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Maternal bereavement: The heightened mortality of mothers after the death of a child

Javier Espinosa [a,*], William N. Evans [b,1]

[a] Department of Economics, Rochester Institute of Technology, 92 Lomb Memorial Drive, Rochester, NY 14623, USA
[b] Department of Economics, University of Notre Dame du Lac, 437 Flanner Hall, Notre Dame, IN 46556, USA

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ABSTRACT

Using a 9-year follow-up of 69,224 mothers aged 20–50 from the National Longitudinal Mortality Survey, we investigate whether there is heightened mortality of mothers after the death of a child. Results from Cox proportional hazard models indicate that the death of a child produces a statistically significant hazard ratio of 2.3. There is suggestive evidence that the heightened mortality is concentrated in the first two years after the death of a child. We find no difference in results based on mother's education or marital status, family size, the child's cause of death or the gender of the child.

1. Introduction

Inquiry into the health outcomes of bereavement has been conducted over the past decades by researchers from a variety of disciplines, including psychology, epidemiology, economics, sociology, and other social and medical sciences.[2] Researchers have examined the physical health (e.g., mortality, onset of cancer) and mental health consequences (e.g., depression, anxiety), as well as the possible pathways linking the grief of bereavement to health outcomes (Gerra et al., 2003; Gündel et al., 2003; Kiecolt-Glaser et al., 2002; Schleifer et al., 1983; Stroebe et al., 2006). Most commonly, studies have addressed the issue of spousal bereavement and health, finding that bereft spouses experience a variety of poor outcomes including excess mortality, a result referred to as the "spousal bereavement effect".[3] Empirical results have also demonstrated negative relationships between bereavement and measures of health for grandparents,[4] parents,[5] children,[6] and siblings,[7] but there remains a paucity of research addressing the bereavement effect for non-spousal relationships.

A child's death is a horribly tragic event for parents, and the mental repercussions are generally thought to be grave. Yet research on parental health after the death of a child represents only a small portion of the health-outcomes research on bereavement (Stroebe et al.,

* Corresponding author. Tel.: +1 202 415 5094; fax: +1 585 475 2510.
E-mail addresses: javier.espinosa@rit.edu, jxegse@rit.edu (J. Espinosa), wevans1@nd.edu (W.N. Evans).
[1] Tel.: +1 574 631 7039.
[2] For a review, see Stroebe et al. (2007).

[3] For examples, see Elwert and Christakis (2008), Espinosa and Evans (2008), Kaprio et al. (1987), Lillard and Waite (1995), Murray (2000), Rogers (1995), and Smith and Zick (1994).
[4] For examples, see Fry (1997), Ponzetti (1992), and Youngblut et al. (2010).
[5] For examples, see Agerbo (2005), Bohannon (1991), Davies (2006), Levav et al. (1988, 2000), Li et al. (2002, 2003a,b, 2005), Manor et al. (2000), Murphy et al. (1999, 2003), Olsen et al. (2005), Ponzetti (1992), and Schwab (1996).
[6] For examples, see Dowdney (2000) and Worden and Silverman (1996).
[7] For examples, see Birenbaum et al. (1989) and Brent et al. (1993).

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2007). Published research on the relationship between parental bereavement from the death of a child and parental health outcome includes two quantitative studies from Denmark showing evidence of a link between parental bereavement and an increased risk for hospitalization for a mental health reason or for a diabetes-related condition (Li et al., 2005; Olsen et al., 2005). Specifically, Li et al. (2005) demonstrated that parents who experienced the death of a child had a higher risk of first-time hospitalization for a psychiatric disorder than parents who did not lose a child. Moreover, mothers had a higher relative risk than fathers, the effect of which was most acute during the first year and significantly elevated for five years or more. Olsen et al. (2005) showed that mothers who lost a child had a statistically significant increase in risk for first-time hospitalization for diabetes type 2 compared with mothers who did not lose a child. Other works have centered on the relationship between parental bereavement and cancer incidence and survival (Li et al., 2002, 2003a), with results showing a slight increase in overall risk for cancer in bereaved mothers (possibly due to stress-induced adverse life styles) but no difference in cancer survival rates between parents who lost a child and those who did not. Finally, a study using data from Denmark showed that parents who experienced the death of a young child demonstrated a statistically significant increase in mortality (Li et al., 2003b), whereas another study from Israel found that parents who experienced the death of an adult child did not show an elevation in mortality (Levav et al., 1988).

In this study, we investigate whether there is heightened mortality of mothers after the death of a child—i.e., a *maternal bereavement effect*—using a nine-year follow-up of 69,224 mothers aged 20–50 from the National Longitudinal Mortality Survey of the United States Census Bureau.[8] This is the first study to empirically analyze this issue with a large, nationally representative U.S. data set. In this case, we exploit a unique feature of the National Longitudinal Mortality Survey that allows us to track the mortality of children even after they leave the household. We find evidence that a mother's mortality is heightened after the death of a child and that the excess mortality is greater in the first two years after the death of a child. The latter finding is especially relevant to public health policy and the timing of interventions that aim to ameliorate the adverse health outcomes mothers experience after the death of a child. In addition to investigating the relationship between maternal bereavement and mortality in the United States, we aim to assess whether there is heterogeneity in the effect based on such characteristics as household income, mother's educational attainment, and the child's cause of death.

2. Data and methods

All the work related to this study has been carried out in accordance with the Uniform Requirements for manuscripts submitted to Biomedical journals, and no human or animal experimentation is associated with this work.

2.1. Data

The primary data source for this analysis is the public-use version of the National Longitudinal Mortality Study (NLMS), a data set constructed by merging person-level responses from the Census long-form data and the Current Population Survey (CPS) with death certificate information on mortality status and cause of death from the National Death Index (NDI). The CPS data provide information on household income, occupation, education, and other demographic variables; for most surveys, the CPS data do not provide detailed information on health status or health behaviors. The public-use version of the NLMS contains data from five monthly CPS samples from 1979 through 1981, and in this version the mortality follow-up is set to nine years from entry into the CPS and captures the number of days from initial survey until death or the end of nine years. The linking of the cross-sectional data to the NDI provides a time-series component to the static CPS data; namely, there is a variable that captures the number of days a respondent is alive after entering the survey. All other NLMS variables, like household income and mother education, are measured only once and that is at the time respondents enter the survey.[9]

The CPS is a household-based survey, and it contains data on all household members. Using household identification numbers and variables in the NLMS that identify a respondent's relationship to the head of household, we are able to link mothers to their children.[10] Because the vast majority of children live with the mother regardless of the family living arrangement, the match of the CPS to the NDI allows us to observe deaths of children even once they leave their original household. A limitation of the household survey is that a household sample that includes older mothers with older children will, by construction, not include children who left the household before the initial interview. Thus, we cannot identify these children or their mortality information.

This limitation in using the NLMS data set to address the maternal bereavement question is illustrated in Fig. 1. The 1980 Census 5-Percent Public Use Micro Samples (PUMS)[11] asks mothers how many children they have had, and using the detailed relationship codes in the census, we can calculate—by age of mother—the fraction of children ever born to the mother who are still living in the household. This number is reported in Fig. 1. Note that through age 35, roughly 90% of children ever born to a mother are in their mother's household at the time of the survey. However, this number drops precipitously after that. The fraction of children still in the mother's household falls to 73% at age 40, 50% by age 45, and 33% by age 50.

[8] We use the nomenclature "mother mortality" instead of "maternal mortality" to avoid confusion, as the latter term is commonly employed by researchers to denote deaths of women due to causes associated with pregnancy (Hill et al., 2007).

[9] See Rogot et al. (1988, 1992) for more information on the NLMS.

[10] We cannot identify if a child is adopted in the NLMS.

[11] We access the data using the Integrated Public Use Microdata Samples, see Ruggles et al. (2010).

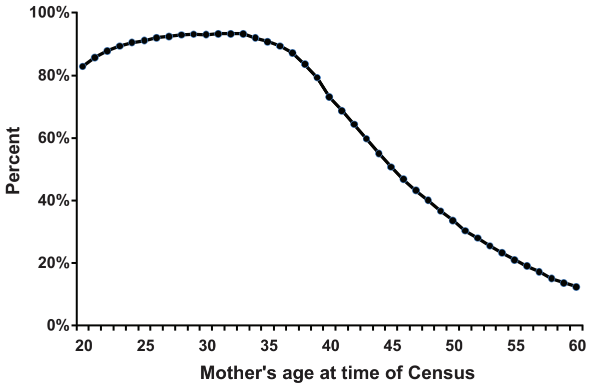

Fig. 1. Percent of children even born still living with mother by mother's age, 1980 Census 5% PUMS.

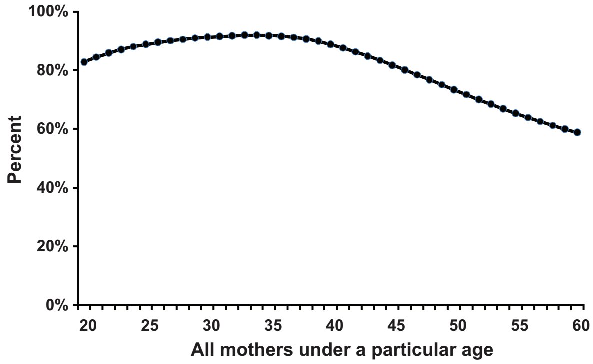

Fig. 3. Cumulative percent of children ever born still living with mother under a particular age, 1980 Census 5% PUMS.

Because many of the older mothers in the sample will have children who already left the household by the time of the initial CPS survey that forms the basis of the NLMS, these women may experience the death of a child during the nine-year follow-up period, but this will not be recorded because that child was not in the household at the time of the initial survey. This should systematically attenuate any bereavement effect on mortality for these mothers. This concern can been seen in Fig. 2 where we graph the fraction of mothers in the NLMS that die within the nine-year follow-up by five-year age groups based on the mother's age at the time of the initial survey. We also include the fraction of mothers in each age group who experienced the death of a child during the nine-year follow-up period. Note, as expected, mother mortality is increasing with age of the mother at the time of the survey. We would expect the fraction of mothers who lost a child to increase with the initial age of mother at the time of the survey, since older mothers tend to have older and more children, yet this rate is relatively flat after age 40.

In general, this problem can be dealt with by restricting the sample to younger mothers. Given the low mortality rates of mothers in this age range and their children, we must trade off expanding the sample to include older mothers in order to increase power against the attenuation bias associated with including older mothers whose children have already left the household. For this reason, we restrict the sample to women who were aged 20–50 at the time of the survey. With this age restriction, we have a large fraction of all children born to mothers in the age range in our sample. This can be seen in Fig. 3, where we report the cumulative percent—by the age of the mother—of all children born to mothers in the 1980 PUMS who remain living with their mothers at the time of the survey. In this figure, we see that 92% of children for mothers aged 35 and under at the time of the original survey are living in their mother's household. This number falls to 88% for mothers 40 and under at the time of the survey, but the number is still 75% for mothers 50 and under.

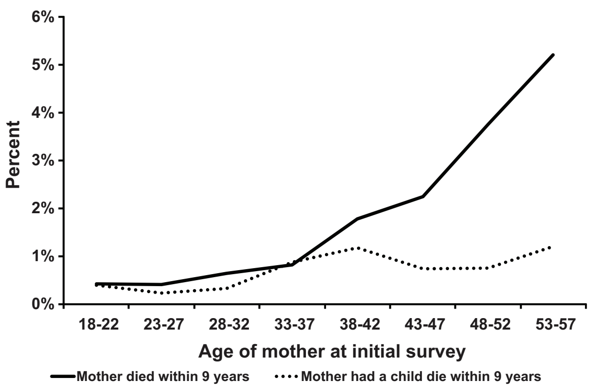

Fig. 2. 9-Year mortality rates for mothers and percent of mothers that lost a child within 9-years, NLMS.

Another limitation of the NLMS is that a mother could have a child that dies before the survey begins; thus, there may be bereft mothers in the analysis sample who are not identified as having experienced the death of a child. Similar to the previous issue, a sample including older mothers will likely have more cases of unidentified bereft mothers than a sample that includes fewer older mother observations. Thus, if the problem is qualitatively important, estimates of the maternal bereavement effect will be attenuated. We examine this issue in detail in Section 3.3.

2.2. Sample characteristics

The NLMS data enable us to identify 69,224 records of mothers aged 20–50 at the initial survey. In the first column of Table 1 we report descriptive statistics for this sample. The nine-year mortality rate for the sample of mothers is low at 1.2%. The incidence of maternal bereavement is also low, at 0.7%. The sample of mothers is largely composed of women who are married (85%), white (87%), and non-Hispanic (91%). Slightly more than a simple majority of the mothers are between the ages of 20 and 34. Approximately one-half have a high school education, and one-third have some college education or have completed college. Less than 20% of mothers in the sample have only some high school education or never attended high school.

There are seven income categories: five categories capture household incomes (1980 dollars) below $25,000, in $5000 increments, and the other two categories are for household incomes between $25,000 and $50,000 and for

Table 1
Sample characteristics of mothers, aged 20–50 years at time of survey, United States 1979–1981, publicly use National Longitudinal Mortality Study.

Variable	All mothers	Mothers with a child that died within 9 years	Mothers without a child's death	H_0: means are equal
	% of sample	% of sample	% of sample	*p*-Value
Mother died within 9 years	1.20	1.90	1.19	0.16
A child died within 9 years	0.70	–	–	–
Marital status				
Married	84.52	81.86	84.53	0.11
Not married	15.48	18.14	15.47	0.11
Race				
White	87.25	83.97	87.28	0.03
Black	9.30	11.18	9.29	0.16
Other race	3.45	4.85	3.43	0.09
Ethnicity				
Hispanic	8.68	11.60	8.66	0.02
Not Hispanic	91.32	88.40	91.34	0.02
Family income				
≥$50 K	4.46	2.74	4.48	0.07
≥$25 K, <$50 K	26.41	25.53	26.41	0.66
≥$20 K, <$25 K	17.03	15.40	17.05	0.34
≥$15 K, <$20 K	16.08	13.92	16.09	0.20
≥$10 K, <$15 K	17.22	20.89	17.19	0.03
≥$5 K, <$10 K	12.08	13.29	12.07	0.42
<$5 K	6.72	8.23	6.71	0.19
Education of mother				
College graduate	13.87	11.19	13.89	0.09
Some college	18.36	13.08	18.39	<0.01
High school graduate	48.78	49.37	48.78	0.80
Some high school	12.82	17.09	12.79	<0.01
<High school	6.18	9.28	6.16	<0.01
Mother's age at time of survey				
20–34	51.45	25.53	51.63	<0.01
35–50	48.55	74.47	48.37	<0.01
Occupation of mother				
Private household worker	1.06	1.69	1.06	0.18
Service	11.00	13.50	10.99	0.08
Laborer	0.80	0.84	0.80	0.91
Operators	7.29	6.12	7.29	0.33
Operate equipment	0.60	0.84	0.60	0.49
Craftswomen	1.23	1.05	1.23	0.73
Clerical	21.76	18.14	21.78	0.06
Sales	3.83	4.64	3.83	0.36
Professional	10.69	7.81	10.71	0.04
Manager	3.97	4.01	3.97	0.97
Farming	1.17	1.27	1.17	0.84
Missing/never worked	36.59	40.08	36.56	0.11
Family size at time of interview				
One child	38.25	22.15	38.36	<0.01
Two children	38.42	35.02	38.44	0.13
Three children	16.11	23.00	16.07	<0.01
Four children	5.52	12.45	5.47	<0.01
Five children	1.69	7.38	1.66	<0.01
Observations	69,224	474	68,750	

household incomes greater than $50,000. Less than 5% of mothers had a household income of $50,000 or higher, and slightly more than one-quarter had a household income between $25,000 and $50,000. Half of the mothers in the sample had a household income between $10,000 and $25,000. Less than 10% of mothers had a household income below $5000. In addition, about one-third of the respondents were stay-at-home mothers at the time of the survey (or reported no occupation) while 22% reported being employed in clerical positions, and 11% in professional occupations and service jobs each.

In Table 1, we also list descriptive statistics for two samples of mothers based on whether a mother experienced the death of a child during the follow-up period. For each variable, we conduct a simple *t*-test where the null hypothesis is that the means are equal across types of mothers. The last column in Table 1 shows the *p*-values for each hypothesis test. Focusing on the *p*-values that are <0.01, we find for several variables that the mean values for mothers who experience the death of a child are statistically different than the means for mothers without the experience. In the data, bereft mothers are less likely to have some college education and are more likely to have some high school or less than a high school education. In terms of age, mothers of a dead child tend to be older than 34 years. Lastly, we see that bereft mothers are more likely to have three or more children in the household at the start of the survey while mothers without a child's death during the follow-up period tend to have only one child in the household.

2.3. Empirical model

All the records in the sample have follow-up information for mother and child, measured in days. The maximum length of the follow-up period is nine years (3288 days) or the time until death if it comes earlier. Given the ages of the mothers in the sample, almost 99% of the observations are right-censored. We use proportional hazard model to estimate the impact of a child's death on the mother's hazard of death (Cox, 1972). In the Cox model, the hazard of death at time t for person i, designated as $h_i(t)$, can be expressed as the product of a baseline hazard common to all people, designated as $h_0(t)$, and a proportional term that is a function of individual characteristics that may or may not be time-variant. The hazard is then defined as

$$h_i(t) = h_0(t) \cdot \exp[X_i\beta + C_i(t)\alpha] \quad (1)$$

The vector X contains a set of time-invariant demographic and socioeconomic characteristics measured at the start of the follow-up period and includes the variables listed in Table 1. The bereavement effect is captured by the coefficient on the indicator $C_i(t)$, which is a time-variant variable that initially equals 0 for all mothers but then turns to the value 1 in all periods after a mother experiences the death of child. The Cox model is estimated as a partial maximum likelihood model, and following standard conventions, we report the hazard ratios (HR) on parameters rather than the parameter estimates themselves. The parameter of interest is the HR on the variable $C_i(t)$ or the value $\exp(\alpha)$. A hazard ratio of 1 would indicate no change in the probability of death for mothers after the death of their child. Estimates were generated using the PROC PHREG command in SAS version 9.2.

3. Results

In this section we present the baseline results for Eq. (1), which are shown in Table 2. Next, we examine the heterogeneity in the results. In Table 3, we examine how the results vary based on characteristics of the child, such as the time since death, the cause of death, plus the age and gender of the child. In Table 4 we examine how the results vary based on characteristics of the mother such as income,

Table 2
Cox proportional hazard models, mothers aged 20–50 at time of survey, United States 1979–1981, publicly available National Longitudinal Mortality Study, hazard ratios (95% confidence intervals).

Covariate	Hazard (95% CI)	Covariate	Hazard (95% CI)
Child dies before mother	2.334 (1.207, 4.514)	Occupation of mother	
Not married	0.953 (0.778, 1.168)	Farming	0.602 (0.298, 1.214)
Hispanic	0.623 (0.460, 0.844)	Pvt. household worker	1.075 (0.667, 1.735)
Black	1.682 (1.382, 2.049)	Service	0.758 (0.606, 0.948)
Other race	1.092 (0.759, 1.572)	Laborer	0.858 (0.406, 1.816)
<HS education	1.303 (0.920, 1.844)	Operator	0.585 (0.434, 0.790)
Some HS education	1.573 (1.168, 2.119)	Operate equipment	0.797 (0.329, 1.928)
High school graduate	1.062 (0.815, 1.384)	Craftswomen	0.567 (0.268, 1.201)
Some college	1.011 (0.755, 1.355)	Clerical	0.759 (0.620, 0.929)
<$5 K	2.356 (1.520, 3.653)	Sales	0.678 (0.447, 1.029)
≥$5 K, <$10 K	1.917 (1.276, 2.880)	Managers	0.789 (0.543, 1.146)
≥$10 K, <$15 K	1.552 (1.048, 2.298)	Professional	0.710 (0.528, 0.955)
≥$15 K, <$20 K	1.516 (1.027, 2.237)	One child	2.987 (1.474, 6.051)
≥$20 K, <$25 K	1.171 (0.790, 1.735)	Two children	2.406 (1.187, 4.880)
≥$25 K, <$50 K	1.121 (0.774, 1.623)	Three children	2.307 (1.126, 4.727)
Age in years	0.768 (0.411, 1.435)	Four children	1.716 (0.800, 3.680)
Age squared × 0.1	1.121 (0.943, 1.333)	−2 log likelihood	17,696.89
Age cubed × 0.01	0.890 (0.762, 1.041)		

Table 3
Heterogeneity in the bereavement effect based on child characteristics, Cox proportional hazard models, mothers aged 20–50 at time of survey, United States 1979–1981, publicly available National Longitudinal Mortality Study.

Covariate	Hazard ratios (95% CI)	−2 log likelihood	p-Value on H_0 that coefficients are equal
Model 1: Child dies before mother	2.334 (1.207, 4.514)	17,696.89	
Model 2: Child dies before mother ×			
≤2 years of a child's death	4.260 (1.905, 9.526)	17,693.53	0.078
>2 years of a child's death	1.225 (0.393, 3.813)		
Model 3: Child dies before mother ×			
Child dies of poisoning/injury	2.309 (1.032, 5.169)	17,696.89	0.963
Child dies of other cause	2.386 (0.767, 7.423)		
Model 4: Child dies before mother ×			
Child is male	2.591 (1.230, 5.460)	17,696.63	0.621
Child is female	1.742 (0.434, 6.991)		
Model 5: Child dies x			
Child ≤19 at time of death	2.536 (1.202, 5.348)	17,696.71	0.683
Child >19 at time of death	1.826 (0.455, 7.331)		

Other covariates include those listed in Table 2.

Table 4
Heterogeneity in the bereavement effect based on mother's characteristics, Cox proportional hazard models, mothers aged 20–50 at time of survey, United States 1979–1981, publicly available National Longitudinal Mortality Study.

Covariate	Hazard ratios (95% CI)	−2 log likelihood	p-Value on H_0 that coefficients are equal
Model 1: Child dies before mother ×			
Household income ≥$20 K	0.820 (0.115, 5.846)	17,694.80	0.218
Household income <$20 K	3.043 (1.510, 6.133)		
Model 2: Child dies before mother ×			
Mother's education ≥some college	2.930 (0.726, 11.8)	17,696.77	0.725
Mother's education <some college	2.207 (1.045, 4.657)		
Model 3: Child dies before mother ×			
Mothers in the workforce	1.538 (0.493, 4.795)	17,695.81	0.313
Mothers not in the workforce	3.152 (1.403, 7.083)		
Model 4: Child dies before mother ×			
Mother has 1 child at time of survey	2.538 (0.814, 7.911)	19,696.86	0.861
Mother has >1 child at time of survey	2.244 (1.000, 5.034)		
Model 5: Child dies before mother ×			
Mother is White	2.882 (1.432, 5.801)	17,695.38	0.285
Mother is not White	0.923 (0.129, 6.603)		
Model 6: Child dies before mother ×			
Mother is married	2.481 (1.175, 5.236)	17,696.79	0.758
Mother is not married	1.935 (0.480, 7.805)		

Other covariates include those listed in Table 2.

Table 5
Examining sample selection and the bereavement effect, Cox proportional hazard models, 5 samples that differ by mothers ages, United States 1979–1981, publicly available National Longitudinal Mortality Study.

Samples	Hazard ratio	95% CI	−2 log likelihood	Observations
1. Mothers ages 20–50	2.334	[1.207, 4.514]	17,696.89	69,224
2. Mothers ages 20–60	2.036	[1.090, 3.802]	23,109.75	73,508
3. Mothers ages 30–50	2.131	[1.059, 4.288]	15,465.09	49,992
4. Mothers ages 30–60	1.855	[0.960, 3.582]	20,826.67	54,276
5. Mothers ages 35–50	2.300	[1.142, 4.632]	12,866.07	33,609

Other covariates include those listed in Table 2.

education, race, marital status and number of other children. Finally, in Table 5, we consider whether the results are sensitive to the sample composition paying particular attention to the age range of mothers included in the analysis sample.

3.1. Baseline estimates

The HRs corresponding to Eq. (1) are reported in Table 2. In the model, the vector X_i has a large number of covariates, including dummy variables for mothers who were not married at start of the survey and Hispanic mothers, 2 covariates for race (with whites being the reference group), a set of 4 dummy variables for education (with college graduates being the reference group), 6 income dummies (with income >$50,000 as the reference group), and 11 occupational dummies (with mother's not reporting any occupation or having never worked as the reference group). There are three covariates for mother's age which include the mother's age in years, the square of mother's age multiplied by 0.1, and mother's age cubed multiplied by 0.01.

The HRs on other covariates are consistent with prior literature. The results suggest that income is protective, with the HRs declining monotonically as incomes rise. Only the coefficients on the four lowest income groups are statistically significant. The HRs are declining in a near monotonic pattern for rising levels of education, and those with some high school education have a statistically significant 60% higher hazard than college-educated women. Black women have a hazard 68% larger than that of white women, which is statistically significant. The hazard ratio associated with the Hispanic variable reproduces the "Hispanic paradox", that this ethnic group tends to be of lower socioeconomic status but has better health and mortality outcomes than their white non-Hispanic counterparts (Franzini et al., 2001; Markides and Eschbach, 2005; Palloni and Arias, 2004).

The HR for the time-dependent variable that is the focus of this paper indicates that the death of a child heightens mother mortality. In Table 2, we estimate an HR of 2.3 (95% CI, 1.1–4.6) on the variable $C_i(t)$, indicating that the hazard of death for a mother doubles after the death of her child. Although we have a large sample, the small number of deaths of children before the death of their mother means that the confidence interval for this HR is wide, but we can still reject the null hypothesis that the HR equals 1 at a p-value of 0.05.

The results for maternal bereavement are much larger than the results for spousal bereavement using the same data set. In a sample of married individuals aged 50–70 from the public-use NLMS, researchers found a statistically significant HR of 1.3 for bereft husbands after the death of their wife and a statistically significant HR of 1.2 for bereft wives after the death of their husband (Espinosa and Evans, 2008). However, this reported statistic does not overlap with the ages of mothers in our sample. To have a more appropriate comparison of the maternal and spousal bereavement effects, we create a sample of spouses between the ages of 20 and 50 using the NLMS and the sample construction methods described in Espinosa and Evans, 2008. The hazard ratio estimate for married women who experience the death of a husband is 1.35 with a 95% CI of 1.01–1.82. Our estimate of the impact of maternal bereavement is larger and may highlight the severity of the grief experienced by mothers after the death of a child.

3.2. Heterogeneity in the maternal bereavement effect

Next, we consider heterogeneity in the bereavement effect along a number of different lines. Given the small number of deaths of children that occur before deaths of mothers, we are limited in our ability to examine this heterogeneity; so we primarily consider broad dichotomous groupings. Let D_i be a dummy variable that equals 1 for the first group and 0 for the second group. In order to examine potential heterogeneity in the bereavement effect across groups, we allow the time time-dependent covariate to vary for the two distinct groups as measured by D_i and $(1-D_i)$ which produces the equation

$$h_i(t) = h_0(t) \cdot \exp[X_i\beta + D_iC_i(t)\alpha_1 + (1-D_i)C_i(t)\alpha_2] \qquad (2)$$

Using this hazard function, the bereavement effects are captured by the coefficients α_1 and α_2. The vector X is the same as it is in Eq. (1).

In Table 3, we consider whether the results vary based on characteristics of the child. In the first row of Table 3, we repeat the hazard ratio for the bereavement effect from Table 2, just for reference. Each subsequent model in Table 3 utilizes the hazard with two time-variant variables.

In model (2) of Table 3 we allow the maternal bereavement effect to vary based on the time since the death of the child. Previous work on the spousal bereavement effect suggests that the risk of death for the surviving spouse is particularly pronounced in the period soon after the death of their spouse, with the bereavement effect decreasing with elapsed time (Kaprio et al., 1987; Schaefer et al., 1995; Lichtenstein et al., 1998; Manor and Eisenbach, 2003; Espinosa and Evans, 2008). We can capture this possibility by allowing the maternal bereavement effect to vary based on the time since the child's death. Specifically, we consider whether the bereavement effect differs in the first two years after the death of a child versus more than two years after the child's death.

Consistent with work on spousal bereavement effect, we find that the impact of maternal bereavement is pronounced soon after the death of a child. In the two years after the death of a child, the HR is 4.26 (95% CI, 1.90–9.53), suggesting a 326% increase in the hazard of mortality during the first two years after the death of a child. We can easily reject the null that this HR is 1. In contrast, the HR more than two years after the death of a child is 1.22; the confidence interval indicates we cannot reject the null that this HR equals 1. These results are suggestive that the bereavement effect is strongest right after the death of a child, and a -2 log likelihood test indicates that we can reject the null hypothesis that these two coefficients are the same at a p-value of 0.10.

These results are consistent with results from the Danish study on maternal bereavement (Li et al., 2003b) which found that mother mortality increases 60% in the first few years following a child's death. Our results that suggest a bereavement effect even after 2 years are consistent with the results from this same study which found persistent elevated mortality for bereft mothers even 10 years after a child's death (Li et al., 2003b). However, the results are in conflict with the Israeli study which found no significant effect at any time after a child's death (Levav et al., 1988).

In model (3), we examine whether the bereavement effect varies based on the cause of death for the child. Previous research on spousal bereavement shows that cause of spousal death has a relationship with the magnitude of the bereavement effect (Li et al., 2003b; Espinosa and Evans, 2008). Given the ages of the children in this analysis, most child deaths are from poisonings and injuries (roughly 70%); so we allow the bereavement effect to vary based on these causes versus all other causes. Dividing the sample in this manner, we find a large bereavement effect for both groups, and there is a slightly larger bereavement effect for deaths from other causes (HR = 2.39 and 95% CI of (0.77, 7.42)) versus poisonings and injuries (HR = 2.31 and 95% CI of (1.03, 5.17)). However, we lose statistical precision so that in only one case, poisonings and injuries, can we reject the null that the HRs equal 1. We cannot reject the null that the bereavement effect is the same across these broad causes of death (p-value of 0.96). These results are suggestive that the bereavement effect is large for both broad classes of causes of death for children.

In model (4) of Table 3, we examine whether the bereavement effect varies by the sex of the dead child. Past research (Werthmann et al., 2010) that has studied the heterogeneity of maternal bereavement based on sex of child finds no evidence of a difference in bereavement effects. Using the NLMS data, we find the hazard ratio for mothers who experience the death of a male child is 2.60 (95% CI (1.23, 5.46)) and who experience the death of a female child is 1.74 (95% CI (0.43, 6.97)). While the bereavement effect is larger for mothers of a dead son, we cannot reject the null hypothesis that the two hazard ratios are equal (p-value = 0.63). Thus, we do not find evidence in the NLMS data that there is a statistically significant difference between the bereavement effects of mothers of dead sons and mothers of dead daughters.

In the final model of Table 3, we examine heterogeneity in the bereavement effect based on the age of the child at the time of death. Past parental bereavement research has conducted similar analysis into the variation of the bereavement effect by the age of the child (Levav et al., 1988; Li et al., 2003b). In the case of Li et al., only children under the age of 18 are in the analysis sample, and their results suggest that the maternal bereavement effect is the same regardless of the age of the child at the time of death. In the case of Levav et al., only adult children are present in the data, and the authors do not find evidence of excess mortality in bereaved parents. In our investigation, we estimate two coefficients for the bereavement effect: one for mothers whose child was 19 or under at the time of death and one for a child who was older than 19 at the time of death. We find a slightly higher bereavement effect for the deaths of younger children with the HR for deaths of children 19 and under being 2.54 and a 95% CI of (1.202, 5.348); so in this case, we can reject the null the HR equals 1. For deaths of children older than 19, the HR is still large at 1.9 but the 95% CI encompasses 1 (0.46, 7.33) which means for this group we cannot reject the null that the HR equals 1. The *p*-value for the test that both coefficients are equal is very large indicating we also cannot reject this null hypothesis. These results are suggestive that the bereavement effect is large regardless of the age of the child at the time of their death.

In Table 4, we consider whether the results vary based on characteristics of the mother. In model (1) of the table, we consider whether the bereavement effect varies by household income, which is measured at the start of the survey. The concern is that mothers with higher income may have access to better health care or other interventions that may reduce the bereavement effect. In this model, the bereavement effects are for mothers with higher income ($\geq$\$20,000) and with lower income ($<$\$20,000). The maternal bereavement effect for higher income mothers (HR = 0.82 and 95% CI of (0.12, 5.85)) is smaller than for lower income mothers (HR = 3.04 and 95% CI of (1.51, 6.13)), and only in the case of lower income bereft mothers can we reject the null that the hazard ratio is 1. We cannot reject the null hypothesis that the two bereavement effects are equal (*p*-value = 0.22). While not conclusive, the results do point to a larger bereavement effect for mothers from lower income households.

In model (2) of Table 4, we allow the hazard ratios for the bereavement effect to vary by the educational status of mothers at the start of the survey. Li et al. (2003b) finds no evidence that the maternal bereavement effect differs based on mother's education. For this study, bereft mothers are pooled into two education groups: mothers with some college or more education and mothers with a high school diploma or less education. The HRs for these two groups of mothers are both greater than 2.0. The HR for higher educated mothers who experience the death of a child is 2.93 (95% CI, 0.73–11.82) while for lower educated mothers it is 2.21 (95% CI, 1.05–4.66). Like the previous models in Table 3, only one hazard ratio is statistically significant, and in this model it is the HR for bereft mothers with no college education. Also, we cannot reject the null hypothesis that the two HRs for the bereavement effects in model (5) are equal. These results are suggestive that the maternal bereavement effect is large for both education groups.

We allow the bereavement effect to vary based on the mother's work history in model (3) of Table 4. Bereft mothers are separated into two groups: those who responded that they work or had worked in a specific occupation (i.e., mothers with occupation) and those who responded that they did not work in one of the occupations listed on the survey or who did not respond to the question (i.e., mothers without occupation). We find that the bereavement effect for mothers without occupations (HR = 3.15 and 95% CI of (1.40, 7.08)) is more than double the magnitude of what we find for mothers with occupations (HR = 1.54 and 95% CI of (0.49, 4.80)). We can only reject the null hypothesis that the hazard ratio is equal to 1 for bereft mothers without an occupation. While the difference in hazard ratios is large, we cannot reject the null hypothesis that the two hazard ratios are equal. Our findings are however suggestive that the bereavement effect is large for both groups.

In model (4) of Table 4, we allow the bereavement effect to vary based on whether the mother reports only one child versus having two or more children. Past research finds larger bereavement effects for mothers with fewer children, but the differences in hazard ratios are not statistically significant (Li et al., 2003b). We find that the mean bereavement effect is large for both groups, but the effect is more pronounced for mothers with only one child (HR = 2.54 and 95% CI of (0.81, 7.91)) versus two or more children (HR = 2.24 and 95% CI of (1.00, 5.03)). As in the previous model, however, we are unable to reject the null that HR is equal to 1 for one estimate—mothers with only one child, but we are able to reject the null for the other estimate—mothers with children. We cannot reject the null that the bereavement effect is the same across these broad family types (*p*-value of 0.75), thus the results are suggestive that the bereavement effect is large for both family types.

In model (5) of Table 4 the bereavement effect varies by the mother's race with one coefficient for bereft mothers who are white and another coefficient for bereft mothers who are not white. The magnitudes of the resulting hazard ratios are quite different. The bereavement effect estimate for white mothers is 2.88 (95% CI (1.43, 5.08)), and for mothers of other races it is 0.923 (95% CI (0.13, 6.06)). We can reject the null hypothesis that the hazard ratio is equal to 1 for white mothers, but not for mothers of other races. The difference in the magnitudes of the bereavement effects is great, but we cannot reject the null that the two are equal (*p*-value = 0.29).

Lastly, in model (6) of Table 4, there are two time-variant covariates that allow us to examine possible heterogeneity in the bereavement effect based on marital status. The hazard ratio for bereft mothers who were married at the start of the survey period is 2.48 (95% CI (1.18, 5.24)) which is greater than the hazard ratio for bereft mothers who were not married, 1.94 (95% CI (0.48, 7.81)). In the case of bereft mothers who were married at the start of the survey, we can reject the null hypothesis that the hazard ratio is 1; however we cannot reject this null hypothesis in the case of bereft mothers who were not married. We also cannot reject the null hypothesis that the two hazard ratios are equal. Similar to past research (Li et al., 2003b), these results suggest that the bereavement effect is large regardless of the marital status of the mother.

The results in Table 4 paint an interesting picture in that in three of the four models that arguably measure socioeconomic status the bereavement effect is smaller for the groups that are generally thought to be disadvantaged (mother's with lower education, non-white mothers, single mothers).

3.3. Sample selection and attenuation bias

As noted earlier, older mothers are more likely to have children who have moved out of the house before the initial survey data is collected for the NLMS. Older mothers are also more likely than younger mothers to have experienced the death of child before entering the survey. In both of these cases, the mother's bereavement would not be captured, and the mother would be incorrectly selected into the sample of mothers who did not experience a child's death. If such cases exist in the data, the excess mortality of bereaved mothers would be attenuated. If these same mothers have children in our sample, there children are likely to be older and hence, they are more likely to have experienced a death. Therefore, as we add older mother's to our sample, we are adding more cases that have experienced the death of a child, but we are also adding observations that have more measurement error in the key covariate of interest.

To examine the sensitivity of our results to the age ranges of the mothers in our sample, we re-estimate the basic results from Table 2 after changing the age restriction for mothers in the sample. If the attenuation bias is present, we expect the hazard ratio estimate of the impact of the death of a child on mother mortality for the ''younger'' samples to be larger than for the ''older'' sample, because the bias in the estimate of the hazard of dying for mothers who have not experienced the death of child is partly mitigated by the exclusion of older mothers.

Following this line of reasoning, we create four samples of mothers, one of which is the main analysis file used in the main text: sample 1 contains moms with ages 20–50; sample 2 with ages 20–60; sample 3 with ages 30–50; and sample 4 with ages 30–60. Using Eq. (1), we estimate the hazard ratio of interest for each of the four samples, and the results are listed in Table 5 along with the related 95% confidence intervals, −2 log likelihood statistics, and number of observations for each sample. Starting with the estimate from sample 2 and moving to the estimate from sample 1, the hazard ratio for the older sample of mothers (2.04) is smaller than for the younger sample (2.33). Comparing sample 4 to sample 3, the hazard ratio for the older sample of mothers (1.86) is smaller than for the younger sample of moms (2.13). These results support our concern that when older mothers are included in the sample there is an attenuation bias.

Another sample selection concern is that a mother may have newborn children after the date the initial survey data are collected, and these children may die during the follow-up period. Such cases would not be captured in the NLMS, and they are more likely to occur to mothers who are younger at the time the survey data was collected. These mothers, then, would be incorrectly identified in the data as never experiencing the death of a child. This issue leads to an overestimate of the hazard of dying for mothers who have not experienced the death of a child which would attenuate the bereavement effect estimated by a proportional hazard model.

Comparing samples where the condition on maximum age remains the same but the minimum age restriction increases, the hazard ratio estimates for the older sample should be larger than for the younger sample if the attenuation bias is significant. Using the results listed in Table 5 and comparing estimates from sample 3 to sample 1, the hazard ratio for the older group of moms (2.13) is smaller than for the younger group (2.33). Similarly, the hazard ratio associated with sample 4 (1.841) is smaller than the hazard ratio associated with sample 2 (2.022). Thus, the findings do not bolster the concern of an attenuation bias for a sample that includes younger mothers.

One theory in psychology of how mothers respond to the death of a child dates back to Poznansky (1972) and is called replacement child syndrome. The theory suggests that bereft mothers may cope with the grief by attempting to become pregnant after the death of a child or by ''replacing'' the dead child with a living sibling. We do not observe any additions to the family after the initial survey; therefore, we cannot determine whether the phenomenon is present. A concern, then, is that bereft mothers have replacement children that may attenuate the impact of bereavement on mother mortality. However, we can eliminate the possibility of the replacement child contaminating our work by reducing the sample to include women aged 35–50, most of which have completed fertility by this time, and therefore there should be little post-mortality adjustment in family size. These results are listed in the final row of Table 5. This sample has 33,609 observations, and the HR on the bereavement effect is 2.30 with a 95% CI of (1.14, 4.63). This is similar in magnitude to the maternal bereavement effect for mothers aged 20–50 suggesting that the replacement child syndrome does not diminish the impact of our results.

4. Discussion and conclusion

To our knowledge, this is the first effort to estimate the maternal bereavement effect using a large, nationally representative US data source. The hazard ratio estimate in the baseline specification (HR = 2.33) suggests that mother mortality increases 133% after the death of a child. This result is larger than that found in a similar study using Danish data, where the hazard ratio was 1.4 (Li et al., 2003b). The confidence intervals for both studies have considerable overlap, so we cannot reject the null that the results across the two studies are the same. Our work and the Danish study run contrary to other maternal bereavement research that uses Israeli data and reports the absence of an elevated mortality risk for mothers after an adult child's death (Levav et al., 1988).

Like most studies on the bereavement, the results in the basic model are potentially subject to an omitted variables bias. If unmeasured characteristics of the mother or child impact the likelihoods of both mother and child dying (e.g., genetic disorders), then the results would overstate the bereavement effect. In the case of maternal bereavement research, Levav et al. (1988) investigates causality by comparing the mortality of mothers bereaving the death of an adult male child in the Yom Kippur War to mothers bereaving the death of an adult male child due to accidents not related to this war. Assuming that mother expectations of child death are greater for children at war, the authors suggest that a causal bereavement effect would result in excess mortality rates for bereft mothers of male children that died due to causes not related to the war. Levav et al. (1988) finds no empiric evidence supporting the causal interpretation of the maternal bereavement effect.

In the case of spousal bereavement, previous research utilized information on the spouse's cause of death to help distinguish between causation and correlation in the bereavement effect (Espinosa and Evans, 2008). That research began with the observation that causes of spousal death such as heart disease, some accidents, and lung cancers are associated with unobserved characteristics that may be shared with the surviving spouse, and other causes of death such as leukemia are not explained by individual characteristics. The authors label the former causes of death informative and the latter causes uninformative, because they do not reveal information about the surviving spouse. If the spousal bereavement effect is simple correlation, we should find heightened mortality for a widow after the spouse dies from an informative death but not an uninformative death. In contrast, heightened mortality after the death of a spouse from either an informative or uninformative death is a result consistent with a causal interpretation. The results Espinosa and Evans report suggest that bereavement does not directly heighten widow mortality, although it does appear to heighten widower mortality.

Cause of child death information is available to our study; so presumably, we could use a strategy similar to the one used for spousal bereavement to help determine the causality of the maternal bereavement effect. However, implementing the same empirical methodology as the previous research is difficult for two reasons. First, the number of child deaths occurring in the data is much smaller than the number of deaths of married people aged 50–70. Second, there is little variation in cause of child death as 70% are due to accident or poisoning, and the lack of variation constrains our ability to identify causes of child death that are not explained by mother or child characteristics.

Despite this limitation, there is some evidence from Tables 3 and 4 that is at least suggestive of a causal interpretation. The results in the table indicate the impact of the death of a child on mother's mortality is the same across various subgroups such as high and low education and type of death. In the case of mother's education, there are large differences in child mortality across groups. The mortality rate for children with mothers who have a high school education or less education is 0.78% versus 0.53% for children with mothers who have some college or more education, a 47% difference in mortality rates. Therefore, if the estimates of the bereavement effect only reflect unobserved heterogeneity, we might expect a larger response in families where mortality is more determined by observed characteristics, a result we do not find. Running counter to this is the fact that we find large qualitative, but not statistical, differences in the HRs for mothers in high and low income families, and not surprisingly, child mortality is correlated with household income. In the main analysis file of mothers aged 20–50, the nine-year mortality rate for children living in low income households (<\$20,000 in 1980) is 0.75% versus 0.64% for children living in high income households (≥\$20,000).

Espinosa and Evans (2008) note that accidents and poisonings are likely correlated with individual behavior that may be shared with the surviving spouse, and Phelan et al. (2004) rate accidents and poisonings as highly preventable causes of death. We take this into account and provide an alternative approach to estimating maternal bereavement effects based on cause of child death in model (3) of Table 3. In this case, the hazard ratio for mothers of children who die of causes other than accidents and poisonings is roughly the same size as these more informative causes of death, possibly indicative of a causal maternal bereavement effect. As noted above, we cannot reject the null hypothesis that the two hazard ratios are equal, but the hazard ratio for these other causes of death is not statistically significant. Our finding is consistent with a causal interpretation but the results are only suggestive.

Finally, if the argument is that there are unobserved factors common to both child and mother that increase mortality for both, the incredibly large mother mortality soon after the death of a child and the large drop off in this effect over time is more consistent with a causal interpretation than an omitted variables bias interpretation. If there is correlation in death rates produced by unobserved heterogeneity, we would not expect the sudden drop off in deaths of mothers in such a short period of time.

Future work should address the issue of unobserved heterogeneity, with a starting point of identifying a larger and more comprehensive data source than the public-use version of the NLMS used in this analysis.

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