Author Manuscript
Author Manuscript
Author Manuscript

Published in final edited form as:

Dev Psychol. 2014 February ; 50(2): 554–565. doi:10.1037/a0033647.

Two Approaches to Estimating the Effect of Parenting on the Development of Executive Function in Early Childhood

Clancy Blair,
New York University

C. Cybele Raver,
New York University

Daniel J. Berry, and
New York University

Family Life Project Investigators

Abstract

In the current article, we contrast 2 analytical approaches to estimate the relation of parenting to executive function development in a sample of 1,292 children assessed longitudinally between the ages of 36 and 60 months of age. Children were administered a newly developed and validated battery of 6 executive function tasks tapping inhibitory control, working memory, and attention shifting. Residualized change analysis indicated that higher quality parenting as indicated by higher scores on widely used measures of parenting at both earlier and later time points predicted more positive gain in executive function at 60 months. Latent change score models in which parenting and executive function over time were held to standards of longitudinal measurement invariance provided additional evidence of the association between change in parenting quality and change in executive function. In these models, cross-lagged paths indicated that in addition to parenting predicting change in executive function, executive function bidirectionally predicted change in parenting quality. Results were robust with the addition of covariates, including child sex, race, maternal education, and household income-to-need. Strengths and drawbacks of the 2 analytic approaches are discussed, and the findings are considered in light of emerging methodological innovations for testing the extent to which executive function is malleable and open to the influence of experience.

Keywords

poverty; parenting; executive function; childhood

In this article, we apply a longitudinal analysis to the development of executive function abilities in early childhood. We contrast residualized change and latent change score (LCS) approaches to estimating the strength of the relation of experience, as indicated by measures

Correspondence concerning this article should be addressed to Clancy Blair, Department of Applied Psychology, New York University, 246 Greene Street Kimball Hall, 8th Floor, New York, NY 10003. clancy.blair@nyu.edu.

Clancy Blair, C. Cybele Raver, and Daniel J. Berry, Department of Applied Psychology, New York University; Family Life Project Investigators. Daniel J. Berry is now at the Department of Educational Psychology, University of Illinois at Urbana-Champaign.

of parenting quality, to executive function development. *Executive functions* refer to domain general cognitive abilities important for reasoning and planning and problem solving. Although sometimes construed as a general supervisory system, executive functions in the cognitive and developmental literatures have come to be defined as specific working memory, inhibitory control, and attention or set-shifting cognitive abilities (Garon, Bryson, & Smith, 2008). Understanding the course of executive function development is important, given the relevance of these aspects of cognitive ability to early school achievement (Blair & Razza, 2007; McClelland et al., 2007) and to emotion regulation and social competence (Raver et al., 2011).

Pioneering efforts in the measurement of executive functioning in young children provide evidence of age-related variation in task performance using cross-sectional samples (Espy, Bull, Martin, & Stroup, 2006; Wiebe, Espy, & Charak, 2008). As well, two groundbreaking longitudinal studies have demonstrated relations of executive function in early childhood, beginning at age 2 years, to the development of theory-of-mind ability (Carlson, Mandell, & Williams, 2004; Hughes & Ensor, 2007). Only recently, however, have researchers examined predictors of change in executive function in early childhood using longitudinal data.

As with many areas of developmental science, a key question for executive function research is to estimate the relation of the environment to its development with as little bias as possible. Using residualized change models, a small number of recent studies indicate that specific aspects of parenting are related to executive function longitudinally (Bernier, Carlson, & Whipple, 2010; Blair et al., 2011; Hammond, Müller, Carpendale, Bibok, & Liebermann-Finestone, 2011). These studies indicated that relations between the quality of early parenting and later executive function were present over and above a measure of child executive function at an earlier time point.

Limitations in Prior Studies

Although these prior studies of executive function development suggest a central role for parenting, they are limited in specific ways. One such limitation concerns the absence of tests of the longitudinal measurement invariance of executive function measures. The absence of invariance testing leads to an unanswered question concerning the extent to which executive function itself is actually changing over time or whether unique constructs were measured at earlier and later time points. One obstacle to drawing inferences about influences on any aspect of development, including executive function, concerns the indication that the same construct is measured across time points. Currently, only two studies have demonstrated invariance of executive function measured longitudinally in early childhood. The first of these demonstrated partial strong invariance of executive function (metric and scalar invariance) in a sample of 191 children seen between 4 and 6 years of age using three tasks (Hughes, Ensor, Wilson, & Graham, 2009). The second—the measure and longitudinal sample that are the focus of this analysis—demonstrated partial strong invariance of a newly developed executive function task battery with 1,292 children seen at ages 36, 48, and 60 months (Willoughby, Blair, Wirth, Greenberg, and the Family Life Project Investigators, 2010, 2012; Willoughby, Wirth, & Blair, 2011). As of yet, no researchers have examined the relation of aspects of children's experience to change in executive function as measured by this battery. Accordingly, a key focus of this article was to anchor our analysis of predictors of change in executive function over time with an assessment that has demonstrated longitudinal measurement invariance.

A second specific limitation in prior studies concerns the absence of a later measure of parenting. In the absence of later measures of parenting, studies have been unable to address whether change in parenting, again adhering to strictures of measurement invariance, was associated with change in child executive function abilities. As well, there are multiple dimensions of parenting that examinations of measurement invariance can help to differentiate. Furthermore, the absence of a later measure of parenting precludes the opportunity to examine possible bidirectional relations between parenting and children's executive function abilities. Indeed, although bidirectional or "cascade" theoretical models of development are pervasive (Ford & Lerner, 1992; Gottlieb, 1991), such models are considered less often empirically (Blair & Raver, 2012; Masten & Cicchetti, 2010). It is logical that higher or lower levels of executive function in children pose specific demands on parenting behavior that will account for unique variance in parenting behavior over and above an earlier measure of the same parenting construct.

Expectations for Change

Relatively few developmental studies have attempted to closely link *changes* in children's typical experience to *changes* in child ability. This may be because past research has tended to be conducted with an assumption of stability in children's normative environments: Recent analysis of data from the National Institute of Child Health and Human Development (NICHD) Study of Early Childcare, however, indicated an association between change in the quality of the home environment and change in child language development from age 36–54 months (Son & Morrison, 2010). Importantly, however, the meaning of change in the home environment over time in that study is indeterminate given the absence of a test of longitudinal measurement invariance of the Home Observation for Measurement of the Environment (HOME) scale (Caldwell & Bradley, 1984) between 36 and 54 months.

A central point for consideration in the analysis of change in children's typical experience as a predictor of child outcomes concerns the theoretical expectation for that change. Relatively few studies have examined normative or typical changes in parenting behavior. A number of studies, however, highlight the ways that levels of sensitivity, harshness, responsiveness, and provisioning can change over time, both as a function of children's temperament and as a function of developmental epoch, or stage. For example, research on parenting in early childhood has identified the ways that emotionally negative and difficult temperament during infancy and toddlerhood elicits increased harshness and/or detachment on the part of parents (Patterson, Reid, & Dishion, 1992; Simons, Chao, Conger, & Elder, 2001). In addition, recent work on child development in the context of poverty has underscored ways that parenting is malleable in the context of economic pressure and poverty-related risks such as higher community violence. Past work has emphasized ways that the quality of parenting is eroded by losses of income, increases in community violence, and higher involvement in low-wage, menial work (Raver, 2003). Fortunately, some evidence also suggests improvement in parenting when family fortunes improve: With moves to safer, higher income neighborhoods, parenting practices among low-income families have experimentally been found to improve (Levanthal & Brooks-Gunn, 2000), and with increases in income, families have been found to increase the level of cognitive stimulation in the home environment (Votruba-Drzal, 2003).

Analytic Methods

Given relatively few experimental studies examining parenting (Landry, Smith, & Swank, 2006) and executive function development (Bierman, Nix, Greenberg, Blair, & Domitrovich, 2008; Raver et al., 2011) and inherent limits on experimental methodology (Bloom, 2005; Ludwig, Kling, & Mullainathan, 2011), alternative approaches are needed that leverage naturally occurring variation to derive causal inference from nonexperimental data. Several such approaches are available, including propensity score matching, fixed effects, instrumental variables, and autolagged path models, as well as residualized change models (Berger, Bruch, Johnson, James, & Rubin, 2009; NICHD Early Child Care Research Network [ECCRN] & Duncan, 2003; Votruba-Drzal, 2003). These models offer the possibility of “controlling” for time-invariant and unobserved confounds and, in most instances, reduce the likelihood that continuity in children’s own functioning may account in part or in whole for any observed association between higher quality parenting and higher child executive function. For instance, when longitudinal data are available, adjustment of the dependent variable by that variable measured at an earlier time point helps to reduce selection bias and unobserved variables bias (Morgan & Winship, 2007; NICHD ECCRN & Duncan, 2003). Notably, however, this residualized change model makes the potentially incorrect assumption that random variation in the earlier measure (executive function at the earlier time point) is unrelated to random variation in the dependent measure (later executive functioning residualized on earlier executive functioning). Despite this potential drawback, residualized change models allow stronger inference about effects of an independent variable, in this case parenting, on a dependent variable (executive functioning) when selection into levels of parenting quality are not likely explained by time-invariant characteristics of the child (Allison, 1990).

Given this strength, however, a potential problem with residualized models is the possibility of not actually modeling change. Conclusions about change in residualized change models using observed variables are dependent on the assumption that “change” in the observed measure over time, in fact, represents change in the construct of interest, rather than changes in the measurement of the construct over time. A key threat to causal inference in nonexperimental data analysis is the possibility that measures of constructs assessed longitudinally may share only conceptual similarity, or what is referred to as *heterotypic continuity*, over time. In the absence of invariance testing, as well as attention to the reliability and validity of measures at each time point, inference about change may be limited. That is, instead of coming to conclusions about change in a demonstrably similar construct over time, findings might indicate the effect of a distinct but similarly named construct at the later time point over and above an earlier construct. Developmental change might be confounded with measurement change. Finally, residualized change models (fitted in the context of ordinary least squares [OLS] regression) are able to consider only a single dependent variable at a time. As noted above, theory suggests the potential role of dynamic relations between children's executive functioning and parenting quality over time. These relations are typically missed by traditional residualized change approaches.

In the current article, we contrast two approaches to estimating the effect of parenting on executive function in young children, residualized change analysis (Allison, 1990) and latent change analysis (McArdle, 2009), in order to better understand ways in which to model influences on child development both outside and inside the home that themselves may be undergoing change over relatively short periods of time. We first use residualized change models to test the respective roles of both early and later measures of parenting quality measured in two different ways on change in executive function from 36 to 60 months. We use residualized change models as they most clearly offer the flexibility of allowing for the inclusion of stage-salient measures of the quality of parenting that are not statistically identical at both time points. Here, measures of parenting are conceptually rather than statistically identical. As such, our assumptions about change are limited. Also, we cannot simultaneously examine bidirectional relations between parenting and executive function in these models. Accordingly, we then carry out a second set of analyses using LCS models. The LCS models allow us to statistically equate our measures over time and allow us to explicitly model change in both executive function and in parenting quality. In addition, these models allow for the examination of bidirectional relations between parenting and executive function and thereby to come to stronger conclusions about the potential dynamics of parenting and children's executive functioning abilities over time. Notably, in doing so, however, we restrict our analyses to only those components of parenting quality and executive function that are statistically identical at earlier and later time points and for which we can adjust for partial measurement invariance across time.

Method

Participants

The Family Life Project (FLP) was designed to study young children and their families in two of the four major geographical areas of the United States with high poverty rates. Specifically, three counties in Eastern North Carolina and three counties in Central Pennsylvania were selected to be indicative of the Black South and Appalachia, respectively. The FLP adopted a developmental epidemiological design in which sampling procedures were used to recruit a representative sample of 1,292 children whose families resided in one of the six counties at the time of the child's birth. Low-income families in both states and African American families in North Carolina were oversampled (African American families were not oversampled in Pennsylvania because the target communities were at least 95% non-African American). FLP recruiters identified 5,471 (59% North Carolina, 41% Pennsylvania) women who gave birth to a child in the 12-month period. A total of 1,515 (28%) of all identified families were determined to be ineligible for participation for three primary reasons: not speaking English as the primary language in the home, residence in a nontarget county, and intent to move within 3 years. Of the 2,691 eligible families who agreed to the randomization process, 1,571 (58%) families were selected to participate using the sampling fractions that were continually updated from our data center. Of those families selected to participate in the study, 1,292 (82%) families completed a home visit at 2 months of child age, at which point they were formally enrolled in the study. Interested readers are referred to other articles summarizing study recruitment strategies and detailed descriptions of participating families and their communities (Burchinal, Vernon-Feagans, Cox, and the Family Life Project Investigators, 2008).

Procedure

Families participating in the study were seen in home visits at approximately annual intervals beginning at child age 7 months and continuing through the second grade. The data for the analyses presented here were collected in the home by highly trained research assistants at child ages of approximately 36 and 60 months. During the visits at each time point, the primary caregiver, in 99% of cases the mother, answered questions about demographics and income and participated in a structured interaction with the child. At both time points, children were administered executive function tasks as well as other child assessments.

Measures

Executive function—At home visits at child ages of 36 and 60 months, children were administered a newly developed and validated battery of six executive function tasks tapping inhibitory control, working memory, and attention shifting. Details on the tasks and administration procedures as well as psychometric characteristics are available in Willoughby et al. (2011) and Willoughby et al. (2010). For each of the six tasks, children were administered training trials and up to three practice trials, as needed. As is standard for executive function measures with children (Zelazo & Müller, 2002), children were required to successfully complete pretest trials in which they clearly demonstrated knowledge of the rules for the task and the ability to successfully complete the pretest trials as instructed. Children were required to complete 75% of test trials in a given task in order to receive a score for that task.

On the basis of prior validation of the executive function construct with these data (e.g., Willoughby, Blair, et al., 2012), we adopted a common factor representation of children's executive function abilities at 36 and 60 months. For OLS analyses including observed scores, we created a variable indicating average percent correct responding across the six tasks at the two time points. For covariance structure analyses, including latent variables, we created a common scale over time derived from longitudinal item response theory (IRT) models (Willoughby, Wirth, Blair, & the Family Life Project Investigators, 2012).

Parenting quality—Parenting quality at both time points was operationalized through two different measures; parent responsiveness and cognitive stimulation measured through the HOME scale (Caldwell & Bradley, 1984) and parenting sensitivity and stimulation measured through observer ratings during child age-appropriate structured parent–child interaction (PCX) tasks (Cox, Paley, Burchinal, & Payne, 1999; see also NICHD ECCRN, 1999). Items from the HOME scale at each time point pertaining to responsiveness and stimulation were identified and averaged to create overall composites of parenting quality. The composite at child age 36 months included 18 items ($\alpha = .79$) from the Infant/Toddler version of the scale. The composite at child age 60 months included 25 items (α = .71) from the Early Childhood version of the scale. Notably, only a total of six items pertaining to parent responsiveness and four items pertaining to cognitive stimulation were common across the two versions of the HOME scale. The majority of items on the two versions of the scale do not overlap.

Observer ratings of parenting quality from the structured parent–child interactions were coded on a 1–7 scale (Cox et al., 1999; NICHD ECCRN, 1999). At child age 36 months, positive parenting was defined by observer ratings of sensitivity, detachment (reverse scored), positive regard, and stimulation. At child age 60 months, positive parenting was defined by sensitivity, detachment (reverse scored), positive regard, support for autonomy, stimulation, and negative regard (reverse scored). These composites were derived from factor analytic models conducted by the FLP data analysis team. These factors at each time point represent age and stage-salient indicators of high-quality parenting. Interrater reliabilities were acceptable; interclass correlations ranged from .91 to .95 at 36 months and from .90 to .94 at 60 months.

Demographic covariates included maternal years of education, child race/ethnicity (as reported by mothers), child sex, and income-to-need ratio (estimated from mothers' report of total family income, divided by the federal poverty threshold, and adjusted for number of persons in the home).

Data Analysis

To test hypotheses concerning the relation of parenting quality to the development of executive function in young children, we first proceeded by regressing executive function at child age 60 months on executive function measured at child age 36 months and included composite measures of parenting quality, as measured by the HOME scale and by observer ratings, measured at 36 and 60 months, in separate equations. The intention of this analysis was to examine the prediction of residualized change in executive function from age 36 to 60 months by the measures of parenting at the two time points. We then examined changes in coefficients occurring with the inclusion of covariates in the model as a further check on possible omitted variable bias. We next established invariance in the measures of parenting quality and executive function and applied LCS analysis to estimate mean-level changes in parenting quality and children's executive function abilities between 36 and 60 months. In these models, we tested bidirectional relations between parenting quality and executive function. All models were fitted using the robust maximum likelihood estimator available in Mplus (Muthén & Muthén, 1998–2011), with numerical integration for the models including HOME items, which are dichotomous. Missing data in all analyses were addressed using full information maximum likelihood estimation.

Results

Means, standard deviations, and number of participants with values for each of the variables in the analysis are presented in Table 1. Correlation across time points for the HOME was $r = .46$, $p < .0001$, and for observed parenting was $r = .62$, $p < .0001$. Correlation between the HOME scale and observed parenting was $r = .42$, $p < .0001$, at 36 months and $r = .46$, $p <$.

0001, at 60 months. On average, there was a small but significant decrease between age 36 and 60 months for the HOME scale, paired $t = -9.34(980)$, $p < .0001$, and a small but significant increase for the observational measure of parenting, paired $t = 4.99(895)$, $p < .0001$. Executive function was correlated $r = .33$, $p < .0001$, between time points and increased significantly between 36 and 60 months, paired $t = 37.89(964)$, $p < .0001$.

Residualized Change Models

For our first set of analyses, we regressed executive function at 60 months on the composites from the HOME scale at 36 and 60 months of age. This model included executive function at 36 months as an additional predictor so that the regression coefficients for measures of parenting could, as is typically done in developmental research, be interpreted as predicting residualized change, or gain in children's executive function from 36 to 60 months. Results using full information maximum likelihood to address missing data indicated that executive function showed moderate rank-order stability over time ($B_{EF36} = .20$, $p < .001$, $\beta = .31$). Moreover, higher quality parenting at both earlier and later time points predicted greater residualized change, that is, more positive gain in executive function at 60 months ($B_{HOME36} = .13$, $p < .001$, $\beta = .17$; and $B_{HOME60} = .11$, $p < .01$, $\beta = .09$, respectively). Findings were robust with the addition of covariates, with effects for child race and sex but not for maternal education or household income-to-need.

We then predicted executive function at 60 months in a separate equation in which we included measures of observed parenting at 36 and 60 months instead of the HOME responsiveness and stimulation scale. Similar to the findings for parenting as assessed by the HOME scale, observed maternal sensitivity at 36 ($B_{PCX36} = .02$, $p < .001$, $\beta = .16$) and 60 months ($B_{PCX60} = .014$, $p < .01$, $\beta = .12$) was positively associated with greater residualized change in child executive function, over and above the effect for executive function at 36 months ($B_{EF36} = .17$, $p < .001$, $\beta = .26$). Findings in this model were largely robust with the addition of covariates, with an effect for child race and sex. With the addition of the covariates, however, the effect for parenting at 60 months became nonsignificant ($B_{PCX60} = .008$, $p = .12$, $\beta = .06$), indicating no unique effect for this variable over and above the effect of its counterpart at 36 months.

As described above, residualized change models provide a relatively conservative specification of the relation of aspects of parenting quality to executive function in early childhood. The models indicate unique and positive effects of parenting both at earlier and later time points. The effects can be interpreted as the independent contribution of unique aspects of parenting at both time points as well as the effect of change in some common aspect of parenting shared by the two measures. Given the moderate level of correlation between the measures of parenting at each time point, we created a *z*-score summary variable and reran the analysis. As expected, the composite was highly stable between time points ($r = .65$, $p < .0001$), and results indicated that the combined variable made unique contributions at both earlier and later time points to executive function at 60 months over and above executive function at 36 months ($B_{PCz36} = .012$, $p < .001$, $\beta = .16$) and 60 months ($B_{PCz60} = .011$, $p < .01$, $\beta = .14$). As noted above, however, the models are indeterminate as to the extent to which change in parenting is being estimated and can be meaningfully interpreted as predicting change in executive function.

Results From LCS Analyses

We first examined longitudinal measurement invariance of executive function and of the measures of parenting quality. On the basis of prior validation of the executive function construct with these data, we adopted a common factor representation of children's executive function abilities at 36 and 60 months using IRT-adjusted scores for each of the tasks. A first set of measurement models indicated that respective within-time confirmatory factor analysis (CFA) models at 36, $\chi^2 = .89(5)$, $p = .97$, comparative fit index (CFI) = 1.00, root-mean-square error of approximation (RMSEA) = .00, and 60 months, $\chi^2 = 4.71(9)$, $p = .89$, CFI = 1.00, RMSEA = .00, fit the data well. As expected, the longitudinal model demonstrated metric invariance and partial scalar invariance, $\chi^2 = 63.36(43)$, $p = .02$, CFI = .97, RMSEA = .02. All of the factor loadings were statistically significant, with the standardized loadings ranging between .26 and .64. The variances for the latent factors were statistically significant at 36 ($\phi = .08$, $p < .001$) and 60 ($\phi = .12$, $p < .001$) months of age. As expected, there was notable rank-order stability in children's executive function abilities over this period ($\phi_{stand} = .79$).

CFA of the six common items on the HOME scale indicated a single factor tapping parenting responsiveness at 36 and 60 months. Independently, the within-time CFA models fit the data well at both 36, $\chi^2 = 11.41(5)$, $p = .04$, CFI = 1.00, RMSEA = .04, and 60 months of age, $\chi^2 = 5.0(6)$, $p = .55$, CFI = 1.00, RMSEA = .00. Tests of longitudinal measurement invariance demonstrated metric invariance as well as partial scalar invariance. The partially invariant longitudinal CFA fit the data reasonably well, $\chi^2 = 54.05(18)$, $p < .001$, CFI = .98, RMSEA = .04. All factor loadings were statistically significant and in the expected direction. The latent variances were statistically significant at 36 ($\phi = .89$, $p < .001$) and 60 ($\phi = .74$, $p < .001$) months. As expected, there was moderate rank-order stability in parent responsiveness ($\phi_{stand} = .35$, $p < .001$). A CFA including the four common items tapping cognitive stimulation on the HOME scale at 36 and 60 months with the parenting responsiveness items significantly degraded model fit. Examination of these items in a separate CFA also did not produce a well-fitting model, suggesting that the items may tap distinct aspects of cognitive stimulation at the two time points. Therefore, no further analysis of the longitudinal relation of this aspect of parenting quality to child executive function was pursued with these data.

The CFA model of observed parenting sensitivity fit the data well at 36 months, $\chi^2 = 10.86(4)$, $p = .03$, CFI = 1.00, RMSEA = .04, and reasonably well at 60 months, $\chi^2 = 83.41(4)$, $p < .001$, CFI = .96, RMSEA = .14. On the basis of preliminary models, we relaxed the constraint that the within-time residual covariance between the sensitivity and negative regard items was zero in these models. This single latent factor defined by observer ratings of parenting at 36 and 60 months demonstrated partial metric and scalar invariance. The partially invariant longitudinal CFA model fit the data reasonably well, $\chi^2 = 259.28(31)$, $p < .001$, CFI = .95, RMSEA = .08. Each of the factor loadings was statistically significant and in the expected direction, with standardized loadings ranging between −.39 and .97.

There was statistically significant variability in the maternal sensitivity factors at both 36 months ($\phi = 1.20$, $p < .001$) and 60 months ($\phi = 1.92$, $p < .001$), and moderate to strong rank-order stability in caregivers' sensitivity levels over time ($\phi_{stand} = .67$, $p < .001$).

Subsequent CFA models indicated that the HOME scale items and observed parenting measures were best represented as two distinct measures, rather than as a common-factor model, in which variation across the HOME responsiveness and observed parenting sensitivity items are explained by a single latent factor. When modeled as distinct factors, the two latent factors were only modestly correlated at 36 and 60 months ($\phi_{stand} = .32$–$.34$, respectively). Furthermore, tests of common-factor models at each point in time (item residuals adjusted for methods variance) indicated that the common factor explained very little variation in several of the observed parenting items, with R^2 estimates ranging from .01 to .09 at 36 months to .02 to .07 at 60 months. Collectively, these findings indicate that, although the HOME scale and the observed parenting measure each tap aspects of parenting, they are best modeled as distinct latent factors.

Mean-level changes in parenting quality and children's executive function abilities—The LCS model for executive function fit the data well, $\chi^2 = 63.18(-43)$, $p = .02$, CFI = .97, RMSEA = .02. As expected, the model indicated that, on average, children showed substantial, statistically significant gains in executive function between 36 and 60 months ($\Delta EF_{\alpha} = .82$, $p < .001$). Using the estimated standard deviation from the executive function factor at 36 months as the scale, children tended to show an approximate 2.83 standard deviation increase in latent executive function. There was statistically significant between-child variation in the amount of executive function change between 36 and 60 months ($\psi = 0.05$, $p = .001$), and executive function at age 36 months was unassociated with the rate of executive function change ($\psi = -.01$, $p = .66$).

The LCS model for the common items on the HOME scale indicated *decreases* rather than increases in the latent variable indicating caregiver responsiveness ($\Delta Resp_{\alpha} = -3.19$, $p < .001$). Scaled on the estimated standard deviation of caregiver responsiveness at 36 months, this corresponded to an approximate −.66 standard deviation decrease. Similarly, the LCS model of observer ratings of parenting sensitivity during the semistructured caregiver–child interaction task, $\chi^2 = 259.27(31)$, $p < .001$, CFI = .95, RMSEA = .08, also fit the data well and showed a similar decline ($\Delta Sens_{\alpha} = -.71$, $p < .001$), corresponding to a −.65 standard deviation decrease between child age 36 and 60 months. These declines would appear to represent normatively expected change in types of parenting behavior (e.g., touching, hugging, and other forms of physical comfort) included in the longitudinally invariant parenting latent constructs and that become comparatively less frequent as children transition from the toddler to the preschool periods.

Across the LCS models for both measures of parenting, there was substantial between-person variation in the extent to which parenting quality changed over time; the variances for the latent change factors were 26.33 ($p < .001$) and 1.10 ($p < .001$), respectively. The respective covariance estimates between parenting quality at child age 36 months and changes in parenting quality were also statistically significant and negative in both LCS models for the HOME scale ($\psi = -19.10$, $p < .001$, $\psi_{stand} = -.77$) and for observer ratings (ψ

= −0.19, p = .001, ψ_{stand} = −.17), indicating that caregivers with lower scores at the first time point, 36 months, tended to show less decline in the respective parenting indicator over time.

Individual Differences in Parenting Quality and Executive Function Development

The bivariate model estimating simultaneous LCS models for parenting as measured by the common items on the HOME scale and child executive function indicated that, adjusting for child executive function at 36 months, less decline in parenting behavior was associated with greater positive gain in child executive function (ψ = .15, p = .03). This corresponded to a standardized residual covariance (i.e., partial correlation) of approximately .22.

A similar relation was evident for observed parenting sensitivity. The bivariate model fit the data reasonably well, χ^2 = 633.72(182), p < .001, CFI = .92, RMSEA = .05. As above, there was a statistically significant covariance between changes in observed parenting sensitivity and changes in executive function (ψ = 0.05, p < .01). Again, adjusting for child executive function at 36 months, less decline in observed parenting sensitivity between 36 and 60 months was associated with greater positive gain in executive function. This corresponded to a standardized residual covariance of approximately .21.

Notably, these preliminary models do not account for potential longitudinal bidirectional relations between parenting and executive function. It may be that higher or lower executive function at child age 36 months elicits changes in parenting at child age 60 months. To test such bidirectionality, we added cross-lagged paths to each of the bivariate LCS models. The inclusion of these cross-lagged parameters improved model fit in the model of parenting assessed by the HOME scale (S-B Δ −2LL = 16.22, Δdf = 2, p < .001). On average, parenting as assessed by the HOME scale at 36 months was associated with greater positive change in executive function between 36 and 60 months (B = .02, p < .01). This corresponded to a standardized regression coefficient of approximately .35. Substantively, a one standard deviation difference in caregiver responsiveness at 36 months is associated with a .35 standard deviation difference in executive function change between 36 and 60 months. Collectively, executive function and parenting at 36 months accounted for approximately 12% of the between-child variation in executive function change rates, with virtually all (99%) of this total explained variation accounted for by variation in early parent responsiveness.

There was, however, also an indication that child executive function at 36 months was associated with a lesser decline in parenting responsiveness as measured by the HOME scale between 36 and 60 months (B = 1.77, p = .04). A one standard deviation difference in executive function at 36 months was associated with a .11 standard deviation difference in parenting change between 36 and 60 months. In combination, child executive function at 36 months and parenting at 36 months accounted for approximately 59% of the between-subjects variation in change in parenting; however, only approximately 1% of this total explained variation was accounted for child executive function at 36 months.

Notably, with the addition of the cross-lagged paths, the residual covariance between the latent change factors was no longer statistically significant. This indicates that the association between change in parenting responsiveness and change in executive function is attributable to bidirectional relations among the variables as well as possible unobservables associated with parenting and executive function. That is, the analysis indicates that residual covariance between change in parenting and change in executive function was largely explained by the cross-lagged relations.

Findings similar to those described above were evident for parenting sensitivity as measured by observer ratings. In this instance, however, cross-lagged paths were not significant in both directions. The addition of the cross-lagged paths did improve model fit ($\Delta S\text{-}B\chi^2$ = 14.60, $\Delta df = 2$, $p < .001$); however, this improvement was driven by the path from child executive function at 36 months to change in observed parenting ($B = 1.08$, $p = .001$, $\beta = .31$). There was no relation between parenting as measured by observer ratings and change in children's executive function abilities ($B = -0.01$, $p = .711$). Similar to the models for parenting as measured by the HOME scale, higher child executive function at 36 months was associated with a less pronounced decline in observed parenting. In combination, observed parenting at 36 months and child executive function at 36 months accounted for approximately 10% of the between-subjects variation in change in parenting. Approximately, 80% of this total variation was accounted for by child executive function at 36 months. Again, with the addition of the cross-lagged paths, the residual covariance between the latent change factors was no longer statistically significant.

Finally, in a last series of models, we included household income-to-need at child age 36 months as a predictor of changes in parenting and executive function along with a set of control covariates. As presented in Table 2 and displayed in Figure 1, family income-to-need and several of the control covariates were associated with both change in children's executive function abilities and changes in parent responsiveness, as indicated by the HOME scale. Income-to-need was positively associated with change in parenting responsiveness as measured by the HOME scale ($B = 0.30$, $p = .02$), as well as change in executive function ($B = 0.03$, $p = .03$). Given the conservative nature of the model, these relations corresponded to rather notable standardized effects, .06 and .12, respectively. The findings indicate that a one standard deviation difference in family income-to-need at child age 36 months is associated with an approximate .07 standard deviation change in parenting and .13 standard deviation difference in executive function change from 36 to 60 months. Similarly, on average, maternal education was positively related to gain in executive function ($B = 0.04$, $p < .001$, $\beta = .24$) and to less negative decline in parental responsiveness ($B = 0.45$, $p < .001$, $\beta = .15$). African American children tended to show less gain in executive function ($B = -0.08$, $p = .05$, $\beta = -.26$) and more negative decline in parental responsiveness ($B = -1.18$, $p = .002$, $\beta = -.20$). Boys tended to show comparatively less gain in executive function than did girls ($B = -0.14$, $p < .001$, $\beta = -.47$); sex was unassociated with changes in parenting. Notably, the cross-lagged relations of interest were largely robust to the inclusion of the control covariates. The positive relation between parenting responsiveness and change in executive function was attenuated slightly ($B = .01$, $p = .02$, $\beta = .19$) but remained significant. The positive relation between executive function and change in parenting responsiveness as measured by the HOME scale was also attenuated slightly and dropped to marginal levels of statistical significance ($B = 1.54$, $p = .08$, $\beta = .07$).

As shown in Figure 2 and reported in Table 3, household income-to-need was similarly associated with change in parenting sensitivity as measured by observer ratings. Higher income-to-need at child age 36 months was associated with a less pronounced decline in parenting sensitivity ($B = 0.13$, $p < .001$, $\beta = .13$). Furthermore, the estimated relations between the control covariates were quite similar to those seen above for parental responsiveness. Higher levels of maternal education were associated with less negative decline in parenting sensitivity ($B = 0.15$, $p < .001$, $\beta = .26$), and, on average, African American children tended to have parents who showed more negative declines in sensitivity ($B = -0.52$, $p < .001$, $\beta = -.44$). Sex was unassociated with changes in parenting sensitivity. Most notably, the cross-lagged association of child executive function at 36 months with change in parenting sensitivity was partly attenuated but remained statistically significant ($B = 0.61$, $p = .02$, $\beta = .15$).

Discussion

This analysis examined the development of executive function abilities between age 36 and 60 months in a large prospective longitudinal sample of children and families in predominantly low-income and rural communities in the United States. Using a newly developed executive function task battery with established psychometric properties, we examined the relation of two theoretically informed aspects of the quality of parental care, sensitivity and responsiveness, to gains in executive function. Building on prior studies suggesting that parenting plays a central role in the development of executive function (Bernier et al., 2010; Blair et al., 2011; Hammond et al., 2011), we pursued two analytical approaches to test ways that parenting quality affects the development of executive function across the preschool period. Using residualized change models, we found that higher parenting sensitivity and responsiveness at 36 months (and for responsiveness, at child age 60 months) was associated with higher child executive function at age 60 months *over and above* executive function at child age 36 months. Following the logic of residualized change models, one can interpret these results as indicating that parenting sensitivity and responsiveness are associated with later child executive functioning, conditioned on an earlier measure of child executive function. Although the residualized change approach has some advantages, adjusting for unobserved variables and selection into parenting quality that may be explained by early executive functioning (Morgan & Winship, 2007), limitations to residualized change models, including questions about developmental change in measures as well as potential bidirectional relations among variables, at a minimum, lead to questions about the interpretation of these effects.

To address issues relating to possible confounding of measurement change with developmental change as well as possible bidirectional relations among variables, we conducted LCS analyses in which all measures demonstrated partial longitudinal measurement invariance. Doing so for the HOME scale, however, necessitated that we restrict the number of items on the scale to six common items. Notably, we found that mean levels of the longitudinally invariant latent measures of parenting responsiveness and sensitivity decreased by approximately two thirds of a standard deviation between child ages 36 and 60 months. This decrease would appear to be a normative and expected decline in the types of behaviors being assessed, such as holding, caressing, and responding to the child verbally, for which parents may have fewer opportunities as children become more autonomous and self-directed. Although perhaps surprising, it is worth noting that this decline is not without precedent. Longitudinal analysis from the NICHD Study of Early Child Care and Youth Development—another large, longitudinal study of children in context—has shown similar normative declines in maternal sensitivity across this period, using a very similar measured of observed parenting behavior (Hirsh-Pasek & Burchinal, 2006).

Prior to adjusting for the cross-lagged relations, for both measures, lesser decline was associated with more positive gain in child executive function from age 36 to 60 months. Notably, the substantial declines in both of the invariant parenting latent measures are in contrast to small changes in the manifest variables in the residualized change analysis. We interpret these differences as illustrative of the distinction between the two approaches. In the residualized change approach, one might conclude that an increase in an underlying parenting construct is associated with change in executive function. From the LCS models, however, we see that such a conclusion is not necessarily correct. The LCS model indicates that what is common between 36 and 60 months in each of the measures of parenting is declining and that a lesser decline in this common construct across the two time points is associated with more positive change in child executive function. As such, the LCS and residualized change models would seem to be in agreement in that a higher level of parenting sensitivity or responsiveness at the later time point, whether a lesser decline in a common component or a higher level of primarily unique, but also some common component is associated with more positive change in executive function.

In addition to providing the basis for unambiguously examining the relation of change in parenting sensitivity and responsiveness to change in child executive function, the LCS models also provide the basis for examining the question of whether children's executive function are predictive of changes in parenting. Findings for both the HOME scale and observed parenting sensitivity indicated that higher executive function at age 36 months predicted lesser decline in the common aspect of parenting tapped by each measure between 36 and 60 months. Indeed, these findings suggest that higher levels of children's executive function abilities elicit higher quality care from caregivers in what may be one component of a genuinely transactional process in development. We found, however, that with the addition of the cross-lagged, bidirectional paths, the association between parenting and change in child executive function was statistically significant for one but not both measures of parenting. Specifically, the model containing parenting as measured by the invariant measure of responsiveness (the HOME scale) indicated bidirectionality; parenting at 36 months predicted change in executive function, whereas executive function at 36 months predicted change in parenting. This bidirectional relation did not hold for the observed measure of parenting sensitivity in which child executive function predicted lesser decline in parenting, but not the other way around.

Strengths, Limitations, and Conclusions

On one hand, it would seem logical that any analysis of change should require the use of demonstrably the same measure across two time points in order to meaningfully interpret any effect of change in the predictor on change in the outcome. That is, it is difficult to meaningfully interpret developmental change when it is confounded with changes in measurement. On the other hand, particularly early in development, measures of relevant constructs may be more similar in name than in content; displaying heterotypic continuity, or similarity in underlying meaning but not in indicators of that meaning. Here, the construct validity of a given measure of environmental quality such as parenting behavior may depend on the measure looking very different at a later developmental stage than it did at an earlier developmental stage. In particular, the HOME scale is composed of very different items at early and later developmental periods, with only six items in common between 36 and 60 months. Here, the statistical requirement of measurement equivalence as implemented in this analysis required that the measures lose some theoretical or conceptual adequacy from a developmental perspective in order to directly address questions about developmental change. As such, it is necessary that longitudinal analyses squarely address potential trade-offs between *conceptual adequacy* and *measurement adequacy* and strive to include measures that meet standards of measurement invariance.

Overall, findings from both the residualized change and latent change models provide generally strong grounds for inference about the relation of parenting behavior to executive function development in early childhood. These findings strengthen the empirical basis for earlier, more exploratory evidence that parenting, specifically maternal scaffolding, maternal sensitivity, and support for autonomy, plays an important role in predicting this key domain of development (Bernier et al., 2010; Hammond et al., 2011). The two approaches, however, are distinct in the inference that they provide about parenting and child development. The residualized change models have several advantages that help to strengthen inference about the relation of experience to executive function development. An advantage is that measures of constructs can be understood to vary with development. Indicators of age-appropriate parenting at child age 60 months are likely to be different from those at age 36 months. Such variation with development is manifestly the case for the HOME scale at child ages 36 and 60 months. Here, one might consider the conceptual adequacy of the measure to be high, but its measurement adequacy, that is, its ability to inform conclusions about change in the outcome from change in the predictor, to be low. Although the measurement adequacy of observed parenting sensitivity and executive function measures is perhaps higher than for the HOME scale, the absence of invariance over time for both parenting measures in the residualized change analysis limits any inferential strength that might be leveraged from the examination of change.

A clear benefit of the LCS approach when combined with the requirement of measurement invariance and examination of bidirectionality is the potential to provide strong causal inference from nonexperimental data. In part, expectations for causal inference from nonexperimental data (Cook, Shadish, & Wong, 2008) center on the potential for establishing equivalence, whether between groups or between measures over time (preferably both). Establishing equivalence between Time 1 and Time 2 measures of our constructs in the LCS models allows for a more direct, albeit imperfect, interpretation of relations between changes in parenting and changes in executive function. The LCS models demonstrated that change in child executive function is accounted for in part by parenting behavior. However, the change in executive function was attributable to parenting at an earlier time point rather being attributable to change in parenting between that earlier time point and a later time point. Covariance between change in executive functioning and change in parenting in these models was no longer significant with the addition of the cross-lagged paths.

Interestingly, although our LCS models allowed use to model true change in parenting and executive function over time, at least in its current form, such measurement approaches also have some substantive weaknesses. There are of course other aspects of parenting that are increasing (as well as decreasing) in the sample, but these are not captured in our LCS models, which were constrained to capture what is common in the respective measures over time. As such, we are unable to come to conclusions about aspects of parenting that may have been increasing or emerging over time as children are developing, as we did in the residualized change models. Notably, our intention here is not to make strong claims for which aspects of parenting are most relevant to child executive function development so much as it is to outline analytical choices that must be made when using repeated measures data to strengthen causal inference about relations between parenting and child development.

The foregoing points illustrate the way in which the essential strength of the LCS approach as implemented in this analysis may also be its primary constraint, if not limitation. The LCS analyses allows one to come to relatively strong conclusions about the way in which specific aspects of parenting are changing, but our interpretations are restricted to the limited aspects of parenting that were measured the same way, over time, which more often than not in developmental research is likely to be a subset of parenting behaviors that we are interested in. Here, the causal inference is necessarily restricted as some of the conceptual adequacy of the assessment of parenting is sacrificed to its measurement adequacy. To this extent, one criticism of any application of the approach, including this one, is that it may be analogous to looking for a lost object in a location where the light is best rather than in the location in which it was actually lost. To this end, it is important to interpret findings from the cross-lagged models with caution. When considering the “bigger picture” of testing theoretically and metrically strong developmental models, our LCS findings suggest that this type of model may be less satisfactory if different types of parenting processes and strategies might be hypothesized to come into play at 60 months than are not detectable at 36 months. This consideration of benefits and trade-offs also suggests that our residualized change models might be a better representation of development, allowing for time-specific and developmentally sensitive measurement of constructs.

The larger implications of both sets of our analyses pertain to the robust evidence they provide for the experiential shaping or canalizing of executive function development for children in poverty. Developmental theory emphasizes both the role of the early environment in shaping child outcomes and the role of child behavior in shaping the environment in which development is occurring. Through feed-forward and feedback processes, development is understood to unfold and to become increasingly stable over time. In this, the LCS analysis would appear to be providing a fundamental glimpse into the developmental process, whereby a developmental system propagates and sustains behavior.

Such continuity in development has been a central tenet of developmental theory, but it has rarely been tested directly. For children developing in the context of poverty, such as the sample included in this analysis, the policy relevance of the models would seem to be in their implications for programs focusing on fostering parenting behavior. The LCS models indicate the need for a developmentally appropriate approach to parenting behavior. These models recommend a focus on developmental change in parenting, in the present case, limiting expected declines in types of behaviors identified in the LCS models as common between child ages 36 and 60 months, while simultaneously increasing those that are emerging as children age, that is, those that are unique to age 36 or 60 months, as indicated in the residualized change analysis. Overall, our findings suggest that developmentally appropriate efforts to enhance or substantially improve the type of care that children receive are likely to alter trajectories of executive function development. Although additional work is needed to fully examine developmental change in parenting in early childhood, longitudinal analyses such as this one are an important step in developing and testing models of intervention to support children's cognitive function in the context of socioeconomic disadvantage. Evidence from randomized control trials of interventions targeting aspects of parenting that decline with age as well as those that increase with age offer the logical next step in establishing the role of parenting behavior in shaping executive function.

Acknowledgments

Support for this research was provided by National Institute of Child Health and Human Development Grants R01 HD51502 and P01 HD39667, with cofunding from the National Institute on Drug Abuse. The Family Life Project Investigators include Lynne Vernon-Feagans, Martha Cox, Clancy Blair, Peg Burchinal, Linda Burton, Keith Crnic, Ann Crouter, Patricia Garrett-Peters, Mark Greenberg, Maureen Ittig, Stephanie Lanza, Roger Mills-Koonce, Debra Skinner, Cynthia Stifter, and Michael Willoughby. We thank the many families and research assistants who made this study possible.

References

Allison PD. Change scores as dependent variables in regression analysis. Sociological Methodology. 1990; 20:93–114.10.2307/271083

Berger LM, Bruch SK, Johnson EI, James S, Rubin D. Estimating the "impact" of out-of-home placement on child well-being: Approaching the problem of selection bias. Child Development. 2009; 80:1856–1876.10.1111/j.1467-8624.2009.01372.x [PubMed: 19930356]

Bernier A, Carlson SM, Whipple N. From external regulation to self-regulation: Early parenting precursors of young children's executive functioning. Child Development. 2010; 81:326–339.10.1111/j.1467-8624.2009.01397.x [PubMed: 20331670]

Bierman KL, Nix RL, Greenberg MT, Blair C, Domitrovich C. Executive functions and school readiness intervention: Impact, moderation, and mediation in the Head Start REDI program. Development and Psychopathology. 2008; 20:821– 843.10.1017/S0954579408000394 [PubMed: 18606033]

Blair C, Granger D, Willoughby M, Mills-Koonce R, Cox M, Greenberg MT. the FLP Investigators. Salivary cortisol mediates effects of poverty and parenting on executive functions in early childhood. Child Development. 2011; 82:1970–1984.10.1111/j.1467-8624.2011.01643.x [PubMed: 22026915]

Blair C, Raver CC. Child development in the context of adversity: Experiential canalization of brain and behavior. American Psychologist. 2012; 67:309–318.10.1037/a0027493 [PubMed: 22390355]

Blair C, Razza RP. Relating effortful control, executive function, and false belief understanding to emerging math and literacy ability in kindergarten. Child Development. 2007; 78:647–663.10.1111/j.1467-8624.2007.01019.x [PubMed: 17381795]

Bloom, HS. Learning more from social experiments: Evolving analytic approaches. New York, NY: Russell Sage Foundation; 2005.

Burchinal M, Vernon-Feagans L, Cox M. the Family Life Project Investigators. Cumulative social risk, parenting, and infant development in rural low-income communities. Parenting: Science & Practice. 2008; 8:41–69.10.1080/15295190701830672

Caldwell, BM.; Bradley, RH. Home Observation for Measurement of the Environment. Little Rock: University of Arkansas at Little Rock; 1984.

Carlson SM, Mandell DJ, Williams L. Executive function and theory of mind: Stability and prediction from age 2 to 3. Developmental Psychology. 2004; 40:1105–1122.10.1037/0012-1649.40.6.1105 [PubMed: 15535760]

Cook TD, Shadish WR, Wong VC. Three conditions under which experiments and observational studies produce comparable causal estimates: New findings from within–study comparisons. Journal of Policy Analysis and Management. 2008; 27:724–750.10.1002/pam.20375

Cox M, Paley B, Burchinal M, Payne C. Marital perceptions and interactions across the transition to parenthood. Journal of Marriage and the Family. 1999; 61:611–625.10.2307/353564

Espy KA, Bull RB, Martin J, Stroup W. Measuring the development of executive control with the Shape School. Psychological Assessment. 2006; 18:373–381.10.1037/1040-3590.18.4.373 [PubMed: 17154758]

Ford, DL.; Lerner, RM. Developmental systems theory: An integrative approach. Newbury Park, CA: Sage; 1992.

Garon N, Bryson SE, Smith IM. Executive function in preschoolers: A review using an integrative framework. Psychological Bulletin. 2008; 134:31–60.10.1037/0033-2909.134.1.31 [PubMed: 18193994]

Gottlieb G. Experiential canalization of behavioral development: Theory. Developmental Psychology. 1991; 27:4–13.10.1037/0012-1649.27.1.4

Hammond SI, Müller U, Carpendale JIM, Bibok MB, Liebermann-Finestone DP. The effects of parental scaffolding on preschoolers' executive function. Developmental Psychology. 2011 Advance online publication. 10.1037/a0025519

Hirsh-Pasek K, Burchinal M. Mother and caregiver sensitivity over time: Predicting language and academic outcomes with variable- and person-centered approaches. Merrill-Palmer Quarterly. 2006; 52:449–485.10.1353/mpq.2006.0027

Hughes C, Ensor R. Executive function and theory of mind: Predictive relations from ages 2 to 4 years. Developmental Psychology. 2007; 43:1447–1459.10.1037/0012-1649.43.6.1447 [PubMed: 18020823]

Hughes C, Ensor R, Wilson A, Graham A. Tracking executive function across the transition to school: A latent variable approach. Developmental Neuropsychology. 2009; 35:20–36.10.1080/87565640903325691 [PubMed: 20390590]

Landry SH, Smith KE, Swank PR. Responsive parenting: Establishing early foundations for social, communication, and independent problem-solving skills. Developmental Psychology. 2006; 42:627–642.10.1037/0012-1649.42.4.627 [PubMed: 16802896]

Leventhal T, Brooks-Gunn J. The neighborhoods they live in: The effects of neighborhood residence on child and adolescent outcomes. Psychological Bulletin. 2000; 126:309–337.10.1037/0033-2909.126.2.309 [PubMed: 10748645]

Ludwig J, Kling J, Mullainathan S. Mechanism experiments and policy evaluations. Journal of Economic Perspectives. 2011; 25:17–38.10.1257/jep.25.3.17

Masten AS, Cicchetti D. Developmental cascades. Development and Psychopathology. 2010; 22:491–495.10.1017/S0954579410000222 [PubMed: 20576173]

McArdle JJ. Latent variable modeling of differences and changes with longitudinal data. Annual Review of Psychology. 2009; 60:577–605.10.1146/annurev.psych.60.110707.163612

McClelland MM, Cameron CE, Connor CM, Farris CL, Jewkes AM, Morrison FJ. Links between behavioral regulation and preschoolers' literacy, vocabulary, and math skills. Developmental Psychology. 2007; 43:947–959.10.1037/0012-1649.43.4.947 [PubMed: 17605527]

Morgan, SL.; Winship, C. Counterfactuals and causal inference: Methods and principles for social research. New York, NY: Cambridge University Press; 2007.

Muthén, LK.; Muthén, BO. Mplus user's guide. 6. Los Angeles, CA: Muthén & Muthén; 1998–2011.

Duncan GJ. National Institute of Child Health and Human Development Early Child Care Research Network. Modeling the impacts of child care quality on children's preschool cognitive development. Child Development. 2003; 74:1454–1475.10.1111/1467-8624.00617 [PubMed: 14552408]

NICHD Early Child Care Research Network. Chronicity of maternal depressive symptoms, maternal sensitivity, and child functioning at 36 months. Developmental Psychology. 1999; 35:1297–1310.10.1037/0012-1649.35.5.1297 [PubMed: 10493655]

Patterson, GR.; Reid, JB.; Dishion, TJ. Antisocial boys: A social interactional approach. Eugene, OR: Castalia; 1992.

Raver CC. Does work pay psychologically as well as economically? The role of employment in predicting depressive symptoms and parenting among low-income families. Child Development. 2003; 74:1720–1736. [PubMed: 14669892]

Raver CC, Jones SM, Li-Grining CP, Zhai F, Bub K, Pressler E. CSRP's impact on low-income preschoolers' pre-academic skills: Self-regulation as a mediating mechanism. Child Development. 2011; 82:362–378.10.1111/j.1467-8624.2010.01561.x [PubMed: 21291447]

Simons RL, Chao W, Conger RD, Elder GH. Quality of parenting as mediator of the effect of childhood defiance on adolescent friendship choices and delinquency: A growth curve analysis. Journal of Marriage and Family. 2001; 63:63–79.10.1111/j.1741-3737.2001.00063.x

Son SH, Morrison FJ. The nature and impact of changes in the home learning environment on development of language and academic skills of preschool children. Developmental Psychology. 2010; 46:1103–1118.10.1037/a0020065 [PubMed: 20822226]

Votruba-Drzal E. Income changes and cognitive stimulation in young children's home learning environments. Journal of Marriage and Family. 2003; 65:341–355.10.1111/j.1741-3737.2003.00341.x

Wiebe SA, Espy KA, Charak D. Using confirmatory factor analysis to understand executive control in preschool children: I. Latent structure. Developmental Psychology. 2008; 44:575–587.10.1037/0012-1649.44.2.575 [PubMed: 18331145]

Willoughby MT, Blair CB, Wirth RJ, Greenberg M. the Family Life Project Investigators. The measurement of executive function at age 3 years: Psychometric properties and criterion validity of a new battery of tasks. Psychological Assessment. 2010; 22:306–317.10.1037/a0018708 [PubMed: 20528058]

Willoughby M, Blair C, Wirth RJ, Greenberg M. the Family Life Project Investigators. The measurement of executive function at age 5: Psychometric properties and relation to academic achievement. Psychological Assessment. 2012; 24:226–239. [PubMed: 21966934]

Willoughby MT, Wirth RJ, Blair CB. Contributions of modern measurement theory to measuring executive function in early childhood: An empirical demonstration. Journal of Experimental Child Psychology. 2011; 108:414–435.10.1016/j.jecp.2010.04.007 [PubMed: 20553690]

Willoughby MW, Wirth RJ, Blair C. the Family Life Project Investigators. Executive function in early childhood: Longitudinal measurement invariance and developmental change. Psychological Assessment. 2012; 24:418–431.10.1037/a0025779 [PubMed: 22023561]

Zelazo, PD.; Müller, U. Executive function in typical and atypical development. In: Goswami, U., editor. Blackwell handbook of childhood cognitive development. Oxford, United Kingdom: Blackwell; 2002. p. 445-469.

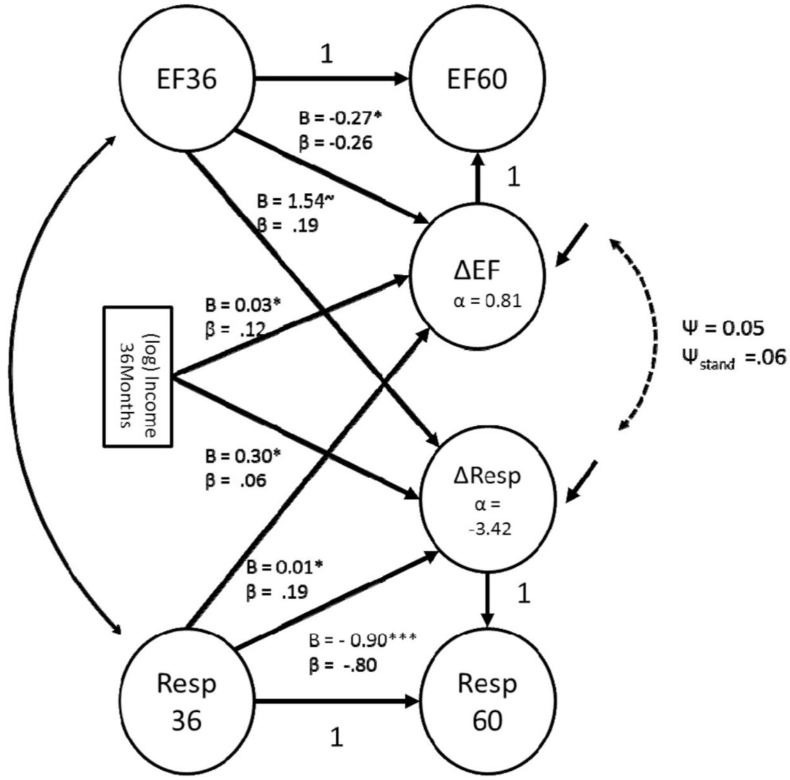

Figure 1.

Bivariate latent change score model estimating the respective cross-lagged relations between parental responsiveness (Resp) and children's executive function (EF) abilities between 36 and 60 months, adjusting for household income and the control covariates (*n* = 1,094). ~$p < .10$. * $p < .05$. *** $p < .001$.

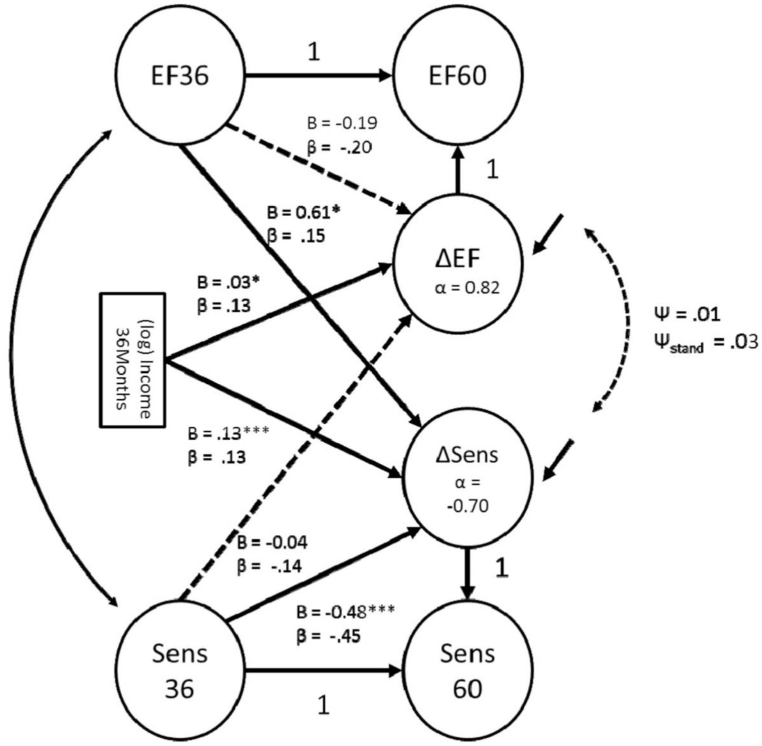

Figure 2.

Bivariate latent change score model estimating the respective cross-lagged relations between parental sensitivity (Sens) and children's executive function (EF) abilities between 36 and 60 months, adjusting for household income and the control covariates (n = 1,097). * p < .05. *** p < .001.

Table 1

Descriptive Statistics for Variables in the Analysis

Variable	***N***	***M***	***SD***
Maternal education	1,098	14.98	2.83
Income-to-need	1,099	1.92	1.61
Positive parenting 36 months	1,055	3.82	1.08
Positive parenting 60 months	963	3.98	1.09
HOME total mean 36 months	1,068	0.80	0.16
HOME total mean 60 months	1,038	0.76	0.10
Age in months 36 months	1,123	37.05	1.76
Age in months 60 months	1,099	60.62	3.26
Male	1,292	0.51	
Black	1,292	0.42	
Executive function 36 months	950	0.49	0.21
Executive function 60 months	1,038	0.72	0.14

Note. HOME = Home Observation for Measurement of the Environment.

Table 2

Fitted Estimates From a Cross-Lagged Latent Change Score Model, Considering Bidirectional Relations Between Parental Responsiveness and Executive Functioning Between 36 and 60 Months (*n* = 1,094)

	ΔEF		ΔParental responsiveness	
Variable	***B***	β	***B***	β
Intercept	0.81***	2.80	−3.41***	−0.58
Responsiveness (HOME 36 months)	0.013	0.19	−0.90***	−0.80
Executive function (36 months)	−0.27*	−0.26	1.54†	0.07
Control covariates				
Income	0.03*	0.12	0.30*	0.06
Maternal education	0.04***	0.24	0.45***	0.15
African American	−0.08†	−0.26	−1.18**	−0.20
Boy	−0.14***	−0.47	−.25	−0.04
Site	−0.02	−0.06	2.77***	0.47
Variance components (standardized)				
Latent covariance		0.05 (.06)		
R^2	.30		.70	
Fit statistics		−2LL = −27718.038		

Note. Absolute indices of model fit are not available with estimates derived from numerical integration. EF * executive function; HOME = Home Observation for Measurement of the Environment; −2LL = −2 log likelihood.

† $p < .10$.

* $p < .05$.

** $p < .01$.

*** $p < .001$.

Table 3

Fitted Estimates From a Cross-Lagged Latent Change Score Model, Considering Bidirectional Relations Between Parental Sensitivity and Executive Functioning Between 36 and 60 Months (n = 1,097)

	ΔEF		ΔParental sensitivity	
Variable	***B***	β	***B***	β
Intercept	0.82***	2.93	−0.70***	−0.60
Parental sensitivity (36 months)	−0.04	−0.20	−0.48***	−0.45
Executive function (36 months)	−0.19	−0.14	0.61*	0.15
Control covariates				
Income	0.03*	0.13	0.13***	0.13
Maternal education	0.03***	0.23	0.15***	0.26
African American	−0.07†	−0.24	−0.52***	−0.44
Boy	−0.13	−0.47	0.02	0.02
Site	−0.00	−0.01	0.12	0.10
Variance components (standardized)				
Latent covariance	0.01 (.03)			
R^2	.28		.34	
Fit statistics	χ^2= 1,339.36, df = 275, p < .001, CFI = .83, RMSEA =.06			

Note. EF = executive function; CFI = comparative fit index; RMSEA = root-mean-square error of approximation.

† $p < .10$.

* $p < .05$.

*** $p < .001$.