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ABSTRACT

Child Health, Remote Work and the Female Wage Penalty

Using data on American women and the health status of their children, this paper studies the effect of remote work on female earnings. Instrumental variables estimates, which exploit a temporary child health shock as exogenous variation in the propensity to work at home, yield an hourly wage penalty of 77.1 percent. Earnings losses together with positive selection, and alternative first stage regressions, suggest that task re-assignment or lack of social interaction are likely mechanisms. The estimates also have implications for the costs of social distancing during a pandemic and may be especially applicable when children must be temporarily quarantined.

JEL Classification: C26, J13, J22, I19

Keywords: female labor supply, female earnings, remote work, fertility, health, instrumental variables, COVID-19, pandemic, quarantine, lockdown

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1 INTRODUCTION

Do women who work at home tend to earn less than women who work exclusively on-site? Theoretically, remote work could lead to lower earnings if jobs that allow for regular working hours at home also offer lower wage rates. Lower wage rates in jobs that permit remote work can arise in a competitive labor market if working at home is considered a job benefit. The benefits can include better coordination of time needed to care for children as well as reduced childcare and commuting costs. Another possibility is that remote work leads to lower earnings because it negatively impacts an individual's productivity. Lower productivity could arise due to shirking or the employer re-assigning the employee to less valuable job tasks that are more suitably performed at home. A desire to work at home might also signal to employers a limited commitment and devotion to the job which then delays or prevents a move up the job ladder. In addition, sociologists and psychologists have long emphasized the importance of social interaction with co-workers as key factors for successful employment outcomes. (Williams, 2000; Blair-Loy, 2006; Williams et al., 2013; Allen et al., 2015).

Despite the various theoretical reasons why a negative association between working at home and earnings may arise, the empirical evidence on remote-work wage losses is sparse and inconclusive. Amongst the most notable previous research, Bertrand et al. (2010) find that when female MBA graduates choose a new job that allows for remote work, these highly educated professional women suffer an earnings penalty of 20 percent. Using NLSY data, Glass and Noonan (2016) find that remote work during regular working hours is associated with only a modest effect on earnings. Bloom et al. (2015) provide evidence of a positive rather than a negative productivity effect of working at home amongst call-center employees in China.

In this paper, the relationship between remote work and labor market earnings is studied by linking data on women in the National Longitudinal Study of Youth (NLSY79) to their children in the NLSY79 Child and Young Adult Survey. Our main contribution to the literature, in addition to using a nationally representative data set, is the presentation of Instrumental Variables (IV) estimates of remote-work earnings losses. The IV framework exploits a temporary child health shock as a source of exogenous variation in the propensity to work at home. Incorporating this potentially plausible source of exogenous variation helps further correct for biases due to unobserved omitted variables that change over time and reverse causality that have not yet been comprehensively addressed in the literature.

It is important to note that working at home was already a major economic phenomenon in the US well before the Covid-19 pandemic emerged. Prior to the start of the pandemic, Dingel and Neiman (2020) estimated that 37 percent of all jobs in the US could be performed entirely at home, and that these jobs accounted for 46 percent of all wages. As states in the US and countries around the the world partially loosen social distancing restrictions, it is likely that the share of all jobs involving at least some remote work will remain at what are now historically unprecedented levels worldwide. Remote work will remain prevalent not only because of continued social distancing restrictions at workplaces, but also as a result of social distancing guidelines in the educational system. These latter guidelines may lead to increased childcare responsibilities and more remote work, especially if children are required to temporarily quarantine. Therefore, our focus in this paper on the relationship between temporary child health shocks, remote work and female wages prior to the pandemic should be relevant for assessing the costs of social distancing during the pandemic era.

The NLSY79 Child and Young Adult Survey (CYA) is particularly useful in this context because it contains assessments of the health of all biological children born to female respondents in the NLSY79. This allows construction of the child health instrument. The logic of the instrument is simply that a temporary child health shock increases the desirability of remote work so that a parent can more flexibly attend to the needs of the child. The key identifying assumption is that after controlling for the standard determinants of wages, the temporary child health shock does not affect the mother's earnings beyond inducing her to work more at home, for a possibly limited period of time. Of course, one cannot completely rule out that the mother's productivity is directly affected by the child health shock beyond just inducing her to work more at home. The IV estimates have added value beyond the Fixed Effects (FE) estimates we present, to the extent that the sick child is not a direct burden on work-task completion at home.

The main results of the study indicate that there is a female wage penalty to working at home which is statistically significant and very large in magnitude. According to our preferred specifications, Ordinary Least Squares (OLS) estimates show that working from home leads to a decrease in the mean hourly wage of 8.7 percent. FE estimates, which take advantage of the longitudinal aspect of the data, yield a larger wage penalty of 12.2 percent. Instrumental Variables (IV) estimates that include fixed effects and exploit the child health instrument produce a wage penalty of 77.1 percent; several orders of magnitude stronger than OLS and FE.

The IV estimates we present are local average treatment effects which pertain to working mothers who are induced to work at home following a temporary child health shock. The much larger magnitude of IV estimates, relative to OLS and FE estimates, indicates positive selection into working at home amongst this particular subgroup of women. The co-existence of positive selection and a substantial wage penalty to remote work, together with alternative first-stage regression results, suggest that some mechanisms are more likely than others. The results are more consistent with women being assigned, or choosing, less "valuable" work assignments, or having less productive social interactions with colleagues, than they are with shirking, a reduction in hours worked, or occupational changes to lower paying jobs that permit work at home (compensating differentials).

The rest of the paper is structured as follows. Section 2 describes the data, provides OLS and FE estimates of the wage penalty and defines the child health instrument. Section 3 outlines the IV approach. Section 4 reports reduced-form and IV estimates of the wage penalty. Section 5 discusses the magnitudes of the estimates, the potential mechanisms underlying the wage penalty, and assesses the validity of the instrument through a placebo test. Section 6 summarizes and concludes.

2 DATA

The NLSY79 is a large nationally representative sample of American young men and women who were 14-22 years old when they were first surveyed in 1979. Data is available at an annual frequency until 1994. The survey became biannual from 1994 onward. The NLSY79 follows the same individuals over time, gathering event histories related to the respondent's labor market experience, education, family background and wages.[1]

The NLSY79 introduced questions on the number of hours per week usually worked at home starting in 1988. This determines the starting year for the analysis. We focus on employed white females who are 24 to 55 years old between the years 1988 and 2012. Only white women who have finished their education are included in the sample. Women with incomplete observations on their marital status and fertility history, inconsistent schooling information, and missing information on occupation (missing census code) and wages are excluded. Because we estimate fixed-effects regressions, we also require more than one year of employment attachment for each individual in the sample. After implementing these standard sample exclusion restrictions, the estimation sample consists of 1,606 women and 17,378 women-year observations.

The CYA surveys contain information on children born to female NLSY79 respondents. NLSY79 children are assessed and interviewed every two years since 1986. For consistency with the NLSY79, children are followed after 1988. Information about a child's health is first provided by the mother. As the child ages, the health information becomes self-reported. The questions on temporary health conditions in the CYA enable the creation of health histories for the children of NLSY79 female respondents. The temporary health problems include limiting health conditions, accidents and injuries requiring medical attention or hospitalization, emotional and behavioral problems, as well as utilization of specialized medical equipment and services. The number of children in the sample is 3,028.

Total annual hours worked for each female in the sample is defined as the sum of weekly hours worked on site (job location is outside of the home) and weekly hours worked at home. Respondents in the NLSY79 can report up to five employers. If more than one employer is identified, only the hourly wage and annual hours worked at the main job are considered. A woman is defined as employed if she reports working at least 10 hours per week, or 520 hours annually. If the sum of annual hours is less than 520, she is also determined to be employed if she worked more than 260 hours in total

[1] The sample originally included 12,686 respondents. It contained a cross-section of 6,111 individuals of which 3,108 were women and 3,003 were men. There was also a set of supplemental samples designed to increase the representation of civilian Hispanics or Latinos, Blacks, the economically disadvantaged, non-Black/non-Hispanic youths (5,295 in total) and a military oversample designed to increase the representation of those serving in the military as of September 30, 1978 (1,280 in total). More information on NLSY79 can be found here.

and reported more than 30 hours weekly.[2]

The overwhelming majority of women in the sample do not work exclusively at home. The distribution of remote work hours shows that 91.3 percent of women work at home less than 1,560 hours per year (30 hours per week on average) and 70 percent less than 520 hours (10 hours per week on average). The mean hours worked at home, excluding zero hours, is 500. The mean is equivalent to a little more than one day a week of remote work.

Table 1 displays the proportion that work at home and the distribution of the number of children under 18 at selected ages of the woman. Pooling over all ages, the work-at-home rate is 17.2 percent. Column (2) shows that the proportion of women who work at home drops steadily from a high of 21.7 percent at age 25 to a low of 13.4 percent at age 55. Column (3) reveals a u-shape in the proportion of women with no children under the age of 18. The proportion is 66.9 percent at age 25, falls to a low of 22.1 percent at age 35, and then rises to reach a high of 92.1 percent at age 55.

The patterns in columns (1) - (3) in Table 1 do not suggest a strong correlation between working at home and the mere presence of children in the household. If this were the case, one would expect to see an increase in the proportion working at home (or at least a slower drop) as the proportion with no child under 18 decreases. The proportion working at home should also decrease more rapidly as the proportion of women with no child under 18 increases again. Neither of these patterns clearly emerge. This leaves room for other explanations for the propensity to work at home, such as the health status of children, which is developed further below.

Table 2 displays sample means and differences in means by work location (at home vs. on site). The figures illustrate that women who work at home tend to be more highly educated, are more likely to work in professional, technical or managerial roles, and work more hours in total. In addition, they have higher wages and are more likely to be married. The unconditional correlation between work hourly wages and work location indicates a 9.6 percent wage premium to working at home, rather than a wage penalty.

[2] To avoid undue influence of outliers, we treat the highest 1‰ and the lowest 2‰ of reported hourly wage observations on the main job as missing.

Table 1: Proportion Working at Home and Distribution of Number of Children < 18 at Selected Ages

Woman's Age (1)	Work at Home (2)	No Child <18 (3)	N (4)
25	.217	.669	290
30	.197	.386	888
35	.166	.221	809
40	.141	.241	611
45	.155	.438	534
50	.158	.734	354
55	.134	.921	89
Total	.172	.388	17,378

Note: N is the number of observations at each age. Total refers to all ages between 24 and 55.

2.1 OLS and FE Estimates

OLS and FE estimates of the impact of working at home on mean wages are presented in Table 3. Column (1) includes only an indicator for working at home. This specification, estimated by OLS, yields a precisely estimated remote-work wage premium of 9.6 percent (as indicated by the unconditional correlation in Table 2).

In column (2), controls are added, including education, annual total work hours, occupation, marital status, spousal income and number of children. The OLS estimates in column (2) now reveal a wage penalty. Working at home is associated with a decrease in mean wages of 8.7 percent. The estimates in column (2) also show that mean hourly wages increase significantly with level of education, total hours worked and being employed in a professional, technical or managerial occupation, as well as being employed in a sales or a clerical position.

Column (3) reports FE estimates which take advantage of the longitudinal aspect of the data. Eliminating time-invariant unobserved individual characteristics and controlling for time-varying observed heterogeneity, a precisely estimated wage penalty of 12.2 percent is obtained. This constitutes a 40.2 percent increase in the wage penalty compared to the OLS estimate in column (2). The other time-varying controls are con-

Table 2: Sample Means

	Full Sample (1)	Work at Home (2)	Work on Site (3)
Age	37.05	36.28	37.21
$hgc < 12$	.060	.019	.068
$12 \leq hgc < 16$	.729	.559	.764
$hgc \geqslant 16$	.211	.422	.167
Professional, Technical and Managerial	.337	.561	.291
Sales and Clerical	.358	.229	.384
Services, Craftsmen, Operatives and Laborers	.305	.210	.325
Total Hours$\leqslant 1,040$	.159	.143	.162
$1,040 <$Total Hours$\leq 1,560$	.166	.146	.171
$1,560 <$Total Hours$\leq 2,080$	.448	.219	.496
Total Hours$> 2,080$	.227	.492	.172
Married	.708	.738	.702
Log Hourly Wage	2.545	2.625	2.529
N	1,606	849	1,562
NT	17,378	2,989	14,389

Note: The figures are averages in the pooled sample. N is the number of women. NT is the number of woman-year observations. *hgc* is highest grade completed. Total hours is the sum of hours worked on site and at home in a calendar year. Wages are hourly wages from the main job earned by an employee in a calendar year. Wages are deflated using the CPI index with a base year of 2005.

siderably reduced in magnitude, in comparison to OLS, though most remain statistically significant.

Table 3: OLS and Fixed Effects Estimates of the Wage Penalty

	Log of Hourly Wage		
	(1)	(2)	(3)
Work at Home	.096	-.087	-.122
	(.028)	(.024)	(.019)
$I(12 \leq hgc < 16)$		.161	
		(.032)	
$I(hgc \geqslant 16)$		.486	
		(.040)	
$I(1,040 <$Total Hours$\leq 1,560)$		.043	.017
		(.017)	(.014)
$I(1,560 <$Total Hours$\leq 2,080)$		.233	.099
		(.018)	(.013)
$I($Total Hours$> 2,080)$		.264	.085
		(.022)	(.016)
Professional, Technical		.383	.145
and Managers		(.022)	(.017)
Sales and Clerical		.186	.059
		(.019)	(.016)
Other regressors	No	Yes	Yes
Fixed Effects	No	No	Yes
Adjusted R^2	.004	.243	.093

Note: Clustered standard errors at the individual level in parentheses. The number of women is 1,606. The number of woman-year observations is 17,378. The dependent variable is the natural log of hourly wage in constant 2005 dollars. Work at home is an indicator for having worked at home during the survey year. *hgc* is the highest grade completed. Total hours refer to the sum of hours worked on site and at home in a calendar year. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income and an indicator for having children under 18.

2.2 The Child Health Instrument

The OLS and FE estimates of the wage effect of working at home may suffer from biases due to unobserved omitted variables that change over time and reverse causality. These biases can be reduced by introducing exogenous variation in the propensity to work at home. We use a temporary child health shock as a source of plausible exogenous variation.

Children's health problems have been used before as instruments in related contexts. For example, Powers (2001) uses 11 impairment categories to instrument the parental assessment of children's functional disability, assuming that the impairments are important determinants of the child care burden but do not directly interfere with parental labor supply. The study finds that the effect of child disability on maternal labor supply is insignificant for married women and negative and more severe for female household heads. The results are substantially different when no instrument is used.

Zan and Scharff (2018) uses a variety of chronic health conditions to instrument the financial and time health-related costs of children under 18 years old. They similarly assume that children's health problems affect their mothers' employment only through health-related financial and time caregiving burdens. Estimates that exploit the exogenous variation show that mothers are more likely to participate in the labor market with a higher monetary caregiving burden, and less likely to participate with a higher time caregiving demand. The effects of caregiving on mothers' employment are underestimated without instrumenting.

The logic of the instrument in our case is simply that a temporary child health shock increases the desirability of remote work in order to more flexibly care for the child. The key identifying assumption is that after controlling for a comprehensive set of standard determinants of wages, including total hours worked and unobserved fixed effects, the temporary child health shock does not affect the mother's earnings beyond inducing her to work more at home, for what might be a limited amount of time. The local average treatment effect that is produced pertains to women who are induced to work at home due to a child's health becoming compromised. It is likely to be a lower-bound estimate since we do not explicitly take into account women who are induced to leave the labour market completely and who may experience more severe losses in human capital and earnings (Eckstein and Wolpin, 1989; Francesconi, 2002; Keane and Wolpin, 2010; Adda et al., 2017).

In constructing the child health instrument, we consider a broad range of health problems causing temporary activity limitations and participation restrictions, as well as injuries and accidents requiring medical attention or hospitalization. Mothers, and later children themselves, are asked in the survey whether the child has a condition that limits school attendance, work and play activities or requires special equipment. The type and duration of the condition is also specified. Mothers are also asked if children had an accident, injury, or illness requiring medical attention or hospitalization and when the three most recent injuries and accidents occurred. Responses to questions about serious behavioral issues, mental or emotional conditions are also used for the construction of the instrument.

The questions about the duration of limitations and the time of injuries or accidents in the CYA allow creation of a continuous child health history. This enables one to distinguish between a permanent and a temporary health problem of the child. A temporary health problem is defined as one which occurs for one year only. Limitations, accidents, injuries and mental conditions, as described above, with a duration of more than one year, or health issues that occur as a result of another disability, in the sense that they coexist with a permanent health condition, are not considered temporary and are not taken into account.

Table 4 presents the proportion of children with a temporary health problem at each child age (less than or equal to 18). A maximum of four children per mother are considered. The overall prevalence of at least one child with a temporary health problem in the household is 12.3 percent. The proportion of mothers in the sample that have at least one child with a temporary health problem is 15.6 percent. Note that pre-school children (less than 7 years of age) are more likely to experience a temporary health problem. This is consistent with evidence for the U.S. and other countries that pre-school children spend more time at home, and the home is the leading location of accidents for young children (Pauline et al., 2007; Phelan et al., 2011).[3]

[3] Prevalence (or prevalence rate) is defined as the proportion of persons in a population who have a particular condition over a specified period of time. The prevalence rate of both the permanent and temporary health conditions in this dataset is 38.7 percent. Different reports use different data and criteria to define the level of limitation or disability. According to Bethell et al. (2011), who use data from the 2007 National Survey of Children's Health, in children younger than 17 , the prevalence of chronic conditions is 43 percent and reaches 49.9 percent for moderate or severe conditions (as rated by parent greater than mild). Data from Survey of Income and Program Participation (SIPP) show that the prevalence of non-severe and severe disability, as defined by the difficulty performing a specific set of functional and participatory activities, for children under 15 is 8.4 percent in 2010. Approximately 50 percent of children with disability were classified with severe disabilities (see Current Population Report). Child Trendsuse National Health Interview Survey data for 1998 - 2013 and a set of questions related to limitations in normal physical activities due to health conditions and impairments, difficulty seeing, difficulty hearing, diagnosed learning disabilities, or difficulty bathing or showering without assistance and find that the proportion of children aged 5 to 17, whose parent or other adult household member reported as having at least one limitation, remained relatively constant from 1998 to 2013, fluctuating between 17 and 20 percent.

Table 4: Proportion of Children with a Temporary Health Problem by Age

Child's Age	Health Problem	*N*
0	.078	1,503
1	.266	1,666
2	.266	1,832
3	.213	2,025
4	.188	2,192
5	.174	2,333
6	.164	2,484
7	.140	2,622
8	.144	2,741
9	.118	2,833
10	.110	2,870
11	.106	2,912
12	.086	2,914
13	.102	2,908
14	.079	2,872
15	.062	2,816
16	.056	2,739
17	.067	2,650
18	.045	2,565
Total	.123	47,477

Note: *N* is the number of children observations at each age.

In the regression analysis that follows, a temporary child health problem is represented by a dummy variable which equals one if at least one child is temporarily afflicted, and equals zero otherwise.

3 Instrumental Variables Framework

The child health instrument is exploited within the framework of a two-stage least squares model that estimates a linear relationship between the log of the hourly wage of woman i at time t, $Y_{i,t}$, and working at home at time t, $F_{i,t}$,

$$Y_{i,t} = \alpha_i + \beta_1 F_{i,t} + \beta_2 X_{i,t} + u_{i,t}, \tag{1}$$

where α_i is an unobserved individual fixed effect, $X_{i,t}$ is a vector of time-varying individual characteristics including age, age squared, regular hours, regular hours squared, overtime hours, overtime hours squared, marital status, the existence of more than two children in the household, and different occupational categories. $u_{i,t}$ is an individual-specific productivity shock in each year t. This is the same set of controls used in the OLS and FE estimations presented earlier.

The first stage equation in the two-stage least squares procedure is

$$F_{i,t} = \gamma_i + \delta_1 H_{i,t} + \delta_2 X_{i,t} + \upsilon_{i,t}, \tag{2}$$

where γ_i is an unobserved individual fixed effect, $H_{i,t}$ is the child health instrument and $\upsilon_{i,t}$ is an individual-specific error term in each year t that may be correlated with $u_{i,t}$ in (1).

As mentioned earlier, the key identifying assumption is that a child's temporary health issue increases the propensity to work at home but does not directly influence wages, after controlling for observable determinants of wages and unobservable time-invariant productivity characteristics. The IV estimates have a causal interpretation as long as the association between children's health and wages is exclusively due to the association between children's health and the decision to work remotely. The main identification challenge arises from the possible impact of children's health on wages through alternative channels such as the choice of working hours and occupation, and through the unobserved determinants of earnings captured by $u_{i,t}$. In order to address these threats to identification, flexible specifications for hours worked are included in the regressions as well as indicators for different occupational categories. A placebo test is also performed to more firmly establish the exogeneity of the instrument.

4 Estimation Results

4.1 Reduced Form Estimates

Table 5 presents reduced form estimates of the effect of a temporary child health problem. The same set of covariates are used as in the OLS and FE regressions in Table 3. Columns (1) and (2) of Table 5 display first-stage estimation results without and with fixed effects, respectively.

In both columns(1) and (2), the coefficient for a temporary child health problem is large in magnitude and statistically significant. A temporary child health problem substantially increases the probability to work at home. The increase in the probability is 5.1 percent without fixed effects and 3.1 percent with fixed effects. These are large magnitudes considering that the the mean proportion that work at home in the sample is 17.2 percent. The point estimates correspond to increases in the probability of working at home of 29.7 percent and 18 percent, respectively. The *F*-statistic in column (1) is 25.54 and in column (2) it is 12.21. This clearly illustrates that the instrument is both relevant and strong.

Columns (3) and (4) show a precisely estimated negative effect of a temporary child health problem on mean hourly wages. Mean wages are lower by 2.7 percent without fixed effects and 2.4 percent with fixed effects. The ratio of the coefficients corresponding to the temporary child health variable in Table 5 already indicate that the IV estimates of the wage effect of working at home will be negative and quite substantial in magnitude.

Table 5: Reduced Form Estimates

	Work at Home		Log of Hourly Wage	
	(1)	(2)	(3)	(4)
Child Health Problem	.051	.031	-.027	-.024
	(.010)	(.009)	(.015)	(.011)
$I(12 \leq hgc < 16)$	.056		.156	
	(.011)		(.033)	
$I(hgc \geqslant 16)$	.205		.468	
	(.020)		(.040)	
$I(1,040 <$Total Hours$\leq 1,560)$	-.001	.000	.043	.017
	(.011)	(.009)	(.017)	(.014)
$I(1,560 <$Total Hours$\leq 2,080)$	-.061	-.030	.237	.103
	(.010)	(.009)	(.018)	(.013)
$I($Total Hours$> 2,080)$	.195	.151	.247	.066
	(.015)	(.013)	(.022)	(.016)
Professional, Technical	.076	.020	.377	.143
and Managers	(.013)	(.014)	(.022)	(.017)
Sales and Clerical	-.021	-.057	.188	.066
	(.009)	(.012)	(.019)	(.017)
Other regressors	Yes	Yes	Yes	Yes
Fixed Effects	No	Yes	No	Yes
F-statistic	25.54	12.21	94.04	48.87
	(.000)	(.001)	(.000)	(.000)
Adjusted R^2	.151	.108	.241	.101

Note: Clustered standard errors at the individual level in parentheses. The number of women is 1,606. The number of woman-year observations is 17,378. The dependent variable in Columns (1) and (2) is a dummy indicating having worked at home during the survey year. The dependent variable in Columns (3) and (4) is the natural log of hourly wage in constant 2005 dollars. *hgc* is the highest grade completed. Total hours worked is the sum of hours worked on site and at home in a calendar year. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income and an indicator for having children under 18. The *F*-statistic is for the test of excluded instruments (*P-values* in parentheses below).

4.2 IV Estimates

IV estimates of the effect of working at home on hourly wages are reported in Table 6. The same set of covariates described earlier for the OLS and FE regressions in Table 3 are included. Working at home is instrumented by a temporary child health problem.

The IV estimates of the wage penalty in Table 6 are large in magnitude and precisely estimated. In Column (1), without fixed effects, the wage penalty is 52.9 percent. Similar to the corresponding OLS estimates in Table 3, mean hourly wages significantly increase with level of education, total hours worked and being employed in a professional, technical or managerial occupation, as well as being employed in a sales or a clerical position.

In Column (2), with fixed effects included, the wage penalty increases substantially in absolute value to 77.1 percent. The effects of the other time-varying covariates are considerably reduced in magnitude, compared to column (1), though most remain statistically significant. This was also the case for the un-instrumented FE estimates reported in Table 3.[4]

5 Discussion

5.1 Magnitudes and Mechanisms

The IV estimates in Table 6 are properly interpreted as local average treatment effects that capture the change in mean hourly wages amongst women who are induced to work at home as a result of at least one child in the household developing a temporary health issue. This subpopulation of women (the compliers (Angrist et al., 1996)) are those who would not have worked at home had the child not become ill.

The local average treatment effects reported in Table 6 are more than 6 times the magnitude of the corresponding OLS and FE estimates, indicating that these latter estimates are substantially biased toward zero (under-estimated) and there is positive selection. The sample means in Table 2 are also highly suggestive of positive selection

[4]No precisely estimated interactions with the indicator for work at home were found, suggesting a lack of substantial heterogenous treatment effects. These results are not reported for sake of brevity but are available upon request.

Table 6: IV Estimates of the Female Wage Penalty

	Log of Hourly Wage	
	(1)	(2)
Work at Home	-.529	-.771
	(.297)	(.389)
$I(12 \leq hgc < 16)$	.185	
	(.036)	
$I(hgc \geqslant 16)$	.576	
	(.072)	
$I(1,040 <$Total Hours$\leq 1,560)$	.043	.019
	(.018)	(.015)
$I(1,560 <$Total Hours$\leq 2,080)$	.205	.079
	(.026)	(.019)
$I($Total Hours$> 2,080)$	.350	.182
	(.060)	(.060)
Professional, Technical	.417	.158
and Managers	(.031)	(.019)
Sales and Clerical	.177	.022
	(.020)	(.028)
Other regressors	Yes	Yes
Fixed Effects	No	Yes
Adjusted R^2	.150	.010

Note: Clustered standard errors at the individual level in parentheses. The number of women is 1,606. The number of woman-year observations is 17,378. The dependent variable is the natural log of hourly wage in constant 2005 dollars. Work at Home is an indicator for having worked at home. *hgc* is the highest grade completed. Total hours worked is the sum of hours worked on site and at home in a calendar year. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income and an indicator for having children under 18.

on unobservables because women who work at home are, on average, more highly educated and more often work in professional, technical or managerial roles. Positive selection of mothers into remote work and flexible jobs was also indicated in the study of the gender earnings gap amongst MBA graduates by Bertrand et al. (2010) and in the study of telecommuting by Glass and Noonan (2016).

Note that wage effects of these magnitudes (52.9 to 77.1 percent) are not unprecedented in the wider literature on flexible working conditions. As mentioned earlier, Bertrand et al. (2010) find a remote-work wage penalty amongst female MBA graduates of 20 percent. However, the wage penalty amongst women that choose a new job with flexible working hours is much higher, reaching 60 percent. These latter estimates are produced from fixed-effects regressions on a selected sample of highly educated women. Our IV estimates are derived from a nationally-representative data set of women that span the education-level spectrum.

In order to assess to what extent the child health problem is leading to less hours worked or occupational changes which could bring about lower wages in return for remote work possibilities (compensating differentials), alternative first stage estimates are displayed in Table 7. The dependent variables are the different indicators for total hours worked and occupation. The results show that the child health shock reduces the proportion of women who work more than 40 hours per week by only 4 percent, while there is a corresponding increase in the proportion of women who work less than 20 hours per week and between 20 and 30 hours per week. There is also only a slight increase (1.9 percent) in the proportion of women that change occupations.

In light of these alternative first-stage results, and the co-existence of positive selection with a large wage penalty, the most likely mechanisms driving the remote-work wage penalty are highly skilled women being assigned, or choosing, less "valuable" work assignments, or less productive social interactions with colleagues. The array of results is less consistent with shirking, or a reduction in hours worked, or occupational changes to lower paying jobs that permit work at home (compensating differentials). Along these lines, Bertrand et al. (2010) speculate that MBA mothers may be forced out, or opt out, of the "fast-track" after choosing more flexible work arrangements. The more there is a tournament, or up-or-out, structure to the occupation, the more "task-shift"

Table 7: Alternative First-Stage Estimates

	Total Hours Worked			Professional, Technical and Managers	Sales and Clerical
	1,040 to 1,560	1,561 to 2,080	>2,080		
	(1)	(2)	(3)	(4)	(5)
Child Health	.019	-.005	-.040	.006	.019
Problem	(.011)	(.013)	(.010)	(.008)	(.009)
I(Total Hours Worked				.006	.006
1,040 to 1,560)				(.010)	(.011)
I(Total Hours Worked				.037	.022
1,561 to 2,080)				(.010)	(.011)
I(Total Hours Worked				.081	.006
>2,080)				(.012)	(.012)
Professional,	-.045	-.005	.098		-.641
Technical, Managers	(.012)	(.017)	(.014)		(.014)
Sales and Clerical	-.009	.042	-.017	-.579	
	(.012)	(.015)	(.011)	(.014)	
Other Regressors	Yes	Yes	Yes	Yes	Yes
Fixed Effects	Yes	Yes	Yes	Yes	Yes
F-statistic	2.79	0.15	16.10	0.60	4.50
	(.095)	(.699)	(.000)	(.440)	(.034)
Adjusted R^2	.015	.015	.055	.298	.287

Note: Clustered standard errors at the individual level in parentheses. The number of women is 1,606. The number of woman-year observations is 17,378. Total hours worked is the sum of hours worked on site and at home in a calendar year. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income and an indicator for having children under 18. The *F*-statistic is for the test of excluded instruments (*P-values* in parentheses below).

there is likely to be when choosing to work at home. In our sample, a substantial proportion of women (33.7 percent) work in the professional, technical and managerial occupational category where tournament and up-or-out employment structures are the most prevalent.

5.2 Instrument Exogeneity

The first-stage estimates in Table 5 illustrated that the child health instrument is relevant and strong. Evidence for the exogeneity of the instrument is clearly more difficult to establish, especially in the absence of over-identification. Nonetheless, we attempt to justify instrument exogeneity, and hence the validity of the instrument, by performing a simple placebo test.

The placebo test "falsely" assigns the temporary child health problem to be two years before it actually occurred, and then runs the same IV procedure with fixed effects as in column (2) of Table 6. The results of the placebo test are presented in Table 8. Note that the effect of working at home on mean hourly wages is positive (43.4 percent) and very imprecisely estimated. Thus, there is little evidence for anticipation of the child health shock, or the existence of unobserved persistent determinants of earnings correlated with the child health shock, that drive the substantial wage penalties found in Table 6.

Moreover, the positive effect of working at home is further evidence in favor of positive selection. Women who later experience a wage penalty for choosing to work at home, following a temporary child health shock, have 43.4 percent higher mean earnings, two years prior to the shock, than women who continue to work on site.

6 Conclusion

Using data on women in the NLSY79 and their children in the NLSY79 Child and Young Adult Survey, this paper estimates the wage effects associated with working at home. There is still no consensus in the literature as to whether a wage premium or a wage penalty to remote work is a more likely outcome. The main contribution

Table 8: IV Estimates of the Female Wage Penalty - Placebo Test

	Log of Hourly Wage
Work at Home	.434
	(.602)
$I(1,040 <$Total Hours$\leq 1,560)$	.000
	(.016)
$I(1,560 <$Total Hours$\leq 2,080)$	.110
	(.021)
$I($Total Hours$> 2,080)$	-.007
	(.093)
Professional, Technical	.129
and Managers	(.023)
Sales and clerical	.089
	(.040)
Other regressors	Yes
Fixed Effects	Yes
Adjusted R^2	.047

Note: Clustered standard errors at the individual level in parentheses. The number of women is 1,606. The number of woman-year observations is 16,484. Total hours worked is the sum of hours worked on site and at home in a calendar year. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income and an indicator for having children under 18.

of this paper is in the presentation of IV estimates of wage effects using nationally representative data on American women and the health status of their children as an instrument. The source of exogenous variation in the propensity to work at home that we use, a temporary child health shock, is shown to be relevant and strong, as well as plausibly exogenous via a placebo test.

The study finds female wage penalties to working at home which are statistically significant and substantial in magnitude. OLS estimates yield a wage penalty of 8.7 percent. Fixed-effects regressions yield a larger wage penalty of 12.2 percent. IV estimates that include fixed effects and exploit the child health instrument produce a much more serious wage penalty of up to 77.1 percent. Even this very large magnitude is likely to be a lower-bound estimate since we do not explicitly take into account women who are induced to leave the labour market completely as a result of a temporary child health shock. These latter woman may experience more severe losses in human capital and earnings.

The large negative IV estimate, relative to OLS and FE estimates, is indicative of positive selection into working at home. The co-existence of positive selection and wage penalties, and the results from alternative first-stage regressions, are suggestive of particular mechanisms that are more likely to underly the remote-work wage penalty than others. The array of results is more consistent with women being assigned, or choosing, less "valuable" work assignments when working at home, or having less productive social interactions with colleagues, than it is with shirking, or a reduction in hours worked, or occupational changes to lower paying jobs that permit work at home (compensating differentials). Better data would enable a more precise disentangling of the various mechanisms and possibly a further comparison to alternative possibilities such as negative signaling and statistical discrimination.

It is likely that as states in the US and countries across the world partially loosen social distancing restrictions, the share of all jobs involving at least some remote work will remain at historically unprecedented levels worldwide. Remote work will remain prevalent not only because of adherence to continued social distancing norms at the workplace, but also due to social distancing guidelines in the educational system. These latter guidelines may lead to increased childcare responsibilities and more remote work, especially when children are required to temporarily quarantine. Thus, the estimates we offer in this paper, which link temporary child health shocks, remote work and female wages prior to the pandemic, may be particularly relevant for assessing the costs of social distancing in the post-pandemic era.

In closing, we also note that public discussions and policy proposals involving remote work often center around how to increase the supply of work-at-home opportunities, rather than directly addressing the wage penalty. The wage penalty aspect of remote work, relative to the supply issue, has now become a much more salient concern in light of the pandemic. Our results suggest that policymakers might consider shifting focus.

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