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Too Early for Early Education? Effects on Parenting for Mothers and Fathers

Objective: *To estimate the effect of children's age of entry into early childhood education and care (ECEC) on parenting quality of mothers*

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Key Words: child care, early childhood education, fathers, mothers, parenting.

and fathers in a context of universal access to subsidized ECEC following a 1 year paid parental leave.

Background: *Children entering non-parental care settings in early childhood may have negative consequences for parenting quality. Yet, current evidence supporting this claim is predominantly from the United States, is focused almost exclusively on mothers, and is predominantly based on statistical approaches that are vulnerable to unobserved selection bias.*

Method: *Data are from a Norwegian longitudinal study, including ratings of observed mother-child (n = 901) and father-child (n = 621) interactions, and children's age of entry into ECEC. Multivariate regression models and instrumental variable models were used to estimate the causal effect of age of entry on parenting quality.*

Results: *There was no support for the hypothesis that an earlier age of entry into ECEC negatively affects parenting quality, for either fathers or mothers. This was true for the sample as a whole, and for different sociodemographic subgroups.*

Conclusion: *In a Norwegian context in which families have universal access to subsidized ECEC from the time their child is 1 year of age, and most children enter ECEC in their second*

year, there is no evidence that an earlier age of entry in ECEC harms parenting quality.

While the proportion of infants and toddlers in early childhood education and care (ECEC) continues to rise in higher-income countries (OECD, 2018), there are concerns and controversies over adverse consequences, both historical and more recent. Most attention has been paid to consequences for children, with the primary concern being that early, extensive, and continuous care may lead to externalizing behavior problems such as aggression (e.g., Belsky, 1999, 2001; Huston, Bobbitt, & Bentley, 2015). Researchers have also been concerned that parent-child relationship quality is disrupted or becomes less close with early entry into ECEC (e.g., Belsky, 1999; NICHD Early Child Care Research Network, 1999, 2003). Studies addressing this latter question have come to conflicting conclusions, as we will review more in detail below. Importantly, research on this topic is primarily from North America, is almost exclusively focused on mothers, and either uses self-reported measures of parenting or statistical methods prone to bias from omitted variables. To build on existing evidence, we examined the relation between earlier entry into ECEC and the parent-child relationship, in terms of both mother's and father's parenting quality, using a quasi-experimental design with longitudinal and observational data from Norway.

BACKGROUND

Classic theoretical work on the parent-child relationship has emphasized the hazards of early separations for the quality of mother-child interactions (Brazelton, 1986; Sroufe, 1988; Vaughn, Gove, & Egeland, 1980). This position was consistent with one of the basic tenets of attachment theory, that a child's early separation from an attachment figure caused despair, anxiety, and withdrawal (Bowlby, 1973). It was also consistent with a more generalized theory of infant development, in which interactions with the primary caregiver are a cornerstone for the child's development (e.g., Stern, 1985). From these theoretical positions, extensive and continuous separations from the mother during the first years of life (e.g., non-parental infant and toddler care) were expected to reduce the time

and opportunities for developing a nurturing and sensitive mother-infant relationship. This would in turn have negative consequences for the child, given the importance of early parenting quality for subsequent child development in multiple domains (e.g., Collins, Maccoby, Steinberg, Hetherington, & Bornstein, 2000; NICHD Early Child Care Research Network, 2004; Wolff & van Ijzendoorn, 1997).

Previous theoretical work on this topic has paid particular attention to separations during the first years of life because of its foundational role in the development of attachment, perhaps also combined with practical concerns in a U.S. context where very early separations are more common. Yet, as noted by Bowlby (1969) and subsequently discussed by Sroufe and Waters (1977), children in their second and third years of life experience separations from their parents as threatening, and are not easily comforted by unfamiliar adults. Around age one, children are supposed to have established an internal representation of their parent's availability and responsiveness (Bretherton & Munholland, 2008). Repeated separations starting at this age, as would be the case if the child enters ECEC between ages one and two, may thus be a great source of distress for the child. For example, 2-year-old children, when distressed, maintain as much proximity to their caregivers as 1-year-old children, and are similarly distressed when faced with separations (Marvin, Britner, & Russell, 2016). A case in point is the fact that the Strange Situation procedure for assessing attachment security is considered valid from age one, when the attachment system is organized, through age 3 (Ainsworth, Blehar, Waters, & Wall, 1978; Solomon & George, 2016). From the perspective of the parent, this means that introducing daily extensive separations in the life of a 1- or 2-year-old child also entails containing the emotional responses (being sadness, clinging, anger, or rejection) which the child may exhibit both at drop-off and pick-up in ECEC, and during the time the child spends at home. Beyond an attachment perspective, having a child entering ECEC leads to disruption of routines and patterns in the relationship established during the child's first year. In cases where parents return to work when the child enters ECEC, balancing stress, demands, reorganization of household routines, and aspirations in multiple arenas may very well also constitute challenges. In sum,

there are multiple reasons why early entry into ECEC, not only in infancy, but also during the child's second year, may have negative consequences for parent-child relations.

The most concerning evidence which may be considered in support of this hypothesis comes from a quasi-experimental evaluation of the roll-out of universal child care subsidies in Canadian province of Quebec (Baker, Gruber, & Milligan, 2008). These authors found the roll-out led to increased hostile and inconsistent parenting, measured by parent self-report. Consistent with our theoretical prediction, this study included families with children entering across the infant/toddler and preschool years. While the authors did not specifically test whether the effect was moderated by child age at entry, the policy most strongly affected children aged 1–3 years (Kottelenberg & Lehrer, 2014), in part because of the expansion of a parental leave insurance made entry into non-parental care during the child's first year less common (Japel, Tremblay, & Cote, 2005). Notable strengths of this study are that it was based on representative national data, and had a research design with very strong internal validity, comparing parenting quality in Quebec before and after the child care expansion, to that observed in other areas of Canada not seeing a similar expansion. These strengths must be weighed against the potential imprecisions in their results due to the use of self-reported measures of parenting, considered to be more prone to bias and less valid than observational measures (Bennett, Sullivan, & Lewis, 2006; Hawes & Dadds, 2006).

Beyond this study, and particularly in studies using observations rather than self-reports of parenting quality, evidence for harmful effects of early and extensive child care on maternal parenting quality, is mixed, even within studies. Two seminal articles used data from the National Institutes of Child Health and Human Development Study of Early Child Care and Youth Development (NICHD SECCYD) to examine consequences of early and extensive non-parental care for the quality of mother-child interactions, focusing on sensitive responsiveness to the young child's needs (NICHD Early Child Care Research Network (ECCRN), 1999, 2003). In their large and rich dataset, conditioning on observed selection factors and additional child and family variables, the ECCRN found support for negative consequences of high amounts of

early non-parental child care for the quality of mother-child interactions up to child age 3 years (NICHD Early Child Care Research Network, 1999). Notably, this evidence was not strong, and there was no evidence that separations during early infancy were more harmful for the quality of mother-child interactions or for the security of attachment (NICHD Early Child Care Research Network, 1997a) than separations later in infancy. In addition, in a study targeting contrasting subgroups from the SECCYD, found that mother-infant interaction quality was not different in the group in which the children spent more than 30 hours in care during their first 6 months compared to infants cared for at home (Booth, Clarke-Stewart, Vandell, McCartney, & Owen, 2002).

In their subsequent study of longer-term effects on mother-child interactions through first grade, applying a similar analytical strategy, the ECCRN found the effect to be moderated by race (NICHD Early Child Care Research Network, 2003), in a way consistent with a non-significant trend seen for maternal sensitivity in the first 3 years. More non-maternal care across the first 3 years was negatively associated with maternal sensitivity in mother-child interactions from age three through first grade for white children, but positively associated for non-white children. (The non-white sample was primarily African-American.)

Increasingly, researchers raise concerns about the internal validity—the ability to make causal inferences—from conditional regression models of the type used by the ECCRN and many other researchers (e.g., Duncan, Magnuson, & Ludwig, 2004; Miller, Henry, & Votruba-Drzal, 2016). Briefly summarized, the concern is that if there are selection factors causing some parents to choose child care for their children and others not to do so, and if these factors also have a causal influence on parenting quality, the estimates will be biased if all of these selection factors are not included (correctly) in the regression model and are not measured well. Notably, such unobserved selection factors may bias the estimates either upwards (making the causal effect seem stronger than it is) or downwards (making it seem weaker). For example, conscientious parents with strong relational skills may be more prone to go back to work early because they feel their skills are needed by their coworkers (leading them to send their

children earlier into ECEC), while also providing high-quality parenting. If not adequately measured and conditioned on in a regression model, this scenario would push a true causal effect down. Nomaguchi and DeMaris (2013) addressed these types of concerns with regard to the findings in the NICHD SECCYD, by reanalyzing the data using fixed-effects panel models. This analytical approach relies exclusively on within-child (and parent) variation over time, essentially using each child (and parent) as its own control. Fixed-effects panel models provide stronger evidence for internal validity than standard conditional regression models, but may still be biased by unobserved time-varying confounding. Taking this approach, Nomaguchi and DeMaris (2013) found no evidence that amount of child care was associated with mother-child interactions, even within the subgroups where this has been previously shown. In addition, there is one notable experimental study randomizing a small sample of at-risk children into early infant care, finding small positive benefits of infant care for mother-child interactions (Burchinal, Bryant, Lee, & Ramey, 1992).

In sum, studies from the United States using elaborate observational measures of parenting quality (in different research designs) fail to support the hypothesis that early entry into ECEC has detrimental consequences for parenting quality. This is in notable contrast to the one population based, quasi-experimental, study from Quebec (Baker et al., 2008), in which the entry into ECEC across early childhood was found to be detrimental. There are many potential reasons for these discrepancies, relating to research designs and specific research question, measures of parenting, and sociopolitical context. In the present study, we add to this cumulative evidence base by combining the strengths of a quasi-experimental design with those of observations of parenting quality across both fathers and mothers, in a context of universal access to ECEC.

A Quasi-Experimental Test of Age of Entry Effects on Parenting Quality in Norway

In the present study, we test the hypothesis that earlier entry into ECEC has negative consequences for observed parenting quality in a large Norwegian sample. Unlike most previous studies, we examine consequences of early entry into ECEC for fathering as well

as for mothering. The literature examining effects of ECEC on parenting has focused almost exclusively on mother-child relations. Father-child relations have rarely been examined in this literature, with the possible exception of early studies of the effects of maternal employment from a family systems perspective that included infant-father attachments (e.g., Chase-Lansdale & Owen, 1987; Easterbrooks & Goldberg, 1985) from a single sample studied at two ages. Although studies have examined linkages between fathers' involvement with the child's care and paternal sensitivity (e.g., NICHD Early Child Care Research Network, 2000) potential associations of fathering quality with ECEC have not been addressed. To the extent that greater use of ECEC has been linked with maternal gatekeeping (e.g., Easterbrooks, Barrett, Brady, & Davis, 2007), and hence, less contact time with the child for fathers, we test the hypothesis that earlier entry into ECEC has negative consequences for observed fathering quality as well as for mothering quality.

In order to test our hypotheses, we take two analytical approaches. First, we use a conventional analytical approach (covariate adjusted regression models). Second, we use an instrumental variable (IV) approach, exploiting the fact that because of an uptake policy in Norway, where children born prior to 1 September were prioritized, with the consequence that birth month influences age of entry into ECEC. This approach is similar to one previously used in the same data, addressing consequences of early entry into ECEC for children's development of aggression (Dearing, Zachrisson, & Nærde, 2015).

Beyond the quasi-experimental methods and the inclusion of fathers, our focus on Norwegian families can provide a valuable addition to the cumulative knowledge. Norway offers universal subsidized ECEC to families beginning when their children are 1 year of age, and prior to this families have paid parental leave. Use of informal care is rare (Dearing, Zachrisson, Mykletun, & Toppelberg, 2018). Children almost exclusively attend full day care, though the amount of hours children actually spend in care varies, averaging 34 hours per week at age 2 in the sample used in the present study (Dearing et al., 2015), a weekly average reflecting that some parents either work part time or flexi-time, thus reducing their children's time in ECEC.

This socio-political context is radically different from that in the United States, where there is not universal paid parental leave or universal access to infant and toddler care, while being more similar to the sociopolitical context in Quebec.

In the early 2000s, Norway set an aim of universal access to publicly subsidized and quality-regulated ECEC from age one through school entry. Today, the vast majority of 1–2-year-old children in the country (over 80%) are in ECEC settings (Statistics Norway, 2018). However, children rarely enter ECEC prior to 9 months of age due to parental leave policy that provides 10 months leave at full pay or 12 months leave at 80% pay (Ministry of Children and Equality, 2014).

At the time when children in our sample entered ECEC (2007 through 2010), municipalities were required to coordinate enrollment into all centers using one primary enrollment date per year (Ministry of Education, 2007a). That enrollment date generally mirrored school enrollment practices. As was true in the municipalities from which we sampled, most municipalities offered well-subsidized care to children who were at least 1 year old by August, with enrollments on August 1 or August 15 (Ministry of Education, 2007b). In turn, remaining free slots were allocated via a waitlist that prioritized 11-month-old and 10-month-old children. Most infants who were not 10 months of age by August enrolled the following year—the following August—to begin ECEC.

These data from Norway thus provide us with a unique opportunity to use a quasi-experimental design to test the hypothesis that entry into ECEC during earlier ages causes poorer quality parent–child interactions (with regard to both mothers and fathers). Drawing on prior evidence from the United States, we will furthermore explore whether this effect differs across demographic groups, specifically non-western immigrants, and families with parents with low education and who experience economic hardship.

METHOD

Sample and Procedures

The data are from the *Behavior Outlook Norwegian Developmental Study* (BONDS). BONDS is a longitudinal study of 1,159 children (559 girls) from five municipalities in southeast Norway. The study was approved by the Regional

Committee for Medical and Health Research Ethics and the Norwegian Data Inspectorate. All parents provided informed written consent. In 2006–2008, families were informed about the project during their 5-month child health clinic visits in five municipalities in south-eastern Norway. These child health clinics are free and attended almost universally. Families were included in the study at least one parent being able to participate without a translator, with no other inclusion restriction. During the visits at the health clinics, a nurse informed the families about the project, and they were provided contact information if they agreed to be contacted. Information was given to the families of 1,931 eligible children. Of these, 1,465 (76%) accepted to be contacted, and 1,159 (79%, or 60% of those originally informed) eventually agreed to participate. The final sample is has slightly fewer mothers with only primary education than the invited sample (3.6% vs 5.9%), but otherwise fairly similar to the eligible families (for complete details, see Nærde, Janson, & Ogden, 2014). Two families later withdrew their participation and had all data deleted, reducing the total N to 1,157.

In the present study, we include three nested subsamples of children. These are (a) children who attended ECEC by age 3 years (1,073 children reported by their parents to attend ECEC by age 3, included for the analyses of fathers' parenting quality, measured at age 3); (b) children who attended ECEC by age 2 (957 children reported by their parents to attend ECEC by age 2, included for the analyses of mother's parenting quality, measured at age 2); (c) children who attended ECEC by child age 2 years, and were born in the months of February through October, were included in the IV analyses.

Data included in the present study were collected by trained assistants. Both parents were invited to participate in the 6-month interview whereas fathers were primarily targeted in the 1- and 3-years waves, and mothers in the 2-years wave. At child ages 6 months, 1, 2, and 3 years, this included an interview part and a computerized questionnaire section completed by the parent. The parents were also observed in play interactions with their children at age 1 (primarily fathers, not included in the current study), at age 2 (primarily mothers), and at age 3 (primarily fathers). In addition, three brief annual telephone interviews (3 months apart) were conducted with parents in between

in-person interviews. During these telephone interviews, the exact starting date in ECEC was confirmed.

Measures

Observations of Parenting Quality. Maternal and paternal parenting qualities were assessed from the structured interactions. Interaction procedures with mothers at age 2 years (16 minutes) and with fathers at age 3 (18 minutes) included five tasks: (a) free play (4 minutes); (b) clean-up (2 minutes); (c) teaching (6 minutes); (d) inhibition (2 minutes); and (e) waiting (2 minutes at 2 years; 4 minutes at 3 years). The current study utilized data from the teaching task, where the dyad was presented with two toys, a puzzle, and a shape sorting toy, chosen for each age to be a bit too difficult for children this age to complete by themselves. Parents were asked to help the child as much as necessary, spending 3 minutes with each toy in the predefined order. The research assistant left the room during the task, but knocked on the door after 3 minutes, and told the parent to switch to the other toy. Parents were informed that they could choose to terminate the task at any time.

For the present study, we use ratings on six parenting rating items, adapted from those used in the SECCYD at ages 15 and 24 months and following training received from the SECCYD investigator who led the rating of that study's parent-child interactions: (a) sensitivity/responsiveness (how the parent is tuned in to and responds to the child's behavior and signals); (b) detachment/disengagement (lack of emotional involvement and/or interest in the child's activities); (c) intrusiveness (attempts to control the child and impose parent's agenda); (d) cognitive stimulation (input to facilitate the child's learning); (e) positive regard (expressions of positive feelings toward the child); and (f) negative regard for the child (expressions of negative affection including disapproval, abruptness, negative corrections, and harshness). A 5-point rating scale from 1 (not at all characteristic) to 5 (highly characteristic) was applied for all items, representing levels of both quantity and quality of the rated item (Owen et al., 2010). Trained coders rated each item after watching the task twice, and 20% of the interactions were blindly assigned and double coded for inter-rater reliability analyses. Reliability coefficients for single items as determined from intra-class correlations

ranged from 0.72 to 0.80 ($M = 0.77$) for mothers and from 0.66 to 0.79 ($M = 0.73$) for fathers.

Following the rationale of Nordahl et al. (Nordahl, Owen, Ribeiro, & Zachrisson, 2020), we took a latent variable approach to the measurement model for parenting quality, using confirmatory factor analyses. A one-factor model including as indicators five of the six parenting items (excluding intrusiveness; Nordahl et al., 2020), fitted the data adequately for ratings of both parents (fathers: Root Mean Square Error of Approximation (RMSEA) = 0.078, Comparative Fit Index (CFI)/Tucker-Lewis Index (TLI) = 0.951/0.902; mothers: RMSEA = 0.097, CFI/TLI = 0.965/0.930), with standardized factor loadings ranging from 0.85 to -0.48 for fathers, and from 0.85 to -0.62 for mothers (all significant at $p < .001$). We saved out factor scores from the CFAs, which we used as dependent variables in our subsequent analyses.

Age of Entry into ECEC. At every contact point with the families (personal interviews or telephone interviews every 3 months), the parents were asked whether they knew exact starting date in ECEC. When information was obtained, we calculated exact age of entry into ECEC based on the child's birth date.

Covariates. Our selection of covariates was based on previous work on selection into early ECEC in Norway (Sibley, Dearing, Toppelberg, Mykletun, & Zachrisson, 2015; Zachrisson, Janson, & Nærde, 2013) as well as previous work on correlates of parenting quality (Nordahl et al., 2020). As for covariates predicting selection into ECEC, these have been shown to be demographic. Thus, from the 6-month interview, we included two dummies for parent's country of birth (i.e., immigrant status, categorized as Norwegian (reference), Western [Europe, North America, and Oceania], or non-Western [Asia, Africa, Latin America, and Turkey]), the proportion of time (across assessments) the parent has not lived in the family, up until and including the time of measurement of parenting quality, and a dummy for presence of a similar-aged sibling or not (i.e., age difference up to 5 years) including half-siblings and biologically unrelated siblings (i.e., children of a parent's new partner). We also included the child's gender, and dummies for site (five municipalities) and birth cohort (2006, 2007, and 2008). Lacking an exact report of family income, we included a dummy based on a question in the 1-year interview about

economic hardship (whether parents struggled keeping up with running expenses).

As covariates predicting parenting quality, we included a measure of parental mental distress (i.e., symptoms of depression and anxiety) measured with a mean score of the 13-item version of the Hopkins Symptom Check List (SCL-13; Strand, Dalgard, Tambs, & Rognerud, 2003), with an alpha of .91 for mothers at 2 years and .88 for fathers at 3 years. Finally, we included the parent's age at the child's birth, and the parent's report of hours spent in ECEC during a typical week (at age 2 years for the analyses including mother, age 3 years for fathers).

Analyses. In our first set of analyses, we fitted regular ordinary least squares (OLS) regression models $y_i = \alpha + \beta x_i + \beta z_i + \varepsilon_i$ for which: y_i is parenting quality for the i th child, α is the sample intercept, x_i is age of entry for the i th child, and z_i is a vector of covariates, and ε_i is the error term. To test differential associations for theoretically predicted moderators w_i , we subsequently included an interaction term ($x \times w$) along with the main effect of w_i . The validity of the causal inference from these models are based on the untestable assumption that our set of covariates fully capture all confounding covariance between parenting quality and age of entry into ECEC (Duncan et al., 2004; Foster, 2010; McCartney, Bub, & Burchinal, 2006).

As a more conservative approach to accounting for potential unobserved bias in our estimates, we used birth month as an instrument in two-stage IV estimations. Given that a number of assumptions are met, IV can be used to test the causal effects of a "treatment" (here: age of entry into ECEC) when assignment to the treatment condition is, at least partly, determined by the instrument in a manner that approximates randomization (see e.g., Angrist & Pischke, 2008, for a technical introduction; and Miller et al., 2016, for an untechnical introduction to causal inference in developmental psychology). We followed a previous argument, based on the same dataset, that birth month—as a linear indicator—provided a plausible instrument for isolating random variation in age of entry, due to ECEC enrollment policy in Norway (Dearing et al., 2015).

IV models rely on one key assumption of instrument strength, which is testable, and two

assumptions of instrument validity, which are in part based on plausibility. Concerning instrument strength, the instrument must be strongly predictive of exposure to the treatment (by convention, this means F -test statistic of ≥ 10). Concerning validity, the instrument should first be arguably independent of factors that influence the outcome. Second, the instrument should influence the outcome only through the treatment (often referred to as the exclusion restriction). In cases like ours, where the predictor, X (age of entry), and outcome, Y (parenting quality), are continuous indicators, both stages are appropriately conducted using two-stage linear estimators. The two-stage estimation entails that the predictor (age of entry) is regressed on the instrument (birth month) (Equation 1) and, second, the outcome (parenting quality) is regressed on that portion of the variability in age of entry that was predicted by birth month (Equation 2).

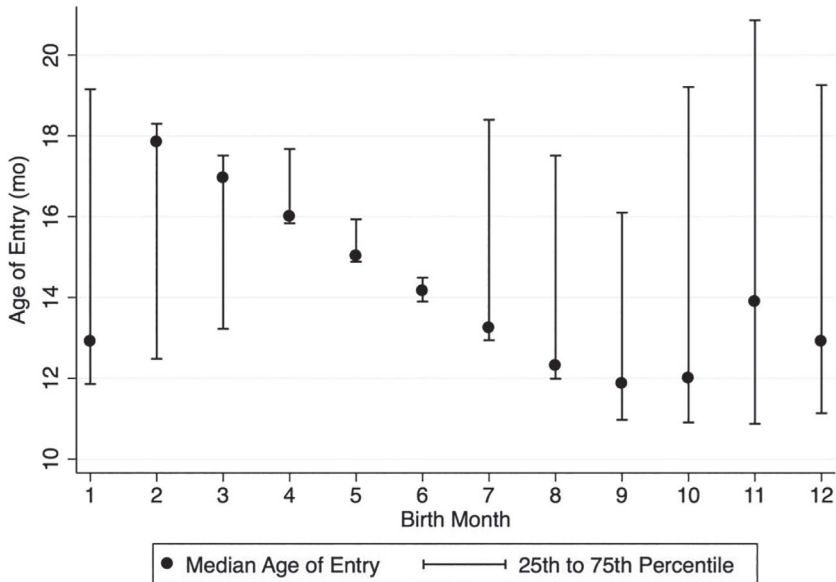
$$X_i = \alpha_0 + \beta_1 IV_i + \mu_i \quad (1)$$

$$Y_i = \alpha_0 + \beta_1 X'_i + \varepsilon_i \quad (2)$$

Because instrument validity is justified in equal parts by empirical evidence and logical argument, it should be considered on a continuum from "less plausible" to "more plausible" (Sovey & Green, 2011). In line with Dearing et al. (2015), we argue that the policy favoring August enrollments into ECEC in Norway allowed us to identify a component of entry age unrelated to any family characteristics. It is notable that the enrollment policy and practice did not restrict age of entry to age one or older, and children may enter ECEC at times other than August. And, it did not require children to enter ECEC at age 1. Birth month is therefore strongly, but not perfectly, related to age of entry for most months of the year. The prediction is also weakest in months furthest from August where parental choice and excess availability seem to play a bigger role (e.g., for children born in January, parents face choices such as enrolling child at 7 months with 3–5 months of paid parental leave remaining if they are able to obtain an open slot versus enrolling at 19 months, 7–9 months after paid leave has ended).

We focused on this group for two reasons. First, it was reasonable to assume that child birth month most strongly determined age of entry into ECEC within the first 2 years of the

FIGURE 1. MEDIAN AGE OF ENTRY PRIOR TO 24 MONTHS BY MONTH OF BIRTH. FIGURE IS TAKEN FROM THE AUTHOR, AND REUSED WITH PERMISSION FROM THE AUTHORS.



child's life, given the Norwegian policy context in which parental leave ended at either 10 or 12 months and subsidized care began at 12 months; all children have the opportunity to enroll by at least their second birthday.

Figure 1 shows that the median age of entry for children born from February through October decreased linearly, which is consistent with the expected effects of the enrollment timing. Children born between November and January demonstrated, in contrast, entry times quite deviant from the linear trend. For children born in September and October, there was likely a strong incentive to enroll in August (just before age 1) because parental leave ends at 10 or 12 months, and their waitlist priority increases their odds of obtaining a slot relative to younger children. Based on the distribution in Figure 1, it is evident that many of these families do, in fact, find an available slot, although this is truer for September births (11-month olds) than October (10-month olds). In turn, for winter births (e.g., November through January), parental choice and luck unrelated to birth month should come increasingly into play, consistent with variation indicated in Figure 1.

There have been concerns that birth months are not independent of family characteristics

(Buckles & Hungerman, 2013), especially in the United States. A review of this debate and an argument for why this is not a plausible threat to the validity of birth month as instrument in our data is presented in Dearing et al. (2015). These authors also detail analyses showing that birth month is balanced on observed background characteristics in this sample.

Missing Data. The percentage of missing data due to item non-response was less than 10 percent across all items, and missing responses were replaced with scale average for all dependent variables. Fathers' education and distress had the highest percentages of missing data, with 17% and 34%, respectively. All other independent variables had less than 10% missing values. For the dependent variables, data was available for 697 fathers at age 3 and 937 mothers at age 2. Missing data on these variables was accounted for with full information maximum likelihood estimation in the CFA models used to provide factor scores. Using multiple imputation with chained equations, we computed 20 datasets based on all variables listed in Table 1, including dependent and independent variables in the imputation model. Results were substantively identical when we used listwise deletion

Table 1. *Descriptive Statistics*

Variable	% Missing	<i>M</i> (<i>SD</i>) or %	Range
Parenting quality ^a			
Mother 24 months	22.13	0 (0.53)	-2.40-1.34
Father 36 months	46.32	0 (0.38)	-2.04-1.22
ECEC			
Age of entry (months)	9.94	14.14 (6.44)	7.00-41.95
Attending ECEC 24 months	5.79	83.09%	
Attending ECEC 36 months	7.52	93.23%	
Weekly hours 24 months	17.55	33.40 (7.72)	5.00-50.00
Weekly hours 36 months	13.22	34.45 (6.53)	5.00-50.00
Child and family covariates			
Gender (boy)	0.00	51.77%	
Father's age at birth (years)	1.82	33.85 (5.41)	21-63
Mother's age at birth (years)	0.09	30.80 (4.89)	16-45
Father's education (years)	17.03	14.09 (2.63)	9-18
Mother's education (years)	1.12	14.34 (2.56)	9-18
Western immigrant background	1.47	6.72%	
Non-Western immigrant background	1.47	7.65%	
Older siblings	0.78	58.45%	
Economic hardship (12 months)	5.27	15.85%	
Mother's distress 24 months	9.42	1.35 (0.44)	1-3.38
Father's distress 36 months	33.97	1.26 (0.34)	1-3.62
Proportion of time with no mom in primary family	2.33	0.01(0.07)	
Proportion of time with no dad in primary family	2.33	0.07(0.23)	
Birth cohort and site			
Born in 2006	0.00	37.42%	
Born in 2007	0.00	45.72%	
Born in 2008	0.00	16.86%	
Tinn	0.00	8.56%	
Bamble	0.00	12.36%	
Porsgrunn	0.00	26.97%	
Skien	0.00	12.88%	
Drammen	0.00	39.23%	

^a Negative values on the parenting quality measures are because they are factor scores, and thus centered.

for missing values, as well as when excluded the dependent variables from the imputation models; we therefore report results from the primary multiple imputation (MI) analyses only.

RESULTS

Table 1 provides descriptive statistics for all variables used, for the full sample ($N = 1,157$). At age 2 years, 83.09% of the total sample had enrolled in ECEC, at 3 years, 93.23% had enrolled. Age 2 and 3 years were the ages at which we collected data on parent-child interactions with mothers and fathers, respectively. Mean age of entry into ECEC was 14.14 months, with a standard deviation of 6.44 months.

Covariate-Adjusted OLS Analyses

To reiterate, our hypothesis was that age of entry into ECEC was associated.

with the quality of parent-child interactions, with mother at age 2 and father at age 3, respectively. Secondly, we explored whether this association varied as a function of a number of moderator variables; parental education, non-western immigrant background, economic hardship in the family, and, as a sensitivity check, the proportion of time the target parent had not lived with the family, which may influence the parenting quality.

As a first set of analyses, we fit linear regression models separately for these two outcomes,

Table 2. Regression Models Predicting Parenting Quality for Mother–Child and Father–Child Interactions from Age of Entry into ECEC

	Mother 24 months (<i>n</i> = 957)		Father 36 months (<i>n</i> = 1,073)	
	Coef (<i>SE</i>)	<i>r_p</i>	Coef (<i>SE</i>)	<i>r_p</i>
Unconditional model				
Age of entry	−0.01 (0.00)	−0.04	−0.00 (0.00)	−0.01
Conditional model				
Age of entry	−0.00 (0.00)	−0.02	0.00 (0.00)	0.00
Child and family covariates				
Gender (boy)	−0.10 (0.04)**	−0.09	−0.08 (0.03)**	−0.10
Mother's age at birth (years)	0.01 (0.00)*	0.06		
Father's age at birth (years)			−0.00 (0.00)	−0.00
Mother's education (years)	0.03 (0.01)***	0.11		
Father's education (years)			0.02 (0.01)**	0.10
Western immigrant background	−0.03 (0.07)	−0.01	−0.01 (0.05)	−0.00
Non-Western immigrant background	−0.18 (0.08)*	−0.08	−0.04 (0.05)	−0.02
Older siblings	0.01 (0.04)	0.01	−0.02 (0.03)	−0.03
Economic hardship (12 months)	−0.06(0.05)	−0.04	−0.09 (0.04)+	−0.07
Mother's distress 24 months	−0.01 (0.04)	−0.00		
Father's distress 36 months			−0.01 (0.04)	−0.00
Mother not in primary family	0.11 (0.26)	0.01		
Father not in primary family			0.00 (0.06)	0.00
Weekly hours 24 months	0.00 (0.00)	0.01		
Weekly hours 36 months			−0.00 (0.00)	0.01
Birth cohort and municipality site				
Born in 2006	0.08 (0.06)	0.05	−0.05 (0.04)	−0.04
Born in 2007	0.07 (0.05)	0.04	−0.02 (0.04)	−0.01
Tinn	−0.11 (0.07)+	−0.05	−0.09 (0.05)+	−0.06
Bamble	−0.08 (0.06)	−0.04	−0.04 (0.04)	−0.03
Porsgrunn	−0.07 (0.04)	−0.05	−0.04 (0.03)	−0.03
Skien	−0.06 (0.06)	−0.03	−0.04 (0.04)	−0.03
Birth month (linear)	−0.00 (0.00)	−0.03	−0.00 (0.01)	−0.02

Note. The covariates that are specific to each parent are included only in models predicting that parent's parenting quality. * $p < .05$. ** $p < .01$. *** $p < 0.001$.

as can be seen in Table 2. There was no evidence that parenting quality, in either mother– or father–child interactions, was associated with age of entry into ECEC. This was true for both unconditional models (without any covariates), and for the conditional (on covariates listed in Table 2). Effect sizes were close to zero across all models (ranging from .00 for father to .04 for mother in the unconditional model), and corresponding p -values were closer to 1 than 0 across all models. More educated parents, and older mothers, were rated as having higher parenting quality ($r_p = .10$, $p < .01$ for father's education; $r_p = .06$, $p < .05$ and $r_p = .11$, $p < .001$ for mother's age and education, respectively). In addition, mothers, but not fathers, with

non-western immigrant background, were rated as having lower parenting quality ($r_p = .08$, $p < .05$). Finally, parents of boys were rated as having lower parenting quality as compared to those of girls ($r_p = .08$, $p < .01$ and $r_p = .10$, $p < .001$, for mothers and fathers, respectively). None of the other covariates were significant predictors of parenting quality, according to conventional standards, although fathers experiencing lower economic hardship were rated as having marginally lower parenting quality ($r_p = .07$, $p < .10$).

In an exploratory set of analyses, we tested whether the association between age of entry and parenting quality varied as a function of mothers' and fathers' education, non-western

immigrant background, economic hardship, and proportion of time the father and mother (respectively) had been absent from the child's primary family. We did this by including interaction terms for each of these moderator variables in separate regression models, conditional on all covariates listed in Table 2. We found no evidence for moderator effect of any of these variables, for either of the parents (results not shown). We also tested a non-linear specification of age of entry (by including an age of entry squared term in the model). This did not yield any evidence for non-linear associations (results not shown).

Instrumental Variable Analyses

Next, we tested whether our null-findings from the OLS models described above were upheld when we used a quasi-experimental approach with IV estimation. To reiterate, we used a two-stage estimation technique where the proportion of variance in age of entry accounted for by birth month (as a function of the main uptake in ECEC as of September 1), was used to predict parenting quality. We restricted our primary IV model to include children born from February through October, as children born during these months entered ECEC at a linearly decreasing age (see Figure 1). The F -value for the first-stage equation was 38.13, indication a strong instrument (well above the conventional cutoff of $F > 10$).

The IV models were consistent with the OLS models, and did not provide any additional evidence that age of entry into ECEC was associated with the quality of parent-child interactions at age 2 with the mother or at age 3 with the father (see Table 3). Moreover, using IV, we found no additional evidence for differential associations for any of the moderators described above.

We also ran a set of sensitivity checks for the IV models. First, we tested the models including a varying range of birth months, as our primary choice of window of birth months was empirically driven by inspection of the data. Our sensitivity checks included children born in March through September (first-stage F -value = 28.10), February through October (first-stage F -value = 37.69), February through November (first-stage F -value = 22.62), and all months (first-stage F -value = 13.48). None of these additional tests provided substantively different results than our primary strategy.

DISCUSSION

The aim of this study was to test the hypothesis that earlier entry into ECEC has negative consequences for observed quality of mother-child and father-child interactions of Norwegian parents. In this study, we address concerns about selection biases in the field, by using an IV approach, which treats birth month as a random variable associated with age of entry into child care. We also expand the evidence base beyond the North American context, to Norway with its unique progressive family policy. Finally, unlike most research conducted so far, we examine the effects of ECEC on parent quality for fathers as well as mothers.

We found no support for our hypothesis that early entry into ECEC has negative consequences for parenting quality, for either of the parents. Across all our statistical approaches, we found no evidence that parent quality, in either mother- or father-child interactions was associated with age of entry into ECEC. Our result is not consistent with the large quasi-experimental study from Quebec, finding universal roll-out of ECEC to have negative consequences for parenting quality (Baker et al., 2008). While this

Table 3. Two-Stage Instrumental Variable Models Predicting Parenting Quality for Mother-Child and Father-Child Interactions from Age of Entry into ECEC

	Mother 24 months ($n = 739$)		Father 36 months ($n = 739$)	
	Coef (SE)	r_p	Coef (SE)	r_p
Unconditional model				
Age of entry	-0.05 (0.04)	-0.05	0.03 (0.03)	0.04
Conditional model ^a				
Age of entry	-0.02 (0.03)	-0.02	0.02 (0.03)	0.03

^aThe models are conditioned on the same covariates as the OLS regressions in Table 2.

study differs in notable ways from ours, in terms of outcomes (parent report vs observations), and scope (policy expansion vs age of entry), both have strong internal validity and take place in contexts of universal child care and in which children tend to enter child care after age 1. If similar effects as those found in Quebec of the child care roll-out were to be found in other contexts as Norway, we believe we would have seen negative effects of age of entry in our study.

Our results are more aligned with previous studies from the United States, having not found consistent evidence that entry into nonparental care during infancy is related to parenting quality, including Booth et al. (2002) and NICHD ECCRN (1999; 2003). Notably, the latter studies found some evidence that quantity of care across the first 3 years of life was associated with the quality of mother-child interactions across infancy and early childhood. Moreover, we did not find evidence for the positive effects of very early nonparental care on mother-child interactions, as shown in the experimental study by Burchinal et al. (1992).

Our results cannot be compared to the North American studies in any straight forward manner. For example, our results should likely be conditioned by the fact that age entry into child care in Norway, as opposed to countries like the United States, is rarely under age 9 months. Notably, a previous study using a similar approach with the same dataset addressed the consequences of early entry into ECEC for children's development of aggression (Author). In line with the present study, the authors found no long-lasting negative effects, albeit some temporary increase in aggression levels upon ECEC entry. It is not surprising that for the quality of parent-child interactions our null results are even more robust. It seems more plausible that ECEC entry would affect children's behavior more directly, whereas parenting quality is a more distal variable affected by multiple processes. These include factors such as, for example, parent ability and mental state, contingency of child's behavior, and a dynamic process involving the two. These include factors such as parent ability and mental state, contingency of child's behavior. For example, parenting quality has been associated with child's temperament (Clark, Kochanska, & Ready, 2000), attachment to the child (Rosen & Rothbaum, 1993), social support (Armstrong, Birnie-Lefcovitch, & Ungar, 2005), marital quality (Benzies,

Harrison, & Magill-Evans, 2004), and depression/anxiety (Lovejoy, Graczyk, O'Hare, & Neuman, 2000). The wider sociopolitical context of child care and parenting should also be taken into account (Scarr & Eisenberg, 1993). The Norwegian context is highly favorable to the conciliation of parent and work domains, providing shorter working days, possibility of reduced working hours and a generally family friendly working environment, which may make the transition into child care a less stressful experience for both parents and children. Moreover, fathers are encouraged to take an active part in their children's upbringing, for example through the father's quota of a 1-year parental leave during the child's first year of life. Thus, compared to other contexts (notably, the United States) parents of both genders in Norway may have ample opportunities to parent their child, and thus be more competent, comfortable, and confident in their parenting, as their child attends ECEC. Even children who are full time in ECEC spend relatively few hours there, again compared to, for example, the United States, allowing more time to continuous interactions with parents.

In our second set of models, we predicted parenting quality from the proportion of variance in age of entry accounted for by birth month, that is, identified as a component of entry age unrelated to family characteristics. Similar to the results from the first set of analyses, we found no evidence that age of entry into ECEC was associated with the quality of parent-child interactions when taking a more conservative approach to selection bias. Our findings, employing a quasi-experimental design, are consistent with previous studies of quantity of care (not age of entry per se) where selection effects were also addressed, namely by fixed-effect models (Nomaguchi & DeMaris, 2013) using the SECCYD data. Keeping in mind that unobserved selection factors can bias an estimate both upwards and downwards, the results from these additional analyses, and their consistency with previous findings in a different context (i.e., the United States), suggests that unobserved selection does not seem to bias our findings.

Yet, we acknowledge that associations in our study not reaching conventional levels of statistical significance is not evidence of the absence of such an association, despite this often being assumed in the psychological research literature (Aczel et al., 2018). Rather, *p*-values greater

than, say, .05, should be interpreted as absence of evidence—not support for the null hypothesis (Cohen, 1994). Thus, our study provides no evidence in support of our null hypothesis in the population of Norwegian children. With this caveat in mind, it is worth considering the standardized effect sizes (partial correlation coefficients) in our study. In the conditional regression models they were, both for mothers (at 24 months) and fathers (at 36 months), zero. In the IV models, the associations ranged from -0.05 to 0.03 . Note that the IV estimates have larger standard errors, which is very common especially in smaller samples; they sacrifice precision to reduce bias. In either case, the effect sizes do not warrant concern over negative effects of age of entry on parenting quality.

This study has several limitations worth pointing out. First, the study did not measure and we therefore did not control for child care quality, which may affect home quality and parenting (McCartney, Dearing, Taylor, & Bub, 2007; NICHD ECCRN, 1999, 2003). Although structural quality in Norwegian ECEC centers is regulated and most of them comply with set quality standards (Gulbrandsen & Eliassen, 2013), observed quality is quite variable (Bjørnstad & Os, 2018). Thus, there may be heterogeneity in the effects we have not detected due to ECEC quality, even though the size of our standard errors renders this unlikely. Second, the fact that in our sample no children started child care before age 8 months makes our study incomparable with research conducted in countries such as the United States, where entry ages are considerably lower on average. In our study, we take parent–child interactions observed in the laboratory context as a sample of behavior normally occurring in their daily interactions, thus a proxy for parenting quality. Although the observational nature of this study is an advantage in relation to other studies using parent self-report (e.g., Baker et al., 2008), we must acknowledge that we have captured a sample of parent–child behaviors that is relatively brief and which might have been influenced by the artificial context or specific isolated factors like mood, amount of sleep, anxiety levels, and so on. Nonetheless, good evidence for validity of the measure of mothering quality, in relation to predictors of individual differences and relations to child outcomes, has been shown (Nordahl et al., in press). Finally, our study had only a 60% baseline participation

rate. While being similar to other comparable studies (e.g., 50% in the NICHD SECCYD; NICHD ECCRN, 1999), we cannot rule out that unobserved selection factors into the study would moderate the association between age of entry and parenting quality.

In spite of these limitations, our study has considerable strengths, including a large sample where both parents' interactions with their children have been rated, in a context which both provides a supportive environment for families, and a design opportunity for a quasi-experiment. We find no support for the hypothesis that earlier entry, and thereby greater quantity of ECEC attendance, entails harmful consequences for children and their parents, when it comes to the quality of their interactions. The most important implication of this finding is that there is one more rigorous research study failing to provide evidence for harms brought on by young children's ECEC experience, at least in the national context studied here. The economic reality of today's families often requires both parents to work if they are able, and thus to have their children cared for by others. We do not find evidence suggesting that parents should be concerned that this will hamper the quality of their parenting.

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