .05$ ** $p < .01$ *** $p < .001$ (one-tailed tests). Standard errors in parantheses.

Table 4. Multilevel Discrete-Time Logit Models of The Timing of First Intercourse. Models Including Neighborhood Effects[a]

Independent Variables	Model 1 Boys	1 Girls	2 Boys	3 Girls	4 Boys	5 Girls	6[b] Boys	6[b] Girls
Log concentrated disadvantage	.423 *	.255	.423 *	.300	.420 *	.307	.495 *	.276
	(.224)	(.210)	(.225)	(.225)	(.227)	(.230)	(.245)	(.221)
Residential stability	.151	-.043	.146	-.019	.148	-.014	.296	.080
	(.147)	(.161)	(.147)	(.164)	(.148)	(.167)	(.155)	(.165)
Immigrant concentration	-.306	-.247	-.314	-.239	-.314	-.236	-.166	-.209
	(.223)	(.177)	(.223)	(.183)	(.224)	(.189)	(.215)	(.222)
Collective efficacy	-.203 *	-.033	-.278	-.312 *	-.270	-.403 *	-.216 *	-.383 *
	(.120)	(.148)	(.185)	(.186)	(.187)	(.206)	(.117)	(.223)
Prevalence of adolescent childbearing	.300 *	.012	.326 *	-.078	.382	-.271	.286	-.324
	(.176)	(.260)	(.187)	(.279)	(.399)	(.353)	(.197)	(.378)
Supervision (high place monitoring)	.246	-.441 *	.225	-.388 *	.247	-.467 *	.109	-.451 *
	(.240)	(.224)	(.236)	(.231)	(.280)	(.246)	(.229)	(.254)
Collective efficacy x Supervision	-	-	.151	.387 *	.129	.509 *	-	.475 *
			(.267)	(.235)	(.285)	(.243)		(.275)
Adolescent childbearing x Supervision	-	-	-	-	-.085	.341	-	.473
					(.457)	(.392)		(.415)
Intercept	-2.415 *	-5.428 ***	-2.404 *	-5.731 ***	-2.407 *	-5.776 ***	-1.335 **	-1.899 ***
	(1.389)	(1.385)	(1.376)	(1.455)	(1.376)	(1.473)	(.499)	(.462)

Note: Models include all family, peer, and individual level covariates (see model 3 from Table 2).

[a] Neighborhood level N=77; Person-period level N=1699.

[b] Includes demographic background covariates and parental supervision only at level one.

* $p < .05$ ** $p < .01$ *** $p < .001$ (one-tailed tests). Standard errors in parantheses.

Figure 1. Cumulative Hazard of First Intercourse by Neighborhood Collective Efficacy (Boys)

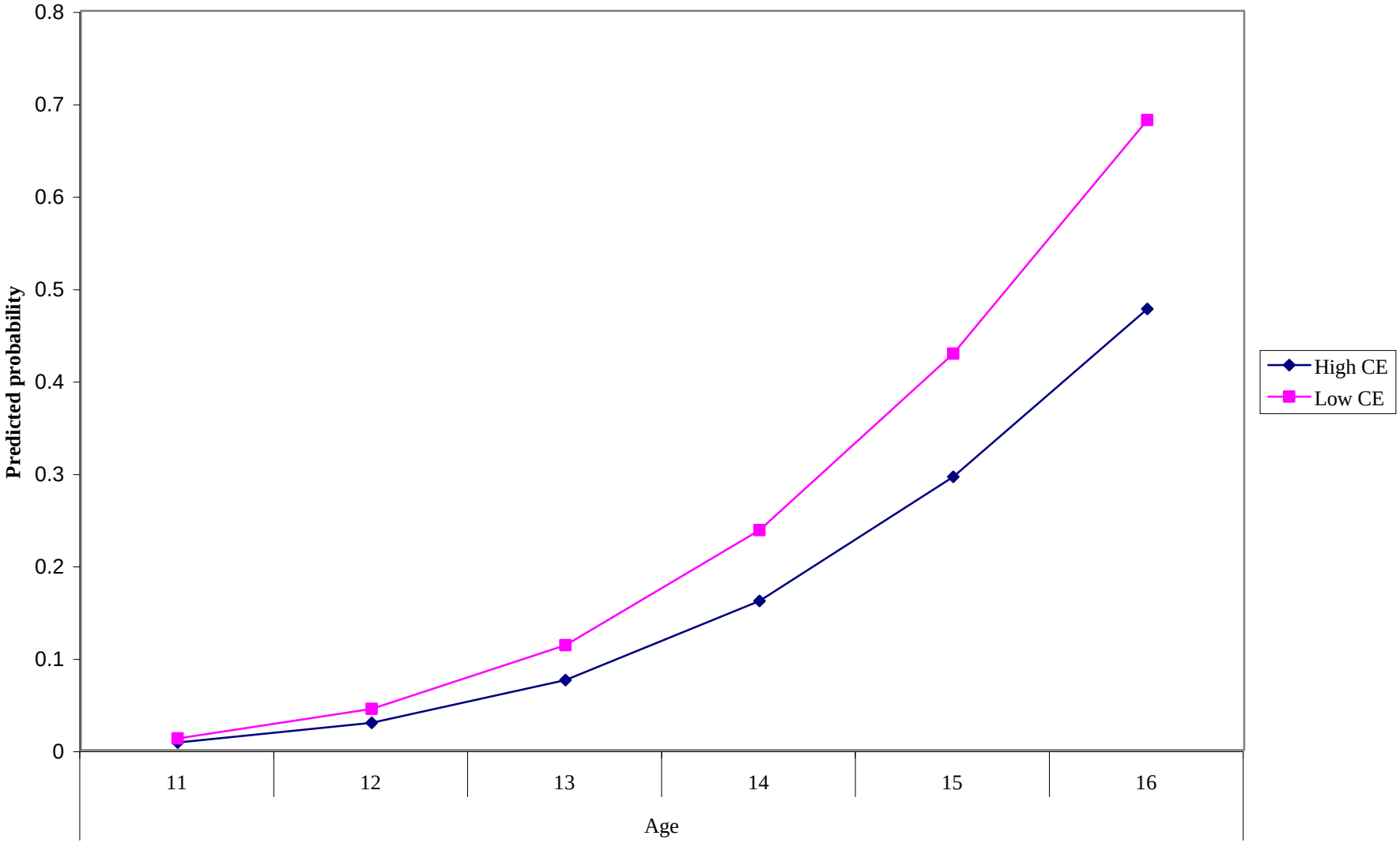
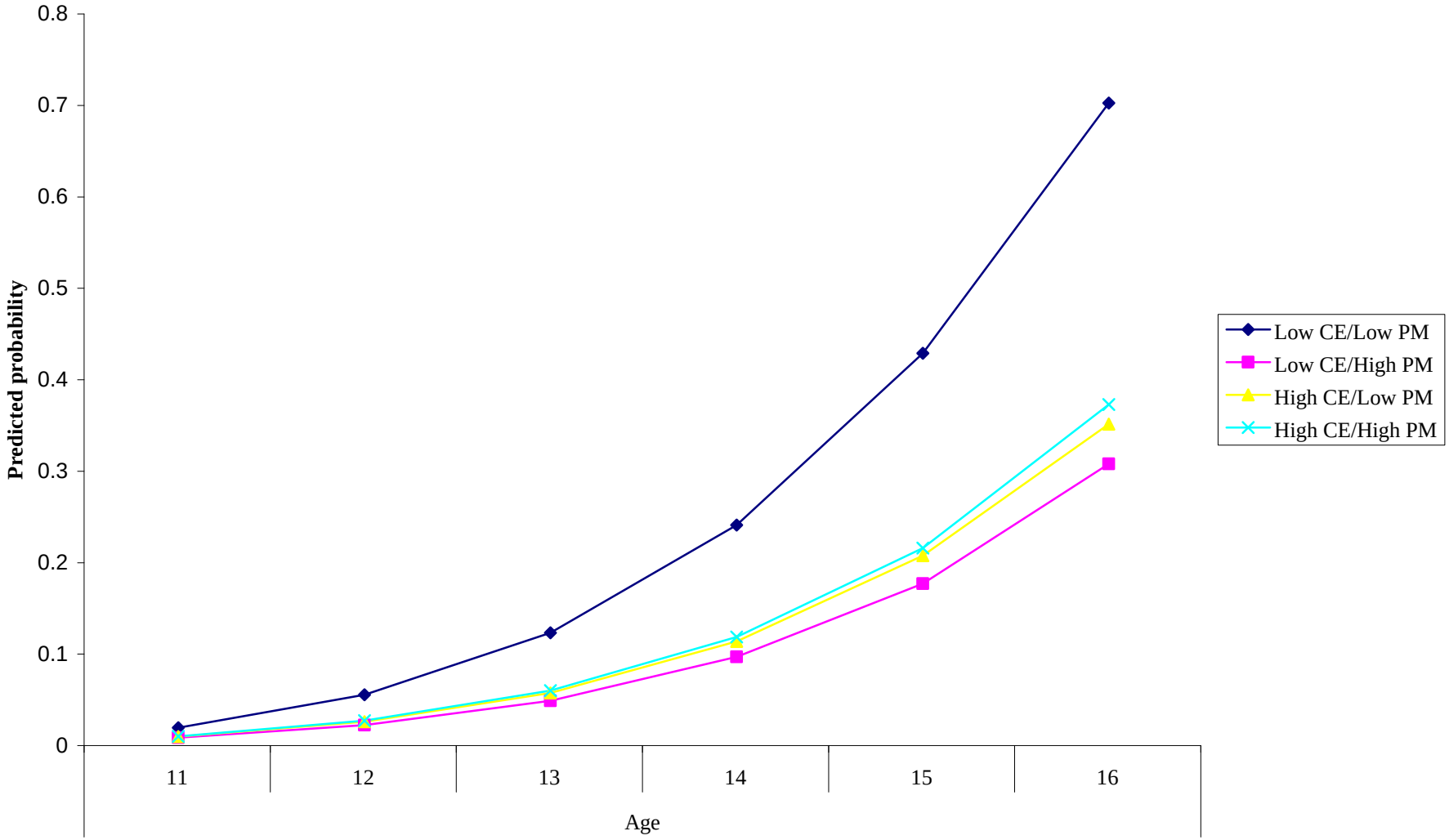


Figure 2. Cumulative Hazard of First Intercourse by Level of Parental Place Monitoring and Neighborhood Collective Efficacy (Girls)



Appendix—Component items of Supervision and Place Monitoring Scales

Item description	Global supervision	Place monitoring
Subject has a set time [curfew] to be home on school nights	✓	
Subject has a set time [curfew] to be home on weekend nights	✓	
PC [primary caregiver] requires subject to sleep at home on school nights	✓	
When PC is not available to subject at home, reasonable procedures have been established for him/her to check in with PC, or their designee, on weekends and after school	✓	
After school subject goes somewhere that adult supervision is provided	✓	✓
PC establishes rules for subject's behavior with peers and asks questions to determine whether they are being followed	✓	
Subject is not allowed to wander in public places without adult supervision for more than 2 hours	✓	✓
PC has had contact with two of the subject's friends in the last 2 weeks	✓	✓
PC talks daily with subject about his/her day	✓	
PC has visited the school or talked to the teacher or counselor within the last 3 months	✓	
PC has discussed television programs with subject during the past two weeks	✓	
PC has discussed current events with subject during the past two weeks	✓	
PC has discussed the hazards of alcohol and drug abuse with subject during the past year	✓	
PC knows signs of drug usage and remains alter to possible experimentation	✓	
Family has a fairly regular and predictable daily schedule for subject	✓	
PC is generally consistent in establishing or applying family rules	✓	