**Article:**
Kouki, A. (2024) Work from home and the racial gap in female wages. European Economic Review, 170. 104864. ISSN 0014-2921

https://doi.org/10.1016/j.euroecorev.2024.104864

European Economic Review 170 (2024) 104864

Contents lists available at ScienceDirect

European Economic Review

journal homepage: www.elsevier.com/locate/eer

# Work from home and the racial gap in female wages

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ARTICLE INFO

*JEL classifications:*
C26
J13
J22
I19

*Keywords:*
Female wages
Female labor supply
Race
Work from home
Telecommuting
Fertility
Child health

ABSTRACT

This paper studies the racial female wage penalty to remote work in the U.S. Instrumental variable estimates yield wage penalties that reach 66.3 % for black women and 33.9 % for white women when hours worked at home increase to 5 per week. Promotion bias, task reassignment and lack of productive social interaction are the most likely mechanisms for the wage losses. The estimates provide rare evidence on the costs of physical distancing due to work from home, particularly for women of different races managing the needs of their sick children.

## 1. Introduction

Do women who work at least some hours from home earn less on average than similar women who work exclusively on-site? Remote work could lead to a reduction in earnings if jobs that permit work from home offer lower wage rates. Lower wage rates may arise if working from home is considered an employee benefit. The perceived benefits can include better coordination of time needed to care for children as well as reduced childcare and commuting costs. Previous research has also shown that remote work is associated with an increase in job satisfaction (Virick et al., 2010) and helps alleviate fears of social stigma (Ghumman and Jackson, 2010; Tislick et al., 2015; Follmer et al., 2018).

Another possibility for the existence of a wage penalty is that remote work negatively impacts an individual's productivity. Lower productivity may result from distractions in the home, shirking due to less effective monitoring, or the employee being reassigned to less valuable job tasks that are more suitably performed at home. A desire to work from 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 (Golden, 2008; Kaplan et al., 2018; Golden and Eddleston, 2020). In addition, sociologists, psychologists, economists and organizational behavior researchers have long emphasized the importance of social capital, network relationships and interaction with co-workers as key factors for successful employment outcomes (Milgrom and Oster, 1987; Williams, 2000; Blair-Loy, 2006; Hersby et al., 2009; Williams et al., 2013; Allen et al., 2015; Cassidy et al., 2016; Bertrand et al., 2019).

While the effect of general work flexibility on earnings has been studied extensively (see e.g., Goldin and Katz (2011); Flabbi and

I am particularly grateful to the Editor and the anonymous referees for their very useful feedback. I declare that I have no relevant or material financial interest that relates to the research described in this paper.

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https://doi.org/10.1016/j.euroecorev.2024.104864
Received 24 November 2023; Received in revised form 6 September 2024; Accepted 9 September 2024
Available online 10 September 2024

Moro (2012); Goldin (2014); Goldin and Katz, 2016; Wiswall and Zafar, 2018; Cortes and Pan, 2019; Ishizuka and Musick, 2021; Cubas et al., 2023), the empirical evidence on remote-work wage losses (or gains) remains sparse and inconclusive. Notable previous research in this specific area includes Bertrand et al. (2010) which finds that female MBA graduates suffer earnings penalties that exceed 20 %. In contrast, Glass and Noonan (2016) find in more representative data that remote work during regular working hours is associated with only a very modest effect on earnings. Bloom et al. (2015) provide evidence of a positive rather than a negative productivity effect of working at home among call-center employees in China. Using a sample of professional employees in a company providing technology services with locations throughout the U.S., Golden and Eddleston (2020) find that telecommuters experience lower salary growth compared to non-telecommuters.

In this paper, the relationship between remote work and wages 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 (NLSCYA). The main contribution to the literature is twofold: (i) this is the first study that focuses exclusively on the racial gap in female wage penalties to remote work and (ii) this data allows to provide a range of estimates that include not only Ordinary Least Squares (OLS) and Fixed Effects (FE) results but also Instrumental Variables (IV) estimates. The IV procedure exploits a temporary child health problem as a source of exogenous variation in the share of hours worked from home. To the extent that a temporary child health issue is a valid instrument, the IV estimates help further correct for biases due to unobserved omitted variables that change over time and reverse causality that have also not been addressed in the literature on remote-work wage losses.

It is important to note that working from home was already a major phenomenon well before the COVID-19 pandemic emerged. Prior to the start of the pandemic, Dingel and Neiman (2020) estimated that 37 % of all jobs in the U.S. could be performed entirely at home, and that these jobs accounted for 46 % of all wages. As there is a growing interest in hybrid work models that allow employees to work both remotely and, in the office, and as metaverse platforms further develop, it is likely that the share of all jobs involving at least some remote work will continue on an upward trend. It is worth noting that about 60 % of American workers whose work can be done remotely are currently working from home either all or most of the time.[1] Although there are not any public health measures in place, children's temporary illnesses can still result in additional caregiving responsibilities, which can be a source of concern for parents, particularly working mothers. Such childcare demands can impact work responsibilities and may have consequences for employment, through the loss of work experience, delayed promotion, a shift to employment in less demanding jobs and prolonged exit from the labor market.

Results from studies that concentrate on the labor market impact of the recent pandemic clearly show that average hours worked have declined more for females than males in percentage terms (Bick and Blandin, 2021). There is also emerging evidence that women are managing to perform less employment tasks from home than men, likely due to their disproportionate responsibility for household production (Adams-Prassl et al., 2022). These imbalances could potentially exacerbate existing inequalities related to race and gender (Alon et al., 2020). The focus in this paper is on the relationship between remote work, temporary child health issues and racial differences in female wage losses prior to the pandemic. However, the findings should be relevant for assessing the differential costs of physical distancing by race for females both in the present and the near future.

The NLCYA dataset is particularly useful in this context because it contains assessments of the health of all biological children born exclusively to NLSY79 female respondents. This allows construction of the child health instrument. The logic of the instrument is that a sudden and temporary child health problem increases the desirability of remote work so that a mother can more flexibly deal with disruption in childcare needs. The IV results add to the OLS and FE evidence to the extent that the temporary child health problem is not a direct burden on work-task completion beyond that caused by the change in work venue alone. Note that OLS and FE estimates are also of considerable value in themselves given the very limited evidence that currently exists on racial gaps in female remote-work wage losses.

The main results of the study indicate that there is a wage penalty to work from home which is substantial in magnitude and statistically significant for both races, and that black women are penalized more than white women. According to the preferred specifications, OLS estimates indicate that when the time worked from home increases to 5 h per week, black women's wages are expected to decrease by 11.5 %, while white women's wages are expected to fall by 3.9 %. This means that there is a 7.6 percentage point race gap in wages due to this increase in hours worked from home. FE estimates, which take advantage of the longitudinal aspect of the data, yield a wage penalty for black women of 9.9 % and a wage penalty for white women of 5.7 % for the same increase in hours worked from home for each race. The percentage point race gap in wages decreases to 4.2 % points. IV estimates that include fixed effects and exploit the child health instrument produce larger wage penalties for both races, reaching 66.3 % for black women and 33.9 % for white women for an increase in hours worked at home to 5 on weekly basis. The percentage point race gap in wages reaches 32.4 % points. Notably, an increase to 5 h from home on average weekly is on par with the 20 % of full workdays that Barrero et al. (2021) suggest being worked from home after the pandemic, compared with just 5 % before. Based on American Time Use Survey (ATUS) data, in 2021, on average, women and men who worked at home did so for 5.5 and 5.8 h on days they worked, respectively; in 2023, both spent about the same of time working at home, with women averaging 5.2 h and men 5.1 h per workday.[2] The absolute magnitudes of the wage penalty are clearly larger in the IV procedure compared to OLS and FE estimates.

One of the main reasons that the IV estimates of the wage penalties are larger than those found using OLS and FE procedures is that they are local average treatment effects (LATE) as opposed to average treatment effects. LATE estimates pertain to working mothers who are just on the margin of working exclusively on-site or working some hours from home, and who are induced to work from home

[1] See how COVID-19 pandemic has reshaped work in America here.
[2] Data files for ATUS are available here.

solely due to the temporary child health problem. The larger IV estimates, relative to OLS and FE, implies that these latter women (the compliers) have higher unobserved earnings potential and there is positive selection into working from home.

The co-existence of positive selection and a substantial wage penalty to remote work, combined with the results of additional reduced-form regressions, suggest that certain mechanisms are more likely to drive these findings than others. The results are more consistent with promotion bias, women performing less "valuable" work assignments from home and having less productive social interactions with colleagues than they are with employee shirking, a reduction in hours worked, or occupational changes to lower paying jobs. The problems of promotion bias, task reassignment and loss in social capital appear to be differentially worse for black women. This is an interesting new finding of the study.

The rest of the paper is structured as follows. Section 2 describes the data. Section 3 presents OLS and FE estimates of the wage penalties for black and white women. Section 4 describes the temporary child health problem instrument and presents its effect estimates in event study graphs. Section 5 explains the IV estimation framework, reports reduced form and IV results for the wage penalties and assesses the validity of the instrument. Section 6 discusses the magnitudes of the estimates and mechanisms underlying the differential remote-work wage penalty. The last section summarizes and concludes.

## 2. Data

The NLSY79 is a large nationally representative sample of American 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 onwards. The NLSY79 follows the same individuals over time, gathering event histories related to the respondent's labor market experience, education, family background and wages.[3] This study focuses on black and white females. Hispanic and Latino females, as well as military personnel, are excluded due to the relatively small number of observations and non-representativeness in these subgroups.

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. Given the focus of this analysis on the relationship of remote work and wages, the sample is limited to employed women who are 24 to 55 years old between the years 1988 and 2012. Since fixed-effects regressions are estimated, more than one year of employment attachment for each female in the sample is required. After implementing additional standard sample exclusion restrictions, the estimation sample consists of 885 black women with 9,645 women-year observations and 1,600 white women with 17,332 women-year observations.[4] While including male workers would allow interesting gender-based comparisons, the NLSCYA surveys contain information on the children of NLSY79 female respondents only.

In the NLSCYA surveys, children are assessed and interviewed every two years since 1986. For consistency with the NLSY79, children are followed only after 1988 when the remote work questions were introduced. Information about a child's health is first provided by the mother. After the age of 15, the health information becomes self-reported by the child. The questions on temporary health conditions in the NLSCYA enable the creation of health histories for the children of NLSY79 female respondents which are used in the instrumental variables analysis.

Various questions in the NLSCYA that refer to the time a specific health problem occurred are used to build a continuous health history for each child and assess the permanent or temporary nature of the health issue. The questions pertain to how long the child has had limitations, date of most recent accident/injury and other more specific questions about the child's health status in the last 12 months. 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 of black women in the sample is 1,782 and the number of children of white women is 2,878.

Weekly employment histories in the NLSY79 are used to construct total annual hours worked. The survey questions ask respondents about (i) the total hours usually worked per week at a specific job; and (ii) the hours per week usually worked at this specific job at home.[5] Total annual hours worked for each female in the sample is defined as the sum of weekly hours worked on site (job location is

[3] The sample originally included 12,686 respondents. It contained a cross-section of 6,111 individuals of which 3,108 were women (2,279 white, 405 black, 226 Hispanic or Latino and 178 white disadvantaged). There was also a set of supplemental samples designed to increase the representation of civilian Hispanics or Latinos, black and the economically disadvantaged non-black/non-Hispanic youths (5,295 in total of which 2,719 were women - 751 Hispanic or Latino, 1,067 black and 901 white disadvantaged), as well as a military oversample designed to increase the representation of those serving in the military as of September 30, 1978 (1,280 in total). In 1985, the military sample was discontinued (1,079 members of the military sample were dropped). In 1991, all members of the economically disadvantaged non-black/non-Hispanic oversample were dropped as well. More information on NLSY79 can be found here. More information on NLSY79 sample design can be found here.

[4] The additional sample exclusion restrictions are as follows. Only women who have completed 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.

[5] For this analysis, I used the NLSY79 questions that ask respondents about the number of hours during the survey week that they worked at all jobs, the number of hours per week usually worked at the CPS job, and the usual hours per week worked at the non-CPS Jobs #1-5. For each job, questions differentiated the number of hours worked at home from the hours worked at the place of employment. I have also used the follow-up questions asked to double check that the at-home hours reported are included in the total hours worked. More information on NLSY79 hours worked can be found here.

outside of the home) and weekly hours worked at home. A woman is defined as employed if she reports working at least 10 h per week or 520 h annually. If the sum of annual hours is less than 520, she is also determined to be employed if she worked more than 260 h in total and reported more than 30 h weekly.[6] For wages, the hourly rate of pay at the time of the interview is used.[7]

The overwhelming majority of women in the sample do not work exclusively at home. The distribution of remote work hours shows that 86.9 % of black women and 91.4 % of white women work at home less than 1,560 h per year (30 h per week on average), while 67.8 % of black women and 70.2 % of white women work less than 520 h (10 h per week on average). The mean hours worked at home, excluding zero hours, is 615.84 for black women and 497.30 for white women. The mean total hours for black and white women that work a positive number of hours at home is 2,141.18 and 1,991.02, respectively. This means that for black women, the mean share of hours worked at home corresponds to 28.76 % of their total hours worked, while for white women, the corresponding proportion is 24.90 %.

Table 1 displays the proportion of women that work at home, the distribution of hours worked at home, as well as the total hours worked and the proportion of women with no children under 18. Pooling over all ages, the work at home rate is 9.8 % for black women and 17.1 % for white women. Column (1) of Table 1 shows that the proportion of black women who work at home drops slightly from 10.6 % at age 24–29 to 9.1 % at age 50–55, while Column (2) shows that the mean hours worked at home increase from 46.3 h at age 24–29 to 52.7 h at age 50–55. Column (6) displays the proportion of white women who work at home. Their work-at-home rate drops more from 18.9 % at age 24–29 to 15.5 % at age 50–55. According to Column (7), their mean hours worked from home increase from 72.3 h at age 24–29 to 80.8 h at age 50–55. Columns (3) and (8) present the mean of total hours black and white women work annually. Black women work on average more than white women. The share of hours worked at home as a proportion of total hours worked for black women is 3.4 %, while for white women, the corresponding share is 4.8 %.

Columns (4) and (9) reveal a U-shape in the proportion of women with no children under the age of 18. For black women, the proportion is 35.2 % at age 24–29, falls to a low of 22.7 % at age 35–39, and then rises to reach a high of 80.2 % at age 50–55. Similarly, for white women, the proportion is 52.2 % at age 24–29, reaches its lowest of 21.6 at age 35–39, and peaks at 79 % at age 50–55.

Table 1 does not suggest a strong correlation between working at home, hours worked at home, total hours worked and the mere presence of children in the household. If this were the case, one would expect to see an increase in both the proportion of those working at home and the hours worked at home or a decrease in the total hours worked as the proportion with no children under 18 decreases. The proportion of those working at home along with their hours worked at home should also decrease more rapidly as the proportion of women with no children under 18 increases again. Neither of these patterns clearly emerge. This leaves room for other explanations for the decision to work at home, such as the health status of children, which is developed further below.

Table 2 displays sample means by race and work location (at home vs. on site). On average, black women earn 15.5 % less than white women. Black women are underrepresented in medium- and high-skilled occupations and overrepresented in low-skilled occupations. Black and white women are also different in terms of family structure; a much higher proportion of white women are married, while the spouses of white women earn more in comparison to the spouses of black women.

The figures illustrate that both black and white women who work at home a positive number of hours tend to be more highly educated, are more likely to work in professional, technical or managerial roles, work more hours in total and are more likely to be married. Black women's wages do not differ significantly by work location, while white women who work at home tend to have higher wages on average. The unconditional correlation between hourly wages and work location indicates a 12 % wage premium to working at home, rather than a wage penalty, for white women only.[8] For the rest of this analysis, instead of using a binary variable indicating whether a woman works a positive number of hours from home or not, I use the share of hours worked from home as a proportion of the sum of total hours worked in a calendar year.

## 3. OLS and FE estimates

OLS estimates of the impact of hours worked at home on mean wages for the combined sample (Columns (1) and (2)), and both OLS and FE estimates by race (Columns (3) - (8)), are presented in Table 3. The main variable of interest is the log of hours worked at home as a proportion of total hours worked. Column (1), estimated by OLS, includes the log of share of hours worked at home as the sole covariate and shows a significant racial penalty for the combined sample. Columns (3) and (6), as Column (1), present this base specification, estimated by OLS. Both coefficients, which can be interpreted as elasticities of wages with respect to the share of hours worked at home, are not consistent with the corresponding correlations of Table 2 which condition only on the work location.

In Columns (2), (4) and (7), confounding determinants of wages are added, including education, part-time status, occupation,

[6] These selection criteria are typical in literature and aim to eliminate outliers. Indicatively, Keane and Wolpin (2010) use a threshold of 260 hours for a six-month period, while Gicheva (2013) drops observations for which weekly hours are less than 15. 16 women are excluded because of this requirement.

[7] Respondents in the NLSY79 can report up to five employers in a survey interview. If more than one employer is identified, only the hourly wage and total annual hours worked at the main job are considered. The main job is defined as the job in which the individual works the most hours in each year. Focusing on the main job is probably not a serious limitation as the annual average multiple job-holding rate for the U.S. was only 4.9 percent in 2012 and had declined gradually or remained flat each year since peaking at 6.2 percent in 1996 (Campolongo, 2013). Deflated hourly wages below $1 and above $200 on the main job are treated as missing. More information on NLSY79 wages can be found here.

[8] NLSY79 respondents were asked whether their employer has made available flexible hours or work schedule. Black and white women in this sample are offered time flexibility in equal proportions and independently of whether they work from home or not.

**Table 1**
Remote work, total working hours, and women without children ≤18.

| Age | Black women | | | | | White women | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Work at home (1) | Hours at home (2) | Total hours worked (3) | No child ≤18 (4) | *N* (5) | Work at home (6) | Hours at home (7) | Total hours worked (8) | No child ≤18 (9) | *N* (10) |
| 24–29 | 0.106 | 46.29 | 1,751.39 | 0.352 | 1,647 | 0.189 | 72.31 | 1,806.77 | 0.522 | 3,085 |
| 30–34 | 0.102 | 65.17 | 1,768.35 | 0.245 | 2,453 | 0.193 | 105.41 | 1,774.06 | 0.312 | 4,427 |
| 35–39 | 0.097 | 67.81 | 1,813.92 | 0.227 | 1,994 | 0.176 | 96.68 | 1,772.43 | 0.216 | 3,476 |
| 40–44 | 0.096 | 65.83 | 1,845.86 | 0.376 | 1,682 | 0.138 | 65.19 | 1,781.72 | 0.302 | 2,902 |
| 45–49 | 0.088 | 54.76 | 1,840.41 | 0.614 | 1,289 | 0.149 | 73.73 | 1,772.75 | 0.528 | 2,341 |
| 50–55 | 0.091 | 52.70 | 1,681.78 | 0.802 | 580 | 0.155 | 80.76 | 1,671.99 | 0.790 | 1,101 |
| Total | 0.098 | 60.47 | 1,792.82 | 0.365 | 9,645 | 0.171 | 85.19 | 1,774.18 | 0.388 | 17,332 |

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

**Table 2**
Sample means.

| | Black women | | | White women | | |
|---|---|---|---|---|---|---|
| | All (1) | Work at home (2) | Work on Site (3) | All (4) | Work at home (5) | Work on Site (6) |
| Age | 37.10 | 36.68 | 37.15 | 37.05 | 36.29 | 37.21 |
| $hgc < 12$ | 0.078 | 0.048 | 0.082 | 0.060 | 0.018 | 0.068 |
| $12 \leq hgc < 16$ | 0.800 | 0.663 | 0.815 | 0.729 | 0.559 | 0.764 |
| $hgc \geq 16$ | 0.122 | 0.289 | 0.104 | 0.211 | 0.423 | 0.167 |
| Professional, technical and managerial | 0.215 | 0.476 | 0.186 | 0.337 | 0.562 | 0.291 |
| Sales and clerical | 0.332 | 0.224 | 0.344 | 0.358 | 0.231 | 0.385 |
| Services, craftsmen, operatives and laborers | 0.453 | 0.300 | 0.470 | 0.304 | 0.207 | 0.325 |
| Part-time | 0.377 | 0.267 | 0.389 | 0.409 | 0.349 | 0.422 |
| Full-time | 0.623 | 0.733 | 0.611 | 0.591 | 0.651 | 0.578 |
| Married | 0.482 | 0.550 | 0.474 | 0.709 | 0.739 | 0.702 |
| Log of spouse earnings | 10.099 | 10.175 | 10.089 | 10.263 | 10.310 | 10.254 |
| Log of hourly wage | 2.400 | 2.421 | 2.397 | 2.555 | 2.654 | 2.534 |
| *N T* | 9,645 | 947 | 8,698 | 17,332 | 2,969 | 14,363 |

Note: The figures are averages. The number of black women is 885 and white women is 1,600. NT is the number of woman-year observations. *hgc* is highest grade completed. Part-time is an indicator for whether the total hours worked on site and at home weekly are less than 35. Spouse earnings are annual earnings of spouse from all jobs before any deductions. Wages are hourly wages from the main job earned by an employee in a calendar year. Missing spouse earnings are not included. Wages and spousal earnings are deflated using the CPI index with a base year of 2005.

marital status, spousal income, and number of children.[9,10] The OLS estimates in Columns (4) and (7) now reveal substantial and statistically significant coefficients for both races. An increase in hours worked at home to 5 on average per week is associated with a decrease in mean wages of 11.5 % for black women, and 3.9 % for white women. The racial gap in the wage penalty is 7.6 % points. The estimates in Columns (4) and (7) show expected wage effects for the level of education, part-time, being employed in a professional, technical or managerial occupation, and being employed in a sales or a clerical position.

Columns (5) and (8) report 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, precisely estimated elasticities with respect to share of hours worked at home are obtained. Notably, compared to the OLS estimates, when accounting for the racial differential in average work-at-home hours and subsequent increases to 5 h per week, FE estimates yield a smaller wage penalty for black women of 9.9 % and a greater wage penalty for white women of 5.7 %. The racial difference in wage penalties is 4.2 % points and decreases compared to OLS estimates. The other determinants of wages, which are time-varying, are reduced in magnitude in comparison to OLS, though most remain statistically significant for both black and white women.[11]

[9] With regards to OLS and combined sample estimates, there is a substantial negative effect of the interaction between hours worked at home and race.

[10] According to the U.S. Bureau of Labor Statistics, part-time is less than 35 hours; full-time is 35 hours or more per week.

[11] I have also estimated additional regressions with alternative specifications, which: (i) instead of an indicator for part-time, included a second-order polynomial of total hours worked, (ii) instead of the actual number of children, included indicators for the number of children (iii) included an additional dummy variable equal to one if spousal income is missing and equal to zero otherwise, (iv) included more variables that may affect spousal income, such as spouse's age, occupation, weekly hours worked. The estimates from these regressions (and IV regressions later on) were not significantly different from those reported.

**Table 3**
OLS and fixed effects estimates of the hourly wage elasticity.

| | Log of hourly wage | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | All | | Black women | | | White women | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Black | -0.159 | -0.082 | | | | | | |
| | (0.019) | (0.015) | | | | | | |
| Log of share of hours worked at home | -0.008 | -0.024 | -0.028 | -0.035 | -0.030 | -0.001 | -0.019 | -0.028 |
| | (0.005) | (0.004) | (0.010) | (0.008) | (0.006) | (0.005) | (0.004) | (0.004) |
| $12 \leq hgc < 16$ | | 0.176 | | 0.195 | | | 0.162 | |
| | | (0.023) | | (0.031) | | | (0.031) | |
| $hgc \geq 16$ | | 0.521 | | 0.534 | | | 0.503 | |
| | | (0.031) | | (0.049) | | | (0.039) | |
| Part-time | | -0.196 | | -0.205 | -0.057 | | -0.191 | -0.060 |
| | | (0.010) | | (0.016) | (0.011) | | (0.013) | (0.009) |
| Professional, technical and managerial | | 0.368 | | 0.328 | 0.078 | | 0.381 | 0.138 |
| | | (0.017) | | (0.027) | (0.018) | | (0.021) | (0.016) |
| Sales and clerical | | 0.193 | | 0.236 | 0.008 | | 0.169 | 0.045 |
| | | (0.014) | | (0.022) | (0.016) | | (0.018) | (0.015) |
| Other regressors | No | Yes | No | Yes | Yes | No | Yes | Yes |
| Fixed Effects | No | No | No | No | Yes | No | No | Yes |
| Adjusted $R^2$ | 0.019 | 0.277 | 0.008 | 0.271 | 0.089 | 0.000 | 0.266 | 0.111 |

Note: Clustered standard errors at the individual level in parentheses. The number of women in the sample is 2,485; the number of woman-year observations is 26,977 (Columns (1) and (2)). The number of black women is 885; the number of black woman-year observations is 9,645 (Columns (3), (4) and (5)). The number of white women is 1,600; the number of white woman-year observations is 17,332 (Columns (6), (7) and (8)). The dependent variable is the log of hourly wage in constant 2005 dollars. The main variable of interest is the log of hours worked at home as a proportion of total hours. Black is a dummy variable equal to one when an individual is black. *hgc* is the highest grade completed. Part-time is an indicator for whether the hours worked on site and at home weekly are less than 35. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income, and the number of children under 18 in the household.

## 4. Temporary child health problems

### *4.1. The instrument*

The OLS and FE estimates of the wage effect of working some hours at home are likely biased due to unobserved, time-varying omitted variables, as well as reverse causality.. These biases can be reduced by exploiting exogenous variation in the hours worked at home in an IV procedure. A temporary child health problem is proposed as a source of plausible exogenous variation with this goal in mind. The instrument is meant to capture a woman's short-term response to an unexpected increase in childcare responsibilities, leading her to choose working at least some hours from home rather than on-site.

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, under the assumption that these impairments are significant determinants of the demands of childcare but do not directlyaffect parental labor supply. The study finds that the effect of child disability on maternal labor supply is not significant for married women but negative and more substantial among female household heads. The results change dramatically when child health instruments are incorporated into the analysis.

Zan and Scharff (2018) use 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 mothers' employment only through health-related financial and time caregiving demands. Estimates that exploit the exogenous variation show that mothers are more likely to participate in the labor market when having higher monetary caregiving responsibilities, but less likely to participate when caregiving demands more of their time. The effects of caregiving on mothers' employment are underestimated when the instrument is not incorporated into the analysis.

The rationale behind the instrument in this study is that a temporary child health shock can increase the desirability of remote work in order to more flexibly care for the child. It is assumed that, after controlling for standard wage determinants and unobserved fixed effects, a temporary child health problem affects mothers' wages solely by prompting them to work more from home, likely only for a short period. Even assuming instrument validity, IV estimates are likely to represent a lower-bound effect, since non-employed women are not working, and therefore, the concept of working from home does not apply to them (Eckstein and Wolpin, 1989; Francesconi, 2002; Neal, 2004; Olivetti and Petrongolo, 2008; Keane and Wolpin, 2010; Adda et al., 2017).[12]

In constructing the child health instrument, a broad range of health problems causing temporary activity limitations and

[12] Neal (2004) includes non-employed women in his study by imputing wages. It is not very practical even with imputation methods for wages to include non-employed women in this study. The remote work offer and hours would also have to be imputed in some way adding more noise to the data.

participation restrictions are considered, 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 in constructing the instrument.

The questions about the duration of limitations and the time of injuries or accidents in the NLSCYA allow for the creation of a continuous child health history that distinguishes between a permanent and a temporary health problem. A temporary health problem is defined as one which occurs for one year only. Limitations, accidents, injuries, and mental conditions with a duration of more than one year, or health issues that occur in a specific year as a result of a permanent or pre-existing health condition can have more persistent effects on a mother's employment outcomes. Such conditions are considered permanent and are not taken into account in this analysis.

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 temporary health issues among childrren is 7.5 % for black-mother households and 12.4 % in white-mother households.[13] According to Table 4, pre-school children (less than 7 years of age) are more likely to experience a temporary health problem. This aligns with evidence from the U.S. and other countries showing that pre-school children spend a significant amount of time at home, which is also the most common setting for accidents involving young children (Pauline et al., 2007; Phelan et al., 2011).

In the analysis that follows, a temporary child health problem is represented by a dummy variable which equals one if at least one child in the household is temporarily afflicted and equals zero otherwise. The proportion of black mothers in the sample that have at least one child with a temporary health problem is 11.6 %. The corresponding proportion for white women is 15.6 %.

### 4.2. *Event study around the child health problem*

When a sudden child health issue occurs, mothers might be expected to adjust their labor supply to accommodate increased childcare needs. In this section, I investigate how a temporary child health problem might affect several labor market outcomes, including hours worked at home, total hours worked, hourly wages, labor force participation and earnings. The temporary child health problem effect estimates are presented in event study graphs in Fig. 1.

Panels A1 - A6 and B1 - B6 illustrate the relationship between the temporary health problem and different outcomes for black and white women, after controlling, additionally for the set of determinants used to produce the OLS and FE results of Table 3, the year the health issue occurs (which I refer to as year zero), up to four years before the health issue, and up to 6 years after the health issue. Each period $t$ (with exception of year zero) represents two years.[14] The treatment effect estimates are normalized to 0 in period $t = -1$ to -2 (i.e., either one or two years before the child health problem occurs).

The event study approach allows to track the magnitude and duration of the effect of the instrument and assess its exogeneity by exploiting individual-level variation in the timing of the child illness. If the temporary child health issue is independent of mothers' labor market outcomes, the estimates should not be significantly different from zero. All the panels of Fig. 1 show that the impact of a temporary child health problem on women's labor market outcomes is zero the years before the issue occurs. This suggests that there is no evidence of anticipation of such a shock and subsequent adjustments, a type of reverse causality, which could eventually bias IV estimates.

Panels A1 and A2 and B1 and B2 illustrate the relationship between a temporary child health issue and black and white mothers' hours worked at-home and total hours worked. Panels A1 and B1 show that a child health problem has a large and positive impact on hours worked at home at the time of the incident. This positive impact declines significantly in subsequent years for both black and white women. Panels A2 and B2 indicate that the child health problem does not have a significant effect on total hours worked. Therefore, any increase in the share of hours worked at home is likely due to an increase in the hours worked at home, rather than a decrease in the total hours worked.

Panels A3 and B3 reveal a negative and significant contemporaneous effect on hourly wages at the time the health problem occurs. Hourly wages tend to recover in the years following the unexpected child health issue. As part of this analysis, I also examined the effect of the child health issue on positive earnings. If the decrease in positive earnings is driven by women working fewer hours, then the earnings of mothers affected by a temporary child health issue should decrease more compared to their hourly wages. Panels A4 and B4 show that the impact of a child health problem on earnings closely mirrors the trajectory of the impact on the hourly wages.

[13] In epidemiology, 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 permanent and temporary health conditions in this dataset is 35.1 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).

[14] NLSY79 wages are recorded at an annual frequency until 1994, and biannually from 1994 onwards.

**Table 4**
Proportion of children with a temporary health problem by mother's race and child's age.

| Child's age | Black women | | White women | |
|---|---|---|---|---|
| | Health problem | *N* | Health problem | *N* |
| 0 | 0.061 | 651 | 0.081 | 1,393 |
| 1 | 0.201 | 725 | 0.270 | 1,553 |
| 2 | 0.140 | 816 | 0.272 | 1,711 |
| 3 | 0.129 | 924 | 0.216 | 1,902 |
| 4 | 0.102 | 1,021 | 0.194 | 2,064 |
| 5 | 0.087 | 1,127 | 0.178 | 2,199 |
| 6 | 0.102 | 1,240 | 0.166 | 2,350 |
| 7 | 0.086 | 1,344 | 0.145 | 2,481 |
| 8 | 0.083 | 1,444 | 0.146 | 2,598 |
| 9 | 0.068 | 1,528 | 0.120 | 2,689 |
| 10 | 0.072 | 1,589 | 0.112 | 2,723 |
| 11 | 0.068 | 1,636 | 0.108 | 2,767 |
| 12 | 0.071 | 1,677 | 0.087 | 2,773 |
| 13 | 0.067 | 1,704 | 0.102 | 2,776 |
| 14 | 0.066 | 1,715 | 0.078 | 2,748 |
| 15 | 0.043 | 1,703 | 0.061 | 2,701 |
| 16 | 0.039 | 1,679 | 0.057 | 2,633 |
| 17 | 0.039 | 1,646 | 0.069 | 2,554 |
| 18 | 0.042 | 1,601 | 0.046 | 2,479 |
| Total | 0.075 | 25,770 | 0.124 | 45,094 |

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

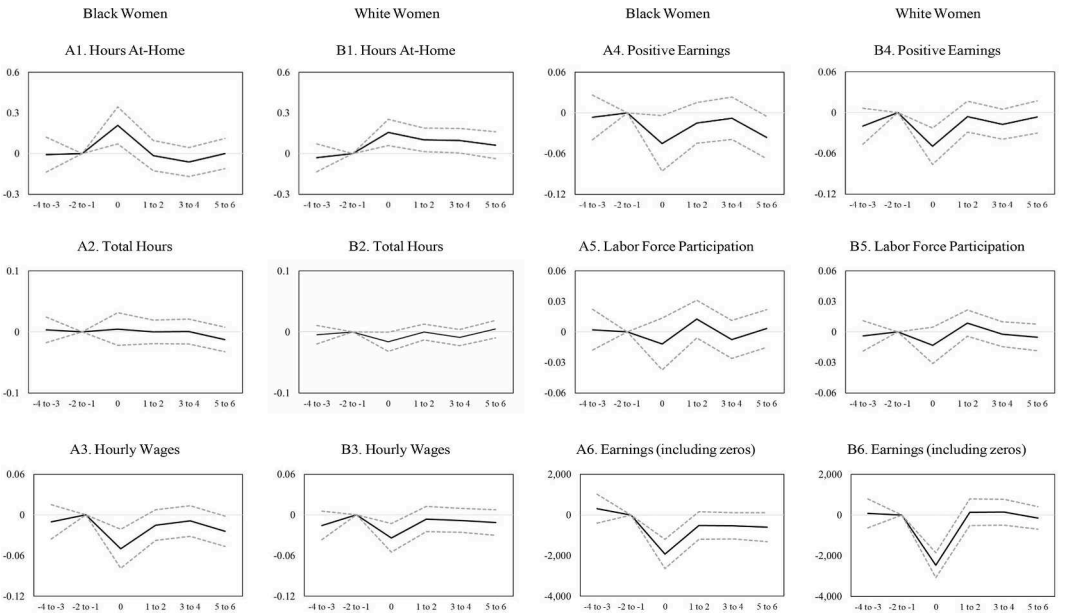


**Fig. 1.** Event study graphs before and after a temporary child health problem.
Notes: The graphs show coefficients from an event-study analysis. Event time is defined as years before and after a temporary health issue. The coefficients in period $t$ =-1 to -2 are normalized to 0. The number of black women is 713; the number of black woman-year observations is 6,104 (with exception of Panel A6, where the number of woman-year observations is 8,078). The number of white women is 1,263; the number of white woman-year observations is 10,565 (with exception of Panel B6, where the number of woman-year observations is 14,358). Hours worked at home is the log of hours worked at home. Total hours is the log of sum of hours worked at-home and on-site. Hourly wage is the log of hourly wage in constant 2005 dollars. Earnings are calculated by multiplying hourly wages (in constant 2005 dollars) by the total hours worked. Positive earnings are the log of earnings. The dashed lines represent 95 % confidence intervals.

I next consider the effect of a child health issue on mothers' labor force participation. Panels A5 and B5 show that women's labor force participation is not significantly affected the year the child health problem occurs. Women continue to participate as much as they did before the child health incident. In turn, Panels A6 and B6 plot the relationship between earnings, including zero earnings for the years mothers are not considered employed, and the temporary child health issue. If a decrease in earnings occurred because women did not participate due to a child health issue, one would expect a more persistent decline in earnings in the years following the

event. However, Panels A6 and B6 contradict such an expectation.[15]

These patterns suggest that a temporary child health issue indirectly influences women's hourly wages through the first-stage impact on the share of hours worked from home (exclusion restriction). As shown in Fig. 1, mothers in this sample receive a lower hourly wage not because they work fewer hours or stop participating due to a temporary child health issue. Any wage losses are potentially due to their choice to work some hours from home to accommodate better the increased children's needs arising suddenly due to a temporary health problem.

## 5. IV estimates

### 5.1. Estimation framework

The temporary child health problem 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 $t$, $Y_{i,b}$ and the log of hours worked at home as a proportion of total hours worked at time $t$, $F_{i,b}$

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

where $\alpha_i$ is an unobserved individual fixed effect, $X_{i,t}$ is a vector of time-varying individual characteristics including part-time, age, age squared, marital status, the number of children in the household, spousal income 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 to produce the OLS and FE results presented earlier.

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

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

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

As mentioned previously, the main assumption in the IV procedure is that a child temporary health issue increases the share of hours worked at home but does not otherwise affect wages beyond the change in work venue, after controlling for standard wage determinants as well as unobservable time-invariant characteristics. The main challenge to a causal interpretation of the effect of remote work on wages in this framework is the possible impact of a temporary child health shock on wages through alternative channels such as the choice of occupation, mental health of the mother and time-variant unobserved determinants of earnings contained in $u_{i,t}$ . In order to address these threats to causality, part-time is included in the regressions as well as indicators for different occupational categories.

Supplementary to the event study analysis, a validity test and alternative first-stage regressions that examine the direct impact of a temporary child health shock on job changes and the mental health of the mother are also presented below.

### 5.2. Reduced form results

Table 5 presents reduced form results for the combined sample (Columns (1) and (2)) and by race (Columns (3) - (10)) of the effect of a temporary child health problem on the log of share of hours worked at home and the log of hourly wages. Columns (1) and (2) present the OLS estimates using the combined sample and show that there is a negative impact of race on both the hours worked at home and hourly wages.[16] Columns (3) and (4) and (7) and (8) of Table 5 display the first-stage results for two-stage least squares, without and with fixed effects for black women and white women, respectively .

The coefficient for a temporary child health problem is large in magnitude and statistically significant for both races with or without fixed effects. A temporary child health problem substantially increases the share of hours worked at home for both black and white women. The increase in this share is 22.2 % without fixed effects and 20.1 % with fixed effects for black women. For white women, the increase in the share of hours worked at home is 26.7 % without fixed effects and 17.9 % with fixed effects.

Columns (5) and (9) and (6) and (10) present the reduced form effect of a child health problem on mean hourly wages without and with fixed effects. With fixed effects included, there is a precisely estimated negative effect of a temporary child health problem on mean hourly wages for both races. Mean wages are lower by 4 % and 2.9 % for black women and white women, respectively. The ratios of the coefficients corresponding to the temporary child health variable in Table 5 indicate that the IV estimates of the wage effect of

[15] I acknowledge that this analysis conditions on employed women, which could result in selection bias with estimates capturing sample composition changes. To gain further insight, I expand the sample to include non-working mothers with zero earnings and perform the same event study analysis. The extended sample includes 738 black women with a total of 8,266 woman-year observations, and 1,325 white women with 14,833 woman-year observations. The regression results from this analysis are not different from those presented in Panels A6 and B6. This increases confidence in the exogeneity of the instrument. While sample selection is a valid concern, it is independent of the temporary child health issue that is the main focus of this study.

[16] With regards to OLS and combined sample estimates, there is a substantial negative effect of the interaction between the health problem and race on the share of hours worked from home.

**Table 5**
Reduced form estimates.

| | All | | Black women | | | | White women | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Log of share of hours worked at home | Log of hourly wage | Log of share of hours worked at home | | Log of hourly wage | | Log of share of hours worked at home | | Log of hourly wage | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
| Black | -0.235 | -0.077 | | | | | | | | |
| | (0.050) | (0.015) | | | | | | | | |
| Child health problem | 0.267 | -0.016 | 0.222 | 0.201 | -0.005 | -0.040 | 0.267 | 0.179 | -0.017 | -0.029 |
| | (0.045) | (0.011) | (0.071) | (0.064) | (0.018) | (0.014) | (0.056) | (0.049) | (0.014) | (0.011) |
| $12 \leq hgc < 16$ | 0.182 | 0.172 | -0.083 | | 0.198 | | 0.362 | | 0.155 | |
| | (0.069) | (0.023) | (0.135) | | (0.034) | | (0.067) | | (0.032) | |
| $hgc \geq 16$ | 1.006 | 0.497 | 0.465 | | 0.518 | | 1.288 | | 0.478 | |
| | (0.109) | (0.031) | (0.193) | | (0.050) | | (0.122) | | (0.039) | |
| Part-time | 0.463 | -0.207 | 0.498 | 0.431 | -0.222 | -0.069 | 0.424 | 0.403 | -0.199 | -0.070 |
| | (0.037) | (0.010) | (0.054) | (0.045) | (0.015) | (0.011) | (0.050) | (0.041) | (0.013) | (0.009) |
| Professional, technical and managerial | 0.411 | 0.359 | 0.518 | 0.137 | 0.310 | 0.074 | 0.350 | 0.007 | 0.374 | 0.138 |
| | (0.064) | (0.017) | (0.107) | (0.086) | (0.028) | (0.019) | (0.081) | (0.084) | (0.021) | (0.016) |
| Sales and clerical | -0.207 | 0.198 | -0.118 | -0.198 | 0.240 | 0.014 | -0.252 | -0.440 | 0.174 | 0.058 |
| | (0.046) | (0.014) | (0.067) | (0.071) | (0.022) | (0.017) | (0.061) | (0.075) | (0.018) | (0.016) |
| Other regressors | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Fixed effects | No | No | No | Yes | No | Yes | No | Yes | No | Yes |
| AR test | 1.97 | | 0.07 | 7.89 | | | 1.51 | 7.64 | | |
| | (0.161) | | (0.797) | (0.005) | | | (0.220) | (0.006) | | |
| Adjusted $R^2$ | 0.079 | 0.271 | 0.054 | 0.024 | 0.258 | 0.075 | 0.081 | 0.028 | 0.261 | 0.094 |

Note: Clustered standard errors at the individual level in parentheses. The number of women in the sample is 2,485; the number of woman-year observations is 26,977. The number of black women is 885; the number of black woman-year observations is 9,645. The number of white women is 1,600; the number of white woman-year observations is 17,332. The dependent variable in Columns (1), (3), and (7) for OLS estimates and in Columns (4) and (8) for FE estimates is the log of hours worked at home as a proportion of total hours worked (see Eq. (2)). The dependent variable in Columns (2), (5), and (9) for OLS estimates and in Columns (8) and (10) for FE estimates is the log of hourly wage in constant 2005 dollars. Black is a dummy variable equal to one when an individual is black. *hgc* is the highest grade completed. Part-time is an indicator for whether the hours worked on site and at home weekly are less than 35. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income, and the number of children under 18 in the household. The AR test is for the Anderson-Rubin test with the p-values in parentheses.

working at home will be negative and significant.

The Anderson-Rubin (AR) tests in Table 5 provide evidence about the strength and relevance of the instrument.[17] The AR test indicates the share of hours worked at home estimates to be significant if the child health problem is significant in the reduced form regression. FE estimates in Columns (4) and (8) show that this is the case. The assumption that the child health issue only affects hourly wages indirectly through its effect on the share of hours worked at home implies that the change in the share of hours worked at home has a causal impact on wages.

### 5.3. IV results

The IV estimates for the elasticity of hourly wage with respect to the share of hours worked at home by race, with fixed effects, are reported in Table 6. As in the FE regressions, education indicators are dropped as controls since they do not vary over time. The IV estimates of the share of hours worked at home are precisely estimated for both races. The elasticity of hourly wages with respect to share of hours worked at home is -0.201 for black women and -0.165 for white women.[18]

An increase in weekly hours worked at home is associated with a significant decrease in hourly wages. Specifically, if work at home increases to an average of 5 h on weekly basis, the wage penalty is 66.3 % for black women and 33.9 % for white women. These results reveal important racial disparities in the negative effects of working from home on hourly wages, with black women experiencing a potential larger decrease in hourly wages, compared to their white counterparts. Possible reasons for larger absolute estimates and wage penalties found in the IV framework, compared to OLS and FE, are discussed in the following section.

[17] Anderson and Rubin (1949) developed an alternative approach to inference. Keane and Neal (2022; 2023) thoroughly discuss the advantages of the AR test by highlighting the limitations of the t-test and F-test in the context of two-stage least squares regression. They argue that it is essential to report the AR test results for two-stage least squares regression, instead of the other tests. They show that the AR test is robust to weak instrument problems, meaning it always has correct size.

[18] No precisely estimated interactions with the share of hours worked at home and part-time, occupation, age of women, marital status, spousal income, number of children under 18, number of children under 6 and above 6 were found, suggesting a lack of substantial heterogeneous treatment effects.

**Table 6**
IV estimates of the hourly wage elasticity.

| | Log of hourly wage | |
|---|---|---|
| | Black women (1) | White Women (2) |
| Log of share of hours worked at home | -0.201 | -0.165 |
| | (0.090) | (0.068) |
| Part-time | 0.018 | -0.004 |
| | (0.041) | (0.030) |
| Professional, technical and managerial | 0.101 | 0.139 |
| | (0.025) | (0.018) |
| Sales and clerical | -0.026 | -0.015 |
| | (0.026) | (0.034) |
| Other regressors | Yes | Yes |
| Fixed effects | Yes | Yes |

Note: Clustered standard errors at the individual level in parentheses. The number of black women is 885; the number of black woman-year observations is 9,645. The number of white women is 1,600; the number of white woman-year observations is 17,332. The dependent variable is the log of hourly wage in constant 2005 dollars (see Eq. (1)). The main variable of interest is the log of hours worked at home as a proportion of total hours. Part-time is an indicator for whether the hours worked on site and at home weekly are less than 35. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income, and the number of children under 18 in the household.

### 5.4. *Instrument validity*

The event study analysis presented in Fig. 1 shows that there is neither anticipation of a temporary child health shock nor persistent subsequent labor market outcome adjustments after its occurrence. The first-stage estimates in Table 5 illustrate that the child health instrument is relevant and strong. Nonetheless, a major concern with the instrument, is the possible violation of exclusion restriction due to a direct effect of a child health shock on mothers' productivity. It may be that caring for a sick child requires more of a mother's effort and attention during or after work hours. This, in turn, may decrease productivity on job, and as a consequence mothers' wages, even when not increasing working from home.

Evidence for the exogeneity of the instrument is clearly more difficult to establish, especially in the absence of over-identification (multiple instruments). In an effort to present further evidence and justify instrument exogeneity, and hence the validity of the instrument, I estimated reduced-form child health problem effects on wages in samples in which a temporary child health issue has little effect on the hours worked from home. Estimating reduced form effects in samples with no-first stage and no substantial reduced form effects should strengthen confidence in instrument exogeneity (Altonji et al., 2005; Angrist et al., 2010). If a sudden child health problem affects productivity in other ways than working some hours at home, then wages in these no-first stage samples should decrease due to this shock.

The no-first stage samples include mothers who are less likely to work from home due to their occupation and the tasks they perform. I use O*NET ratings about the type and frequency of interactions with other people that are required as part of the job, as well as the work settings that influence the nature of work. More precisely, I use means of available O*NET ratings on occupations for physical proximity (i.e. to what extent does this job require the worker to perform job tasks in close physical proximity to other people?), face-to-face discussions (i.e. how often does the worker have to have face-to-face discussions with individuals or teams in this job?) and importance of working with others in a group or team (i.e. how important is it to work with others in a group or team in this job?) for 3-digit classification of occupations.[19] Examples of occupations with higher ratings in this sample are registered nurses, social workers, schoolteachers, salespersons, production checkers, graders, and sorters. Examples of occupations with lower ratings are financial specialists, accounting, auditing and other clerks, housekeepers, bill and account collectors, machine operators.

The no-first stage samples include mothers with job tasks that are more difficult to be performed remotely, as they require higher levels of physical proximity, face-to-face discussions, and work with others in a group or team. As the estimates in Table 7 show, there are no significant effects of a temporary child health shock in any of the subsamples. That is, no effect for the hours worked at home or wages of black and white mothers. This lends additional support to the main assumption that the temporary child health shock does not have a direct effect on wages but is rather channeled through working some hours from home.[20] Additional analysis related to both the

[19] The occupational classification used to construct consistent occupational codes for the period under consideration is provided in Autor and Dorn (2013). The thresholds for classifying the constructed occupations by the difficulty of performing tasks remotely were determined using the sample's median ratings.

[20] I have also estimated reduced-form child health problem effects on labor force participation. The child health problem coefficients are not statistically significant different from zero.

**Table 7**
Alternative reduced form estimates - validity test.

| | Child health problem | |
|---|---|---|
| | Black Women (1) | White women (2) |
| Log of share of hours worked at home | 0.025 (0.146) | 0.039 (0.120) |
| Log of hourly wage | -0.056 (0.039) | 0.002 (0.023) |
| Other regressors | Yes | Yes |
| Fixed effects | Yes | Yes |

Note: Clustered standard errors at the individual level in parentheses. The number of black women is 327; the number of black woman-year observations is 945. The number of white women is 550; the number of white woman-year observations is 1,686. The dependent variables are the log of share of hours worked at home and the log of hourly wage in constant 2005 dollars. The main variable of interest is the child health problem. Other regressors include part-time status, indicators for occupations, age, age squared, spousal income, and the number of children under 18 in the household.

validity of the instrument and mechanisms for the effect are discussed in the next section.[21]

## 6. Magnitudes and mechanisms

The IV estimates in Table 6 should be interpreted as local average treatment effects (LATE) that capture the change in mean hourly wages among the subset of women who are just on the margin between working exclusively on-site and working some hours at home, and who are prompted to do so as a result of at least one child in the household developing a temporary health issue. This subsample, referred to as the "compliers" (see Angrist et al. (1996)), contains the subset of women who would not have worked at home if their child had not become sick.

The LATE estimates reported in Columns (1) and (2) of Table 6 are significantly larger in magnitude - more than 6 to 7 times - compared to the corresponding FE estimates for white and black women in Table 3, respectively. The patterns of results suggest that OLS and FE estimates are substantially biased toward zero (under-estimated) and there is positive selection.

The sample means in Table 2 are already suggestive of positive selection on unobservables because women who work at home are, on average, more highly educated and more often work in professional, technical, or managerial roles. That is, there is clear positive selection on observables so it may not come as a surprise that there is positive selection on unobservables as well. Positive selection of mothers into more flexible jobs was also evidenced in the study of the gender earnings gap among MBA graduates by Bertrand et al. (2010) and in the study of telecommuting by Glass and Noonan (2016). In agreement with Neal (2004), estimates in this study may indicate that black women are more positively selected into work than white women, albeit in the more limited context of remote work only.

Note that negative wage effects of such magnitudes produced within an IV framework are not unprecedented in the wider literature on flexible working conditions. This is true even for estimates that represent average treatment effects (ATE) as opposed to LATE estimates. Bertrand et al. (2010) find a remote-work wage penalty among female MBA graduates of over 20 %. However, the wage penalty among women that choose a new job with flexible working hours is much higher, reaching over 60 %. These latter estimates are produced from fixed-effects regressions on a selected sample of highly educated women. The IV estimates with fixed-effects in the present study are derived from nationally representative data of black and white women across different education levels.

In order to examine the mechanisms and the extent to which a child health problem is leading to other changes besides an increase in hours worked at home, which could also bring about wage penalties to remote work, alternative first-stage estimates are presented in Table 8 for black mothers and in Table 9 for white mothers. The alternative first-stage estimates examine the effect of the temporary child health problem on job changes and deterioration in the mother's mental health.[22]

In Columns (1) - (3) of Tables 8 and 9, employer and occupation information in the NLSY79 is exploited to construct three different measures of a job change (see also Kambourov and Manovskii (2009)). The first definition of a job change is an indicator for a woman switching employer during the year the temporary child health issue occurs. The second definition takes into account a change of employer and occupational category in the time frame of the temporary child health problem. The third definition tracks whether there was a job change as in the second definition but does not consider the switch to be a change if the employer or occupational category reverts to the previous one in a later period. This third definition avoids biases due to cycling between jobs (see Neal (1999)). The

[21] For robustness, I have also estimated IV regressions that include the measures for physical proximity, face-to-face discussions, and the importance of working with others in a group or team, in addition to the variables included in the regressions in Table 6. The estimates for the share of hours worked at home are not significantly different from those reported in Table 6.

[22] Further to these alternative first-stage estimates, I performed event study analyses similar to the ones described in Section 4.1. No indications of either anticipatory, contemporaneous, or more persistent effects of a temporary child health issue on mothers' job mobility or mental health were found.

**Table 8**
Alternative first-stage estimates - black women.

| | Job change | | | Mental |
|---|---|---|---|---|
| | Definition 1 (1) | Definition 2 (2) | Definition 3 (3) | Health (4) |
| Child health problem | -0.016 | 0.002 | 0.010 | -0.002 |
| | (0.017) | (0.012) | (0.011) | (0.004) |
| Part-time | 0.111 | 0.018 | 0.018 | 0.002 |
| | (0.014) | (0.010) | (0.009) | (0.005) |
| Professional, technical | 0.006 | 0.068 | 0.054 | 0.000 |
| and managerial | (0.021) | (0.018) | (0.017) | (0.007) |
| Sales and clerical | 0.010 | 0.009 | 0.007 | -0.012 |
| | (0.018) | (0.016) | (0.016) | (0.006) |
| Other regressors | Yes | Yes | Yes | Yes |
| Fixed effects adjusted $R^2$ | Yes0 | Yes0 | Yes0 | Yes0 |
| | .022 | .011 | .009 | .033 |

Note: Clustered standard errors at the individual level in parentheses. The number of black women is 713; the number of black woman-year observations is 6,104. Definition 1 of a job change is an indicator for a woman switching employer during the year the temporary child health issue occurs. Definition 2 takes into account a change of employer and occupational category in the time frame of the temporary child health problem. Definition 3 tracks whether there was a job change as in definition 2 but does not consider the switch to be a change if the employer or occupational category reverts to the previous one in a later period. Part-time is an indicator for whether the hours worked on site and at home weekly are less than 35. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income, and the number of children under 18 in the household.

**Table 9**
Alternative first-stage estimates - white women.

| | Job change | | | Mental |
|---|---|---|---|---|
| | Definition 1 (1) | Definition 2 (2) | Definition 3 (3) | Health (4) |
| Child health problem | 0.018 | 0.006 | 0.008 | 0.008 |
| | (0.012) | (0.008) | (0.008) | (0.005) |
| Part-time | 0.068 | 0.039 | 0.032 | 0.004 |
| | (0.010) | (0.008) | (0.007) | (0.006) |
| Professional, technical | -0.020 | 0.021 | 0.021 | 0.003 |
| and managerial | (0.015) | (0.013) | (0.012) | (0.009) |
| Sales and clerical | 0.002 | 0.018 | 0.017 | 0.005 |
| | (0.015) | (0.013) | (0.012) | (0.008) |
| Other regressors | Yes | Yes | Yes | Yes |
| Fixed effects adjusted $R^2$ | Yes 0 | Yes 0 | Yes 0 | Yes 0 |
| | .011 | .006 | .005 | .093 |

Note: Clustered standard errors at the individual level in parentheses. The number of white women is 1,263; the number of white woman-year observations is 10,565. Definition 1 of a job change is an indicator for a woman switching employer during the year the temporary child health issue occurs. Definition 2 takes into account a change of employer and occupational category in the time frame of the temporary child health problem. Definition 3 tracks whether there was a job change as in definition 2 but does not consider the switch to be a change if the employer or occupational category reverts to the previous one in a later period. Part-time is an indicator for whether the hours worked on site and at home weekly are less than 35. Other regressors include age, age squared, an indicator for whether the woman is married, spousal income, and the number of children under 18 in the household.

estimates in Columns (1) - (3) do not suggest that there is any substantial job mobility due to a temporary child health issue that could drive the wage penalties for black women or white women.[23,24]

In Column (4), the effect of a temporary child health problem on the mental health of the mother is examined. Mental health is measured by a diagnosis of depression, emotional, nervous, or psychiatric problems. A decrease in mental health could lead to less productivity and lower wages. The results do not indicate a significant increase in the probability of such a diagnosis following a temporary child health problem for either race.

These alternative first-stage regression results, along with the event study analysis and the no-first stage estimates presented in Table 7 rule out the main threats to causal interpretation in the IV procedure. Moreover, these additional results eliminate several of

[23] The same analysis was performed using a finer 3-digit classification of occupations. The occupational classification used to construct consistent occupational codes for the period under consideration is provided in Autor and Dorn (2013). The conclusions are the same.

[24] Arntz et al. (2022) present descriptive evidence for Germany about mothers' wages that remain unchanged after taking up work from home unless they change employers. I have estimated IV regressions with fixed effects only for women who do not change job. The estimates for the share of hours worked at home are not significantly different from those reported in Table 6.

the possible mechanisms that might drive the large wage penalties to remote work. By process of elimination, the most plausible mechanisms driving the remote work wage penalties are therefore promotion bias, performing less "valuable" work assignments, and a decrease in productive social interactions with colleagues.

Supporting the promotion bias mechanism in particular, Bloom et al. (2015) find that telecommuting applies to a range of tasks including jobs in sales and secretarial assistance and that work from home is associated with reduced rates of promotion of about 50 %. Note that in the sample of this study a substantial proportion of black women (33.2 %) and white women (35.8 %) work in sales and clerical occupations. Golden and Eddleston (2020) also find that the extent of telecommuting is negatively related to promotions. Among those working in higher skill jobs, Bertrand et al. (2010) argue that MBA mothers are likely to be forced out, or opt out, of the "fast-track" when they choose more flexible work arrangements. The more there is a tournament or up-or-out structure to the occupation, the more "task-shift" there is likely to be when choosing to work remotely. In the current sample, 21.5 % of black women and 33.7 % of white women work in professional, technical, or managerial occupations where tournament and up-or-out employment structures are most prevalent.

It is generally thought that promotion requires not only expertise, but also social capital and interaction (Williams, 2000; Blair-Loy, 2006; Golden, 2008; Williams et al., 2013; Allen et al., 2015; Bloom et al., 2015). It has been argued that women especially need to demonstrate their skills, take up more prominent tasks and have access to networks of decision makers that enhance their visibility in order to increase opportunities for workplace progression (Milgrom and Oster, 1987; Hersby et al., 2009; Cristea and Leonardi, 2019; von Essen and Smith, 2023). Remote work may render this much more difficult to achieve.

In addition, racial and gender homophily - the tendency of workers to associate and form ties with similar others - can create substantial divisions and undermine networking and the workplace information that is received (McPherson et al., 2001). A woman's lack of white male connections has been argued to be a reason for worse positioning in job ladders, lower promotion probabilities and wages compared to men (Milgrom and Oster, 1987; Blair-Loy, 2001; Cassidy et al., 2016; Bertrand et al., 2019). Recent evidence confirms that a male-to-male promotion advantage exists, especially in positions where the employee works in close physical proximity to the manager (Cullen and Perez-Truglia, 2023). This emphasizes the crucial role of face-to-face interaction and socialization that may be severely compromised by working from home.

The results in this study suggest that the loss of networking and visibility opportunities may hurt black women even more than white women. This is an interesting new finding of the study, considering that the previous literature on racial and gender wage gaps has concentrated mostly on differences in hours and occupations (see Huffman and Cohen, 2004; Alonso-Villar et al., 2012; Gicheva, 2013; Cha and Weeden, 2014; Goldin, 2014; Del Rio and Alonso-Villar, 2015; Erosa et al., 2022).

## 7. Conclusion

This paper examines the racial gap in wage penalties among women due to remote work by linking pre-pandemic data on women in the National Longitudinal Study of Youth (NLSY79) to their children in the NLSY79 Child and Young Adult Survey (NLSCYA). It is the first study that focuses exclusively on racial differences accounting for the impact of remote work on wages using data that allows to provide a range of estimates, including further to OLS and FE, IV results. The findings should be relevant for assessing the differential costs of physical distancing by race for females both in the present and the foreseeable future.

OLS estimates indicate that when time worked from home increases to 5 h per week, black women's wages are expected to decrease by 11.5 %, while white women's wages are expected to fall by 3.9 %. This results in a 7.6 % point racial gap in wages due to the increase in hours worked from home. FE estimates, which take advantage of the longitudinal aspect of the data, yield a wage penalty for black women of 9.9 % and a wage penalty for white women of 5.7 % for the same increase in hours worked from home for each race. The racial wage gap decreases to 4.2 % points compared to the OLS estimates. IV estimates that include fixed effects and use the child health instrument indicate wage penalties of 66.3 % for black women and 33.9 % for white women when hours worked from home increase to 5 per week. This results in a 32.4 % point racial gap in wages. The IV procedure produces larger absolute magnitudes in the remote work wage penalty compared to OLS and FE estimates.

The IV estimates should be interpreted as LATE that pertain to working women who are induced to work from home due to a temporary child health problem. The larger IV estimates also imply positive selection into working at home. The co-existence of positive selection and a substantial wage penalty to remote work, combined with the event study analysis and results of additional reduced-form regressions, suggest that the most likely mechanisms underlying the wage penalties are promotion bias, women being assigned less "valuable" work assignments, and women having less productive social interactions with colleagues.

The problem of promotion bias, task reassignment and loss in social capital appears to be differentially worse for black women. Why that may be the case requires further research. The results in this study suggest that the loss of networking and visibility opportunities may be the main mechanisms explaining why black women are hurt relatively more than white women. Better data would enable a more precise disentangling of the various mechanisms and a further comparison to alternative possibilities such as negative signaling, screening and statistical discrimination.

The OLS, FE and IV estimates in this study offer rare evidence on the racial gap in remote-work wage penalties. Only similar future attempts can help establish a body of evidence that would allow for a more accurate assessment of the plausibility of the range of estimates in this particular study. Note that the results are likely to be lower-bound estimates since it is not straightforward to account non-working women in a remote work context. These latter women may experience more severe losses in human capital and earnings.

In closing, it is also important to note that public discussions and policy proposals involving remote work often center around how to increase the supply of work-from-home opportunities, rather than directly addressing the wage penalty. The results in this study suggest that wage penalties should also become a major focus. As hybrid offices are on the rise and new technologies for remote work,

including the emerging metaverse, are unveiled, it will be important to monitor how race and gender inequality evolve due to the increase in working from home.

## Data availability

Data will be made available on request.

## Supplementary materials

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.euroecorev.2024.104864.

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