When Does Labor Market Flexibility Reduce Gender Wage Gaps?

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Abstract

In this paper we document a new empirical fact about the U.S. labor market: over the past four decades, the gender wage gap among part-time workers was 10-30 percentage points smaller than the gender wage gap among full-time workers. This is a new empirical fact is predicted by the theory that increased labor market flexibility reduces gender wage gaps (Goldin, 2014). We refine this prediction by providing evidence that flexibility that is demanded endogenously, as is the case with pursuing part-time work, reduces gender wage gaps; by contrast flexibility that is supplied exogenously, specifically through state-mandated family leave policies like the Family and Medical Leave Act (1993) and its antecedents, are capitalized into slower wage growth for women when compared to men. Using an event-study design, we show that mandated family leave laws were partially responsible for the stagnation of women's wages in the mid 1990s, bring resolution to an important puzzle in labor economics.

JEL Classification: D13, D14, J12, J15, J16, J24, K11, K12, and K36.

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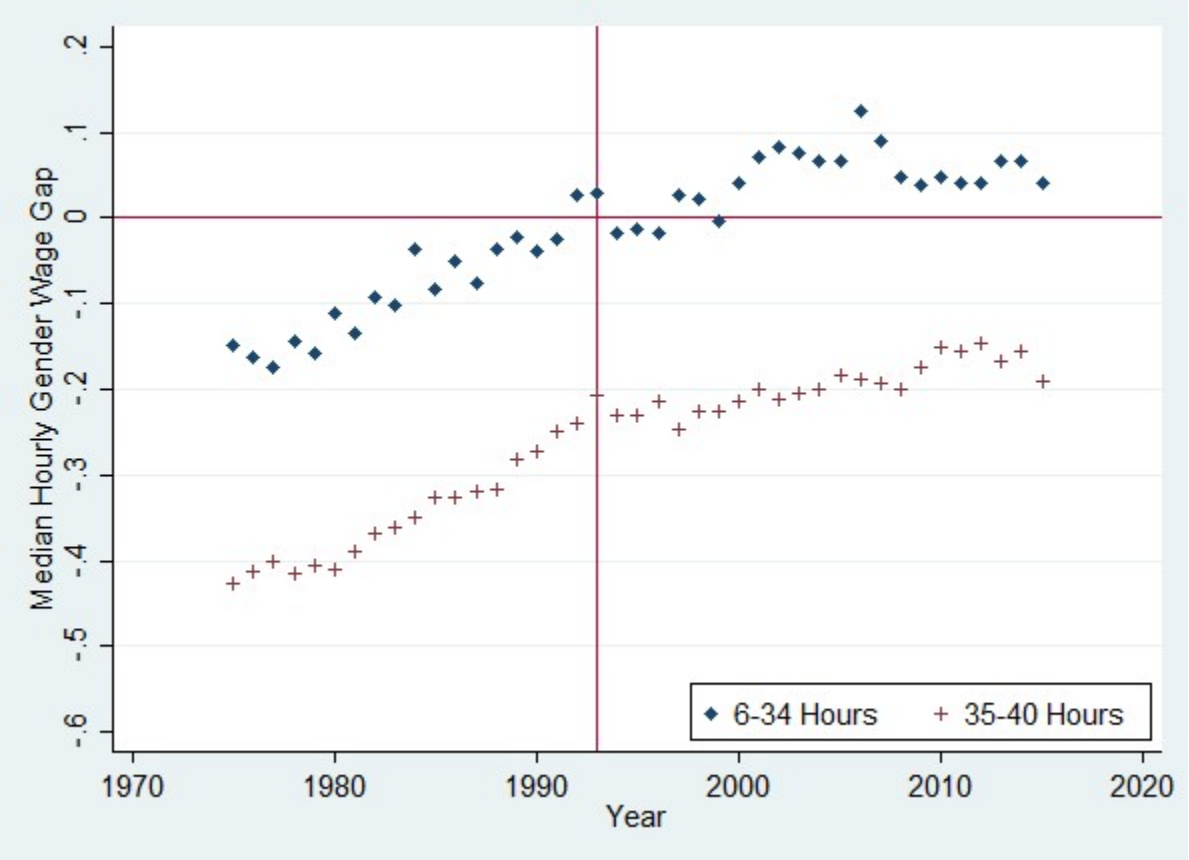
1 Introduction

At the 126th annual meeting of the American Economic Association, Claudia Goldin delivered the presidential address entitled "A Grand Gender Convergence: Its Last Chapter." In the address, Goldin foreshadowed that the final frontier in gender wage convergence "must involve alterations in the labor market, in particular changing how jobs are structured and remunerated to enhance temporal flexibility" (Goldin, 2014). In this paper we study two aspects of labor-market flexibility. First we study the extent to which existing temporal flexibility in the labor market, through part-time employment, affects gender wage gaps and gender wage convergence. Second, we study how state-mandated policies promoting temporal flexibility in the labor market, notably the Family and Medical Leave Act (1993) and its antecedents, affected gender wage gaps and gender wage convergence.

There is simple reason to believe that temporal flexibility that arises endogenously in the market through firms and workers matching to more flexible work contract could have a different impact on wages when compared to temporal flexibility that is exogenously supplied through state and federal mandates. The work of Antecol et al. (2018) showing that family-friendly tenure policies at universities produce the unintended consequence of advantaging men relative to women in the tenure process is one such example. An earlier example comes from the work of Gruber (1994), in which he shows that state maternity-leave laws lower the wages of women early in their career. These two examples, among many others, provide the motivation for this paper in which we provide evidence on the wage impacts of endogenous versus exogenous temporal flexibility in the labor market.

We start our analysis by looking at the gender wage gaps between part-time workers to see whether they are indeed smaller than the gender wage gaps faced by women in full-time work. The existence of smaller gender wage gaps in part-time work is a natural implication of Goldin (2014), which we test by documenting the part-time and full-time gender wage gaps over the past four decades (1975 – 2015), using CPS and ACS data. Most work that considers long-term trends in gender wage gaps either ignores part-time workers or does not separately consider part-time workers. Two studies that do consider part-time gender wage gaps Blank (1990) and Hirsch (2005) focus on short and more recent time windows, 1987 and 1995–2002, respectively. Going back as far as 1975 we find that the gender wage gap among part-time workers has consistently been 10-30 percentage points lower than the gender wage gap among full-time workers. The difference in the gender wage gaps between part-time and full-time workers is illustrated in Figure 11,

Figure 1: **Median Wage for Part-time and Full-time Workers (1975 - 2015)**



Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

in which we we plot the unconditional gender wage gap using the raw data from the CPS covering 1975–2015. This is a new striking empirical fact which is robust to conditioning on observables, instrumenting for labor supply using state variation in income taxes, and excluding workers below the minimum wage.[1] Further, it is consistent with the results in Blank (1990) that men face a larger part-time wage penalty than women, hence the smaller gender wage gap between part-time women and men than full-time women and men. In pointing out the fact that gender wage gaps are smaller among part-time than full-time workers we are not saying that part-time work is superior to full-time work but rather we are establishing a new fact about the U.S. labor market at the intersection of gender and inequality.

In the second part of this work we focus on the effect of state-mandated policies promoting labor market flexibility on both the gender wage gap and the rate of gender wage convergence. From Figure 11 it is evident that there is a break in the trend towards smaller

[1] Summers (1989) argues that government employer mandates may not be fully capitalized into wages in the presence of wage rigidities like minimum wage laws.

gender wage gaps in the 1990s. Prior to this trend break, the gender wage gap is decreasing by roughly 1 percentage point per year; following the trend break, gender wage convergence stalls. We first establish descriptively that this trend break occurs concurrently with the passage of the Family and Medical Leave Act (FMLA) in 1993. The FMLA required large employers (> 50 employees) to provide 12 weeks of job-protected leave to workers for family and medical reasons. The federal policy, which applied only to workers who worked $>$ 1,250 hours in the previous year. Following Kleven et al. (2013), we fit a piece-wise linear regression model to the data and showing that a trend break in the FMLA year maximizes its explanatory power. Next we leverage variation in the timing of state laws that were antecedents to the FMLA in an event-study design to show both econometrically and descriptively that these type of legislation also stagnated gender wage convergence. According to Blau and Kahn (2017) gender wage gaps have reduced over time both in the US and many other developed countries (Blau and Kahn, 2007). A key puzzle in this literature relates to the observation that most of this convergence happened during the 1980s and stagnated thereafter.

Taken together, our results suggest that the temporal flexibility that arises naturally in labor markets reduces wage inequality. These results are consistent with Goldin (2014) who states that "the explanation [of the residual gender wage gap] will rely on labor market equilibrium with compensating differentials and endogenous job design." By contrast government-imposed "flexibility" can slow the rate of wage convergence (Gruber, 1994). We estimate that in the absence of the FMLA, white women and black women would have achieved gender wage parity with white men by 2016 and 2023.

2 Literature Review

According to Bertrand et al. (2010) and Noonan et al. (2005) gender differences in wages among workers in the U.S. are small at the onset of a worker's career and that they grow over time. Albrecht et al. (2018) shows that this pattern is similar in Sweden. Kleven et al. (2019) provide evidence that a key driver gender wage gap is the "mommy penalty": across many different countries women's wages fall by 20% following the birth of their first child. [2] Consistent with this finding, Goldin (2014) argues that preferences for work flexibility can explain these gender differences in wage via a Rosen (1986), Pitts (2003), and Mas and Pallais (2017) compensating wage differential story in which women who value flexibility pay for it in lower wages.

[2]This stands in contrast to the statistical discrimination results of Altonji and Pierret (2001), who show that statistical discrimination on race decreases with labor market experience.

The sorting in the labor market on flexibility occurs on both sides of the labor market. On the labor supply side, women's labor force participation is more sensitive than men's to the availability of affordable childcare (Cortes and Pan, 2017). Similarly, Altonji and Usui (2007) show that women are more likely to take unpaid leave from work than men while. Correspondingly, Acemoglu and Pischke (1999) find that women who value flexibility do not sort to firms with high human capital requirements, and that firms, separately, weary of workplace interruptions by women, shy away from hiring women. On the intensive margin of hours worked, Rebitzer and Taylor (1995) show that highly competitive jobs are unlikely to offer short-hour jobs due to the signal they send about the workers willingness to respond to financial incentives to working longer hours.[3] Perhaps these responses by firms and workers explain why Averett and Hotchkiss (1996) find that the full-time wage premium kicks in earlier for men than for women: 33 hours worked versus 37-39 hours and Cook et al. (2018), using the universe of Chicago Uber drivers, estimate a gender pay gap of 7%, which is fully explained by preferences for flexibility, preferences for driving speed and worker experience.

An important gap in the literature is an application of this flexibility hypothesis to the study of part-time workers. Most work that considers the long-term trend in gender wage gaps either ignores part-time workers or includes wage-gap results that do not separately consider part-time workers. The two studies that do consider part-time gender wage gaps, Blank (1990) and Hirsch (2005), focus on specific windows of time - 1987 and 1995 to 2002, respectively. Our paper contributes to the literature by documenting the part-time gender wage gap and by showing that it has been historically lower than the full-time gender wage gap. This finding is consistent with the Blank (1990) result of a larger part-time wage penalty for men than women (35% versus 24%), which implies a smaller gender wage gap.

The closest paper in the literature to our study regarding the effects of the FMLA on women's wages is Waldfogel (1999), which finds no immediate impact of the FMLA on relative female wages, using a difference-in-difference estimate. Our results also show no significant change in wages after the passage of the FMLA, but, because of our larger sample period, we can test for whether the FMLA affects the rate of gender wage convergence separately from it's immediate effect on the wage level. We find evidence for a slow down in the rate of gender wage convergence due to the FMLA and similar family leave policies that passed in several states prior to the FMLA.

[3] Altonji and Paxson (1986) show that firms and employee preferences are important when setting work hours.

3 Data

We use data from the Annual Social and Economic Supplement (ASEC) of the Current Population Survey (CPS) covering 1975-2015.[4] This data-set includes information on total wage income earned during the previous year, usual hours worked per week during the previous year, and number of weeks worked during the previous year, which allows us to construct our key dependent variable – hourly wages; our key explanatory variable is usual hours worked per week[5]. In our empirical analysis, we use the usual hours worked per week variable to split our sample by worker *employment status*, which we partition in three categories: (i) *part-time worker* ($<$ 35 hours per week), (ii) *full-time worker* ($35-40$ hours per week) and (iii) *over-time worker* ($>$ 40 hours per week). In addition to this wage and employment data, the CPS also reports key demographic variables including gender, race, occupation, educational attainment, state of residence, *etc.*, which we use for the dual purpose of exploring heterogeneity in our data and serving as controls in our Mincer regression of wages on employment status. We restrict our sample to black and white workers aged 18-65 who reported greater than 5 hours worked per week, and positive earnings.

We supplement our forty years of CPS data with data from four other sources: Consumer Price Index data available from the St. Louis Federal Reserve Economic Database (FRED), O*Net, the TaxSim9 calculator available from the National Bureau of Economic Research (NBER) (see: National Bureau of Economic Research, 2013; Feenberg and Coutts (1993)), and the American Community Survey (ACS) (Ruggles et al., 2017). We use the FRED data to adjust the nominal wages reported in the CPS for inflation, taking 2000 as our base year, and as a source of state minimum wages over time. From the O*Net data base we merge in flexibility characteristics of each occupation in the CPS by SOC code. For example, O*Net reports the extent to which a job has structured versus unstructured work, which we take as a continuous measure of the occupation's flexibility. We use the ACS, which has a large sample size of 3 million observations in conjunction with the NBER TaxSim9 calculator to obtain precise instrumental variables estimates of the effect of employment status on wages. From TaxSim9, we obtain the marginal state and federal tax rates of the individuals in the ACS. In turn we use these tax rates as an instrumental variable for the labor supply decision.

[4]This data is available from the Integrated Public Use Microdata Series (IPUMS) at the Minnesota Population Center (Flood et al., 2017).

[5]Hours worked per week and total wage income have top-coded values. In the presented analysis we do not adjust the data for the top coding. However, based upon additional analysis, the top-code does not affect the results

Table 1: Descriptive Statistics for White Men (1975 - 2015)

	Part-time		Full-time		Over-time	
	mean	sd	mean	sd	mean	sd
Employment Characteristics						
Hourly Wage (2000 $)	15.45	65.30	18.91	18.09	21.13	20.35
Hours Worked Per Week	23.04	6.92	39.71	1.10	52.77	9.18
Educational Characteristics						
Less than High School	0.24	0.43	0.31	0.46	0.21	0.40
High School	0.21	0.40	0.22	0.42	0.18	0.39
Some College	0.34	0.47	0.22	0.42	0.22	0.42
College Degree	0.14	0.35	0.17	0.37	0.23	0.42
Post Graduate	0.07	0.25	0.08	0.27	0.16	0.37
Personal Characteristics						
Age	32.30	14.88	38.14	12.05	39.49	11.02
Children	0.42	0.92	1.03	1.23	1.20	1.26
Married	0.35	0.48	0.67	0.47	0.76	0.43
Flexibility Characteristics						
Level of Competition	49.82	9.67	51.73	8.71	53.68	9.31
Time Pressure	68.83	10.77	72.24	6.89	72.25	6.45
Duration of Workweek	58.49	17.47	66.05	14.03	70.14	13.73
Structured vs. Unstructured	27.63	7.87	26.15	7.74	22.73	8.58
N	95782		753606		445254	

*Source: ASEC Supplement of the CPS available through IPUMS, O*Net*

Table 2: Descriptive Statistics for White Women (1975 - 2015)

	(1)		(2)		(3)	
	mean	sd	mean	sd	mean	sd
Employment Characteristics						
Hourly Wage (2000 $)	12.35	19.30	13.79	20.17	16.00	15.08
Hours Worked per Week	22.44	7.13	39.31	1.59	50.86	8.38
Education Characteristics						
Less than High School	0.25	0.43	0.25	0.44	0.14	0.35
High School	0.20	0.40	0.22	0.41	0.16	0.36
Some College	0.32	0.47	0.27	0.44	0.23	0.42
College Degree	0.17	0.38	0.18	0.39	0.27	0.45
Post Graduate	0.06	0.24	0.08	0.27	0.20	0.40
Personal Characteristics						
Age	36.66	13.00	38.15	11.91	39.62	11.15
Children	1.17	1.24	0.97	1.12	0.87	1.08
Married	0.64	0.48	0.60	0.49	0.58	0.49
Flexibility Characteristics						
Level of Competition	47.77	9.44	48.23	9.26	51.06	9.91
Time Pressure	66.73	11.94	71.17	8.68	71.04	7.89
Duration of Workweek	53.24	16.32	58.98	15.22	66.03	15.23
Structured vs. unstructured	24.70	7.60	23.22	7.17	20.72	6.86
N	273986		675870		164363	

*Source: ASEC Supplement of the CPS available through IPUMS, O*Net*

Table 3: Descriptive Statistics for Black Men (1975 - 2015)

	(1)		(2)		(3)	
	mean	sd	mean	sd	mean	sd
Employment Characteristics						
Hourly Wage (2000 $)	12.89	42.16	15.31	15.66	16.61	17.57
Hours Worked per Week	23.28	6.73	39.67	1.17	53.33	9.61
Educational Characteristics						
Less than High School	0.31	0.46	0.32	0.47	0.22	0.41
High School	0.26	0.44	0.27	0.44	0.23	0.42
Some College	0.31	0.46	0.24	0.43	0.26	0.44
College Degree	0.08	0.28	0.12	0.33	0.18	0.39
Post Graduate	0.03	0.18	0.05	0.21	0.11	0.31
Personal Characteristics						
Age	32.26	13.89	38.38	11.91	39.27	10.99
Children	0.43	1.00	0.91	1.28	1.00	1.28
Married	0.25	0.43	0.53	0.50	0.59	0.49
Flexibility Characteristics						
Level of Competition	48.23	9.38	48.82	8.32	50.38	9.20
Time Pressure	69.04	10.77	72.00	7.28	72.09	6.98
Duration of Workweek	57.05	17.70	62.45	14.84	65.30	14.92
Structured vs. Unstructured	28.63	7.06	27.64	7.21	25.59	8.51
N	12356		88641		23927	

*Source: ASEC Supplement of the CPS available through IPUMS, O*Net*

Table 4: Descriptive Statistics for Black Women (1975 - 2015)

	(1)		(2)		(3)	
	mean	sd	mean	sd	mean	sd
Employment Characteristics						
Hourly Wage (2000 $)	10.54	16.85	13.18	15.94	14.39	13.23
Hours Worked per Week	23.13	6.94	39.38	1.52	52.38	9.25
Education Characteristics						
Less than High School	0.33	0.47	0.24	0.43	0.15	0.36
High School	0.25	0.43	0.24	0.43	0.18	0.39
Some College	0.30	0.46	0.30	0.46	0.28	0.45
College Degree	0.09	0.28	0.15	0.36	0.22	0.41
Post Graduate	0.03	0.18	0.06	0.25	0.16	0.37
Personal Characteristics						
Age	35.94	13.94	38.29	11.48	39.37	10.96
Children	1.10	1.33	1.15	1.24	1.00	1.20
Married	0.32	0.47	0.39	0.49	0.36	0.48
Flexibility Characteristics						
Level of Competition	46.44	8.80	46.96	8.56	49.41	9.46
Time Pressure	66.63	11.50	71.06	8.36	71.06	7.98
Duration of Workweek	50.99	15.77	57.13	15.01	62.30	15.92
Structured vs. Unstructured	26.17	6.85	24.58	7.04	22.48	7.22
N	26617		111827		14967	

*Source: ASEC Supplement of the CPS available through IPUMS, O*Net*

3.1 Summary Statistics

In Table 1, 2, 3, and 4 we report summary statistics of the data for 4 groups: white men, white women, black men, and black women, respectively. For each group, we provide summary statistics for three types of workers based on employment status: part-time, full-time, and over-time. When compared to full-time workers and over-time workers, part-time workers tend to be younger and are less likely to have a college or graduate degree. They also earn lower wages on average.

Using full-time workers as our baseline, part-time workers of the same race and gender earn between 84% and 91% of the hourly wages of full-time workers. Overtime workers earn 108% - 116% of the hourly wage of their full-time counterparts. Changing the reference group to be white men of a given employment status, we find that part-time white women experience a smaller gender wage gap than full-time women or over-time women (22% relative to 28% and 24%, respectively). We find a similar pattern in the raw data for black women and black men, with part-time black women facing a gender wage gap of 29% versus 31% and 32% and part-time black men facing a racial wage gap of 17% versus 19% and 22% for full-time and over-time workers, respectively.

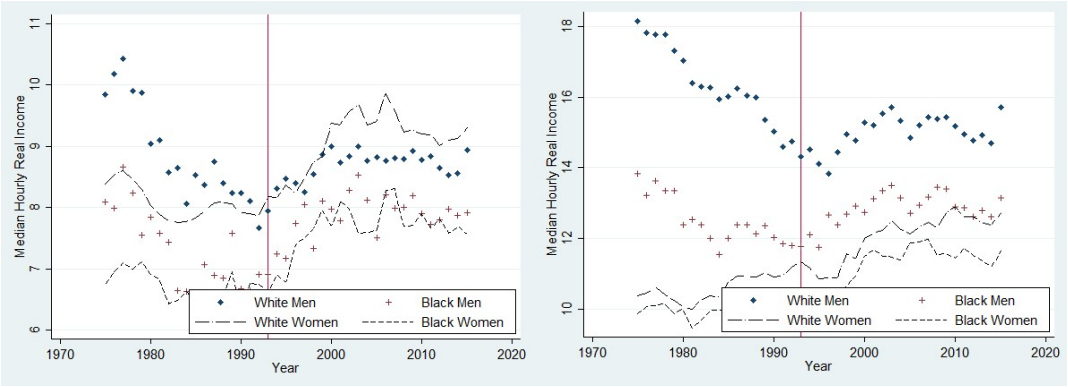
White men who work part-time have on average 0.42 children compared to 1.17 children for white women working part-time. There are similar patterns for black men and women who work part-time (0.43 and 1.10 children, respectively). The picture is starkly different for white men who work full-time (1.03 children) and white women who work full-time (0.97 children). In fact, as white men work more hours per week, they tend to have more children, on average. However, this is the exact opposite of the pattern for white women, who tend to have fewer children on average when they work longer weeks. A similar pattern appears for black men and women.[6]

White men tend to be married at greater rates as they work more hours (part-time white men are married 35% of the time compared to 67% for full-time working white men and 76% for over-time working white men). This is the exact opposite of the pattern for white women (part-time white women are married 64% of the time compared to a 60% marriage rate for full-time working white women and a 58% marriage rate for over-time working white women). A similar pattern is present for black men but not black women. However, these patterns are consistent with an increase in demand for flexibility caused by the increased household production of married women compared to married men.

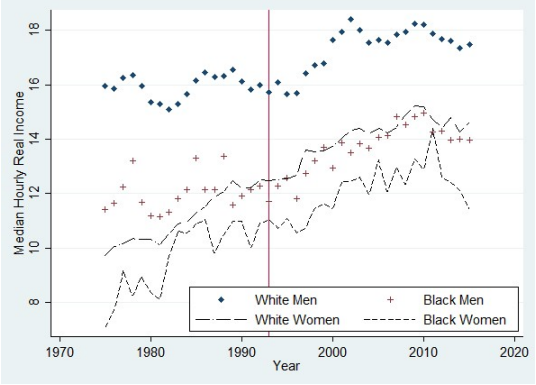
[6]Black women have more children when they work full-time than part-time. However, over-time working black women have the fewest children among black women workers - a return to the general pattern.

3.2 Trends in the Data

Figure 2: **Median Hourly Wages (1975 - 2015)**



(a) (b)



(c)

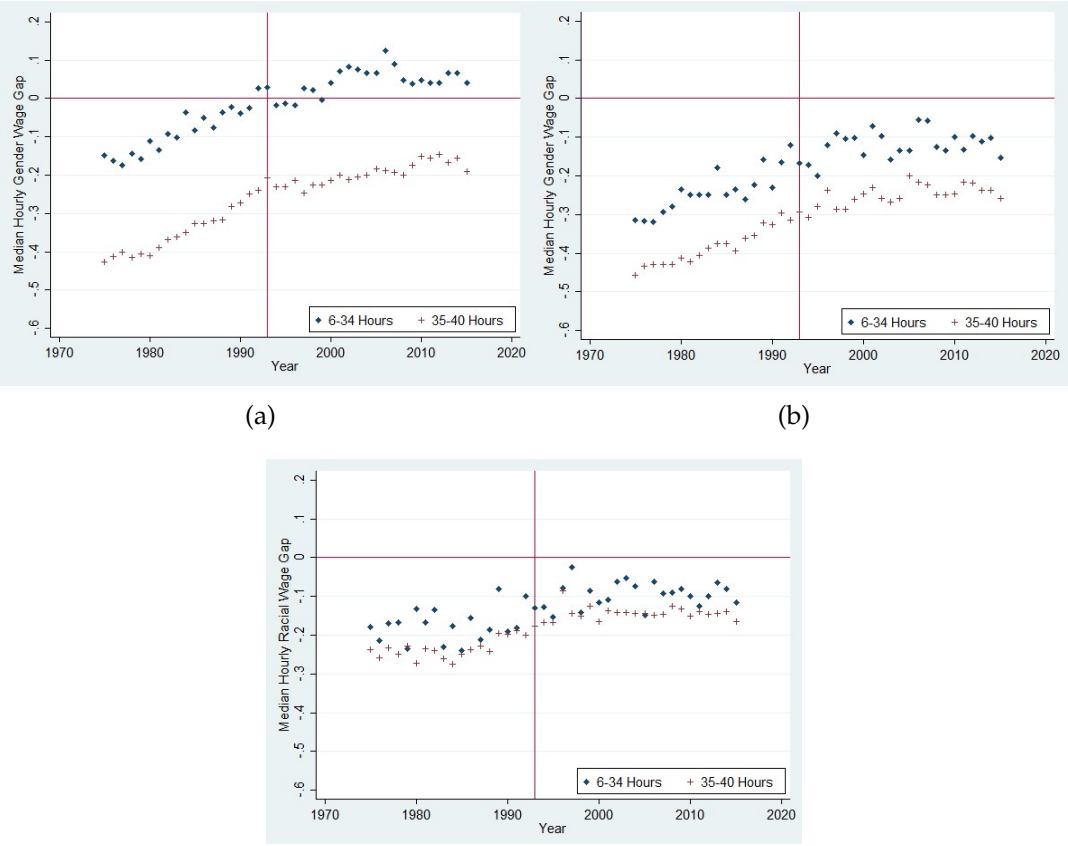
Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the raw median hourly wages for white men, white women, black women, and black men for part-time (a), full-time (b), and over-time (c) workers. The key finding of these graphs is that the wages of men were decreasing earlier in the time period before reversing and increasing beginning in the early to mid 1990s. Moreover, women began to experience wage growth in part-time jobs for the first time in this same time period, which is coincident with a change in labor demand due to the passage of the Family and Medical Leave Act.

In Figure 2 we present the time-series of raw hourly wages for four demographic categories of workers: white men, white women, black men, and black women. The results show that, as men (both white and black) faced increased labor-market competition from women and changing demands for labor in the part-time and full-time sector, the wages of men dropped while the wages of women tended to stay the same, a result consistent with Juhn and Kim (1999). However, in the early 1990s this pattern changed. Specifically, hourly wages of male workers increased in both the part-time and full-time sectors. Moreover, after experiencing a stagnation in part-time hourly wages, women began to

experience growth in part-time wages. Finally, over-time working men experienced a growth in their wages for the first time in our sample in the early 1990s. The timing of these changes is coincident with a lagged response to labor demand after the passage of the Family and Medical Leave Act.

Figure 3: **Median Hourly Wages (1975 - 2015)**



Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the raw median wage gaps between white men and workers from three demographic categories: white women (a), black women (b), and black men (c). The key finding of these graphs is that, in line with expectations, the gender wage gaps are smaller for part-time workers than full-time workers. However, we do not see this pattern for black men. Therefore, the flexibility hypothesis is supported because women have higher roles in household production and therefore demand more flexibility than their male counterparts.

In Figure 3 we present the time-series of raw wage gaps between part-time workers of different demographic groups relative to white men and the time-series of wage gaps between full-time workers of different demographic groups relative to white men. The results are both stark and important to our hypothesis of the supply and demand of flexibility in the labor market. The plot of the median wage gap between white women and white men in Figure 3(a) shows that white women working part-time have faced wage

gaps that were 20-30 percentage points smaller than their full-time counterparts as far back in the data as 1975. This result is predicted by our hypothesis. Since part-time jobs have a cheaper supply of flexibility, the wage gaps caused by using the flexibility should be smaller than the wage gaps caused in full-time work, a setting with a relatively more expensive supply of flexibility, *ceteris paribus*. Moreover, this result is persistent despite a trend in the data towards gender wage parity between white women and white men in both the part-time and full-time sectors. Finally, a break in the trend to raw median wage parity occurred in 1993 - coincident with the Family and Medical Leave Act.

The plot of the median wage gap between black women and white men in Figure 3(b) show a very similar picture to the plot for white women. As a type of placebo test, we generate a similar time-series for the wage gaps faced by black men. In Figure 3(c), we do not see the clear pattern of lower wage gaps for black men relative to white men in part-time work than full-time work that we saw for both black and white women. This suggests that the main effects that we are documenting are due to factors related to gender.

4 Empirical Specification

Our goal in this paper is to analyze the theory that the gender wage gaps are smaller for women in part-time work because of the flexibility of part-time work. To this end, we segment the data into three hours-worked categories: part-time (less than 35 hours worked per week), full-time (35 - 40 hours worked per week), and over-time (greater than 40 hours worked per week).[7] For each segment of the labor market s – part-time workers, full-time workers and over-time workers - we estimate the following Mincer wage equation:

$$\log(w_{i,t,s}) = \alpha_{0,t,s} + \beta_{1,s,t} WW_{i,t,s} + \beta_{2,s,t} BW_{i,t,s} + \beta_{3,s,t} BM_{i,t,s} + \vec{\gamma}_{i,s,t} \cdot \vec{X}_{i,s,t} + \epsilon_{i,t,s}. \quad (1)$$

The dependent variable in our specification is the natural log of the implied real hourly wages (measured in 2000 dollars) for worker i at year t who is observed to work in labor market segment s; $WW_{i,t,s}$, $BW_{i,t,s}$, and $BM_{i,t,s}$ are indicator variables equal to 1 if the individual is a white woman, black woman or black man, respectively, and 0 otherwise; $\vec{X}_{i,t,s}$ is a vector of observable worker characteristics: a quadratic in age, 3-digit occupational

[7] This classification of *employment status* arises from some states categorizing 35 hours of work as full-time whereas other states use 40 hours as the threshold for full-time work.

fixed effects, state fixed effects, and dummy variables for the level of education; and ϵ_i is the error term. We run the weighted regressions in equation (1) separately for each year t and for each labor market segment s.[8] In doing so, we are conditioning on both year and work status, which allows for us to abstract from including time fixed effects in our regression. Running separate regressions also allows us to have occupation fixed effects that vary by year and labor market segment.

Our coefficients of interest are $\beta_{1,s,t}$ and $\beta_{2,s,t}$, which represent the percentage gap in earnings between white (black) women and white men in labor market segment s at time t. In our main results, we plot these coefficients over time to map out how the gender wage gap differs across labor market segments and how this difference in the gender wage gap between part-time and full-time workers has changed over time. For comparison purposes we will also report the time-series for $\beta_{3,s,t}$, which is the percentage gap in earnings between black men and white men in labor market segment s at year t. By comparing whether the racial wage gap among men resembles the gender wage gap, we can determine whether our results capture time trends that are common among all groups of non-white male workers or if the gender wage results that we document are driven by the relationship between gender, employment flexibility and wages.

The key threat to identification in our setting is that workers endogenously sort into their observed labor market segment by choosing how many hours to work. While our OLS procedure will produce unbiased estimates if this selection is fully captured by the observed characteristics, it is susceptible to bias arising from selection on unobservables. To address selection on unobservables, we instrument for the labor supply of a given worker, and hence her or his observed labor market segment, using the average total of the state and federal marginal tax rates faced by other workers of a similar age, same gender and marital status, who also live in the same state. This is a valid instrument due to our use of pre-tax income and the findings of Deslauriers et al. (2018) who show that payroll taxes are passed entirely onto workers. Assuming the same is true for income taxes, the income tax will only affect the pre-tax income through the choice of hours worked. The first stage of our IV procedure is:

$$h_{i,g,a,d,t} = \theta_{0,t} + \sum_k \theta_{k,t} \text{marginal tax}_{-i,g,a,d,t} Z_{i,k} + \vec{\rho}_{g,t} \cdot \vec{X}_{i,g,a,d,t}, \tag{2}$$

where $h_{i,g,a,d,t}$ is the hours worked by individual i, in age group a, of marital group g,

[8] The weighting scheme in the CPS changed in 2014 and the scheme remained changed in subsequent years. In 2014, the sample was surveyed two different ways and had two different sets of weights. Since there is not a clear way to choose weights for 2014 given our sample period, we use all individuals and both sets of weights in that year.

living in state d at time t; marginal $\text{tax}_{-i,g,a,d,t}$ is the mean total of the federal and state marginal tax rates faced by type g individuals (excluding worker i), in age group a, in state d in year t; and $X_{i,g,a,d,t}$ is a vector of observable characteristics of individual i, of group g, living in state d at time t. The variables $Z_{i,k}$ allow for heterogeneity in the labor supply response to the marginal tax rate by gender, marital status and the number of children.

This labor supply instrument is similar to the labor supply instrument in Angrist (1991) and Blau and Kahn (2007), who instrument for an individual's hours worked using the mean hours worked by similar individuals. Our instrument is different in two dimensions. First we use the jackknife or leave-out mean, which reduces the possibility for a mechanical correlation between the individual hours worked and the average hours worked.[9] Second, instead of using average hours worked to construct our leave-out labor supply IV, we use the average total federal and state marginal tax rate. This allows us to exploit a different source of variation – heterogeneity in tax rates – which affects the labor supply decision in a micro-founded model of labor supply and only affects wages through its effect on hours worked. Moreover, our IV is interacted with marital status, which has been shown to cause spouses to respond differently to taxation than single adults Isaac (2018).

5 Results

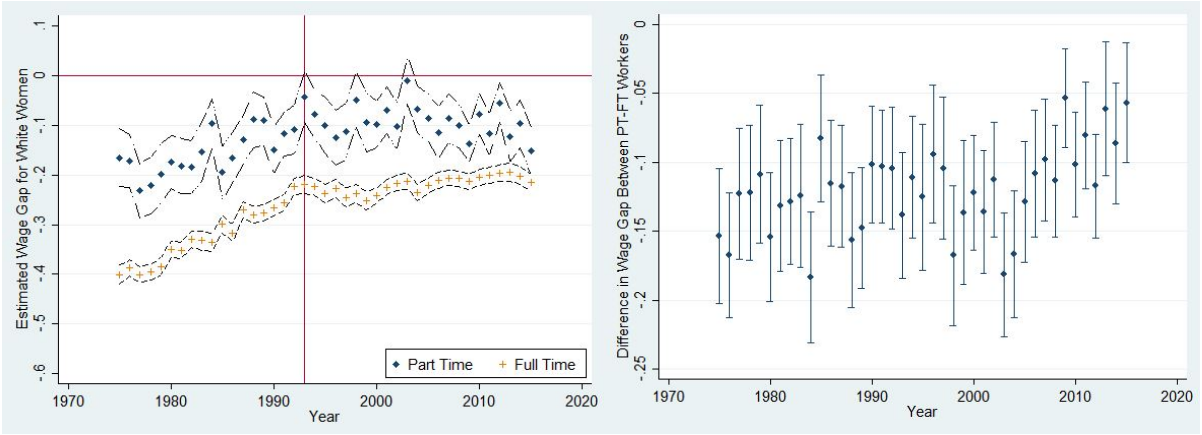
To test whether selection on observables is driving the patterns that we document in the raw data, we regress log wages on observable worker characteristics, including a set of 3-digit occupation fixed effects. This involves estimating equation (1) separately for each year by employment status and recording the estimated wage gap coefficients for women and black men. In Figure 4 and 5, we report the estimated coefficients along with 95% confidence intervals for each year for both full-time workers and part-time workers.

Controlling for selection on observables, as shown in Figure 4(a), reveals that the gender wage gap for part-time white women is larger than what we find in the raw data. For example, in the raw data we observe a part-time gender wage gap of 12% in 1976, whereas here we estimate a wage gap of 18%. Moreover, during the entire 40-year period part-time white women do not achieve wage parity with men, although the raw data suggests this. The estimated wage gaps for full-time white women, by contrast, look remarkably similar to those in the raw data throughout the time series. The net effect

[9] In large data samples, the effect of this mechanical correlation converges at rate $\frac{1}{N}$, making this a minor concern for large N.

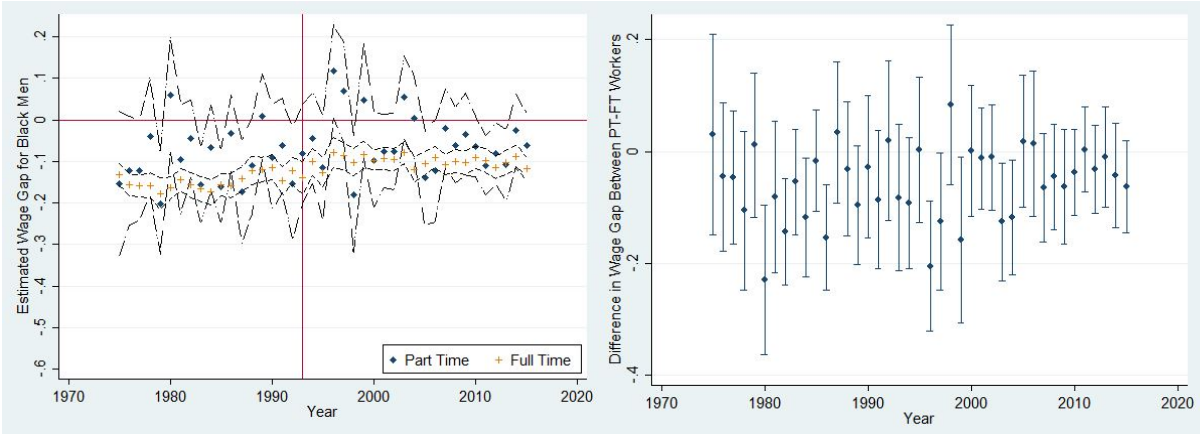
is that the part-time gender wage gap is smaller than the full-time gender wage gap by 10-20 percentage points, as opposed to the apparent 20-30 percentage point gap in the raw data. Moreover, throughout our sample period, the 95% part-time gender wage gap confidence interval is non-overlapping with that of the full-time gender wage gap. In Figure 4(b), we show that the differences between wage gaps for part-time and full-time workers are significant at the 95% level. Further, we observe gender wage convergence for both part-time and full-time women until the mid 1990s, after which there is stagnation in gender wage convergence. Finally, the results for black women are very similar in patterns and trend to the results for white women and can be seen, along with all other subsequent results for black women, in the Online Appendix.

Figure 4: **Gender Wage Gap for White Women Controlling for Observables**



(a) (b)

Figure 5: **Racial Wage Gap for Black Men Controlling for Observables**

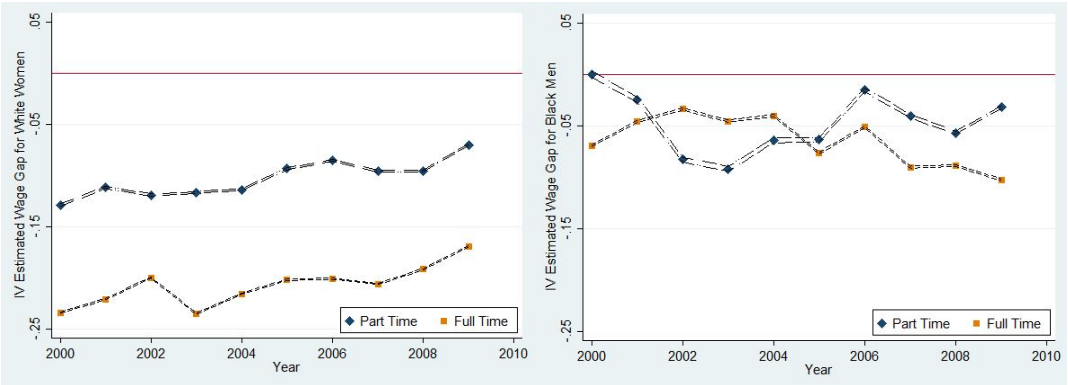


(a) (b)

As a placebo we perform the same exercise for black men, to see if selection on observables is responsible for our failure to find the same pattern for black men in the raw data as we saw for women. As shown in Figure 5(a) we do not find that the part-time racial wage gaps for black men are smaller than their full-time counter parts. Instead, the confidence intervals are overlapping everywhere and in some cases the part-time point estimates are more negative that the full-time point estimates. Moreover, the differences between the part-time and full-time gaps are now insignificant based upon Figure 5(b). This suggests that the part-time wage premium we are observing for women is, in fact, gender-specific.

While the results we have described so far suggest that the part-time gender wage premium is a persistent phenomenon, we were concerned that selection on unobservables could be driving our results. We run our IV on the data from the ACS spanning the decade 2000-2009. Each year of the ACS contains 3 million observations. By contrast, each year of the CPS contains sixty thousand observations. By using the ACS, we forego having a longer sample period for having a drastically larger sample. In Figure 6(a) we report the estimated wage gaps for full-time white women and part-time white women. As before, we observe a part-time gender wage gap which is smaller than the full-time gap. In absolute terms, the part-time gender wage gap is consistently 10 percentage points smaller than the full-time gender wage gap. This was the same parallel shift between the two gender wage gaps in the OLS and the raw data. Moreover, the confidence intervals are tight - to within a single percentage point. Finally, we see a very similar picture for black women.

Figure 6: **Gender and Racial Wage Gaps Controlling for Selection on Unobservables**



(a) (b)

In Figure 6(b), we plot an analogous graph for black men to Figure 6(a) for white women. Here we also see very tight confidence intervals on our point estimates - owing

to the size of the ACS data. One noticeable difference in the results for black men is that we observe much smaller differences between the part-time and full-time racial wage gaps. Which wage gap is smallest also changes year-to-year. And for some years there is no difference at all between the part-time and full-time wage gaps between black men and white men. This result stands in contrast to the results for white women, where we see a clear hierarchy of magnitudes that is both large and persistent over time in favor of women in part-time work arrangements.

Besides the potential for selection on observables, which we address with our instrumental variables approach, there are other concerns that may affect our results. Chief among these concerns is the possibility that the minimum wage truncates the lower end of the wage distribution more heavily for part-time workers than full-time or over-time workers and that this differential truncation may mechanically produce smaller gender wage gaps among part-time workers than among full-time workers (Summers, 1989). To address this concern, we drop workers earning the minimum wage from our sample and re-estimate our models.[10] The results from the non-minimum wage sample are reported in the online appendix (see: the raw median gaps in Figure 12, the estimated wage gaps when controlling for observables in Figure 13, the differences in the estimated wage gaps in Figure 15(a), and the estimated wage gaps when controlling for unobservables in Figure 15. The key finding is that the original results are robust to this alternative sampling technique. Therefore, the striking results presented in Section 5 are not being driven by the minimum wage, but appear to be driven by the forces detailed in our hypothesis.

6 Flexibility and the Family and Medical Leave Act

According to usage data, women account for 58% of leave takers under the FMLA (Tompson and Werner, 1997). The different rates of take-up by gender is unsurprising given the work of Antecol et al. (2018) that shows the differential effects of supposedly gender neutral policies. One way to rationalize the slowdown in gender wage convergence is that the cost of expected leave is capitalized into wages and hiring decisions, consistent with the findings of Das and Polachek (2015) and Bana et al. (2018). We provide three pieces of evidence for this hypothesis.

First, we run a linear regression of the gender wage gap on a time trend, where we allow for a discontinuous change in the slope of the gender wage gap and the intercept at

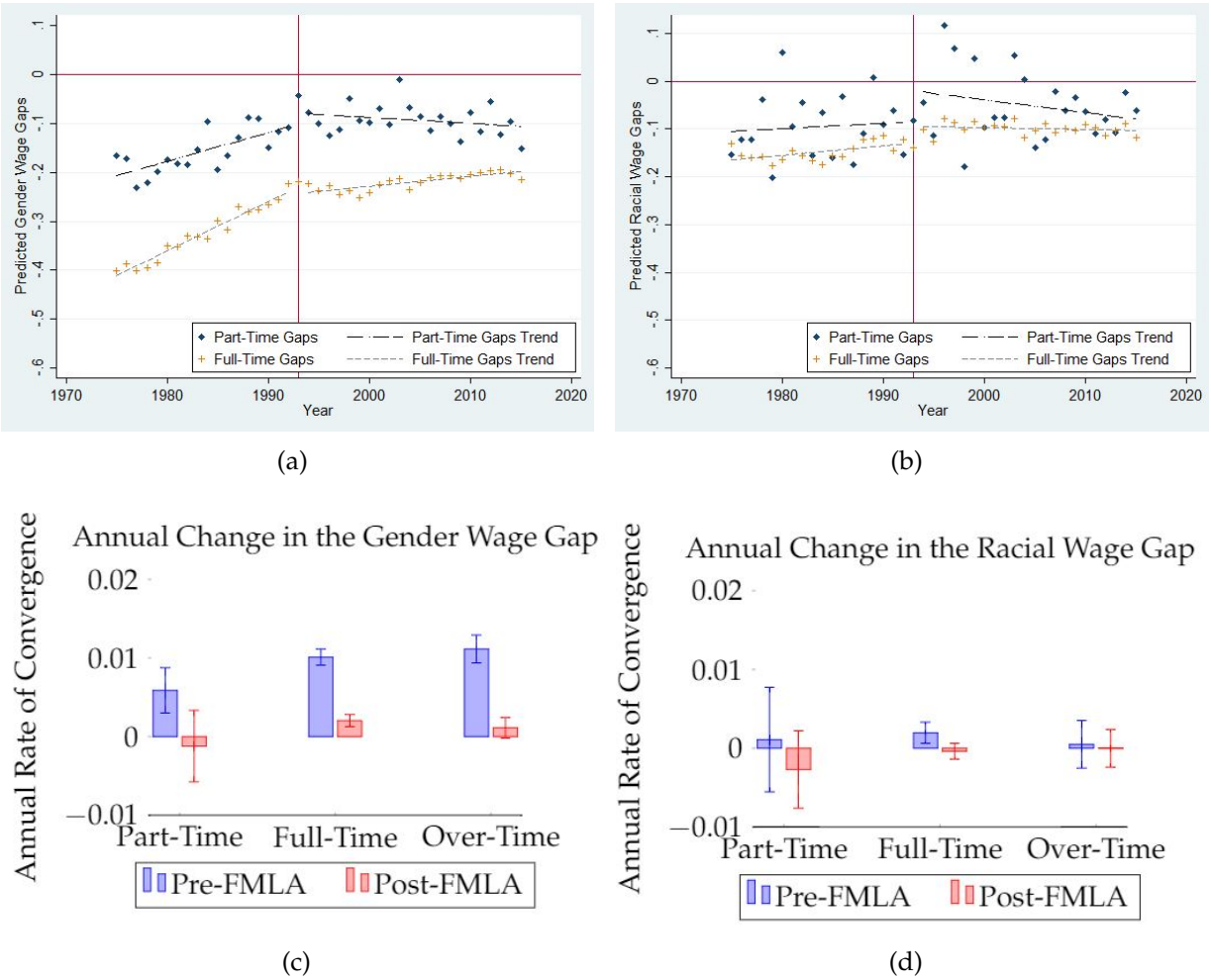
[10] We use state minimum wages when available and federal minimum wages only when state minimum wages are not available.

the FMLA year to see if the break in trend is statistically significant:

$$\text{wage gap}_{t,g} = \alpha_{0,g} + \beta_g Trend + \gamma_g \mathbb{1}(FMLA) + \phi_g \mathbb{1}(FMLA) \times Trend + \epsilon_{t,g}. \quad (3)$$

The variable *Trend* is the year centered around 1993 and $\mathbb{1}(FMLA) = 1$ if the year is in the post-FMLA time period. Therefore, the pre-FMLA rate of gender wage convergence is β_g and the post-FMLA rate of gender wage convergence is $\beta_g + \phi_g$ for group g.

Figure 7: **Estimated Trends in National Wage Gaps**



Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the results of linearly combining the beta coefficients for Trend and Trend-FMLA reported in Table 2. From the graphs, it is clear that the trend in the gaps is upward sloping prior to the date for the passing of the FMLA law. Moreover, it is also clear that the trend line shifts to a slope of zero after the passing of the law. The regression can be expressed as $gap = \beta_0 + \beta_1 Trend + \beta_2 TrendxFMLA + \beta_3 FMLA + \epsilon$ where *Trend* is simply a linear trend counter equal to zero for the year 1993, *FMLA* is a dummy equal to one if the year is after 1993, *TrendxFMLA* is the interaction of *Trend* and *FMLA*. Since the policy change occurred in our survey year 1993, we do not include 1993 in the regressions.

Table 5: **Trend Regression Results for Wage Gaps (1975-2015)**

	White Women			Black Women			Black Men		
VARIABLES	Part-Time	Full-Time	Over-Time	Part-Time	Full-Time	Over-Time	Part-Time	Full-Time	Over-Time
Trend	0.00586***	0.0101***	0.0111***	0.00650***	0.00849***	0.0113***	0.00110	0.00196***	0.000505
	(0.00142)	(0.000514)	(0.000874)	(0.00223)	(0.000702)	(0.00215)	(0.00327)	(0.000665)	(0.00149)
FMLA	0.0219	-0.0136*	-0.0213	-0.00124	-0.00687	-0.0653**	0.0650	0.0350***	-0.00142
	(0.0207)	(0.00747)	(0.0127)	(0.0325)	(0.0102)	(0.0313)	(0.0475)	(0.00968)	(0.0217)
Trend*FMLA	-0.00708***	-0.00808***	-0.0100***	-0.00720**	-0.00743***	-0.0109***	-0.00378	-0.00234***	-0.000499
	(0.00177)	(0.000639)	(0.00109)	(0.00278)	(0.000873)	(0.00268)	(0.00407)	(0.000828)	(0.00185)
Constant	-0.101***	-0.229***	-0.219***	-0.122***	-0.249***	-0.230***	-0.0850**	-0.129***	-0.141***
	(0.0154)	(0.00556)	(0.00947)	(0.0242)	(0.00759)	(0.0233)	(0.0354)	(0.00720)	(0.0161)
Observations	40	40	40	40	40	40	40	40	40
R-squared	0.623	0.973	0.921	0.356	0.927	0.504	0.125	0.771	0.006

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the table above, we have reported the results of a regression on the estimated wage gaps for the average, controlled white woman, black woman, and black man using linear trends. Here, we have assumed that FMLA took affect in 1993 and caused the change in the trend of the wage gap convergence. From the graphs, it is clear that the trend in the gaps is upward sloping prior to the date for the passing of the FMLA law. Moreover, it is also clear that the trend is flat after the passing of the law. The regression can be expressed as $gap = \beta_0 + \beta_1 Trend + \beta_2 TrendxFMLA + \beta_3 FMLA + \epsilon$ where *Trend* is simply a linear trend counter equal to zero for the year 1993, $FMLA$ is a dummy equal to one if the year is after 1993, $TrendxFMLA$ is the interaction of Trend and Post-Trend. Since the policy change occurred in our survey year 1993, we do not include 1993 in the regressions.

In Table 5 we report the results of the regression in equation (3) for white women, black women, and black men. For each of these groups we report regression results for part-time, full-time, and over-time workers. First we notice that women experienced shrinking gender wage gaps regardless of employment status. Prior to 1993, the gains for full-time and over-time women are roughly twice the size of those for part-time women (0.6% per year versus 1.2% per year) during this period. Following the passage of the FMLA the rate of gender wage convergence is close to zero across the board for women. The interaction FMLA*Trend coefficient is roughly equal in magnitude and opposite in sign to the trend coefficient. Since the FMLA affects a full-time or over-time working woman's outside option, part-time workers also see a stagnation of the rate of convergence (Gruber, 1994).

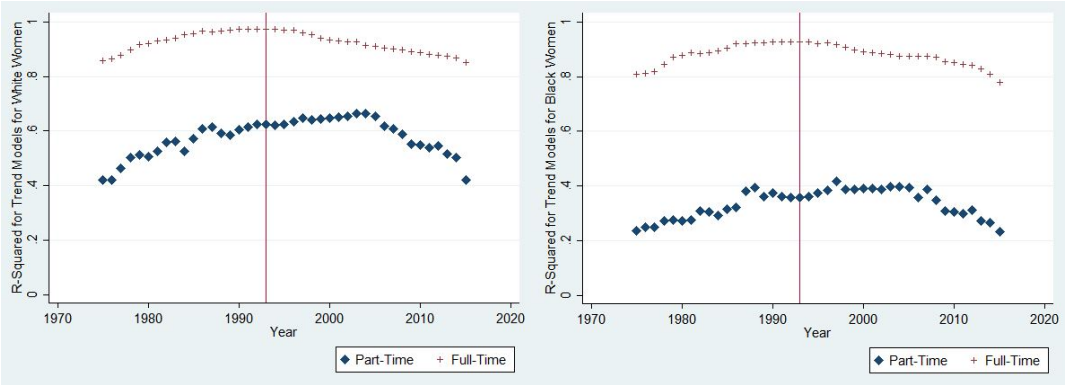
Moreover, there is a level drop in wages for full-time white women and over-time black women that is marginally significant at the 10% level. By contrast, the time trends in the racial wage gap for black men prior to the FMLA range from 50% to 97% *smaller* than the corresponding time trends for the gender wage gaps. Moreover, these pre-trends are only statistically significant for black men in full-time roles and even there it is 1/6 the size of the companion time trend in the gender wage gap for women. The passage of the FMLA does not have significant effect on the convergence rate of the racial wage gap among men except for full-time workers, who experience a shift from a very small positive pre-trend to a very small negative post-trend. This negative post trend for full-time black men is also offset by a 3.6 percentage-point shift in wages. The pre-FMLA and post-FMLA rates of wage convergence are available for all three demographic groups in Figure 7.

For full-time workers in all groups, we find that the regression fit is very good. The R^2 for full-time white women is 0.973, full-time black women 0.927, and full-time black men is 0.771. The fact that our fit is markedly better for women than for men is consistent with the FMLA having a stronger effect on the rate of gender wage convergence than racial wage convergence among men. The fit for women in over-time roles is also quite good $R^2 = 0.921$ (white women) and $R^2 = 0.504$ (black women). By contrast, the model has no explanatory power for black men in over-time roles ($R^2 = 0.006$). For part-time female workers, our piecewise linear model does not do nearly as well in explaining the time series in the gender wage gap $R^2 = 0.623$ (white women) and $R^2 = 0.356$ (black women). The fit is also relatively poor for black men in part-time work ($R^2 = 0.125$), which itself is much worse than the part-time fit for women. These results are consistent with the FMLA having a larger policy effect on women than on men, and within gender having a larger effect on full-time and over-time workers than part-time workers.

While this evidence is consistent with our hypothesis that the workplace regulations

imposed by the FMLA retarded female wage growth, it is not dispositive. We test this hypothesis in two ways. First we alter our model in equation (3) by moving the break from 1993 to alternative placebo years. We then plot the regression R^2 for each group as a function of the break year. Following the approach in Kleven et al. (2013), we look for the break year that maximizes the R^2 as the most likely break year that is consistent with a break in trend for the wage gaps that we estimated. For full-time women, the R^2 is maximized with the break at the FMLA year, which is denoted by the vertical line in Figure 8. For full-time men R^2 is maximized away from the FMLA year. For part-time workers, the results are less clear because the R^2 curve is relatively flat. This is not surprising given that the FMLA predominantly applied to full-time workers and was taken up differentially by women.

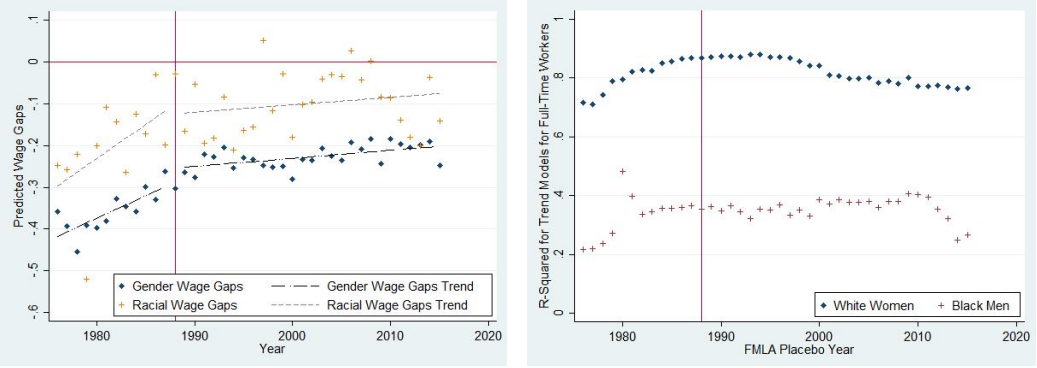
Figure 8: **Testing for Structural Break at FMLA Year (1993)**



(a) (b)

We have a second tool at our disposal to test our hypothesis that government-imposed, labor-market flexibility contributed to the slowdown in gender wage convergence. Twelve states and the District of Columbia enacted maternity leave laws *prior* to the FMLA (Table 6). Of these states close to half enacted their policy in either 1987 or 1988 - namely Minnesota, Rhode Island, Maine, Oregon, Tennessee, and Wisconsin. In Figure 9(a) we plot the estimated gender wage gaps for full-time white women in these states with pre-existing leave laws. We find a similar pattern to the pattern in the overall data. Prior to the enactment of these laws we observe a steady reduction in the gender wage gap. In Figure 9(b), we show that the R^2 for the break in the time trend is also maximized near the time the FMLA-like laws are passed. In Figures 9(c)-(d), we show that the rate of gender wage convergence is positive and significant prior to the FMLA-like law, and small and statistically indistinguishable from 0 afterwards. By comparison, there is no pre or post trend for black men, as was the case with the actual FMLA.

Figure 9: **Estimated Trends in State Wage Gaps**



(a) (b)



(c) (d)

Not only do we find that a break in the time series at the FMLA year best explain the data for women, as measured by the R^2, but it is also associated with large significant pre-trends of gender wage convergence across the board and a flat post-trend to describe gender wage stagnation. Furthermore, we observe a similar pattern in gender wage convergence for states that adopted FMLA-like laws prior to the FMLA. Notably we observe a steady reduction in the gender wage gap prior to the implementation of the leave policy and a stagnation in the gender wage gap following the implementation of the policy. The weight of the evidence points towards a rate of family leave policies in stymieing gender wage convergence.

As a thought experiment, we extrapolate the pre-FMLA trends in the gender wage gap to determine when gender wage parity would have occurred in the absence of the FMLA. At the time that the FMLA is enacted, the gender wage gap is given by the constant term in the regression equation (3), and the pre-FMLA rate of convergence in the gender wage gap is given by the coefficient on trend, β_g. Under a linear extrapolation, white women in part-time work would have achieved wage parity with white men in 2011, followed by part-time black women in 2012, then over-time white women in 2013, then over-time

Table 6: **Summary of State Family Leave Laws Prior to 1993**

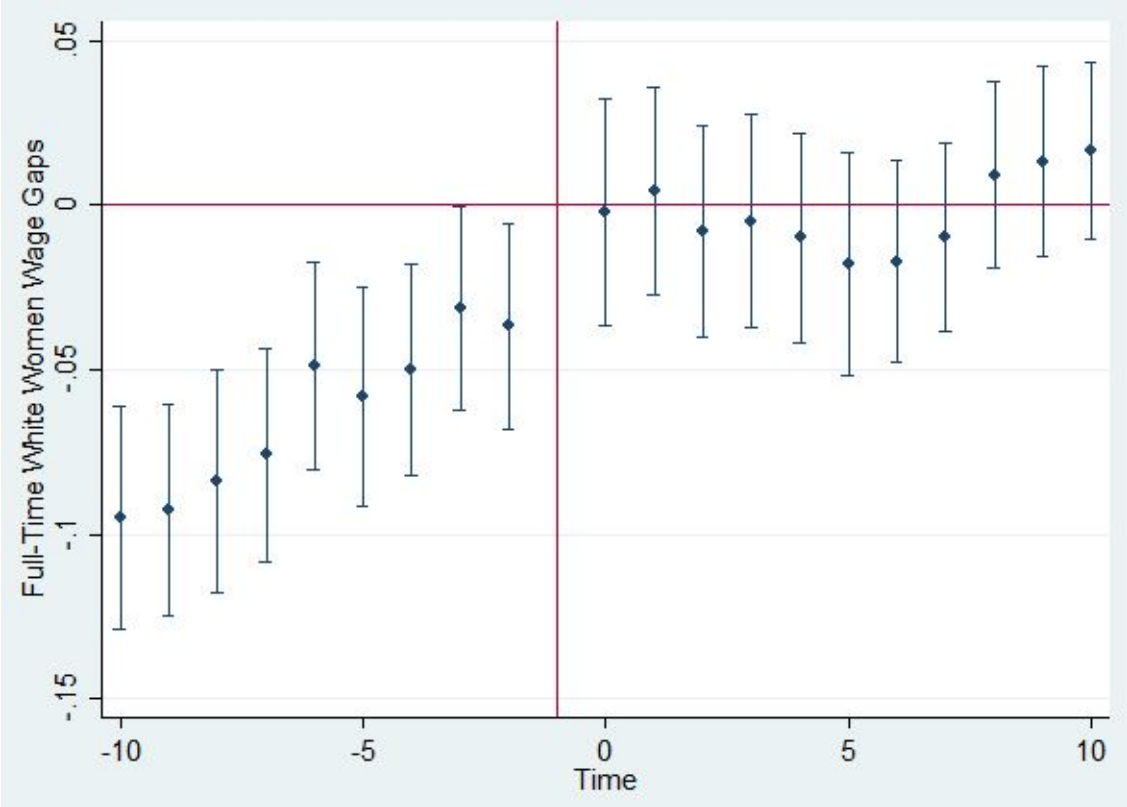
State	Maternity	Paternity
Massachusetts	1972	-
Connecticut	1973	1990
Washington	1973	1989
California	1980	1992
Minnesota	1987	1987
Rhode Island	1987	1987
Maine	1988	1991
Oregon	1988	1988
Tennessee	1988	-
Wisconsin	1988	1988
New Jersey	1990	1990
Washington, DC	1991	1991
Vermont	1992	1992

black women in 2014, full-time white women in 2016, and full-time black women in 2023.

To leverage the full sample of states that pass pre-FMLA maternity leave laws, we employ an event study design. For our event study design, we define an event as the implementation of a family leave policy. Since 12 states implemented the FMLA prior to 1993 we have 12 events prior to 1993 and a remaining 38 events in 1993. Our event study estimates derive from a two-stage process, similar to Charles and Guryan (2008). In the first stage we obtain the state wage gaps by race and gender for each labor market segment by regressing the log of wages on race and gender dummies for each state-year, as we have done previously. In our second step, we regress these state wage gaps on event time dummies ranging from $T = -10$ to $T = 10$. We weight each state wage gap observation by the inverse of standard error from the first stage in order to down-weight noisy state observations.

In addition to testing for breaks in trend, we also test for the effect of the FMLA on gender-wage convergence using a difference-in-difference model. The dependent variable for the model is the wage gap between white men and a specific demographic group, such as white women. The wage gaps are estimated for young and old workers by year following Equation 1 for two groups of states: those states who passed a maternity-leave law before the FMLA and those who did not. There are 12 states including the District of Columbia that passed maternity-leave laws before the passage of the FMLA at the federal level. In our DID specification, these states are used as one of two control groups. The second control group we use is older workers, which we define as those workers who are at least 35 years of age. We use the older workers as a control group because it is unlikely

Figure 10: **Event Study: Family Leave Policy on White Female Wage Gap**



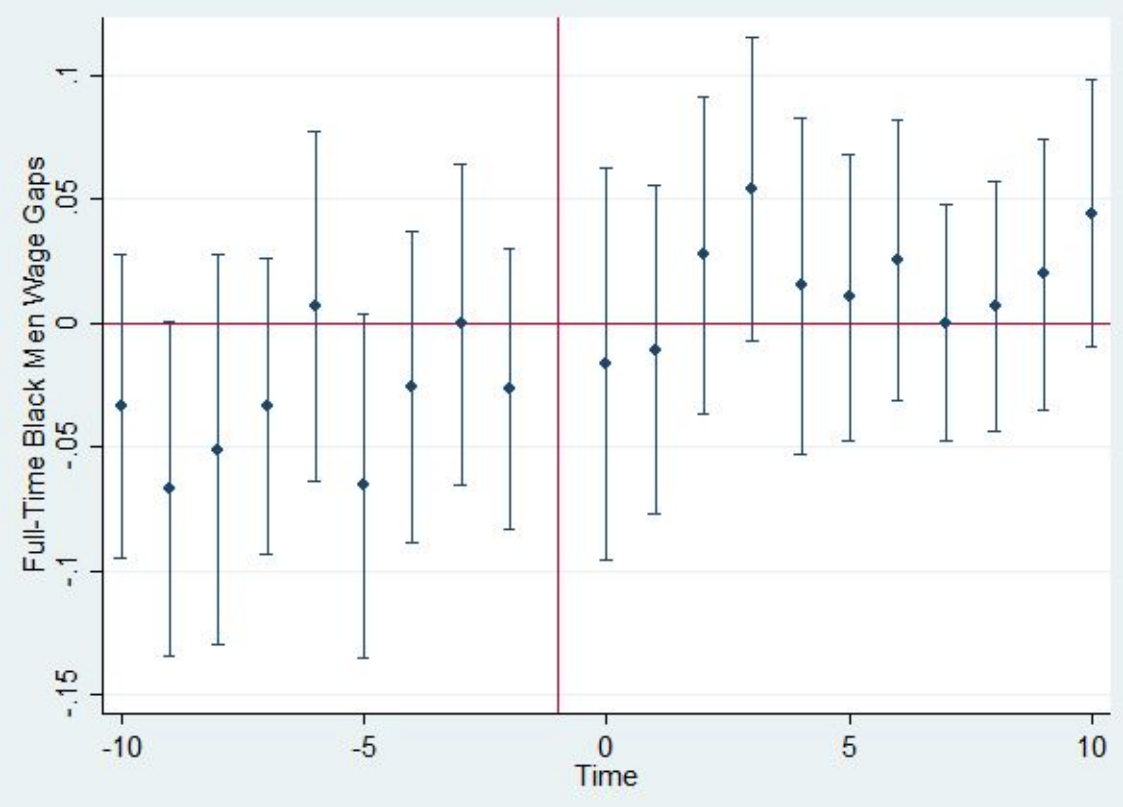
that a woman would have a baby after the age of 35 and therefore would not be affected by a maternity-leave law such as the FMLA. Finally, we use only one observation per demographic and state group per year and omit the passage year of the FMLA, 1993, from the estimation.

The DID model is estimated separately for each demographic group and each hours-worked segment of the economy. The model is given by

$$Gap_{ats} = \alpha_0 + \beta_1 Trend_t + \beta_2 FederalOnly_s + \beta_3 FMLAEnacted_t + \beta_4 Young_a + \beta_5 DID_{ts} + \beta_6 DIDYoung_{ats} + \epsilon_{ats} \quad (4)$$

where Gap_{ats} is the estimated wage gap for the demographic group of age a in year t in state group s. $Trend_t$ is a simple trend counter centered around 1993, the year the FMLA was passed. $FederalOnly_s$ is an indicator for the experimental group, the group that is only affected by the FMLA and not a prior state law. $FMLAEnacted_t$ is an indicator equal to one if the year is after 1993. $Young_a$ is an indicator if the gap is for young workers. DID_{ts} is the interaction between $Trend_t$ and $FMLAEnacted_{ts}$; it controls for a change in

Figure 11: **Event Study: Family Leave Policy Wage Black Male Wage Gap**



Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

the convergence rate for all workers after the passage of the FMLA. $DIDYoung_{ats}$ is the interaction between DID_{ts} and $Young_a$; it controls for a potential change in convergence for young workers that may occur after the passage of the FMLA. If the FMLA is causing gender wage convergence to stall, we should see a negative coefficient on $DIDYoung_{ats}$ that is equal to the prior $Trend_t$ coefficient. We should see the negative coefficient because the workers most likely to bear the cost of the FMLA are young workers, due to their increased probability of having a child and using the leave compared to older workers. Moreover, we would expect to see the negative coefficient only in full-time and over-time wage gaps for women. The reason we would not expect to see convergence for part-time women is that the FMLA is not targeted at part-time workers. Finally, we would not expect to see the convergence for black men because women tend to make more FMLA claims than men (Tompson and Werner, 1997). The racial gaps between white men and black men are used as a placebo test.

The results from estimating the DID specification in Equation 4 are given in Table 7. The results show exactly what we would expect to see if the FMLA were causing the slow-down in gender wage convergence. For white women, we see that there is a

statistically significant negative point estimate for the coefficient of $DIDYoung$ for full-time and over-time workers. We do not see a statistically significant negative estimate for part-time white women. The finding confirms our prior that the effect of FMLA would be concentrated on full-time and over-time workers. A null hypothesis for $\beta_1 + \beta_6 = 0$ cannot be rejected for any of the demographic groups or time segments. The result confirms our prior about the effect of the FMLA stalling gender wage convergence. We also do not find that there is a statistically significant effect for workers as a whole, given by the coefficient of DID. The finding confirms our second prior that the effect of FMLA is concentrated on those who may use the maternity-leave and therefore bear the costs of such leave. Finally, we do not see a statistically significant coefficient on either DID or $DIDYoung$ for black men. The finding confirms our third prior that men are unlikely to be affected by the FMLA because they are less likely to use the FMLA than women. The results of the DID estimation confirm that the FMLA retarded gender wage convergence in the early 1990s.

These results show that maternity-leave laws and the FMLA were associated with a stagnation in the gender wage gap - a new and surprising finding. What is left to explain is how the FMLA may be causing stagnation in different segments of the economy than we anticipated seeing. The mechanism causing the stagnation associated with the FMLA may be running through the labor supply and labor demand for workers in different segments of the economy, as shown by Gruber (1994). To further examine the changes in labor supply and demand, the proportion of workers in a particular labor-market segment by demographic grou.

The labor force participation for white men does not appreciably change during the sample period, which would be expected if the FMLA were mainly affecting women instead of men. Where the biggest difference is seen is in the proportion of women who are working over-time jobs. For both white and black women, the proportion of workers who were working over-time was steadily increasing prior to the passage of the FMLA. However, in the post-FMLA period, the proportion growth has stagnated. This may be caused by a reluctance of firms to hire women in the highly-competitive and time-sensitive roles associated with working over-time due to the increased probability the women would take extended leave from work. Moreover, the proportion of black men in over-time work stagnated during this period as well, confounding the starkness of the results for women. Even with the confounding factor of results for black men, this finding deserves further consideration in future work on flexibility as well as on work regarding family-leave policies.

Table 7: Difference-In-Difference Estimates for the Effect of FMLA (1975–2015)

VARIABLES	White Women Part-Time	White Women Full-Time	White Women Over-Time	Black Women Part-Time	Black Women Full-Time	Black Women Over-Time	Black Men Part-Time	Black Men Full-Time	Black Men Over-Time
Trend	0.004***	0.005***	0.006***	0.002	0.005***	0.002*	-0.000	0.002***	-0.000
	(0.001)	(0.000)	(0.000)	(0.001)	(0.000)	(0.001)	(0.002)	(0.000)	(0.001)
Federal Only	-0.035*	-0.027***	-0.014	-0.080***	-0.036***	-0.059**	-0.052	-0.017	-0.011
	(0.019)	(0.010)	(0.012)	(0.029)	(0.010)	(0.028)	(0.047)	(0.011)	(0.019)
FMLA Enacted	0.069**	0.038**	0.039**	0.051	0.044***	0.018	0.080	0.047***	0.023
	(0.030)	(0.015)	(0.020)	(0.045)	(0.016)	(0.044)	(0.074)	(0.017)	(0.031)
Young	0.139***	0.168***	0.179***	0.085***	0.112***	0.081***	0.062*	0.039***	0.050***
	(0.015)	(0.008)	(0.010)	(0.022)	(0.008)	(0.022)	(0.037)	(0.008)	(0.015)
DID	-0.003	-0.000	-0.002	-0.001	-0.002	-0.001	-0.002	-0.002*	-0.000
	(0.002)	(0.001)	(0.002)	(0.004)	(0.001)	(0.003)	(0.006)	(0.001)	(0.002)
DID*Young	-0.003	-0.005***	-0.005***	-0.000	-0.002**	0.003	-0.001	-0.001	0.000
	(0.002)	(0.001)	(0.001)	(0.003)	(0.001)	(0.003)	(0.005)	(0.001)	(0.002)
Constant	-0.207***	-0.340***	-0.342***	-0.177***	-0.323***	-0.306***	-0.092***	-0.140***	-0.167***
	(0.012)	(0.006)	(0.008)	(0.018)	(0.007)	(0.018)	(0.030)	(0.007)	(0.012)
F-Stat for Trend-DID*Young = 0	0.13	0.45	0.13	0.23	2.73	2.03	0.02	0.85	0.00
Observations	156	156	156	156	156	156	156	156	156
R-squared	0.504	0.860	0.801	0.198	0.749	0.201	0.035	0.342	0.092

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

7 Conclusion

The work of Claudia Goldin has alerted the profession to the way in which labor market flexibility impacts wage inequality in the labor market. Inspired by this work, we set out to see if part-time work was historically more equal than full-time work. We find that indeed this is the case and that this finding is robust to different sampling and empirical specifications. Our result is consistent with Goldin's finding that women earn less per hour because they work fewer hours than men (Goldin, 2014). The result is also consistent with theoretical work on flexibility in the workplace and contract design (Goldin, 2014; Goldin and Katz, 2011). We also document evidence that imposed labor-market flexibility, through the introduction of family-leave laws, was coincident with the depression of gender wage convergence, which is consistent with prior work (Gruber, 1994). The key lesson from this work is that flexibility is priced-in to wage and compensation contracts for workers and firms. Therefore, flexibility that arises due to a workplace's environment, culture, and organization can play a constructive role in reducing gender wage inequality. Flexibility that arises in an occupation due to government mandates and programs, on the other hand, can have the opposite effect of slowing down the steady march towards gender wage parity in the labor market.

References

ACEMOGLU, D. AND J.-S. PISCHKE (1999): "The structure of wages and investment in general training," *Journal of political economy*, 107, 539–572.

ALBRECHT, J., M. A. BRONSON, P. S. THOURSIE, AND S. VROMAN (2018): "The career dynamics of high-skilled women and men: Evidence from Sweden," *European Economic Review*, 105, 83–102.

ALTONJI, J. G. AND C. PAXSON (1986): "Job characteristics and hours of work," .

ALTONJI, J. G. AND C. R. PIERRET (2001): "Employer learning and statistical discrimination," *The Quarterly Journal of Economics*, 116, 313–350.

ALTONJI, J. G. AND E. USUI (2007): "Work hours, wages, and vacation leave," *ILR Review*, 60, 408–428.

ANGRIST, J. (1991): "Grouped-data estimation and testing in simple labor-supply models," *Journal of Econometrics*, 47, 243–266.

ANTECOL, H., K. BEDARD, AND J. STEARNS (2018): "Equal but Inequitable: Who Benefits from Gender-Neutral Tenure Clock Stopping Policies?" *American Economic Review*, 108, 2420–41.

AVERETT, S. L. AND J. L. HOTCHKISS (1996): "Discrimination in the Payment of Full-Time Wage Premiums," *Industrial and Labor Relations Review*, 49, 287–301.

BANA, S., K. BEDARD, AND M. ROSSIN-SLATER (2018): "The impacts of paid family leave benefits: regression kink evidence from California administrative data," Tech. rep., National Bureau of Economic Research.

BERTRAND, M., C. GOLDIN, AND L. F. KATZ (2010): "Dynamics of the gender gap for young professionals in the financial and corporate sectors," *American Economic Journal: Applied Economics*, 2, 228–255.

BLANK, R. M. (1990): "Are part-time jobs bad jobs?" *A future of lousy jobs*, 123–155.

BLAU, F. D. AND L. M. KAHN (2007): "Changes in the Labor Supply Behavior of Married Women: 19802000," *Journal of Labor Economics*, 25, 3143–3259.

——— (2017): "The Gender Wage Gap: Extent, Trends, and Explanations," *Journal of Economic Literature*, 55, 789–865.

CHARLES, K. K. AND J. GURYAN (2008): "Prejudice and wages: an empirical assessment of Beckers The Economics of Discrimination," *Journal of political economy*, 116, 773–809.

COOK, C., R. DIAMOND, J. HALL, J. LIST, AND P. OYER (2018): "The Gender Earnings Gap in the Gig Economy: Evidence from over a Million Rideshare Drivers," Tech. rep., The Field Experiments Website.

CORTES, P. AND J. PAN (2017): "When Time Binds: Substitutes for Household Production, Returns to Working Long Hours, and the Skilled Gender Wage Gap," .

DAS, T. AND S. W. POLACHEK (2015): "Unanticipated effects of California's paid family leave program," *Contemporary Economic Policy*, 33, 619–635.

DESLAURIERS, J., B. DOSTIE, R. GAGNÉ, J. PARÉ, ET AL. (2018): "Estimating the Impacts of Payroll Taxes: Evidence from Canadian Employer-Employee Tax Data," Tech. rep., Institute for the Study of Labor (IZA).

FEENBERG, D. AND E. COUTTS (1993): "An introduction to the TAXSIM model," *Journal of Policy Analysis and management*, 12, 189–194.

FLOOD, S., M. KING, S. RUGGLES, AND J. R. WARREN (2017): "Integrated Public Use Microdata Series, Current Population Survey: Version 5.0. [dataset]," .

GOLDIN, C. (2014): "A grand gender convergence: Its last chapter," *The American Economic Review*, 104, 1091–1119.

GOLDIN, C. AND L. F. KATZ (2011): "The cost of workplace flexibility for high-powered professionals," *The Annals of the American Academy of Political and Social Science*, 638, 45–67.

GRUBER, J. (1994): "The incidence of mandated maternity benefits," *The American economic review*, 622–641.

HIRSCH, B. T. (2005): "Why do part-time workers earn less? The role of worker and job skills," *ILR Review*, 58, 525–551.

ISAAC, E. (2018): "Suddenly Married: Joint Taxation and the Labor Supply of Same-Sex Married Couples After U.S. v. Windsor," .

JUHN, C. AND D. I. KIM (1999): "The effects of rising female labor supply on male wages," *Journal of Labor Economics*, 17, 23–48.

KLEVEN, H., C. LANDAIS, J. POSCH, A. STEINHAUER, AND J. ZWEIMLLER (2019): "Child Penalties across Countries: Evidence and Explanations," *American Economic Review (Papers and Proceedings)*, 109, 122–126.

KLEVEN, H. J., C. LANDAIS, AND E. SAEZ (2013): "Taxation and International Migration of Superstars: Evidence from the European Football Market," *American Economic Review*, 103, 1892–1924.

MAS, A. AND A. PALLAIS (2017): "Valuing Alternative Work Arrangements," *American Economic Review*, 107, 3143–3259.

NATIONAL BUREAU OF ECONOMIC RESEARCH (2013): "Internet TAXSIM Version 9," .

NOONAN, M. C., M. E. CORCORAN, AND P. N. COURANT (2005): "Pay differences among the highly trained: Cohort differences in the sex gap in lawyers' earnings," *Social forces*, 84, 853–872.

PITTS, M. M. (2003): "Why choose womens work if it pays less? A structural model of occupational choice," in *Worker Well-Being and Public Policy*, Emerald Group Publishing Limited, 415–445.

REBITZER, J. B. AND L. J. TAYLOR (1995): "Do labor markets provide enough short-hour jobs? An analysis of work hours and work incentives," *Economic Inquiry*, 33, 257–273.

ROSEN, S. (1986): "The theory of equalizing differences," *Handbook of labor economics*, 1, 641–692.

RUGGLES, S., K. GENADEK, R. GOEKEN, J. GROVER, AND M. SOBEK (2017): "Integrated Public Use Microdata Series: Version 7.0 [dataset]," .

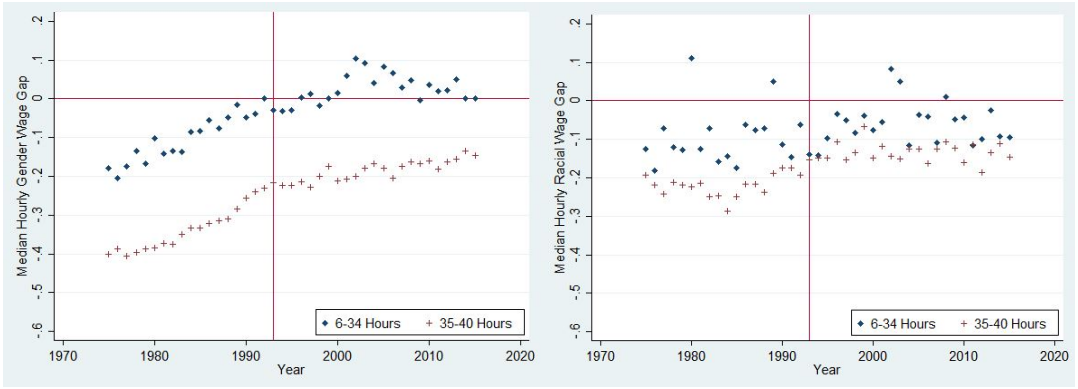
SUMMERS, L. H. (1989): "Some Simple Economics of Mandated Benefits," *American Economic Review (Papers and Proceedings)*, 79, 177–183.

TOMPSON, H. B. AND J. M. WERNER (1997): "The Family and Medical Leave Act: Assessing the Costs and Benefits of Use," *Emp. Rts. & Emp. Pol'y J.*, 1, 125.

WALDFOGEL, J. (1999): "The impact of the family and medical leave act," *Journal of Policy Analysis and Management*, 281–302.

A Appendix

Figure 12: **Uncontrolled Median Wage Gaps Excluding Minimum Wage Workers (1975 - 2015)**

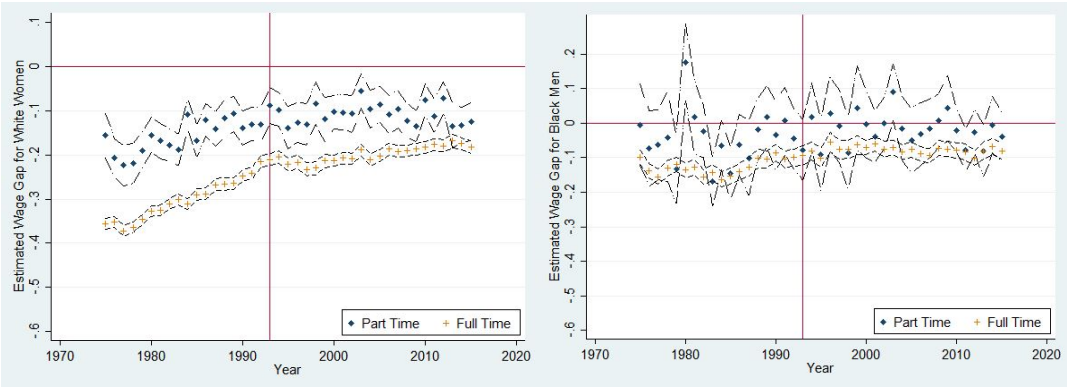


(a) (b)

Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the raw median wage gaps between white men and workers from two demographic categories: white women (a) black men (b) omitting workers who fail to earn more than the minimum wage. The key finding of these graphs is that, in line with expectations, the gender wage gaps are smaller for part-time workers than full-time workers even when omitting workers who fail to earn more than the minimum wage. However, we do not see this pattern for black men. Therefore, the flexibility hypothesis is supported because women have higher roles in household production and therefore demand more flexibility than their male counterparts.

Figure 13: **Controlled Estimated Wage Gaps Excluding Minimum Wage Workers (1975 - 2015)**

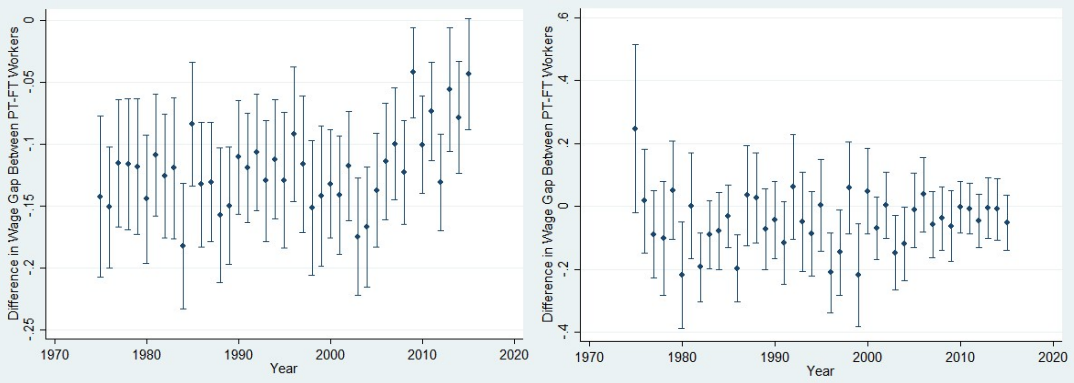


(a) (b)

Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the controlled, estimated wage gaps between white men and workers from Equation 1 from two demographic categories: white women (a) and black men (b) omitting workers who fail to earn more than the minimum wage. The key finding of these graphs is that, in line with expectations, the gender wage gaps are smaller for part-time workers than full-time workers. However, we do not see this pattern for black men. Therefore, the flexibility hypothesis is supported because women have higher roles in household production and therefore demand more flexibility than their male counterparts.

Figure 14: **Difference in Wage Gaps Excluding Minimum Wage Workers (1975 - 2015)**

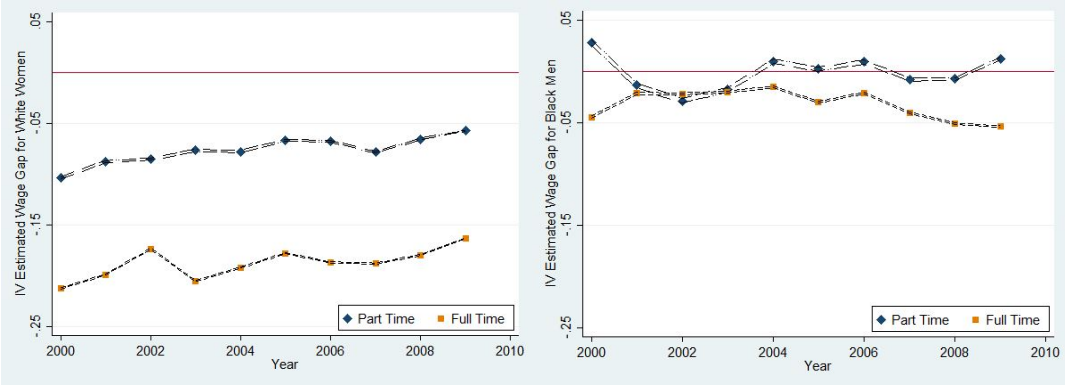


(a) (b)

Source: Annual Social and Economic Supplement of the Current Population Survey available through IPUMS

Notes: In the figures above we have reported the difference in the controlled, estimated wage gaps between white men and workers from a modified version of Equation 1 from two demographic categories: white women (a) and black men (b) omitting workers who fail to earn more than the minimum wage. The modified Equation 1 includes all workers and indicator variables which are the interaction between the demographic categories and the hours-worked status. For instance, there are two variables for white men: white man part-time and white-man over-time, where white man full-time is omitted. The key finding of these graphs is that, in line with expectations, the gender wage gaps are smaller for part-time workers than full-time workers. However, we do not see this pattern for black men. Therefore, the flexibility hypothesis is supported because women have higher roles in household production and therefore demand more flexibility than their male counterparts.

Figure 15: **Instrumented Estimated Wage Gaps Excluding Minimum Wage Workers (1975 - 2015)**



(a) (b)

Source: American Community Survey available through IPUMS

Notes: In the figures above we have reported the controlled, estimated wage gaps between white men and workers from the IV procedure outlined in Equation 1 and Equation 2 from two demographic categories: white women (a) and black men (b) omitting workers who fail to earn more than the minimum wage. The key finding of these graphs is that, in line with expectations, the gender wage gaps are smaller for part-time workers than full-time workers even when omitting workers who fail to earn more than the minimum wage. However, we do not see this pattern for black men. Therefore, the flexibility hypothesis is supported because women have higher roles in household production and therefore demand more flexibility than their male counterparts.