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ASSESSING CAUSALITY AND PERSISTENCE IN ASSOCIATIONS BETWEEN FAMILY DINNERS AND ADOLESCENT WELL-BEING

Kelly Musick and
Cornell University

Ann Meier
University of Minnesota

Abstract

Adolescents who share meals with their parents score better on a range of well-being indicators. Using three waves of the National Longitudinal Survey of Adolescent Health (*N* = 17,977), we assessed the causal nature of these associations and the extent to which they persist into adulthood. We examined links between family dinners and adolescent mental health, substance use, and delinquency at wave 1, accounting for detailed measures of the family environment to test whether family meals simply proxy for other family processes. As a more stringent test of causality, we estimated fixed effects models from waves 1 and 2, and we used wave 3 to explore persistence in the influence of family dinners. Associations between family dinners and adolescent well-being remained significant, net of controls, and some held up to stricter tests of causality. Beyond indirect benefits via earlier well-being, however, family dinners associations did not persist into adulthood.

Keywords

family demography; family structure; National Longitudinal Study of Adolescent Health (ADD Health); parental investment/involvement; well-being

In recent years, the search for ways for families to connect in an increasingly complex and fast-paced world has led back to the dinner table. The allure of the family meal has captured the attention of the popular press (David, 2010; Gibbs, 2006; Hoffman, 2009), policy groups (CASA, 2010; Child Trends, 2010), and researchers (for a review, see Fiese & Schwartz, 2008) in the U.S. and abroad (Ermisch, Iacovou, & Skew, 2011; Ross, 2011). The literature shows that adolescents who share family meals have healthier eating habits and body weight (Fulkerson, Kubik, Story, Lytle, & Arcan, 2009; Hammons & Fiese, 2011; Neumark-Sztainer, Hannan, Story, Croll, & Perry, 2003; Taveras et al., 2005; Videon & Manning, 2003), higher academic achievement (CASA, 2010; Council of Economic Advisers, 2000; Eisenberg, Olson, Neumark-Sztainer, Story, & Bearinger, 2004), better psychological well-being (Council of Economic Advisers, 2000; Eisenberg et al., 2004; Fulkerson et al., 2006, 2009), and lower risk of substance use and delinquency (CASA, 2010; Council of Economic Advisers, 2000; Eisenberg, Neumark-Sztainer, Fulkerson, & Story, 2008; Eisenberg et al., 2004; Fisher, Miles, Austin, Camargo, & Colditz, 2007; Fulkerson et al., 2006; Griffin, Botvin, Scheier, Diaz, & Miller, 2000; Sen, 2010). Associations span a range of teen outcomes and appear substantively significant, for example, the National Center on

CONTACT INFORMATION: Kelly Musick, Department of Policy Analysis and Management, Cornell University, 254 Martha Van Rensselaer Hall, Ithaca, NY 14853-4401 (musick@cornell.edu). Ann Meier, Department of Sociology, University of Minnesota, 1127 Social Sciences, Minneapolis, MN 55455 (meierann@umn.edu).

Addiction and Substance Abuse found that teens who ate fewer than three family dinners per week were about twice as likely to smoke, drink, and get poor grades as compared to teens who ate five to seven family dinners per week (CASA, 2010).

Given the strength of these associations and the fact that eating is universal and routine, family meals offer the potential to significantly influence child behavior and development. Indeed, Fiese and Schwartz (2008, p. 7) maintained that there are "few other collective settings in family life that have this potential across the child's early years into adolescence." Yet much of the literature on family meals is based on point-in-time study designs with limited accounting of other aspects of the family environment. Key questions remain unresolved: Are associations between family meals and child well-being causal? Do they persist over time? We address these questions using rich, nationally representative panel data on adolescents, following their mental health, substance use, and delinquency into young adulthood.

BACKGROUND

There are various arguments for a protective effect of family meals on adolescent mental health and risk-taking. Children thrive on routine and stability (Fiese, 2000; Fomby & Cherlin, 2007), and meals are an important part of what organizes a child's daily activities (Fiese et al., 2002). But more than just routine, mealtime may entail patterned, symbolic practices for many families, including favorite foods, structured roles, and expressions of gratitude. These rituals may be comforting, promoting feelings of closeness and belonging and providing a break from daily stressors (Fiese et al., 2002; Larson, 2008). Family meals may thus afford a regular and positive context for parents to connect with children emotionally, to monitor their social and academic activities, and to convey values and expectations – mechanisms that may directly foster children's well-being (Resnick et al., 1997). Other shared parent–child time or activities may not involve the same potential for regularity, ritual, or focused family time. For example, when teens were asked when, apart from dinner, they talk to parents about their life, over 70% responded they did so driving to or from school or activities (CASA, 2010), a context in which parents and children may be on average more distracted or hurried.

Popular and policy discussions of the link between family dinners and child well-being often assume a causal relationship. But the frequency of family meals is undoubtedly related to family resources, relationships, and other characteristics that contribute to well-being, and these may confound associations between family meals and child outcomes. For instance, lower socioeconomic status, maternal employment, single parenthood, and poor quality family relationships have all been linked to less frequent family meals (CASA, 2010; Cawley & Liu, 2007; Eisenberg et al., 2004; Fulkerson et al., 2006; Neumark-Sztainer et al., 2003; Stewart & Menning, 2009). These factors are in turn closely associated with child well-being (Duncan, Ziol-Guest, & Kalil, 2010; Fertig, Glomm, & Tchernis, 2009; Morrisey, Dunifon, & Kalil, 2011; Musick & Meier, 2010; Resnick et al., 1997). Family meals may thus be a marker for other aspects of the family, with little additional value to child development in and of themselves.

Only a handful of studies control for detailed characteristics of the family that may jointly account for family meals frequency and child well-being. Moreover, the bulk of the work on family meals relies on cross-sectional data (e.g., CASA, 2010; Council of Economic Advisers, 2000; Fulkerson et al., 2006, 2009; Griffin et al., 2000; Neumark-Sztainer et al., 2003), providing no leverage in sorting out temporal order and further constraining assessments of causality. The point-in-time or limited time frame of work in this area offers little insight into the association between family meals and child well-being as adolescents age. A small crop of recent studies on family meals and adolescent well-being have begun to address these issues by expanding their range of controls and incorporating longitudinal designs; we discuss these below.

Using population-based data from Project EAT on children ages 11 – 18 from Minneapolis and St. Paul area schools, Eisenberg and colleagues (2004) examined associations between family meals and a range of outcomes, controlling for family connectedness and race and ethnicity, parental education, and parents' marital status. Of the 8 outcomes examined in the domains of substance use, academic achievement, and psychological well-being, 6 had statistically significant associations with family meals among boys, and 3 remained statistically significant, net of family connectedness (frequent meals were inversely associated with smoking, drinking, and depressive symptoms). Among girls, all 8 bivariate associations were statistically significant, and 7 remained so net of family connectedness (frequent meals were inversely associated with smoking, drinking, and drug use; low grades; depressive symptoms; and suicide ideation and attempts). In a Project EAT follow-up, Eisenberg et al. (2008) examined the association between family meals in grades 7 and 8 and substance use 5 years later. Net of family connectedness and baseline use, female respondents who had frequent family meals had lower odds of subsequent substance use. Similarly, data from the Growing Up Today study (children of a study of nurses) showed that female respondents who ate frequent family meals had lower odds of initiating alcohol use one year later (Fisher et al., 2007).

Controlling for a broader set of family factors, Fulkerson et al. (2006) found an inverse association between the frequency of family dinners and high-risk adolescent behaviors (substance use, sexual activity, depression/suicide, antisocial behaviors, violence, school problems, binge eating/purging, and excessive weight loss). This study was based on nearly 100,000 surveys distributed to 6th to 12th graders from schools (self-selected into administering the survey) in 215 cities and 25 states. Family functioning was assessed with questions tapping parental support, communication, involvement in school, monitoring, role modeling, and academic expectations. Relationships between family dinners and adolescent outcomes were attenuated, but remained statistically significant for both girls and boys after adjusting for family factors and race and ethnicity, maternal education, and family structure.

Sen (2010) examined adolescent problem behaviors using data from the 1997 National Longitudinal Survey of Youth (NLSY97) and a similar set of controls, including indices for family connectedness, parental awareness, and family activities. Net of these controls, family dinners were negatively associated with substance use and running away among girls, and with drinking, physical violence, property destruction, stealing, and running away among boys. This analysis included 12 – 17 year-olds from the 1997 – 1999 rounds of the NLSY97 and used family dinners at $t + 1$ to address potential reverse causality and endogeneity, foregoing other panel methods and limiting the window of observation to the adolescent years.

In sum of extant studies, many associations between family meals and well-being in adolescence held up to controls for various aspects of family functioning. Further, family meals in adolescence remained a significant correlate of substance use 1 and 5 years later. Gender differences emerged, although the nature of these was not entirely consistent across studies. Recent research incorporating richer family controls and longitudinal designs has thus shed light on the importance of family meals, but there are gaps to be filled. Of the longitudinal studies addressing outcomes relevant to our analysis, only Eisenberg et al. (2008) controlled for earlier measures of the outcome. None of these studies used fixed effects approaches to address time-invariant individual-level heterogeneity or followed respondents beyond adolescence, and all but Sen (2010) relied on convenience or community samples. Controls for background factors tended to be limited, some studies including only crude measures of family structure and socioeconomic status, and none including maternal employment, leaving potentially important confounding family features unaccounted for.

We examine the links between family dinners and mental health, substance use, and delinquency, advancing the literature in important ways. We use three waves of the National Longitudinal Survey of Adolescent Health (Add Health), a nationally representative panel data set that follows adolescents into young adulthood. The Council of Economic Advisers (2000) utilized the first wave of Add Health to examine a related set of questions adjusting only for gender, poverty status, and family structure; ours is the first study of family meals and well-being to capitalize on the detailed and panel nature of these data. These features of Add Health allow us to move beyond much prior research in gaining leverage on causality and addressing the extent to which family dinners continue to matter into young adulthood. Our outcomes have potential implications for long-term health and attainment, and they tap both internalizing (i.e., depressive symptoms) and externalizing (i.e., substance use and delinquency) responses. We draw on various methodological approaches to take full advantage of data availability across waves, including fixed effects as our most stringent test of causality, and we test the sensitivity of our findings to the measurement of key variables. Finally, we assess the magnitude of family dinners associations relative to a commonly-studied correlate of child well-being, namely family structure.

Our controls include a rich set of variables potentially confounding prior estimates of the link between family meals and child outcomes. Families with older teenagers, fewer socioeconomic resources, and lower quality family relationships have been found to eat together less frequently (CASA, 2010; Cawley & Liu, 2007; Eisenberg et al., 2004; Fulkerson et al., 2006; Neumark-Sztainer et al., 2003; Stewart & Menning, 2009); these factors also relate to teen well-being and risk-taking (Duncan et al., 2010; Fertig et al., 2009; Morrisey et al., 2011; Musick & Meier, 2010; Resnick et al., 1997). We account for basic demographics, including age and racial and ethnic background; a host of variables tapping family resources, including family structure, number of children in the home, family income, parental education, and maternal employment; and a broader array of family relationship measures than most prior research, including global family relationship quality, parent–child relationship quality, activities with a parent, arguments with a parent, and parental control. Prior research was mixed in finding gender differences in associations between family meals and adolescent outcomes (e.g., Eisenberg et al., 2004, and Fisher et al., 2007, reported stronger associations among girls; Sen, 2010, found limited gender differences and Fulkerson et al., 2006, found no important differences). We tested gender by family dinners interactions and found little evidence of gender differences. We thus control for gender in our multivariate models but show all analyses pooled over gender.

Our analysis proceeds in three steps. First, we investigate the extent to which the association between family meals and adolescent well-being can be explained by sociodemographic characteristics of the family and the quality of family relationships (at wave 1) – factors potentially associated with both the frequency of family dinners and adolescent well-being. Second, to gain additional leverage on the causal relationship between family dinners and adolescent well-being, we use fixed effects models to estimate change in our outcomes as a function of change in family dinners over the course of one year. These models capitalize on the longitudinal nature of the data to control for pre-existing, stable individual differences, for example, temperament or family ideology that may influence both the family meal frequency and adolescent well-being. Finally, we examine persistence in the influence of family meals into adulthood, when children are 18 to 26 years of age (at wave 3).

METHOD

National Longitudinal Survey of Adolescent Health (Add Health)

Add Health is a nationally representative survey of U.S. adolescents who were in grades 7 to 12 in 1994 – 95. In 1995, more than 90,000 adolescents in 80 schools completed a self-administered, in-school questionnaire and more than 20,000 students and one of their parents completed an in-home interview. The Add Health cohort has been followed into young adulthood with a total of four in-home interviews; our study relied on the first three. The wave 2 in-home interview was conducted in 1996 and was limited to the 14,736 students who had not yet graduated high school. The wave 3 in-home interview was fielded in 2001 – 02 and included all wave 1 respondents who could be located in the United States, for a total of 15,170 respondents ranging in age from 18 to 26 (Harris et al., 2009). We drew primarily from the adolescent in-home questionnaires, although some information (i.e., parental education and family income) was taken from the resident parent questionnaire (fielded only in wave 1). In all analyses, we adjusted for Add Health's complex sampling design (Chantala & Tabor, 2010).

We relied on data from Add Health's probability sample, which includes 18,924 adolescents at wave 1 (this excludes 1,850 or 9% of respondents who were either not in the original sampling frame but added in the field or selected as part of a pair where both were not interviewed [Chantala & Tabor, 2010]). We excluded adolescents not living with a parent at the first wave (388 or 2% of cases), leaving a baseline sample of 18,536. For the first step of our analysis, we further excluded adolescents missing information on dependent variables of interest (559 or 3% of cases), for an analysis sample of 17,977. In the second step of our analysis (i.e., examining change between waves 1 and 2), we lost 3,394 cases (18%) of the baseline sample of 18,536 due to nonfollowup; we lost an additional 1,750 (12%) due to nonresponse, 383 (3%) with no parent or parent figure in the household, and 563 (4%) with missing data on dependent variables at wave 1 or 2, for a sample of 12,446. Finally, for the third step of our analysis, focusing on persistence in the association between family dinners and well-being into young adulthood, we relied on those who completed a wave 3 interview. From the baseline sample of 18,536, we lost 4,471 (24%) due to nonresponse and 584 (4%) due to missing data on wave 3 dependent variables, for a sample of 13,481.

Only two of our variables were missing data for more than 5% of our analysis samples: family income was missing for 25% of cases (this was ascertained only in the parent interview, which was not completed for 14% of our wave 1 sample); and father's education was missing for 6% of cases (collected from both parents and adolescents, but disproportionately missing for respondents living apart from a father). We imputed missing data using chained equations in Stata informed by our analysis variables and a small number of auxiliary variables. All cases were used in the imputation, although we excluded those with imputed dependent variables from our analysis (von Hippel, 2007). We generated 25 datasets and combined estimates from the multiply imputed data using Stata's MI prefix. Our key findings appeared robust to variations in imputation model and number of datasets, consistent with Johnson and Young's (2011) sensitivity analysis showing few differences in findings based on a range of imputation strategies applied to large-scale family data.

Outcomes

We examine three outcomes capturing well-being in adolescence and young adulthood: depressive symptoms, substance use, and delinquency. *Depressive symptoms* were assessed at waves 1, 2, and 3 using nine items from the Center for Epidemiological Studies Depression Scale (CES-D). Respondents were asked, "How often was each of the following things true during the past week?" 1) you were bothered by things that usually don't bother you; 2) you felt that you could not shake the blues, even with help from family and friends; 3) you felt that you were just as good as other people; 4) you had trouble keeping your mind on what you were doing; 5) you felt depressed; 6) you felt that you were too tired to do things; 7) you enjoyed life; 8) you felt sad; and 9) you felt that people disliked you. Response options were 0 = *never or rarely*, 1 = *sometimes*, 2 = *a lot of the time*, and 3 = *most or all of the time*. We reverse-coded items 3 and 7 and averaged over all items for a scale ranging from 0 to 3, with higher scores indicating more depressive symptoms (α = .79 at wave 1 and .80 at waves 2 and 3).

We used six questions to construct our measure of *substance use*. At waves 1 and 2, items pertained to the use of alcohol, cigarettes, marijuana, cocaine products, inhalants, and other illegal drugs. At wave 3, inhalants were grouped with other illegal drugs, and respondents were instead asked separately about crystal methamphetamine. The time referent was ever using the substance at wave 1, since the date of last interview at wave 2, and in the past year at wave 3. To recognize the social and legal acceptability of alcohol consumption in adulthood and to maintain consistency in measurement, at each of the three waves we indexed whether the respondent engaged in binge drinking (five or more drinks in one sitting) as opposed to any drinking. Due to the dissimilarity in response options for the substances, we followed the example of McCarthy and Casey (2008) and constructed a dichotomous indicator of any use across the six items.

To assess *delinquency*, we generated an index based on self-reports of participation in 14 delinquent activities in the past 12 months (following Pearce & Haynie, 2004). These included painting graffiti, damaging property, shoplifting, stealing something worth less than $50, stealing something worth $50 or more, burglarizing, using a car without the owner's permission, selling drugs, getting into a serious physical fight, seriously injuring another person, threatening to use a weapon on someone, getting into a group fight, pulling a knife or gun on someone, or shooting or stabbing someone. Adolescents reported on the same items at waves 1 and 2; at wave 3, 3 items were changed to reflect more age-relevant behaviors (see Table 1A, on-line supplement). We counted the number of delinquent activities from 0 to 14, creating an index that is highly skewed, with 43% reporting 0 delinquent acts at wave 1, 54% at wave 2, and 72% at wave 3. The α reliability coefficients were .80 at waves 1 and 2 and.72 at wave 3 (these are equivalent to Kuder-Richardson coefficients for assessing the reliability of dichotomous items [Cortina, 1993]).

Our measures of substance use and delinquency included items that range from reasonably common to much more deviant. For example, between a quarter and over a half of respondents reported smoking, binge drinking, and using marijuana, whereas 10% or less reported using cocaine, inhalants, or other illicit drugs. Similarly, over 20% of respondents reported getting into a physical fight or shoplifting, whereas less than 5% reported using a weapon or breaking into a house (see Table 1A, on-line supplement). We tested the sensitivity of our results to examining single items representing the most common behaviors. Our key findings remained the same. This may reflect the tendency of respondents who engage in the more deviant behaviors to also engage in the more common ones. For example, fully 99% of teens who reported using cocaine also reported smoking cigarettes, binge drinking, or using marijuana.

Family dinners

Our key explanatory variable is *family dinners*. At waves 1 and 2, adolescents were asked, "On how many of the past 7 days was at least one of your parents in the room with you while you ate your evening meal?" Existing literature offers no standard way of coding family meals, although when researchers use a categorical measure, a typical cut-off indexing "frequent" family dinners is 5 or more per week (e.g., CASA, 2010; Council of

Economic Advisers, 2000). For descriptive analyses, we categorized the frequency of family meals into low (0 – 2), medium (3 – 4), and high (5 – 7). We retained the full count (0 – 7) in our models.

Family environment and other controls

We included five measures of the quality of family relationships, assessed at waves 1 and 2: global family relationship quality, parent–child relationship quality, activities with a parent, arguments with a parent, and parental control. *Global family relationship quality* is an average of responses to three questions: 1) how much does your family understand you; 2) how much fun does your family have together; and 3) how much attention does your family pay to you. Response options are on a scale from 1 = *not at all* to 5 = *very much*, with higher scores indicating better family relationships (α = .79 at wave 1 and .70 at wave 2).

Parent–child relationship quality is based on questions addressing in turn adolescents' relationships with their resident mothers and fathers, assessed on a scale from 1 = *strongly agree* to 5 = *strongly disagree*: 1) how close do you feel to your mother/father; 2) how much does your mother/father care about you; 3) how warm and loving is your mother/father towards you; 4) how satisfied are you with your communication with your mother/father; and 5) how satisfied are you with your relationship with your mother/father. We reverse-coded items so that higher values represent better relationships and separately averaged items pertaining to mothers and fathers (α = .85 and .84 for mothers' scores and .89 and .87 for father' scores at waves 1 and 2, respectively). We took the higher of the two scores (or just one in the case of single parents).

We generated a count of *activities with a parent* based on adolescent reports of whether they had done any of the following with their resident mother or father in the past 4 weeks: 1) gone shopping; 2) played a sport; 3) gone to a religious or church event; 4) gone to a theater or museum; or 5) worked on a school project. Questions were asked separately in reference to mothers and fathers and, as above, we generated counts for mothers and fathers (α = .39 for mothers' scores at both waves and .43 and .42 for fathers' scores at waves 1 and 2, respectively) and took the higher score to represent activities with a parent. (α's were weak, but separate scale items were related in similar ways to our outcomes; see Goncy & van Dulmen, 2010, and Jordan & Lewis, 2005, for more on the scale's properties.) In this same question series, adolescents were asked whether they had gotten into a serious argument about their behavior with their resident mother or father in the past 4 weeks. We generated a dichotomous indicator for *arguments with a parent* coded "1" if the adolescent reported a serious argument with either their mother or father. Whereas our other family relationship measures index positive family interactions, assessing arguments explicitly taps negative interactions.

Our last indicator of family relationships is a measure of *parental control*, based on adolescent reports of whether parents let them make their own decisions about: 1) weekend curfew; 2) friends; 3) clothing; 4) television time; 5) television programming; 6) weeknight bedtime; and 7) what they eat. We generated a count from 0 – 7 of the number of "no" responses. Higher values indicate more parental control (α = .63 at wave 1 and .65 at wave 2).

We controlled for variables tapping family resources at waves 1 and 2 that may affect both the management of a regular family dinner and child well-being, including family structure, number of children in the household, family income, parental education, and maternal employment. Family structure was coded "two biological parents" when two biological or adoptive parents were in the household; "stepparent" when there was a biological or adoptive parent and the parent's spouse or cohabiting partner; "single parent" when there was a biological or adoptive parent and no spouse or cohabiting partner; and "other" in the case of households with no biological or adoptive parents present, for example, children living with grandparents or other adults. Family income was constructed from a question on the wave 1 parent interview (and was ascertained only at wave 1 from the parents). Parental education and maternal employment were generated based on adolescent reports at waves 1 and 2 and supplemented with data from the wave 1 parent interview. Parental education refers to the resident parent if the parent lives with the adolescent and to the nonresident parent otherwise, and it is categorized into "less than high school," "high school," "some college," and "college or more." Combining questions on mother's work for pay and work hours as of the interview date, maternal employment is categorized into "full-time employment" (35 or more hours per week), "part-time employment" (less than 35 hours), and "not employed" (where there was no resident mother, we filled in with data on paternal employment). All multivariate models controlled for the adolescent's gender (female vs. male respondent), age at wave 1, and racial and ethnic background (non-Hispanic White, non-Hispanic Black, Hispanic, Asian, and other race). Weighted descriptive statistics of all variables appear in Table 2A of the on-line supplement.

Models

The first step of our analysis asks whether family meals are associated with our outcomes at baseline, and whether any association can be explained by other factors. We used ordinary least squares (OLS) to model depressive symptoms, logistic regression to model substance use, and negative binomial regression to model delinquent acts. As noted, our delinquency count was highly skewed to 0, and we found a statistically significant over-dispersion parameter in delinquency models indicating a conditional variance greater than the conditional mean. The over-dispersed nature of the delinquency data suggested that the negative binomial would provide a better approximation to the data than the Poisson model (Long & Freese, 2006).

We estimated three models for each outcome, beginning with family dinners as the only covariate, then adding sociodemographic controls (age, gender, racial and ethnic background, family structure, parental education, and family income), and finally including our five measures of family relationships. The goal was to assess the extent to which the family dinners association can be accounted for by related aspects of the family environment. Only in the case of our OLS model of depressive symptoms, however, can change in the family dinners coefficient across models be straightforwardly attributed to the inclusion of additional covariates. In the logit and negative binomial regressions, the change in our family dinners coefficient across nested models will reflect both the confounding due to other covariates and rescaling (Karlson, Holm, & Breen, Forthcoming). Rescaling arises in limited dependent variable models because estimated coefficients depend on the error variance of the model, which in turn depends on other covariates in the model. We applied the method proposed by Karlson et al. for addressing this scaling problem and testing the statistical significance of change in coefficients across nested models, implementing their approach using the Stata command –khb– (Kohler & Karlson, 2011). This methodology should apply to the entire family of generalized linear models, although the result has been formally derived only for binary nonlinear probability models (i.e., logits and probits) and cumulative probability models (Karlson, June 2011, e-mail correspondence). We are thus more tentative in interpreting the change in family dinners coefficients across negative binomial regression models of delinquency.

The second step of our analysis capitalized on the longitudinal nature of the data and provided a more stringent test of causality. Using the same measures of mental health, substance use, delinquency, and family dinners at waves 1 and 2, we estimated first difference models, which are equivalent to fixed effects models in the two period case. The fixed effects approach, here, regressing change in our outcomes assessed at waves 1 and 2 on change in family dinners assessed at waves 1 and 2, has two principal advantages (Allison, 1990; Liker, Augustyniak, & Duncan, 1985; Winship & Morgan, 1999). First, regressing Δy on Δx eliminates bias due to time-invariant unobserved factors that might jointly determine family dinners and adolescent well-being (e.g., temperament or family ideology). Second, by modeling changes as opposed to levels, it reduces bias due to persistent reporting errors, for example, any tendency to misreport depressive symptoms, substance use, or delinquency. The estimated effect of family dinners nonetheless potentially suffers from bias due to time-varying unobservables. To reduce this possibility, we ran models controlling for changes in the quality of family relationships, family structure, and maternal work. Unobserved changes remain a potential source of bias, as does any reverse causation, or change in children's behavior that might affect family dinners.

The suitability of first difference models for change in substance use and delinquency may be more questionable than for change in depressive symptoms (which may be reasonably approximated by a normal distribution). The linear probability model is not entirely appropriate for modeling substance use because it has nonnormal error terms and does not bound the probabilities between 0 and 1. And in terms of delinquency, the distinction between uptake and desistance may be more salient than a one-unit change in delinquent acts. We thus ran sensitivity tests based on nonlinear, fixed effects models of uptake. We used conditional logits, which are estimated only on respondents who change status in between waves (Freeman, 1978). This means comparing respondents who take up substance use or delinquency ($t1 = 0$ and $t2 = 1$) to respondents who quit or desist in these behaviors ($t1 = 1$ and $t2 = 0$). We modeled the conditional probability of uptake using logistic regression. Our key findings were robust to our method of analysis (results available upon request).

The third and final step of our analysis addressed whether or not associations between family dinners and well-being persisted into adulthood. We began by regressing our wave 3 outcomes on wave 1 family dinners and controls. As in step 1, we estimated models of depressive symptoms using OLS, models of substance use using logistic regression, and models of delinquent acts using negative binomial regression (our delinquency findings were robust to dichotomizing our measure and counting as delinquent those convicted of an adult crime or those in prison at wave 3; results available upon request). We included controls for sociodemographic characteristics and the quality of family relationships. We then added a control for the wave 1 assessment of the outcome. Because most respondents were no longer in the home in young adulthood, we could not link change in family dinners to change in outcomes. Controlling for an earlier measure of the outcome is a way of accounting for unobserved person-level variation and thus offers a test for the effect of family dinners into young adulthood (Johnson, 2005).

RESULTS

Descriptive Patterns

Table 1 shows descriptive statistics on key measures at wave 1, for the full sample and by family dinners frequency. For descriptive purposes, we categorized family dinners into low (0 – 2 days), medium (3 – 4 days), and high (5 – 7 days) frequency. The first row shows that 23% of adolescents reported family dinners less than 3 times a week, and 60% reported 5 or more weekly shared dinners, results similar to those from other national studies (CASA, 2010; Sen, 2010).

The second set of rows in Table 1 shows that more frequent family meals were associated with fewer depressive symptoms, lower substance use, and fewer delinquent acts among adolescents. Differences in our outcomes by family dinners were statistically significant at the 5% level and appeared reasonably large, for example, 0.78 versus 0.58 depressive symptoms among teens eating less than 3 versus 5 or more meals a week with their parents (a difference of over a half a standard deviation in the depressive symptoms scale). Differences in the quality of family relationships also varied significantly by family dinners frequency (third set of rows), with adolescents frequently sharing meals reporting higher global family relationship quality; better relationships, more activities, fewer arguments with a parent; and greater parental control. Families with both biological parents present, a nonemployed mother, and higher income were overrepresented among those frequently eating together, and stepparent families, single-parent families, and those with a full-time employed mother were underrepresented (final set of rows).

Step 1: Multivariate Analysis of Baseline Associations

Table 2 reports the first step of our multivariate analysis, that is, the wave 1 associations between family dinners and our outcomes, successively adding controls. For each of our three outcomes, Model 1 includes only family dinners, Model 2 adds sociodemographic controls, and Model 3 adds our five measures of the quality of family relationships. The final 2 rows of the table indicate the percent change in the coefficient on family dinners as controls are added. These changes are estimated using a method that allows for the unbiased comparison of coefficients across nested, nonlinear models (Karlson et al., Forthcoming; Kohler & Karlson, 2011). They show the extent to which the family dinners associations can be accounted for by sociodemographic characteristics and family relationships.

The first set of columns in Table 2 reports results of our OLS model of depressive symptoms. In Model 1, an increase of one family dinner per week was associated with a statistically significant reduction of 0.04 on the depressive symptoms scale. Adding sociodemographic controls in Model 2 reduced the family dinners coefficient by 19%, and including controls for family relationship quality in Model 3 reduced the bivariate association by a total of 61%. The family dinners coefficient remained statistically significant net of controls, but was much attenuated by factors that vary both by family dinners and adolescent well-being.

In the full model (Model 3), the family relationship measures were associated with depressive symptoms largely as expected. To assess the relative magnitude of key associations, we used parameter estimates from Model 3 to generate predicted depressive symptom scores, adjusted according to measures of interest. Varying the number of family dinners from 2 to 7 (about a +/− 1 standard deviation change in family dinners) and holding all other covariates at mean values, we estimated predicted depressive symptom scores of 0.67 and 0.60, respectively (relative to an overall mean of 0.63). This translates into a difference of 0.07 points, or about 15% of a one standard deviation change in the depression scale (i.e., 0.47; see Table 1). The difference in predicted symptoms for adolescents from single-parent and two-biological parent families was 0.03 (0.65 vs. 0.62, respectively). This suggests that family dinners are substantively important for depressive symptoms, on a scale larger than family structure (when assessed at 2 vs. 7 dinners per week).

The second set of columns in Table 2 reports exponentiated coefficients, or odds ratios, from our logistic regression models of substance use. In Model 1, an increase of one family dinner per week was associated with a statistically significant reduction in the odds of substance use of .85 times or 15%. Adding sociodemographic controls in Model 2 reduced the size of the family dinners coefficient by 24%, and including controls for family relationship quality in Model 3 reduced the bivariate association by a total of 56%. The family dinners coefficient remained statistically significant, with an increase of one family dinner per week in the full model associated with an 8% reduction in the odds of substance use.

In the full model, all family relationship variables were associated with the risk of substance use in the expected direction. To put the magnitude of results in context, we again generated predicted probabilities based on Model 3, varying key contrasts and holding other variables at their mean values. The overall predicted probability of substance use was .65. The predicted probability of substance use dropped from .70 to .61 with an increase in family dinners from 2 to 7. This association was similar in magnitude to the difference in substance use between adolescents from single-parent and two-biological parent families (i.e., .71 vs. .61).

The third set of columns in Table 2 reports results of our negative binomial regression of delinquency, which models the log of the expected count of delinquent acts. We show exponentiated coefficients, or incidence rate ratios. In Model 1, an increase of one family dinner per week was associated with a statistically significant decline in delinquent acts by a factor of .94 or 6%. Adding sociodemographic controls in Model 2 did not attenuate the size of the family dinners coefficient; including controls for family relationship quality in Model 3 reduced the coefficient by an estimated 49% (recall that the estimated change in coefficients across nested negative binomial models should be interpreted with caution). The coefficient remained significant, implying a 3% reduction in delinquent acts. As with depressive symptoms, the family dinners association was more sensitive to the addition of family relationship variables than other family features like family structure or maternal employment.

In the full model all of the family relationship variables were associated with the risk of delinquency in the expected direction. We again estimated predicted counts of delinquent acts based on Model 3, varying key contrasts. With all covariates held at their mean values, we estimated an overall count of 1.62 delinquent acts. Increasing family dinners from 2 to 7, we estimated a drop in the count of delinquent acts from 1.67 to 1.41, which translates into 0.26 points or 11% of a one standard deviation change in delinquency (i.e., 2.35; see Table 1). This association was somewhat smaller than the difference in predicted delinquent acts between adolescents from single-parent and two-biological parent families (i.e., 1.75 vs. 1.41).

To summarize results so far: First, bivariate associations between family dinners and each of our outcomes were statistically significant and reasonably large. Second, despite variation in family structure, family income, and maternal employment by family dinners, the addition of sociodemographic controls accounted for a relatively modest share of the family dinners association with depressive symptoms and substance use and none of the association with delinquency. The family dinners coefficient was reduced more substantially, however, when the quality of family relationships was controlled. Family dinners appeared to partially proxy for more general aspects of the family environment. Together, family resources and relationships accounted for half or more of the bivariate association between family dinners and our adolescent outcomes. Nonetheless, even in our full models, family dinners were both statistically and substantively important. Associations between family dinners and our outcomes were similar in magnitude to differences by family structure.

Step 2: Fixed Effects Models of Change Across Waves

Having access to a rich set of potentially confounding covariates, in particular, detailed aspects of the family environment, provides leverage on questions of causality. We went a step further by estimating fixed effects models of change in our outcomes on change in family dinners. This approach accounts for unobserved time-invariant factors that might jointly account for family dinners and adolescent well-being. We also controlled for changes in the quality of family relationships, family structure, and maternal employment to limit the

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extent to which changes in family dinners may be picking up other changes related to adolescent well-being.

The first set of columns in Table 3 shows results for depressive symptoms. We estimated the association between change in family dinners and change in symptoms over the course of a year (between waves 1 and 2), without controls (Model 1) and with controls for other observed changes (Model 2). The coefficients on change in family dinners were statistically significant at the $P < .05$ level in both models, although controls reduced the coefficient by about 40%, indicating that an increase of one family dinner decreased depressive symptoms by 0.005. Increases in parent–child and global family relationship quality reduced depressive symptoms, and increases in arguments with parents increased them. Predicted depressive symptoms based on Model 2 showed a slight upward trend across waves, with an estimated mean change in the depression scale of 0.10. A change in family dinners of 5 (an approximately +/− one standard deviation change) reduced depressive symptoms by 0.03. The family dinners effect represented less than 10% of a standard deviation change in our outcome (the standard deviation of our depressive symptoms change scores = 0.44; see Table 2A, on-line supplement). Changes in family structure were not statistically significantly related to changes in depressive symptoms.

The second set of columns in Table 3 report on models of change in substance use. In Model 1, the coefficient on family dinners was −0.004 and statistically significant at $P < .05$. When we added controls (Model 2), the coefficient dropped to −0.003 and lost statistical significance ($P = .12$). Of the family relationship variables, only global family relationship quality and arguments with parents were statistically significant, with increases in relationship quality reducing substance use and increases in arguing elevating it. Mothers moving into full-time employment elevated the risk of substance use, but again changes in family structure were not statistically significant. We tested the sensitivity of results to modeling this process in a nonlinear fashion. Results based on a conditional fixed-effect logit indicated no statistically significant relationship between change in family dinners and smoking uptake.

The final set of columns in Table 3 report results for delinquency. We found no effect of change in family dinners on change in delinquency. We found changes in family relationship variables largely associated with changes in depressive symptoms in the expected directions with two exceptions: increases in activities with parents were *positively* associated with delinquency and changes in family structure were *negatively* associated. We tested whether this process could be better modeled as uptake versus desistance, applying a conditional logit model. The family dinner coefficient was close to zero and statistically nonsignificant in this model, as well.

In sum, change in family dinners was significantly related to change in depressive symptoms. Changes in family dinners were also statistically significant predictors of changes in substance use in a model without controls, but adding changes in other aspects of the family environment reduced the statistical significance of family dinners to $P = .12$, above conventional levels of significance. We found no evidence of a causal effect of family dinners on delinquency, irrespective of how we modeled the process. Our findings suggest a causal relationship between family dinners and some dimensions of adolescent well-being, or at least a relationship that is not completely driven by time-invariant variables, observed or unobserved.

Step 3: Persistence of Effects to Young Adulthood

Table 4 reports on whether any effects of family dinners persist into young adulthood. Model 1 includes all wave 1 sociodemographic controls and the quality of family relationships. Model 2 includes the wave 1 value of our outcomes, serving as a control for person-level, time-invariant heterogeneity. The first set of columns show OLS coefficients for depressive symptoms. Results for Model 1 indicate that wave 1 family dinners were significantly related to wave 3 depressive symptoms, controlling for sociodemographic characteristics and the quality of family relationships. When we added the wave 1 measure of depressive symptoms in Model 2, however, the coefficient on family dinners dropped to zero and lost statistical significance. This was the same story for the logistic regression model of substance use, that is, a statistically significant association between family dinners and substance use (Model 1) that disappeared when the wave 1 measure of substance use was included (Model 2). Family dinners was statistically nonsignificant in both negative binomial regression models of delinquency. We ran a logistic regression on delinquency, also counting as delinquent anyone convicted of an adult crime or in prison at wave 3. Results were not sensitive to these changes.

DISCUSSION

There has been much attention in recent years to the social and health benefits of the family meal, but questions remain about the nature of this association, in particular, whether it is causal and the extent to which it persists. We set out to address these questions, capitalizing on rich, panel data from Add Health to examine links between family dinners and multiple indicators of well-being from adolescence into young adulthood. Consistent with prior literature, we found strong bivariate associations between the frequency of family dinners and lower levels of depressive symptoms, substance use, and delinquent acts among teens.

Eating together may protect children from depression and risky behaviors by providing a regular and comforting context to check in with parents about their day-to-day activities and to connect with them emotionally. The ability to manage a regular family dinner, however, may also be facilitated by family resources like time and money, or it may simply be a proxy for other, more affective, dimensions of the family environment. Indeed, we found more frequent family dinners among families with both biological parents present, a nonemployed mother, and higher income. Controlling for these factors explained a relatively modest share of the family dinners association with depressive symptoms and substance use and none of the association with delinquency. Teens scoring higher on indicators of family quality also ate more frequently with their parents, and these factors in turn explained between a quarter to half the association between family dinners and adolescent outcomes (net of sociodemographic controls). Accounting for the rich set of controls available in Add Health, wave 1 family dinners associations were attenuated but remained statistically and substantively significant (for example, increasing family dinners from 2 to 7 per week was associated with a 0.07 decline in depressive symptoms, or 15% of a one standard deviation change in the depression scale). Examining wave 1 associations leaves open threats to causal inference, but it also provides a more detailed account of the processes linking family dinners and adolescent well-being.

As a more stringent test of causality, we estimated first difference models from waves 1 and 2, accounting for time-invariant unobserved factors that might jointly determine family dinners and adolescent well-being. Change in family dinners was significantly related to change in depressive symptoms, it was significantly related to change in substance use only prior to the addition of controls for other observed changes in the family environment, and it was not related to change in delinquency. The final step of our analysis examined whether any effect of family dinners persisted into young adulthood. Whereas adolescent family dinners were significantly associated with wave 3 measures of depressive symptoms and substance use controlling for wave 1 sociodemographic characteristics and the quality of family relationships, these associations did not stand up to controls for earlier observations on our outcomes. Evidence for family dinners effects appear weaker based on our analyses of the panel data, suggesting that the seemingly large effects of family dinners as estimated at the cross section (and in much prior research) are due mostly to unobserved factors. Alternatively, or also in part, observing change over one year may not be long enough for family dinner routines and rituals to take hold to generate positive feedback. Further, whereas family dinners do not exert independent effects into young adulthood, they may have indirect effects on young adults via earlier well-being.

Given somewhat mixed results and various ways of assessing their substantive implications, it helps to compare findings on family dinners to those on family structure, a commonly studied correlate of child well-being. In our analysis of wave 1, relative to differences associated with living with a single parent versus both biological parents, differences associated with an increase from 2 to 7 family dinners were larger in depressive symptoms, similar in substance use, and smaller in delinquency. Fixed effects models provided consistent evidence of family dinner effects only in the case of depressive symptoms. Likewise in these models, changes in family structure were inconsistently related to our outcomes. Taken together, associations between family dinners and well-being seemed to manifest in ways similar to the oft-studied association between family structure and child well-being.

This work advances the literature on family meals, tapping rich measures of family resources and relationships and examining a longer window into adulthood with longitudinal methods that have not been previously applied to investigations of family meals and well-being. Even with detailed, prospectively collected data and rigorous methods, however, it remains difficult to disentangle family processes. Our estimates of family dinner effects may be upwardly biased by reverse causation, for example, capturing adolescents engaging in risky behaviors who skip the family dinner in order to avoid contact with parents. At the same time, our approach may underestimate the total effect of family dinners. Having measures of the quality of family relationships allowed us to unpack their association with child well-being. But if family dinners are a building block of closeness with parents or feelings of support within the family, family dinner effects are in part indirectly channeled through the quality of family relationships.

As noted at the outset, the universal and routine nature of eating affords mealtime the potential to influence child behavior and development. Our finding associations in adolescence that hold up to individual fixed effects lend some support to a causal interpretation of the link between family meals and at least some aspects of child well-being. But we continue to have very little understanding of the mechanisms behind this relationship (Fiese & Schwartz, 2008). Add Health asks whether “at least one of your parents” was “in the room” during mealtime, not accounting for other family members, or requiring that family members be seated around a table or even interacting. Further, it refers to the “evening meal,” although the social benefits of family meals may not be limited to dinners. In data from Project EAT, more frequent family meals, high priority for family meals, and a positive and more structured atmosphere at meals were all associated with lower risk of disordered eating among adolescents (Neumark-Sztainer et al., 2004). CASA (2010) reported that 75% of teens say they talk to parents about their life at mealtime, and these teens are both more likely to rate shared meals as important and have lower levels of substance use. Future work needs to go further in assessing which elements of mealtime are most salient, looking beyond how often families eat together to examine whether talking, television, texting, eating the same food, or helping in the kitchen mediate or moderate the potential benefits of mealtime.

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What, then, does our study say about the "magic" (Gibbs, 2006) of the family meal? We end on a cautionary note to parents and policy makers eager to find ways to improve family functioning and the fortunes of children. We know from research on children's school outcomes that there is at best a modest match between observational studies showing powerful influences of the family and experimental studies designed to put the findings of this literature into practice (Furstenberg, 2011). Furstenberg argues that many family factors contributing to children's positive developmental trajectories are overlapping and mutually reinforcing, and that it is naïve to expect to alter family practices "by isolating one particular element of family life without acknowledging how families operate as social systems" (p. 465). Family dinners may be part and parcel of a broader package of practices, routines, and rituals that reflect parenting beliefs and priorities. Interventions aimed at increasing the frequency of family meals may be successful only if they can change the family habits that tend to go along with eating as a family.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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Table 1

Means on Key Variables by Frequency of Family Dinners at T1, Add Health (N = 17,977)

		Frequency of Family Dinners (0 – 7 days/week)		
	All	**Low: 0 – 2**	**Medium: 3 – 4**	**High: 5 – 7**
Proportion	1.00	.23	.17	.60
Outcomes				
Depressive symptoms (0 – 3)	0.63 (0.47)	0.78[bc] (0.53)	0.65[ac] (0.45)	0.58[ab] (0.44)
Substance use (0/1)	.63	.75[bc]	.71[ac]	.57[ab]
Delinquency (0 – 14)	1.77 (2.35)	2.14[bc] (2.56)	1.93[ac] (2.44)	1.58[ab] (2.21)
Family relationships				
Global family relationship quality (1 – 5)	3.76 (0.82)	3.44[bc] (0.90)	3.67[ac] (0.77)	3.91[ab] (0.77)
Parent–child relationship quality (1 – 5)	4.51 (0.57)	4.28[bc] (0.74)	4.47[ac] (0.54)	4.61[ab] (0.48)
Activities with a parent (0 – 5)	1.74 (1.14)	1.24[bc] (1.02)	1.63[ac] (1.03)	1.96[ab] (1.15)
Arguments with a parent (0/1)	.38	.40[c]	.40[c]	.36[ab]
Parental control (0 – 7)	1.87 (1.58)	1.47[bc] (1.54)	1.65[ac] (1.51)	2.08[ab] (1.58)
Family structure, income, work				
Two biological parents	.56	.43[bc]	.54[ac]	.62[ab]
Stepparent	.16	.19[c]	.18[c]	.15[ab]
Single parent	.23	.32[bc]	.25[ac]	.20[ab]
Other family structure	.04	.06[bc]	.04[a]	.04[a]
Family income (in thousands: 0 – 999)	45.36 (52.48)	42.11[bc] (53.45)	46.94[a] (53.35)	46.12[a] (52.35)
Mother employed full time	.58	.64[c]	.62[c]	.55[ab]
Mother employed part time	.19	.15[bc]	.18[a]	.20[a]
Mother not currently employed	.23	.21[c]	.20[c]	.25[ab]

Note: Descriptives are weighted and design adjusted using *svy* commands in Stata 12.0. Standard deviations are in parentheses.

[a]Significantly different from low family dinners ($p < .05$).

[b]Significantly different from medium family dinners ($p < .05$).

[c]Significantly different from high family dinners ($p < .05$).

Table 2

Regression Models of Adolescent Depressive Symptoms, Substance Use, and Delinquency at T1 (N = 17,977)

	Depressive Symptoms OLS Coefficients			Substance Use Logistic ORs			Delinquency Negative Binomial IRRs		
	M1	M2	M3	M1	M2	M3	M1	M2	M3
Family dinners	−0.04***	−0.03***	−0.01***	0.85***	0.88***	0.92***	0.94***	0.94***	0.97***
Two bio parent (ref.)									
Stepparent		0.06***	0.03*		1.70***	1.53***		1.22***	1.12**
Single parent		0.06***	0.03*		1.64***	1.52***		1.32***	1.24***
Other family structure		0.11***	0.10***		1.60***	1.53***		1.36***	1.33***
Number children in hh		0.01**	0.01*		0.95**	0.95**		1.01	1.00
Mother employed FT (ref.)[a]									
Mother not employed		−0.01	−0.01		0.81**	0.81**		0.94	0.94
Mother employed PT[a]		−0.01	0.00		0.85*	0.85*		0.97	0.98
Global family relationship			−0.15***			0.69***			0.72***
Parent–child relationship			−0.06***			0.90†			0.92**
Activities with parent			−0.01			0.86***			0.97*
Arguments with parent			0.12***			1.85***			1.52***
Parental control			0.02***			0.91***			0.97**
Constant	0.80***	0.45***	1.25***	3.72***	0.15***	2.35†	2.34***	6.78***	40.04***
ln alpha							0.33	0.19	0.00
Alpha							1.39	1.21	1.00
M2-M1 % Δ fam din coef.	−19.02***			−24.10***			8.09		
M3-M1 % Δ fam din coef.	−61.08***			−55.51***			−48.73***		

Note: Analyses are weighted and design adjusted using *svy* commands in Stata 12.0. OLS coefficients are not standardized. Controls for gender, age, race and ethnicity, family income, and parents' education are included but not shown. The --khb—command estimates and tests statistical significance of differences in family dinners coefficients across nested models, net of rescaling (Karlson et al., 2011; Kohler & Karlson, 2011). Caution is warranted in interpreting change in family dinners coefficients across nested negative binomial models as the test has not been formally derived for these models (Karlson, June 2011 e-mail correspondence).

[a] FT denotes full-time employment, PT denotes part-time employment.

† $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$. (Two-tailed tests.)

Table 3

First Difference Models of Depressive Symptoms, Substance Use, and Delinquency T2 - T1 (N = 12,446)

	T2 - T1		*T2 - T1*		*T2 - T1*	
	Depressive Symptoms		**Substance Use**		**Delinquency**	
	M1	**M2**	**M1**	**M2**	**M1**	**M2**
Change in family dinners	−0.009***	−0.005*	−0.004*	−0.003	−0.003	0.002
Change in family structure		−0.012		−0.003		−0.160*
Change in number of children		0.004		−0.013		0.022
Mother starts full-time employment		0.001		0.049**		0.076
Mother quits full-time employment		−0.030		0.010		0.049
Change in global family relationship		−0.103***		−0.026*		−0.254***
Change in parent–child relationship		−0.073***		−0.017		−0.154**
Change in activities with parent		0.002		0.005		0.074**
Change in arguments with parent		0.040***		0.025*		0.228***
Change in parental control		−0.004		−0.004		−0.034†
Constant	0.101***	0.095***	−0.054***	−0.064***	−0.432***	−0.445***

Note: Analyses are weighted and design adjusted using *svy* commands in Stata 12.0. OLS coefficients are not standardized.

† $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$. (Two-tailed tests.)

Table 4

Regression Models of Young Adult Depressive Symptoms, Substance Use, and Delinquency at T3 (N = 13,481)

	Depressive Symptoms OLS Coefficients		Substance Use Logistic ORs		Delinquency Negative Binomial IRRs	
	M1	**M2**	**M1**	**M2**	**M1**	**M2**
Family dinners	−0.01*	−0.003	0.98†	0.99	0.99	1.00
Two bio parent (ref.)						
Stepparent	0.01	0.003	1.19*	1.08	1.03	1.04
Single parent	0.05**	0.04*	1.12	1.03	1.23**	1.16†
Other family structure	0.08*	0.05†	1.03	0.90	1.13	1.08
Number of children in hh	0.00	0.00	0.95†	0.97	1.01	1.01
Mother employed FT (ref.)[a]						
Mother not employed	0.04**	0.04**	0.87*	0.91	1.02	1.04
Mother employed PT[a]	0.03*	0.04*	0.94	0.97	1.03	1.05
Global family relationship	−0.07***	−0.03***	0.84***	0.91*	0.87***	0.96
Parent–child relationship	−0.03*	−0.01	1.02	1.04	0.90*	0.93
Activities with parent	0.00	0.003	0.91***	0.94*	1.01	1.02
Arguments with parent	0.07***	0.04***	1.48***	1.30***	1.27***	1.11*
Parental control	0.01***	0.01**	0.92***	0.94**	0.99	1.00
T1 measure of outcome		0.27***		3.17***		1.17***
Constant	0.99***	0.69***	36.40***	21.44***	28.48***	8.23***
ln alpha					0.78	0.64
alpha					2.17	1.90

Note: Analyses are weighted and design adjusted using *svy* commands in Stata 12.0. OLS coefficients are not standardized. Controls for gender, age, race and ethnicity, family income, and parents' education are included but not shown.

[a]FT denotes full-time work, PT denotes part-time work.

†$p < .10$.

*$p < .05$.

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** $p < .01$.

*** $p < .001$. (Two-tailed tests.)