Understanding the Mechanisms of Parental Divorce Effects on Child's Higher Education

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Abstract

In this paper we evaluate the degree to which the adverse divorce effect on the child's higher education outcome operates through deprivation of economic resources. Using one million siblings from Taiwan, we first find that parental divorce occurring at ages 13-18 led to a 10.6 percent decrease in the likelihood of university admission at age 18. We then use the same sibling sample to estimate the effect of parental job loss occurring at the same ages, and use the estimated job-loss effect as a benchmark to indicate the potential parental divorce effect due to family income loss. We find the job-loss effect very little. Combined, these results imply a minor role played by reduced income in delivering the parental divorce effect. Thus, non-economic mechanisms, such as psychological and mental shocks, are more likely to dominate. Further examinations show that boys and girls are equally susceptible, and younger teenagers are more vulnerable than the more mature ones, to parental divorce.

JEL: I20, J12, J64

Key words: Parental divorce, parental job loss, higher education

1 Introduction

Research across economic and demographic fields has consistently found that parental divorce adversely affects children's education outcomes. Little evidence, however, has been provided to shed light on the underlying mechanisms behind these effects. Theoretically, parental divorce may affect education through reduced economic resources that limit the investment in children's education (the economical mechanism), or through diminished psychological well-being of children (the psychological mechanism). Evaluating the relative importance of the two mechanisms connotes crucial policy implications. If the economical mechanism dominates, a provision of financial support to single-parent families should be remedial; however, if the psychological mechanism dictates, professional counseling and other supportive services should be more productive than financial aids. Empirically, however, estimating the two effects separately is very challenging because data rarely provide orthogonal variations in income and psychological shocks caused by parental divorce. Indeed, we know of no rigorous study that distinguishes between the two mechanisms.

This paper fills this evidence gap by investigating the extent to which the adverse divorce effect operates through loss of economic resources. The goal of the paper is twofold. First, using a sample of around one million siblings in Taiwan, we estimate the effect of parental divorce occurring at the age of 13 to 18 on the likelihood of university admission at the age of 18. We employ the mother fixed-effects (FE) model to mitigate selection. Second, we then move on to investigating the extent to which the adverse divorce effect operates through loss of economic resources, which has been widely considered as a major mechanism that delivers the divorce effect. To this end, we use the same sibling sample to estimate the effect of *parental job loss* (due to firm closure) on

university admission. The job-loss effect is then used as a benchmark to assess the extent to which the adverse divorce effect was caused by family income reduction.

For such a benchmark to be valid, however, it needs to meet at least two requirements. First, teenagers who experienced parental job loss in our sample are similar to those experiencing parental divorce. This is to ensure that any difference in the two estimated treatment effects was not caused by any difference in characteristics between the two treated populations. In Section 4.2, we show the similarity between the two treated groups. Furthermore, in Section 5.5 we simultaneously estimate the effects of both parental divorce and parental job loss in one regression, controlling for a rich set of covariates. We find that both of the treatment effects estimated from the pooled regression are virtually identical to the effects estimated from the separate regressions.

Second, both treatments should cause a significant family income reduction, and the psychological impact on children caused by job loss is little or at least less severe as that caused by divorce. In Section 4.2, we exploit additional data to show that both divorce and job loss lead to a significant income loss. We also provide evidence showing that job loss due to firm closure did not predict family dissolution; the implied psychological impact on children is therefore unlikely to be comparable to that caused by parental divorce.

Parents' decision on divorce is potentially endogenous. In this paper we adopt two strategies to address the endogeneity. First, we estimate the mother FE model, which relies on comparing the university admission outcomes between siblings who experienced parental divorce before age 18 and those who experienced it after age 18. Under Taiwan's education system, the age of 18 is the expected timing for individuals to make the first attempt for university admission by taking the national university entrance test. While the within-mother comparison removes confounders that are invariant across siblings, it

cannot fully eliminate the OVB if the confounders vary across siblings. Thus, our second strategy is to follow Sigle-Rushton et al. (2014) to control for a set of idiosyncratic characteristics in our mother FE model. These include birth weight, birth parity, year of birth, county at birth, mother's age at birth, and gender.

Another concern about the validity of our FE model arises because the expected timing for the child to take the university entrance test might be an important consideration for parents contemplating divorce. For example, parents might strategically postpone divorce until a specific child or all their children complete the test to avoid any negative impact on test performance. In either case, our FE model is unable to return an unbiased estimate of the divorce effect. In Section 3.3 we show that the monthly divorce rate does not exhibit any sharp change between right before and right after the children′s test dates. In addition, the demographic characteristics of divorced parents are balanced between right before and right after the test dates. Both findings imply little parental manipulation of the divorce timing.

When estimating the job-loss effect, we follow Charles and Stephens (2004), Doiron and Mendolia (2012), Eliason (2012), and Chen, Liu, and Wang (2016) to exploit firm closure as the variation source of job loss. We apply the same mother FE model used for estimating the parental divorce effect to estimate the job-loss effect. Here, our FE model exploits the difference in university admission between siblings who experienced parental job loss before age 18 and those who experienced it after age 18.

Our estimation results suggest that children who experienced parental divorce at ages 13 to 18 suffered a 10.6 percent decrease in the likelihood for university admission at age 18. However, parental job loss occurring at the same ages did not have any negative effect as such. Combined, the two findings imply that the adverse divorce effect was unlikely to be driven by reduction of economic resources for children. Rather, non-economic

mechanisms, such as psychological trauma, might have played a more important role. Further evidence suggests that younger adolescents are more vulnerable to parental divorce: Every 100 days younger in the age at parental divorce implies a 1.2 percent decrease in the likelihood of university admission at age 18. Finally, we estimate the parental divorce effects for sons and daughters separately; the gender difference is found to be little.

Our findings make several contributions to the literature on the effects of marriage dissolution or family structure on children's outcomes. The principal contribution is the adding of causal evidence regarding the divorce effect on children's higher education outcomes. Previous studies tackling endogenous divorce mainly used three different strategies – (1) a family or mother fixed-effects model that relies on within-sibling comparison (Ermisch and Francesconi, 2001; Page and Stevens, 2004; Bjorklund and Sundstrom, 2006; Bjorklund et al., 2007; Steele et al., 2009; Francesconi et al., 2010), (2) a differences-in-differences method with the design relying on amendments of marriage law (Gruber, 2004; Cáceres-Delpiano and Giolito, 2012), and (3) an instrumental variable (Frimmel et al., 2016).[1] Our findings add to the existing evidence in showing significant effects of parental divorce on children's outcomes, although some prior studies found little or insignificant such effects.[2]

To our knowledge, we are the first to assess the weight of the economical mechanism behind the adverse divorce effect on children's education outcome. The opportunity to make this assessment is granted by our unique administrative data that provide both marriage and job history of the parents in our sample. Our finding of non-negative effects

[1] One strand of studies estimated the effect of parental death on children and considered the effect as a benchmark for exogenous parental absence with a bereaved background. See Corak (2001), Lang and Zagorsky (2001), Fronstin, Greenberg, and Robins (2001), and Amato and Anthony (2014).
[2] Examples are Lang and Zagorsky (2001), Bjorklund and Sundstrom (2006), Bjorklund et al. (2007), and Francesconi et al. (2010).

of parental job loss casts a doubt on the effectiveness of financial aids that are targeted toward single parent families. Alternative measures, such as professional counseling and other supportive services, may be more productive.

We are also the first to meticulously examine parental selection on the timing of divorce, which poses a threat to the validity of the family or mother FE model. Previous FE studies ignored or just mentioned the potential selection without carrying out any empirical analysis to assess the extent of such a threat. Our finding of little parental manipulation over the timing of divorce lends support to the validity of the FE models.

This paper is also related to a prior literature that examines whether the exact age at which a child experienced parental divorce matters. The age at parental divorce should be a vital factor determining the impacts, but the existing statistical evidence is extremely scarce. As rare examples, Kravdal and Grundy (2019) and Ermisch and Francesconi (2001) both found that preschool ages are more vulnerable than later ages to the shocks from parental divorce. Focusing on teenagers, our findings are generally consistent with these earlier works by showing that younger adolescents are exposed to more impairment caused by parental divorce.

The rest of the paper proceeds as follows. Section 2 overviews the university admission system and the laws concerning divorce in Taiwan. Section 3 describes the data, sample selection, and sample statistics. Section 4 introduces the mother FE model and then discusses its validity. We also provide evidence to justify the comparison of the two treatment effects in this Section. Section 5 presents estimation results and examines the robustness. Finally, we draw some conclusions in Section 6.

2 Background

2.1 Higher education in Taiwan

In Taiwan, students begin their primary school education at the age of six. Primary school takes 6 years to complete, followed by junior high school that takes 3 years. Thereafter, the path goes into two tracks – academic versus vocational. Academic-track students study at an academic senior high school for three years, whereas vocational-track students study at a vocational high school for three or five years.

The sole starting date for primary schools in Taiwan is September 1st. A child is not permitted to enter elementary school until reaching the 6th birthday prior to September 1st of the relevant year. Throughout this paper, the age of an individual is measured by *years since the date of birth*, rather than the calendar year. An individual born in year y is considered to be n years old during the period from the nth birthday to the date just before the $(n+1)$th birthday.

During the third year of academic high school, students can make two attempts for university admission through taking two different university entrance tests.[3] The first is the General Scholastic Ability Test (GSAT), which is scheduled in late January or early February during the third year of high school; the second is the Advanced Subjects Test (AST), which is scheduled for July 1 to July 3 of the same year. All third-year high school students are required to take the GSAT, but only a portion of them will qualify for the first round of university admissions, which are based on GSAT scores and high school performance. Unsuccessful applicants and those who were not satisfied with their admission outcomes move on to taking the AST in July for the second round of admissions, which are exclusively determined by the AST scores. Each student, admitted or not, is free

[3] Very few students who graduated from vocational high school were admitted to university.

to take both tests again and apply for admission in the following year and any later year.

Given the timing of the elementary school starting date and the age requirement for starting school, the age of 18 is the *expected* age for an individual to make the first attempt for university admission. To see this, we consider children born between September 1979 and August 1980. This cohort was expected to start elementary school in September 1986, after they reached the age of 6. If they followed the academic track in high school, they would have been expected to take the GSAT and AST for the first time in 1998 and, if successful, been enrolled in a university when they were age 18.[4]

To make within-sibling comparisons, we need a clear cutoff date to identify whether a child experienced parental divorce before or after the completion of the university application process. For the purpose of this paper, we consider the timing of the AST (July 1-3) at the *expected* age for university admission as the cutoff. It is a better cutoff than that of the GSAT (late January or early February) for two reasons. First, the number of students admitted through the AST is significantly more than those admitted through the GSAT. Luoh (2018) shows that, prior to 2002, over 90 percent of all fresh university students were admitted through the AST. The proportion decreased but remained high in 2002 (78.4%) and 2003 (80.7%).[5] Because our sample consists of students taking the GSAT and AST between 1998 and 2003, these statistics imply that the AST was the major university admission channel for students in our sample. Second, regardless of the first-round admission outcome after taking the GSAT, all students, admitted or not, are free to move on to take the AST. For all students, therefore, the period between the two tests

[4] To be sure, the expected age for university admission might be different from the *actual* age for university admission due to an early start (entering elementary school before reaching age 6), a late start (entering elementary school after reaching age 7), grade retention, or grade skipping. These cases are uncommon in Taiwan. The proportion of early starters during 1999 to 2006, for example, was 2.42 percent, whilst late starters accounted for only 2.20 percent during the same period of time.

[5] The GSAT was established in 1994, but it was operating as a test run before 2002.

remains susceptible to parental divorce. The dates of the AST provide a 'clean' cutoff that marks the end of any potential divorce effect on university admission.

2.2 Divorce in Taiwan

Traditionally, married couples in Taiwan were reluctant to divorce due to common beliefs in Confucianism, which emphasizes family values. However, the crude divorce rate increased from 1.01 in 1984 to 2.51 in 2010, which was already higher than the average crude divorce rate among all OECD countries around 2010.[6]

Divorce in Taiwan takes two forms – consensual and court-granted. Divorce by mutual consent is made in writing, witnessed by two people, and registered with the Household Registration Bureau. No court action or cooling-off period is required for mutual-consent divorces. Any couple failing to reach an agreement on divorce can resort to the court system. However, divorce will only be granted by a court on grounds such as bigamy, adultery, and intolerable ill-treatment and humiliation. Compared to consensual divorce, court-granted divorce is uncommon. In 2010, for example, the number of court-granted divorces (4,898 cases) accounts for only 8.4% of the entire divorce cases (58,118).

3 Data and samples

3.1 The four administrative datasets

We use four different national administrative datasets. The subjects are males and females born between September 1980 and August 1985. These subjects were linked to their parents' divorce record, and siblings are identified by mother's citizen identification (ID) number. All four administrative datasets are population wide and thus are virtually free

[6] The statistics are from the Department of Statistics in the Ministry of the Interior, Taiwan. The crude divorce rate is the number of divorces per 1,000 mid-year total population during a given year.

from the problem of sample attrition. This constitutes a strong advantage of administrative data over longitudinal survey data, which often suffer from serious attrition.

The first administrative dataset is Birth Registry records, ranging from 1978 to 1999. The dataset reports ID number, birth date, birth parity, birth type (singleton or multiple births), birth weight (in grams), and both parents' ID numbers and characteristics at the time of the childbirth, including age, marital status, education level, and residential county. We link siblings by their mother's ID.

Our second dataset is the University Entrance Test records from 1998 to 2003. They report ID, GSAT and AST scores, and, if successfully admitted, the name of the university admitted to. We are interested in two margins of higher education outcomes – admission to *any* university at age 18, and admission to a *public* university at age 18 (we use two separate dummy variables to indicate the two different outcomes). In Taiwan, with better education quality and more financial resources, public universities are more selective than private ones. We utilize these two outcome variables to shed light on the effects of parental divorce on the likelihood of being admitted to any university at all, as well as the likelihood of being admitted to a university of better quality.

We obtain information of parental divorce from Divorce Registry records (1998-2003), our third administrative dataset. The records report each divorced person's ID number, date of divorce, date of marriage, and date of birth. Using this dataset, our treatment variable is defined as a dummy variable indicating the experience of parental divorce prior to the AST at age 18 ($= 1$) or otherwise ($= 0$). Given the time windows of our data, 13 is the youngest age at which a child in our sample experienced parental divorce.

The last set of administrative data is the Unemployment Insurance (UI) payment records (1998-2003). The dataset reports ID number, the exact date, and three different

job-loss reasons that qualify a worker for the UI – firm closure (due to relocation, suspension, or bankruptcy), layoff, and involuntary job leave due to other reasons. To reduce the concern about endogenous job loss, we limit our job-loss indicator to only cases due to firm closure, and remove cases due to layoff and other reasons from the sample. We use these data to define another treatment variable as a dummy variable indicating whether an individual experienced parental job loss prior to the AST at age 18 ($= 1$) or otherwise ($= 0$).

3.2 The sibling sample

We restrict our sample to siblings born between September 1, 1979 and August 31, 1985. These individuals were selected because they were expected to take the university entrance tests for the first time between 1998 and 2003, the time window of the University Entrance Test records.

In order to ensure that our estimates could be more readily interpreted, our sample excludes families with any sibling born to an unmarried mother (0.5%), families with half-siblings born to different fathers (2.5%), and all step-siblings. In our sample, half-siblings born to different mothers are considered as belonging to different families. Our sibling sample is finalized with 1,073,833 individuals from 481,459 families.

We carry out the main estimation using a pooled sample that combines both male and female siblings, with a gender dummy being included in the regression as a control to allow the education outcome to differ between genders. Later in Section 5.5 we estimate the parental divorce effects separately for males and females.

3.3 The Endogeneity of Parental Divorce

Parents may choose the timing of divorce. Particularly, parents might deliberately delay

divorce until the child completes the AST to avoid any negative influence on the child's performance. In such case, the timing of parental divorce relative to the expected AST dates is endogenous, and the resulting selection bias cannot be removed by within-sibling comparison. To gauge the degree of such selection, we examine whether parental divorce risk increased shortly after the child completed the AST. Such an increase signals parental selection. We use our sibling sample to calculate monthly divorce rates, measured as the number of divorce cases per 100,000 individuals in a month, over the entire sample period (from 1998 to 2003). We adjust the divorce rates by multiplying the rates by 30/31 for 31-day months, by 30/28 for 28-day months, and by 30/29 for 29-day months. To remove the seasonality and time trend of divorce rates, we obtain the residuals from regressing the adjusted divorce rates on the full set of calendar-month fixed effects and a time trend variable measured in months to the AST dates at age 18.

Figure 1 presents the residual divorce rates, with $t = 0$ indicating July of the year the child is expected to take the AST. The following months, August, September, and October, are respectively indicated by $t = 1$, $t = 2$, $t = 3$, and so on; the months prior to July – June, May and April – are respectively indicated by $t = -1$, $t = -2$, $t = -3$, and so on. The curve does not exhibit any sharp increase in the residual divorce rate right after $t = 0$. If anything, the rate follows a mild increasing trend before $t = 0$, and a deceasing trend after $t = 0$.

Furthermore, we examine the mean of observable parental characteristics right before and right after $t = 0$, using all divorced parents in our sample. Any significant change in these characteristics right after $t = 0$ implies that parents with particular characteristics are able to manipulate the timing of divorce around $t = 0$. The examination results are presented in Figure 2, where Figure 2-A shows the average ages for parents who divorced between $t = -12$ and $t = 12$. The two curves show that both

mother's (blue curve) and father's (red curve) ages at the time of divorce follow a smooth, increasing pattern throughout the time period. No significant change can be detected right after $t = 0$. Figure 2-B presents the proportions of mothers (blue curve) and fathers (red curve) with a high school diploma. Although both curves fluctuate across months, neither curve exhibits any obvious change right after $t = 0$. Finally, Figure 2-C shows no abrupt change in the proportions of parents living in urban areas at $t = 0$.

Combined, Figures 1 and 2 suggest that the expected timing for the child to take the AST is unlikely an important consideration when parents are contemplating divorce. At least within the time window between one year before and one year after expected timing of the AST, parents do not seem to strategically delay divorce.

3.4 Summary statistics

We present the descriptive statistics of our sibling sample in Table 1. Columns 1 and 2 show the means and standard deviations of variables in the entire sample. The observations are of 1,073,833 children, both males and females. Columns 3 and 4 present the corresponding statistics for children experiencing parental divorce prior to age 18, which amounts to 22,811 individuals or 2.1 percent of the sample.

Column 1 suggests that around 14.5 percent of the children in the sample were admitted to a university at age 18 and 5.4 percent were admitted to a public university at age 18. Both figures are much lower for children experiencing parental divorce: 8 percent and 2.7 percent (column 3), respectively. For the entire sample, around 1 percent of the children were exposed to parental job loss due to firm closure during the sample period. The proportion is nearly doubled (1.8 percent) for children experiencing parental divorce. For other variables, a comparison between columns (1) and (3) shows that children exposed to parental divorce are more likely to live in urban areas, have younger and less

educated parents, and be higher order births, but they are not much different from the main group in terms of sex ratio, birthweight, year of birth, and family size.

4 Empirical strategy

4.1 The mother fixed-effects model

Our regression is specified as follows:

$$Y_{ijt} = \alpha_1 + \beta_1 D_{ijt} + \boldsymbol{X}_{ijt}\gamma_1 + \boldsymbol{F}_j\delta_2 + a_j + u_{ijt}, \quad (1)$$

where Y_{ijt} is a dummy variable, taking the value of one if individual i of family j is admitted to any university (or a public university) at age 18 in calendar year t ($t = 1998, 1999, \ldots, 2003$). $D_{ijt} = 1$ indicates that individual i experienced parental divorce prior to the last day of the AST (July 3) at age 18. $\boldsymbol{X}_{ijt}$ is a vector of idiosyncratic characteristics comprised of gender, birthweight, birth parity, and school year of birth.[7] $\boldsymbol{F}_j$ is a vector of observable family-level characteristics, including parental characteristics (mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term) and household variables (county of residence and number of siblings). a_j measures the unobserved and invariant characteristics across siblings born to the same mother. Finally, u_{ijt} is the individual specific error term.

We start with estimating Equation (1) using the OLS model. The OLS estimation relies on comparing those who experienced parental divorce prior to the AST at age 18 and those without such experience. The coefficient of interest is β_1, which indicates parental divorce effect. We undertake two different specifications of the OLS regression.

[7] A school year in Taiwan operates from September 1 to August 31 in the next year. A child born in school year t is expected to start schooling in year $t + 7$. We control for school year rather than calendar year because the former is more relevant to yearly university admission rates.

The first is a baseline regression which only controls for the individual characteristics $\boldsymbol{X_{ijt}}$ (hereafter, the 'OLS' model). The second is the regression extended to control for both the individual characteristics $\boldsymbol{X_{ijt}}$ and the family-level characteristics $\boldsymbol{F_j}$ (hereafter, the 'OLS+controls' model). A comparison of the estimates of β_1 obtained from the two OLS regressions sheds light on the impact of the observed family-level controls on the estimated divorce effects.

Next, we estimate Equation (1) using the mother fixed-effects model (hereafter, the FE model), and compare the FE estimate to the OLS estimates. The FE estimation relies on comparing siblings who experienced parental divorce *before* versus *after* the AST at age 18. The advantage of using the FE model is that a_j of Equation (1) will be differenced out by the within-mother comparison. Thus, to the extent that all confounders are at the family level and invariant across siblings, the FE model produces an unbiased estimate of β_1.

One potential limitation of our mother FE model is that some confounders may vary across siblings, resulting in the FE regression not fully eliminating the omitted variables bias. In addition, there might be time-varying factors, and the two treatments of interest (parental divorce and job loss) might associate with different such factors. To mitigate this concern, we follow Sigle-Rushton et al. (2014) to control for a set of individual characteristics described as $\boldsymbol{X_{ijt}}$. Among those controls, birth weight and gender are intended to capture the differences in health endowment between siblings; birth parity and gender are potentially important determinants for differences in family resources (finance, nutrition, and parents' quality time) allocated to different siblings; school year of birth is designed to measure differences in yearly university admission rates, which mainly depend on the population size of the relevant school year cohort. To be sure, this strategy is unable to address *unobservable* time-varying factors that affect siblings' outcome

differently.

The OLS and FE estimates of β_1 may be different for the following reasons. First, the OLS estimates may suffer from omitted variables biases caused by confounders that are invariant across siblings, whilst the FE estimate is unaffected by such confounders. Second, parental marriages usually dissolve after a long period of marital discord, which likely has an adverse impact on children′s education outcomes. If so, both our OLS and FE estimates capture such a marital discord effect, but the FE estimate takes the effect to a lesser degree. To see this, consider a family of two children whose parents divorced when the first child was 19 years old and the second child was 17 years old. Further assume that the divorce put an end to a bad relationship that lasted for 3 years. Before attaining the age of 18, the second child experienced 3 years of parental marriage discord, whilst the first child was only exposed to the first 2 years of the discord. Our OLS estimates, which compare education outcomes between individuals with and without experiencing parental divorce prior to age 18, are confounded with the entire 3 years of marital discord effect. Alternatively, by comparing siblings who experienced parental divorce before and after age 18, the FE estimate captures only the first 2 years of the marital discord effect. Making a precise assessment of the effect from pre-divorce parental conflicts is beyond the scope of this study. However, since the FE estimate captures such effects less, it is expected to be less biased than the OLS counterparts.

To estimate the effects of parental job loss on children's university admission, we simply replace the covariate of interest (D_{ijt}) of Equation (1) for a dummy variable indicating that individual *i* experienced parental job loss prior to July 3rd of the year the individual turns the age of 18 ($D_{ijt} = 1$) or otherwise ($D_{ijt} = 0$). We will first estimate the effect without distinguishing between maternal and paternal job loss; namely, D_{ijt} is set to be 1 when either mother or father was laid off. We then move on to separately

estimate the effects brought about by maternal and paternal job loss, using two different dummy variables to indicate the two cases.

4.2 Comparison of the job-loss and divorce effects

The second goal of the paper is to evaluate the degree to which the divorce effect operates through family income loss. As we have mentioned earlier, our strategy focuses on a comparison of the parental divorce effect and the effect of another treatment – parental job loss, which will be shown in this section to cause a severe and long-lasting reduction of family income. The job-loss effect is used as a benchmark to shed light on the role of family income fluctuation in determining the university admission outcome of children who experienced parental divorce.

In this section we conduct two examinations to support the legitimacy of using the job-loss effect as a benchmark. First, we show that children who experienced parental job loss are relatively similar to those experiencing parental divorce in terms of observed characteristics. The similarity alleviates the concern about context specificity of the two treatment effects, and thus enhances the comparability. Columns (3) and (5) present the means of a set of characteristics for children experiencing parental divorce prior to age 18 and those experiencing parental job loss prior to age 18, respectively. The two groups are very similar in gender composition, birth weight, age (reflected by year of birth), and number of siblings they have. The children who experienced parental divorce tend to have younger mothers and less educated parents than those who experienced parental job loss, but the differences do not seem substantial: The mean maternal age at birth is 21.8 for the former and 23.6 for the latter; the proportions of fathers attaining college or professional education are respectively 3.0 and 1.3 percent for the former, compared to 4.3 and 2.1 percent for the latter. The proportion of fathers acquiring a high school (academic or

vocational) diploma is 27.7 percent for the former, compared to 34 percent for the latter.

Out of caution, we simultaneously estimate both treatment effects in a single regression, which controls for a rich collection of individual and family characteristics. The two treatment effects estimated from this regression, therefore, are marginal effects holding constant all the characteristics. The regression results, as presented later in Section 5.5, suggest that the two effects estimated from the pooled regression are virtually the same as their counterparts estimated from the separate regressions.

Next, we examine the degrees of income loss caused by parental divorce and parental job loss (due to firm closure). We start with evaluating the adverse effects of firm closure on employees' earnings. First, firm closure leads to unemployment, and the resulting income loss can only be partially offset by Unemployment Insurance benefits during unemployment.[8] Second, when a laid-off worker finds a new job, the reemployment wage is often lower than the wage prior to firm closure. Figure 3 presents the effects on these two different margins. Using the sibling sample, we observe all the parents who experienced a job loss due to firm closure and track their employment status and wage for 6 years after the firm closed. Because our data only report labor market outcomes at the end of each calendar year, for those who lost their jobs at any time during year *t*, we calculate the average employment rate and wage at the end of years $t, t+1, t+2, \ldots,$ and $t+5$. Here, we mark these years as year 1, year 2, …, and year 6, respectively.

Figure 3-A shows the average reemployment rates for fathers (blue curve) and mothers (red curve), separately. The blue curve suggests that, at the end of year 1 only 70 percent of the laid-off fathers were reemployed. The rate kept decreasing in the following 5 years, showing no sign of bouncing back. For laid-off mothers, 67 percent found a new

[8] The Unemployment Insurance program provides 60 percent of the wage prior to the job loss. The benefits terminate at the end of the first six months of unemployment or whenever the workers are reemployed.

job by the end of year 1. The proportion climbed back to 72 percent in year 2 and further to 75 percent in year 3, then started declining for the remaining years. These dynamic patterns indicate a severe, long-lasting family income loss caused by firm closure. The picture is even dimmer if we weigh in reemployment wage. Figure 3-B presents laid-off fathers' and mothers' reemployment wages from the year they were reemployed to year 6.[9] The blue curve suggests that the reemployed fathers suffered a NT$10,000 (35 percent) reduction from their monthly wage just prior to firm closure. Worse, the average wage continued to fall in the following 5 years. The wage loss for laid-off mothers (red curve) is less severe and seems mitigated over time, but the reemployment wage failed to return to the prior-to-firm-closure level even at the end of year 6.

It is rather difficult to precisely measure the income loss resulting from parental divorce, given that our main data lack income, assignment of custody, alimony obligation and enforcement, and any instrument for parental divorce. Thus, we resort to additional data from the *Survey of Family Income and Expenditure* (SFIE), conducted annually based on a sample of randomly selected 13,000-16,000 households (or about 52,000-68,000 individuals). The SFIE reports socio-demographic characteristics (including marital status) and income of each member of the sampled households. For our sample period 1998-2003, we obtain teenagers aged 13 to 18 and categorize them into three different groups according to their parent's marital status being (1) divorced mother, (2) divorced father, or (3) married. We then calculate each household's per capita income adjusted by equivalence scale, which assigns a value of 1 to each household member aged 18 or older and a value 0.6 to each member younger than 18. We find that the average equivalence-scaled income per capita is NT$313,733 (standard deviation is NT$204,252) for children

[9] Note that the observations here are not panel-balanced: Reemployed workers enter the sample on a rolling basis, starting at the time of reemployment.

living in an intact family, which is 12.8 percent higher than the counterpart for children living with the divorced mother (NT$278,127), and 21.9 percent higher than the figure for children living with the divorced father (NT$257,336).[10] Although these income differences do not indicate the causal effects of parental divorce, they do suggest that children living with a divorced parent had significantly lower economic resources than what children living with married parents had.

We now turn to the potential impact of parental job loss on the child's psychological well-being. Theoretically, the impact can be both positive and negative. On the one hand, the financial and mental pressure caused by job loss may intensify conflicts among family members and increase the risk of family dissolution. On the other hand, laid-off parents may be able to spend more time with their children, which may have positive effects on the children's development and well-being. Moreover, children exposed to family adversity may learn to take more responsibility for the family and turn more serious about school education and time use. The combined effect, therefore, is ambiguous and may differ across children with different characteristics.[11]

Because our data lack any information to measure psychological status, we are unable to estimate the psychological impact of parental job loss. Alternatively, we take a glance at this potential impact by examining whether parental job loss is associated with a higher risk of parental divorce. Obviously, divorce indicates the existence of conflict in the family, which adversely affects children's psychological well-being. We use parents in our sibling sample to run an OLS regression of divorce on past experience of job loss due to firm closure. The estimation results are presented in Table A1 in the Appendix. Column 1 shows

[10] The numbers of observation are 34,926, 1,224, and 1,169 for the intact family group, divorced-mother group, and divorced-father group, respectively. All incomes are deflated to 2001 value using annual CPI.
[11] For example, Nikolova and Nikolaev (2018) have recently shown that individuals who experienced parental job loss at ages 0-5 reported a lower level of life satisfaction at ages 18-31. The effect, however, was little or even positive for those exposed to parental unemployment at ages 6-10.

that, when no covariates are controlled for in the regression, experiencing job loss is associated with a 0.103 percentage point increase in the probability of subsequent marriage dissolution. The coefficient is minimal and statistically insignificant. In columns 2 and 3, we present the estimates when we add individual-level and family-level controls to the OLS regressions.[12] These results do not support the hypothesis that parental job loss due to firm closure predicts marriage dissolution. To be sure, we are unable to rule out other channels through which parental job loss would impede children's psychological well-being. However, such potential impact is unlikely to be comparable to that caused by parental divorce.

To summarize, the evidence above supports the hypotheses that both parental job loss and parental divorce caused a significant reduction of family income, and that the psychological impact of parental job loss on teenagers in our sample was unlikely to be significant. These findings strengthen the validity of our empirical strategy to use the parental job-loss effect as a benchmark to assess the extent to which the divorce effect operates through reduction of family income.

5 Results

5.1 The effects of parental divorce

Table 2 presents the OLS, OLS+controls, and FE estimates of the divorce effects on children's university admission obtained from estimating Equation (1). In column 1, when only individual characteristics are controlled for, the OLS estimate of the divorce effect (represented by β_1 in Equation (1)) suggests that experiencing parental divorce prior to

[12] Individual controls are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings.

the AST at age 18 is associated with a decrease of 7.68 percentage points (ppts) in the likelihood of university admission at age 18. The OLS estimate greatly decreases to 4.82 ppts when parental characteristics and household variables are added to the controls, as shown in column 2.[13] Column 3 presents the mother FE estimates. The estimate of β_1 suggests that parental divorce leads to a 1.56 ppt decrease in the odds of university admission at age 18. Although the FE estimate is markedly smaller than its OLS counterparts, it is statistically and economically significant, accounting for a 10.8 percent decrease from the average university admission rate (14.5 percent) for our entire sample.

The difference between the OLS and FE estimates implies that the negative correlation between parental divorce and education outcomes estimated by the cross-sectional comparison may be seriously biased by unobserved family-level confounders. Another implication is that controlling for a rich set of individual and family characteristics in the OLS+controls model is unlikely to return an unbiased estimate of the divorce effect.

The results for public university admission paint a similar picture. The OLS estimate in column 4 suggests that experiencing parental divorce prior to the AST at age 18 caused a decrease of 3.13 ppts in the likelihood of admission to any public university. The estimate declines to 1.85 ppts when more controls are added to the regression (column 5), and further reduces to 0.86 ppt as estimated using the FE model (column 6). All three estimates are statistically significant. The FE estimate accounts for a 15.9 percent decrease from the average admission rate (5.4 percent) for public universities.

[13] Individuals characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings.

The FE estimates in columns 3 and 6 of Table 2 suggest that the negative impact of parental divorce on a child's higher education outcome may be twofold: it impeded not only the admission to any university, but the admission to a university with better education quality and more abundant resources. Given the recent evidence suggesting a positive return to university selectivity (Hoekstra, 2009; Dale and Krueger, 2014; Li et al, 2012; MacLeod et al., 2017), our finding of the twofold effect of parental divorce is particularly alarming.

5.2 The role of income loss

Table 3 presents the estimated effects on university admission of parental job loss due to firm closure. For brevity, we present only the OLS+controls and FE estimates. Ignoring the OLS estimates without additional controls does not suffer a loss because Table 2 (Section 5.1) has shown that such OLS estimates are not very informative. In Table 3, column 1 shows the coefficient estimate of a dummy variable indicating whether the child experienced parental job loss prior to the AST at age 18 $(= 1)$ or not $(= 0)$, without distinguishing between maternal and paternal job loss. The estimate suggests that being exposed to prior-to-AST parental job loss is associated with a significant decrease of 0.96 ppt in the likelihood of university admission at age 18. The mother FE estimate is presented in column 2. In sharp contrast with the OLS+controls estimate, there is a positive FE estimate of 0.83 ppt, but it is statistically insignificant.

The difference between the OLS+controls and FE estimates implies that parental job loss due to firm closure might not be exogenous. Indeed, it is plausible that individuals who experienced firm closure may have different characteristics than those without such experience. Before a firm reaches the shutdown point, it might have experienced hardship in business and even suffered a long-run loss in profit. Naturally, the employees will catch

the drift of the firm's dim future, but they may respond in different ways. Those who have better outside opportunities might choose to find another job before the firm closure; the remaining workers stay until being laid off at the time of closure. If these two types of workers are different in unobserved characteristics that affect adversely their children's education, our estimated OLS effect of parental job loss is potentially biased. This highlights the advantage of our mother FE model that removes family-level confounders and controls for a rich set of idiosyncratic characteristics of siblings.

Next, we replace the job loss dummy variable for two dummy variables that separately indicate maternal and paternal job loss. We then repeat the estimation of the mother FE model, with the results being presented in column 3 of Table 3. Whilst the estimated effect of paternal job loss is minimal and insignificant, the estimated effect of maternal job loss is significantly positive – experiencing maternal job loss prior to the AST at age 18 increased the probability of university admission at age 18 by 1.67 ppts (accounting for 11.5 percent of the sample mean). Columns 4 to 6 present the estimated effects on public university admission, and the pattern of the three estimates is similar to those shown in columns 1 to 3. The OLS estimate is negative (column 4), but the FE estimate turns positive (column 5), and the positive effect was mainly driven by maternal job loss (column 6).

Why would parental job loss have a positive effect on children's academic performance? Particularly, why is the positive effect of job loss driven by job losses of mothers? One possible explanation is that parents, especially mothers, who turn unemployed may increase time spent on their children. Consequently, parents enhance care of their children and develop effective communication with them.[14] To shed some

[14] Our findings about the differential effects of unemployment between fathers and mothers are in line with some previous relevant studies. Kalil and Ziol-Guest (2008) found that fathers' involuntary job loss lead to a high probability of children's grade repetition and school suspension or expulsion, but the corresponding

light on this point, we use data from the 2000 and 2004 versions of Taiwan's *Time Use Survey*, which reports the exact time individuals spent on various household activities, including housekeeping, cooking, parenting, and other activities. Here, we observe married men and women who were at the ages between 40 and 55, parallel to the range of parents' ages in our sibling sample. For employed and non-employed females (males) separately, we calculate the average amount of time they spent on housekeeping and parenting from 6pm to 9am the next morning, during which all family members are likely be at home. Table 4 presents the statistics, which indicate that employed females spent an average of 82.1 minutes on housekeeping and parenting from 6pm to 9am per weekday (column 1). The average length of time for non-employed females is 106.4 minutes (column 2), which is around 30 percent higher than that for employed females. The margin is statistically significant, as suggested by the result of the *t* test on equality of the two average numbers (column 3). For males, however, the picture is utterly different. Employed males spent only 18 minutes per weekday on housekeeping and parenting, and the time only marginally increased to 21.2 minutes for non-employed males. Being employed does not seem to make a difference on fathers' housekeeping and parenting time.

The statistics in Table 4 have two implications. First, non-employed married women spent more time than employed ones on housekeeping and parenting, supporting our parental-time-at-home hypothesis as a possible explanation for our finding of the positive effect of maternal job loss on university admission. Second, and more important, we now obtain a more profound understanding of the little effect a father's job loss has on university admission. One possible scenario is that the adverse income effect caused by a father's job loss was offset by gain in the father's time at home. This scenario is

effect is insignificant for mothers, even in households with the mother being the major income earner. Rege, Telle, and Votruba (2011) found that paternal job loss due to plant closure has a negative effect on children's GPA, but the effect of maternal job loss is not significant.

implausible because Table 4 shows that employment status does not make a large difference between the time that married men spend on housekeeping and parenting. This makes another scenario more plausible: income loss due to a father's job loss does not adversely affect a child's university admission.

5.3 Discussion

In the preceding subsection, we have shown that parental job loss did not hinder children's university admission, despite the sizable and long-lasting loss of family income (Section 4.2). One may be wondering why children's higher education outcomes were impervious to negative family income shocks – a finding generally at odds with previous studies that mostly found adverse effects of parental job displacement on children's outcomes (Oreopoulos et al., 2008; Stevens and Schaller, 2010; Rege et al., 2011; Lindo, 2011). Here, we provide a discussion on the inconsistency.

In our sample, more than 90 percent of the parental divorces took place when the child was between the ages of 15 and 18 (the ages for high school education). It is possible that parents experiencing firm closure may choose to cut consumption without sacrificing education investment on children at these ages for high school. It is also possible that education investment was not as productive for children at ages 15-18 as for younger children, so whether family income loss causes a reduction in education investment does not matter. Furthermore, as we have mentioned earlier, it is possible that teenagers exposed to a family income shock may learn to be more responsible for the family and become more diligent on school work, offsetting the negative income effect. Finally, higher education is fairly cost friendly for parents in Taiwan. The average tuition fee was less than US$2,000 per year for public universities and less than US$4,000 per year for private universities. In addition, university students from low income families are granted access

to multiple financial aids. It is therefore unlikely that the cost of higher education would prevent single-parent students from enrolling.

Our finding of a non-negative effect of family income shocks on the odds of children's university admission implies that the adverse parental divorce effect found in Section 5.1 is unlikely to operate through the economical mechanism. Non-economic mechanisms, such as psychological and emotional shocks, are more likely dominant.

5.4 Does age at parental divorce matter?

By far we have focused on estimating the effect of parental divorce on the child′s university admission without considering the exact age at which the child experienced parental divorce. The age at parental divorce should be an important factor in determining the magnitude of the divorce effects on the child′s outcomes, but the existing evidence on the effect is extremely scarce.

The age-at-parental-divorce effect is complex for at least two reasons. First, younger children may be more susceptible to environmental changes caused by parental divorce. On the contrary, it is plausible that older children may be more capable of acknowledging the downside of parental marriage dissolution, and thus are exposed to more adverse shocks. Second, when the child′s outcome is measured at a fixed age, age at parental divorce is perfectly collinear with another factor – time lapse since divorce. This makes the picture more complicated because the effect of time lapse since divorce itself may be twofold. It is plausible that the divorce effect is acute in the short run, but it fades out as children becomes accustomed to the status of parental divorce. It is also plausible that the adverse impact of parental divorce would accumulate over time.

Our study adds valuable evidence to the literature. Our sibling data provide a unique opportunity to investigate whether children experiencing parental divorce at different ages

would exhibit any differences in university admission. The availability of the exact date of birth and the exact date of parental divorce allows for a precise measure of time elapse since divorce, denoted by *TED*. Specifically, *TED* is measured by the number of days between the date of parental divorce and the last date of the AST (July 3) at age 18 for children experiencing parental divorce. For those experiencing parental divorce after the date, *TED* is set to be zero.

To conduct the estimation, we extend Equation (1) by incorporating the interaction term of TED_{ijt} and D_{ijt} as an additional covariate:

$$Y_{ijt} = \alpha_1 + \beta_1 D_{ijt} + \beta_2 D_{ijt} \times TED_{ijt} + \boldsymbol{X}_{ijt}\gamma_1 + \boldsymbol{F}_j\delta_2 + a_j + u_{ijt}, \quad (2)$$

where β_2 is the coefficient of interest, which is expected to be zero if *TED* does not matter. We present the estimation results in Table 5, and, for a more effective comparison, we only present the FE estimates. Column 1 duplicates the estimates from column 3 of Table 2, which shows the FE estimate of the divorce effect (β_1) on university admission. Column 2 of Table 5 presents the estimates of β_1 and β_2 from estimating Equation (2). The estimate of β_1 is -0.79 percentage point, which is markedly lower than the corresponding estimate (-1.56 percentage points) in column 1. More importantly, the estimate of β_2 is significantly negative, implying that the adverse divorce effect was stronger if the parental divorce occurred at a younger age. Quantitatively, every 100 days younger led to a 0.17 ppt decrease in the likelihood of university admission at age 18. This implies a 1.19 ppt decrease if the effect is measured at the mean of *TED* (698 days).

Columns 3 and 4 of Table 4 present the FE estimates for public university admission. Again, column 3 copies column 6 from Table 2, while column 4 shows the results from estimating Equation (2). After inserting the interaction term of the divorce dummy and time length since parental divorce, β_1 decreases from 0.86 to 0.67 percentage point and remains statistically significant. Meanwhile, the estimate of β_2 is

negative but insignificant, suggesting that the age at parental divorce does not matter much for public university admission.

We do not have a clear answer as to why age at parental divorce has different effects on the margins of university admission and public university admission. Our conjecture is that higher achieving students, who are more likely to gain admission to a public university, adapted better to the adverse impacts from parental divorce.

5.5 Robustness examinations

We consider two robustness checks. First, we simultaneously estimate the effects of both parental divorce and parental job loss in one regression. Second, we estimate the divorce effects for males and females separately.

To complete the first task, we extend Equation (1) to incorporate a dummy variable (denoted by J_{ijt}) on the right-hand side to indicate experiencing parental job loss prior to the AST at age 18 ($J_{ijt} = 1$) or not ($J_{ijt} = 0$). The regression is specified as

$$Y_{ijt} = \alpha_1 + \beta_1 D_{ijt} + \beta_2 J_{ijt} + \boldsymbol{X}_{ijt}\gamma_1 + \boldsymbol{F}_j\delta_2 + a_j + u_{ijt}. \quad (3)$$

Table 6 presents the estimates of β_1 and β_2. Column 1 suggests that the OLS+controls estimate of the parental divorce effect on university admission is -0.0482, which is nearly identical to the corresponding estimate when the divorce effect is estimated separately (see column 2 of Table 2). The fact that adding the job loss dummy variable (J_{ij}) to the controls of Equation (1) does not make any difference implies that the divorce dummy variable (D_{ijt}) and the job loss dummy variable (J_{ijt}) are uncorrelated. It is therefore not surprising to find that the estimate of the job-loss effect (-0.0092) is very similar to the corresponding estimate (-0.0096) when the effect is estimated independently in column 1 of Table 3. For the remaining five columns of Table 6, the similarity can also be seen between the pooled-model estimates and the corresponding split-model estimates, no matter whether the

comparison is made for the estimated divorce effects (between Tables 6 and 2) or job-loss effects (between Tables 6 and 3). These results clearly indicate robustness of our main estimates.

Next, we investigate whether the parental divorce effects differ between sons and daughters. For the estimation of our mother FE model, we construct a 'male sibling' sample comprised of males for whom there was at least one observation of a brother in the sample period (regardless of any sisters he may have), and a parallel 'female sibling' sample constructed in the symmetric way. Note that individuals who were the only son or only daughter of a family are excluded from the two samples, so the sample representativeness changes. Using the two split samples, we repeat the OLS, OLS+controls, and FE estimations of Equation (1), and the results are presented for males (upper panel) and females (lower panel) separately in Table 7. In each of the six columns, the estimated divorce effect for the males is fairly close to the corresponding estimate for the females. These results suggest that parental divorce caused similar damages in the likelihood of university admission for sons and daughters.

6 Conclusions

We have studied the extent to which parental divorce would affect a child's university admission. The estimation results from our mother FE model, which relies on within-sibling comparisons, suggest that children exposed to parental divorce at the ages of 13-18 suffered a 10.6 percent decrease in the likelihood of university admission at age 18, and a 15.7 percent decrease in the likelihood of public university admission at the same age. Our second finding indicates non-negative effects of parental job loss on university admission, although parental job loss had led to a sizable and prolonged loss of family income. This finding implies that the detrimental effects of parental divorce were unlikely

to operate through deprivation of economic resources. A policy implication, therefore, is that public financial aids provided to single parent families may not be as remedial for the child as we thought. Alternative supports, such as professional counseling, academic advising, and other supportive services should not be ignored.

Our findings also suggest that younger adolescents appear to be more susceptible than older adolescents to parental divorce. This may be because younger children are mentally more vulnerable to family dissolution, or the detrimental effect of parental divorce accumulates over time. In either case, more attention should be paid to younger children with divorced parents.

The existing evidence relating to parental divorce and its impacts on children has been predominantly focused on high income countries. However, divorce has increasingly become a major concern for middle- and low-income countries, including those in East Asia where divorce was once culturally forbidden due to Confucianism. Using Taiwanese data, our findings of significantly negative effects of parental divorce and the implied mental shocks are alarming. Future research that probes into this direction would likely make valuable contributions towards filling the current gap in the evidence.

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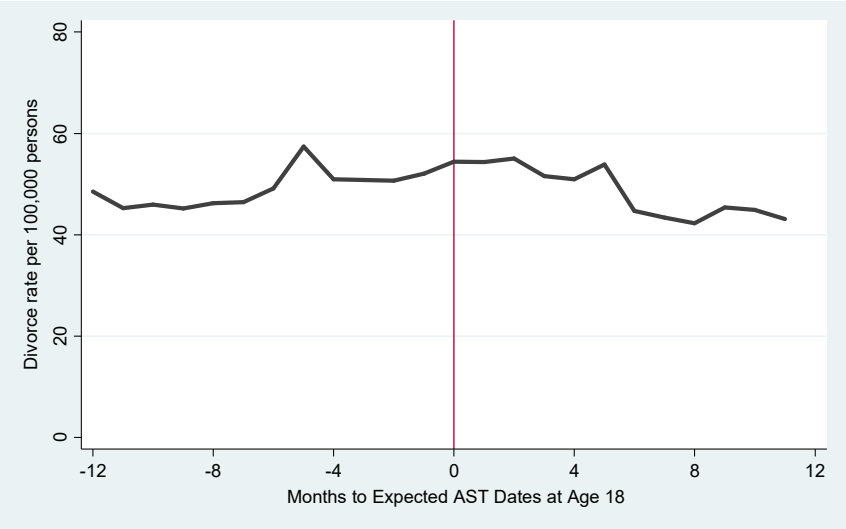
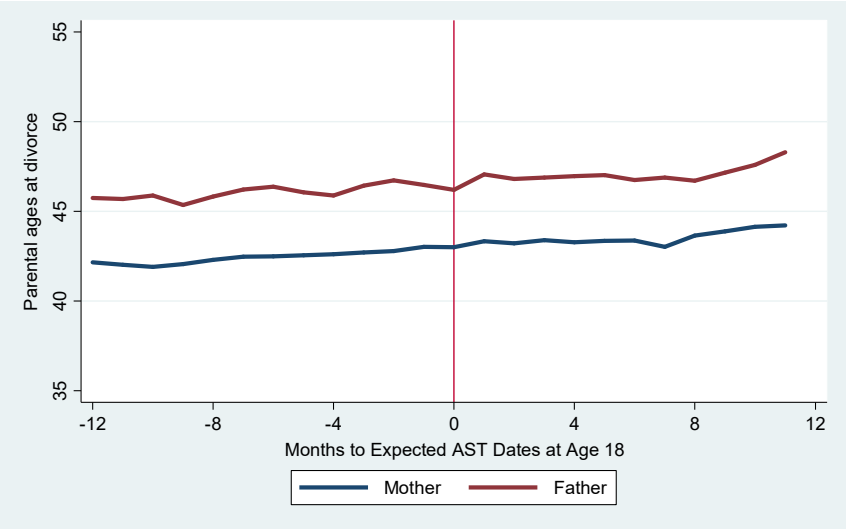
Figure 1: Monthly parental divorce rate



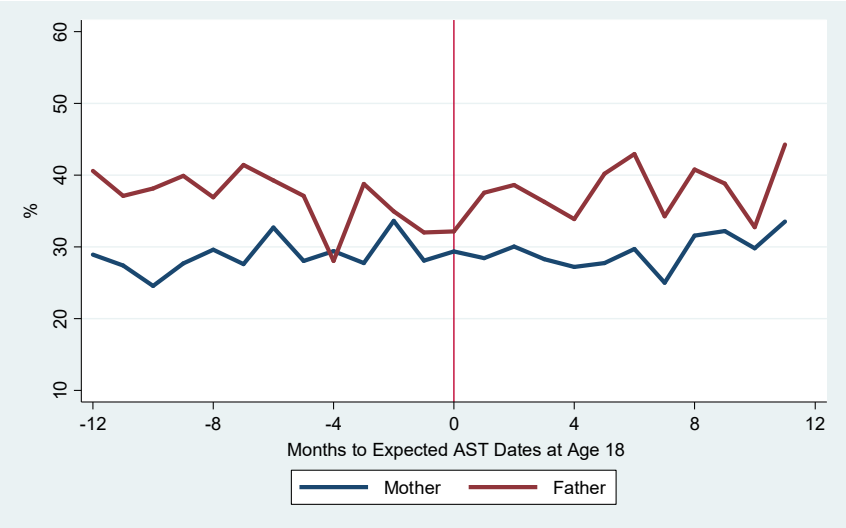
Notes: Divorce rate is defined as number of divorce cases per 100,000 individuals in a month. The curve presents residuals obtained from regressing the monthly divorce rates on the full set of calendar-month fixed effects and a time trend variable measured in months to the AST dates at age 18. AST refers to the Advanced Subjects Test (AST), the national university entrance test scheduled for July 1 to July 3 every year.

Figure 2: Means of observable characteristics

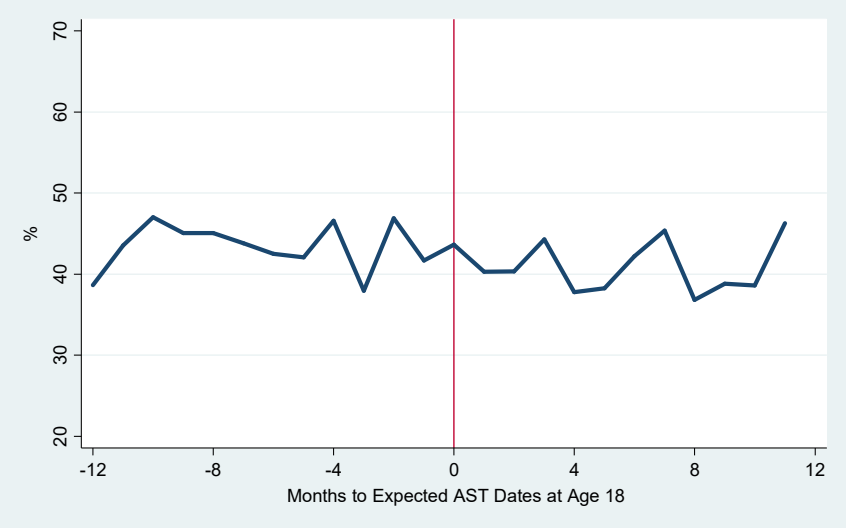
(A) Parental ages at the time of divorce



(B) Percent of parents with a high school diploma



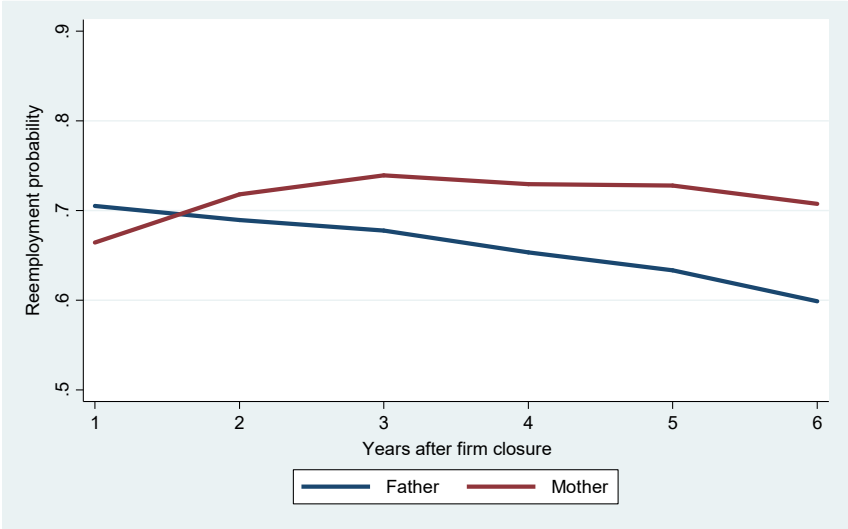
(C) Percent of parents living in urban areas



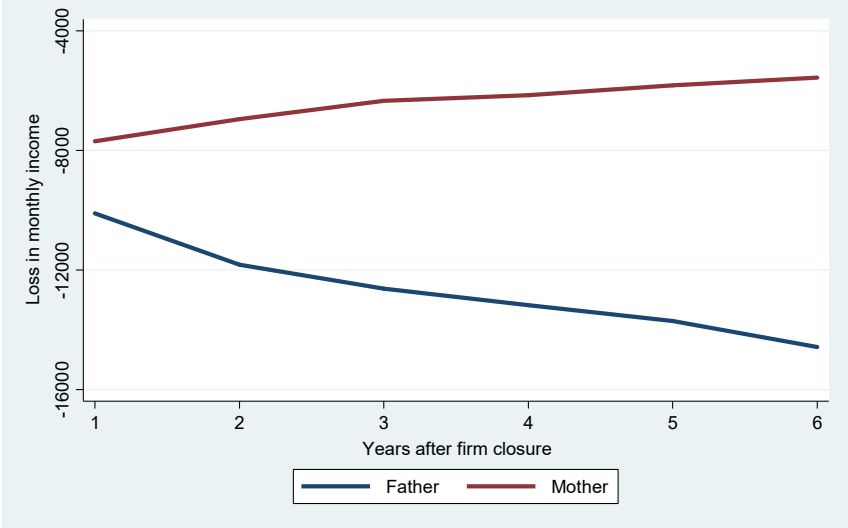
Note: AST refers to the Advanced Subjects Test (AST), the national university entrance test scheduled for July 1 to July 3 every year.

Figure 3: The impacts of firm closure on employment and wage

(A)



(B)



Notes: Samples are parents who experienced a job loss due to firm closure. Loss in monthly income is calculated as monthly wage after reemployment minus the monthly wage just prior to firm closure.

Table 1: Summary statistics

	All children		Children experiencing parental divorce		Children experiencing parental job loss	
	Mean	S.D.	Mean	S.D.	Mean	S.D.
	(1)	(2)	(3)	(4)	(5)	(6)
Children's characteristics						
Sex (Male = 1; Female = 0)	0.510	(0.500)	0.513	(0.500)	0.528	(0.499)
Second birth	0.445	(0.497)	0.534	(0.499)	0.576	(0.494)
Third or later birth	0.183	(0.387)	0.202	(0.401)	0.269	(0.444)
Birth weight (kg)	3.263	(0.455)	3.266	(0.462)	3.284	(0.453)
Year of birth	1982	(1.681)	1983	(1.427)	1983	(1.680)
Demographic controls						
Sibling size	2.923	(0.848)	2.927	(0.921)	2.836	(0.804)
Mother's age at first birth	23.261	(3.220)	21.755	(3.225)	23.574	(3.142)
Mother's year of birth	1957	(3.422)	1959	(3.423)	1957	(3.325)
Father's year of birth	1954	(3.827)	1955	(3.942)	1954	(3.577)
born in urban area	0.361	(0.480)	0.393	(0.488)	0.412	(0.492)
Maternal education level						
college degree+	0.021	(0.144)	0.013	(0.112)	0.016	(0.125)
professional degree	0.010	(0.099)	0.007	(0.081)	0.009	(0.093)
academic high school diploma	0.047	(0.212)	0.048	(0.214)	0.052	(0.222)
vocational high school diploma	0.189	(0.392)	0.175	(0.380)	0.229	(0.420)
junior high school diploma	0.257	(0.437)	0.308	(0.462)	0.288	(0.453)
Primary school diploma	0.475	(0.499)	0.449	(0.497)	0.406	(0.491)
Paternal education level						
college degree+	0.048	(0.214)	0.030	(0.172)	0.043	(0.203)
professional degree	0.018	(0.134)	0.013	(0.115)	0.021	(0.142)
academic high school diploma	0.075	(0.264)	0.072	(0.258)	0.083	(0.276)
vocational high school diploma	0.221	(0.415)	0.205	(0.404)	0.267	(0.443)
junior high school diploma	0.235	(0.424)	0.299	(0.458)	0.252	(0.434)
Primary school diploma	0.401	(0.490)	0.379	(0.485)	0.334	(0.472)
University admission at age 18	0.145	(0.352)	0.080	(0.272)	0.156	(0.363)
Public university admission at age 18	0.054	(0.227)	0.027	(0.161)	0.056	(0.002)
Parental divorce	0.021	(0.144)	-	-	0.041	(0.197)
Parental job loss due to firm closure	0.010	(0.096)	0.018	(0.131)	-	-
Mother	0.005	(0.071)	0.010	(0.097)	-	-
Father	0.005	(0.070)	0.009	(0.095)	-	-
Number of children	1,073,833		22,811		9,888	
Number of mothers	481,459					

Note: Children are selected from families with at least two children born between September 1979 and August 1985. We exclude children from families with multiple births, displaced parents due to other reasons than firm closure, mothers who were not married when giving birth, half-siblings. Families with children who do not have birthweight information are also excluded. Urban areas are Taipei, Hsingchu, Taichung, Tainan, and Kaohsiung cities.*** p<0.01, ** p<0.05, * p<0.1.

Table 2: The effects of parental divorce on university admission at age 18

	All university admission			Public university admission		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS+controls	Mother FE	OLS	OLS+controls	Mother FE
Divorce before 18	-0.0768***	-0.0482***	-0.0156***	-0.0313***	-0.0185***	-0.0086***
	(0.0018)	(0.0018)	(0.0037)	(0.0011)	(0.0011)	(0.0024)
Individual characteristics	X	X	X	X	X	X
Parental characteristics		X	--		X	--
Household variables		X	--		X	--
Number of observations:	1,073,832					
Number of mothers:	481,459					

Notes: The dependent variable is a dummy variable indicating admission to a university (or a public university) at age 18. The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 3: The effects of parental job loss on university admission at age 18

	All university admission			Public university admission		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS+controls	Mother FE	Mother FE	OLS+controls	Mother FE	Mother FE
Job loss (either parent)	-0.0096***	0.0083		-0.0029	0.0075**	
	(0.0035)	(0.0053)		(0.0023)	(0.0037)	
Job loss (father)			0.0004			0.0036
			(0.0075)			(0.0052)
Job loss (mother)			0.0167**			0.0101**
			(0.0070)			(0.0049)
Individual characteristics	X	X	X	X	X	X
Parental characteristics	X	--	--	X	--	--
Household variables	X	--	--	X	--	--
Number of observations:	1,073,832					
Number of mothers:	481,459					

Notes: The dependent variable is a dummy variable indicating admission to a university (or a public university) at age 18. The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Average time (in minutes) per weekday spent on housekeeping and parenting

	(1)	(2)	(3)
	Employed	**Not employed**	**(2)-(1)**
Females	82.1	106.4	24.3***
	N = 1,041	N = 963	
Males	18.0	21.2	3.2
	N = 1,637	N = 335	

Notes: The data are from Taiwan's Time Use Survey, 2000 and 2004. We calculate the average amount of time spent on housekeeping and parenting from 6pm to 9am the next morning for married men and women who were at the ages between 40 and 55 during the sample period. N refers to number of observations. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: The effects of parental divorce duration on university admission at age 18

	All university admission		Public university admission	
	(1)	(2)	(3)	(4)
	Mother FE	Mother FE	Mother FE	Mother FE
Divorce before 18	-0.0156***	-0.0079*	-0.0086***	-0.0067**
	(0.0037)	(0.0043)	(0.0024)	(0.0027)
Divorce before 18 x time length x 100		-0.0017***		-0.0004
		(0.0005)		(0.0003)
Individual characteristics	X	X	X	X
Parental characteristics	--	--	--	--
Household variables	--	--	--	--
Number of observations:	1,073,832			
Number of mothers:	481,459			

Notes: The dependent variable is a dummy variable indicating admission to a university (or a public university) at age 18. The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: The effects of parental divorce and job loss on university admission at age 18

	All university admission			Public university admission		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS+controls	Mother FE	Mother FE	OLS+controls	Mother FE	Mother FE
Divorce before 18	-0.0482***	-0.0157***	-0.0157***	-0.0184***	-0.0086***	-0.0086***
	(0.0018)	(0.0037)	(0.0037)	(0.0011)	(0.0024)	(0.0024)
Job loss (either parent)	-0.0092***	0.0083		-0.0027	0.0075**	
	(0.0035)	(0.0053)		(0.0023)	(0.0037)	
Job loss (father)			0.0004			0.0036
			(0.0075)			(0.0052)
Job loss (mother)			0.0168**			0.0101**
			(0.0070)			(0.0049)
Individual characteristics	X	X	X	X	X	X
Parental characteristics	X	--	--	X	--	--
Household variables	X	--	--	X	--	--
Number of observations:	1,073,832					
Number of mothers:	481,459					

Notes: The dependent variable is a dummy variable indicating admission to a university (or a public university) at age 18. The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7: The effects of parental divorce on university admission at age 18 by gender

	All university admission			Public university admission		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS+controls	Mother FE	OLS	OLS+controls	Mother FE
Males						
Divorce before 18	-0.0733***	-0.0460***	-0.0172***	-0.0283***	-0.0161***	-0.0059
	(0.0030)	(0.0030)	(0.0060)	(0.0019)	(0.0019)	(0.0038)
Individual characteristics	X	X	X	X	X	X
Parental characteristics		X	--		X	--
Household variables		X	--		X	--
Number of observations:	355,300					
Number of mothers:	151,936					
Females						
Divorce before 18	-0.0733***	-0.0480***	-0.0194***	-0.0293***	-0.0181***	-0.0067*
	(0.0033)	(0.0032)	(0.0063)	(0.0019)	(0.0019)	(0.0037)
Individual characteristics	X	X	X	X	X	X
Parental characteristics		X	--		X	--
Household variables		X	--		X	--
Number of observations:	345,107					
Number of mothers:	139,943					

Notes: The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A1: Parental job loss and risk of parental divorce

	(1)	(2)	(3)
Job loss (either parent)	0.103	0.166	0.125
	(0.128)	(0.128)	(0.128)
Individual characteristics		X	X
Parental characteristics			X
Household variables			X
Number of observations: 1,073,446			

Notes: The dependent variable is a dummy variable indicating parental divorce. All estimates are obtained from estimating the mother fixed-effects model. The indicated individual characteristics are gender, birthweight, birth parity, and school year of birth. Parental characteristics are mother's year of birth, father's year of birth, mother's education level, father's education level, and mother's age at first birth and its squared term. Household variables are county of residence and number of siblings. Robust standard errors are in parentheses. *** p<0.01, ** p<0.05, * p<0.1.