RIETI Discussion Paper Series 19-E-024

Costs of Employment and Flexible Labor Demand: Evidence from Maternity and Parental Leave Reforms

ASAI, Yukiko
Waseda University

The Research Institute of Economy, Trade and Industry
https://www.rieti.go.jp/en/

RIETI Discussion Paper Series 19-E-024
March 2019

Costs of Employment and Flexible Labor Demand: Evidence from Maternity and Parental Leave Reforms[1]

Yukiko Asai[2]
(Waseda University)

Abstract

This study examines the effects on workers from the labor-demand response to the costs of mandated maternity and parental leave programs. Japan introduced generous parental leave policies in the 1990s, but for many years, firms still had to pay the social insurance payments during leave, amounting to 13 percent of earnings. A series of reforms occurring in 2000, 2001 and 2014 which gradually reduced these costs to zero. This paper uses this quasi-experimental variation in the cost of female employment to measure the labor demand response. I focus on two key outcomes: starting wages of women of child-bearing age and the probability they are hired on permanent contracts. I find that a 100 thousand yen (approximately $1,000 USD) decrease in the costs of employment during leave increases the probability of starting on a permanent contract by 1.6 percentage points, and increases starting pay by 3.3 percent. In contrast with previous studies, the universal social insurance program setting I study allows me to separate the effects of changes in costs from endogenous responses by workers and firms to the available benefits. These findings have important implications for other countries mandating similar benefit schemes.

Keywords: Maternity Leave, Female Employment
JEL classification: J13, J21, J22

[1]This study is conducted as a part of the Project "Employment System Reform to Use Human Resources More Efficiently (PI: Hideo Owan)" undertaken at Research Institute of Economy, Trade and Industry (RIETI). This study is supported by the Grants-in-Aid for Scientific Research (Grant no.:18H03632) and Institute of Economic Research, Hitotsubashi University. The author would like to thank Hideo Owan, Dmitri Koustas and seminar participants at RIETI for their comments and suggestions, which contributed to significantly improving the paper. I thank the Statistics Bureau of the Ministry of Internal Affairs and Communications, and the Ministry of Health, Labor and Welfare for making data available for research.
[2] Waseda University, E-mail: y.asai3@kurenai.waseda.jp

1 Introduction

Persistent gender gaps in employment and wages exist in many OECD countries. While some of this reflects a personal choice of women to remain at home with their young children, part of this is due to labor market institutions, government policies and a corporate environment that make it difficult to stay employed during pregnancy and to return to work after childbearing. This paper studies one aspect of this: government-mandated maternity and parental leave programs. On one hand, these programs encourage job and employment continuity of new mothers; at the same time, they impose costs for employers and introduce incentives for firms to reduce these costs, which can manifest in lower starting wages and reduced employment. The equilibrium implications for wages and employment of women of child-bearing age, the key target of these policies, can therefore be ambiguous.

Depending on a country's institutions, there are different types of costs employers face from maternity and parental leave. In an institutional setting where the firm provides leave benefits, the firm is on the hook for income replacement during leave. In many countries, maternity and parental leave is a mandatory benefit financed by the social insurance system. This is also how California's maternity leave in the U.S. is financed. The workers' income replacement functions similarly to unemployment insurance while on leave: the firm is responsible for social insurance payments, borne equally by men and women, to finance the system. In addition to these direct costs, there is the effect on firm output from losing a worker. These costs depend on whether the job can be easily replaced by hiring a temporary worker or whether tasks can be covered by coworkers. In addition, in almost all countries, mandated

benefits only apply to workers on full-time permanent employment contracts, and not workers doing part-time and temporary work.

In this paper, I study a unique series of reforms that gradually lowered costs to firms of female workers who go on maternity and parental leave. The particular context is Japan, a country that with relatively generous mandatory leave policies of 13.5 months with income replacement financed via the social security system. However, for many years, firms still had to pay for the worker's social insurance payments during leave (namely medical and pension contributions), amounting to 13 percent of earnings. A series of reforms occurring in 2000, 2001 and 2014 gradually reduced these costs to zero.

Using matched worker-firm data, I study the effects of these reforms on employment and starting wages of new hires. In particular, I examine the differential effects of these reforms on women of child-bearing age. I also examine another channel that has received less attention in the literature, substitution towards part-time/ temporary work contracts. Specifically, I investigate whether younger women received lower starting pay and were more likely to be hired on temporary contract in years where firms faced higher costs of maternity and parental leave.[1]

My findings are as follows: I find large effects from reducing employment costs during leave. Reducing costs by the equivalent of a 100 thousand yen (≈$1,000) increased the probability of being hired as a permanent worker by 1.7 percent and increased starting wages by 2.5 percent. These effects are larger for large firms and

[1]The share of part-time and temporary employment among women has been increasing in Japan. While more women participate in the labor market in 2015 than 1990, among those in the labor market, 42% of women were on temporary contracts in 2015, compared with just 21% in 1990.

firms that employ many women.

This study contributes to the previous literature on mandated benefits in three key ways. Whereas other studies have investigated the introduction or expansion of mandated benefits and the effects of their costs on employment and wages (Scott et al., 1989; Baicker and Chandra, 2006), I focus on an experiment that reduced these costs. Second, I am able to separate the effects of changes in costs from endogenous responses by workers and firms to the menu of benefits. In the U.S. case, employers can switch the type of plan or quality of the insurance program. In addition, high ability workers are likely to self select into better health/pension insurance programs. In the Japanese case, health and pension insurance programs are universal, so all persons enroll in the same plan, and all patients are free to choose any of the medical providers. The insurance rate is determined by the national agencies, does not vary by the frequency of usage nor demographic backgrounds of the employees, and hospital fees are highly regulated. Finally, the contingent workforce in Japan is large and used alongside permanent workers on a temporary basis, while entire occupations, such as janitors, are more typically outsourced or subcontracted in the U.S. and Germany to avoid benefits (Goldschmidt and Schmieder, 2017). It is therefore common to observe substitutability of temporary and permanent workers in the Japanese case.[2]

My results have a number of implications. First, if costs are differentially borne

[2]Workers on permanent contracts have no predetermined period of employment, work full-time and are covered by all social insurance programs. Conversely, temporary employees work part-time or fixed/short term employees, are paid less and are eligible for fewer social insurance programs. There are two main types of temporary employment contracts in Japan: temporary workers directly hired by companies and so-called "dispatch workers," contracted through a temporary staffing agency. In this paper, I focus on temporary workers directly employed by companies.

across different types of workers (men versus women, temporary versus permanent contracts), firms have incentives to change their employment composition. To the extent that the incidence of taxes during leave are paid by firms, this incentivizes firms to substitute towards temporary employment contracts as maternity/parental leave programs become more generous over time.[3] This has relevance for the design of optimal social insurance programs from the perspective of a social planner who wants to maximize equity across groups. These same incentives are present in the California system, for instance, which excludes independent contractors contracted by firms. Moreover, this has important implications for gender wage and employment gaps.

This paper proceeds as follows: Section 2 discussed related literature. Section 3 provides an overview of the institutional background and discusses the relevant policy reforms. Section 4 provides a brief theoretical overview with testable predictions. Section 5 discusses the datasets used in this paper. Section 6 lays out the empirical strategy, the results of which are presented in Section 7. Section 8 concludes.

2 Related Literature

This study is closely related to the literature investigating the effects of the costs of employer benefits on employment and wages. These studies tend to find either little effect or a small decrease in employment of those affected, while the effects on wages

[3]The income replacement rate during parental leave was changed several times during this period. However, there was little change in female labor force participation after the increase in replacement rate. see Asai (2015) for details.

are mostly negative.[4] For example, Gruber and Krueger (1991) investigate the introduction of workers' compensation insurance. They find 86% of the insurance costs are shifted onto workers, with little effect on employment. Gruber (1994) studies increased insurance costs from adding maternity benefits to health insurance plans as mandated by legal changes. Comparing states who passed mandates with states who did not, he finds that the increase in the cost reduced hourly wages of women of childbearing age by 5.4 percentage points, but had little effects on employment. Baicker and Chandra (2006) exploit variation in health insurance premiums coming from shocks to medical malpractice payments. In contrast to earlier work, this study did find employment effects: a 10% increase in health insurance premiums reduced the probability of employment by 1.2pp, hours worked by 2.4%, and the probability of working full time by 1.9%. It also reduced the wages of workers who are covered by employer-provided health insurance by 2.3%.

This study is also related to the broader literature on maternity/parental leave policies and labor market outcomes. Maternity and parental leave programs allow a woman to spend time with her newborn child when the value of time with the child is relatively high, and facilitate returning to work after childbearing.[5] Previous studies have exploited three main sources of policy variation: (1)implementation of leave programs (Baum, 2003; Klerman and Leibowitz, 1997), (2) extensions of leave (Baker and Milligan, 2008; Schonberg and Ludsteck, 2007; Lalive and Zweimuller, 2009; Lalive et al., 2013; Gallen, 2019), and (3) increases in cash benefits during leave

[4]See Currie and Madrian (2000) for literature review before 2000.

[5]Consider a model where a woman makes a discrete choice of whether or not to work in each period. The value of staying at home relative to working is posited to decline as the child grows older. (Heckman (1980); Klerman and Leibowitz (1997)).

(Lalive et al., 2013; Asai, 2015).[6] These studies are typically framed as investigating the labor *supply* response. Labor supply outcomes typically studied include returning to work (at any employer) and job continuity (returning to the *same* employer). The literature finds that *unpaid* leave has little impact on mothers' labor supply. For example, Klerman and Leibowitz (1997) and Baum (2003) investigate the labor supply effect of the US's Family and Medical Leave Act and found no statistically significant effect on employment or leave. On the other hand, studies examining the effect of job-protected leave with cash benefits find an increase in job continuity. Baker et al. (2008) find that extending leave in Canada induced women to return to their previous employers. Schonberg and Ludsteck (2007) find similar results looking at a similar policy implementation in Germany.

I contribute to this literature by investigating whether the cost of leave affects employees' wages, employment, and the incidence of temporary contracts. As far as I know, this is the only study to examine the relationship between the cost of maternity leave and labor demand for female workers. My study is also the first to examine implications in a country with an universal leave system.

3 Institutional Background and Policy Reforms

3.1 Social Insurance Programs in Japan

All Japanese citizens are covered by three social insurance programs; universal health insurance, pension insurance, and elderly care insurance. Insurance rates are deter-

[6] See Olivetti and Petrongolo (2017) for a recent review of literature.

mined by the government and do not vary by the frequency of usage nor demographic backgrounds of the persons (i.e. there is no risk or experience ratings). All firms are responsible for paying 50% of social insurance payments for employees if they work 30 or more hours. Short-time or temporary workers working fewer hours are entirely responsible for paying for their social insurance contributions by themselves.

Figure 1 shows the changes in the premiums of the different components of the Japanese social insurance system over time. The largest component is pension insurance. From the mid 1990s, the pension insurance rate was 17.34% of earnings excluding bonus payments, until a reform in April 2003 included these bonus payments while at the same time reducing the rate to 13.58%. Another pension reform in 2004 reversed this, mandating an increase in the rate by 0.354% every year from October 2004 until the rate reached at 18.30% in September 2017. The second largest component is health insurance. The premium remained at around 8% until 2009, when the average rate started to increase from September 2009 after the program was changed to allow for different regional rates. The rate shown in the figure is the national average in September 2009; however, the local rates only vary by approximately 0.5% depending on the health care spending in each region. In addition to the health and pension insurance rates, persons who are age 40 or older are required to pay for elder care insurance. Mandatory elder care insurance was introduced in 2000 to cover the cost of home care and nursing home for elderly persons. This rate was 0.6% when it first introduced and gradually increased to 1.58% in later years. Not shown in the figure is employment insurance, financing parental leave and unemployment, which amounts to less than 1%.

3.2 Policy Reforms to Social Insurance Costs During Leave

Although worker's income during maternity/parental leave is replaced by employment insurance, firms (and workers) were initially responsible for paying these social insurance rates described in the previous section during the leave. These costs are gradually eliminated in a series of reforms I describe in this section. These reforms are the key variation I will be using in this paper.

Figure 2 summarizes the social insurance costs during leave over this time period with an illustrative example of the costs faced by employers. All female workers who give birth are qualified to take 3.5-months maternity leave; female workers on permanent contracts are entitled to another 10-months parental leave.[7] The horizontal line in the figure denotes months since the childbirth, and the vertical line represents the insurance rate as a percent of earnings. The insurance rate shown in the figure is as of 2000; the health insurance was 8.5%, and pension insurance was 17.35%. The employers were responsible for paying for 50% of the social insurance payment during leave, which is $\sim$ 12.9% of income (depicted as the dot "Before reform" line). Note that this figure describes the social insurance rate for women under 40; therefore, elder care insurance is not pictured here.

In April 2000, the pension insurance payment during parental leave was eliminated, which is depicted as the solid red line in the figure. In January 2001, the health insurance payment during parental leave was removed, which is depicted by

[7]Maternity Leave is the gray area in the figure, which is 42 days before and 56 days after childbirth. It is mandated by the Labor Standards Act (1947), and it is illegal for new mothers to work within 42 days of childbirth for maternal health reasons. After 3 months maternity leave, mothers are entitled to up to 10 months of parental leave under the Child Care and Family Care Leave Act (1992). See Asai (2015) for more details on the leave programs in Japan.

the dashed line in the figure. After the 2000 and 2001 reforms, firms were still responsible for the social insurance payment during maternity leave until April 2014, when the social insurance payment during maternity leave was completely eliminated (depicted as the solid blue line). Importantly, these reforms were not accompanied by an increase in the length of leave that could have the potential to significantly shift the labor supply curve.

In this paper, I measure the cost of employment based on these changes in social insurance rates described in Figure 1 and the changes in the social insurance rates during leave described in Figure 2. The cost of employment during leave is defined as zero for part-time and temporary workers because the firms typically do not pay social insurance costs for them, and the lack of tenure and long-term employment contracts make it difficult for part-time and temporary workers to take leave.[8] In fact, the take-up rate of leave among temporary workers is low compared to the rate among permanent workers.[9] Note that fathers are also eligible to take leave, but the

[8]All working mothers are eligible to take maternity leave. However parental leave is only available to workers who have contributed to the employment insurance at least for a year before taking the leave. Before the reform in April 2005, part-time and temporary workers are not qualified to take the parental leave. The reform in April 2005 made the leave available to part-time and temporary workers as long as they pay to the employment insurance cost and meet the requirement. Even though some part-time and temporary workers are technically qualified to take the leave, the employment insurance program states that persons with less than one year of continuous employment, with a contract that will terminate in less than a year, and with less than two days of work per week can be excluded from the parental leave program. This requirement and their limited contract length make it difficult for them to take leave.

[9]According to the Annual Population and Social Security Surveys from 2005 to 2009, only around 18% of temporary workers remain employed after the first childbirth. 14% of temporary workers return to work after three months of maternity leave (i.e., without parental leave) and only 4% of temporary workers take parental leave followed by maternity leave. On the other hand, around 50% of permanent workers remain employed after the first childbirth. 43% of permanent workers take both maternity leave and parental leave. See Work-life Balance Report 2011 by the Cabinet Office Japan for more details.

take-up rate by fathers over this period was less than 1.0%.

4 Theoretical Framework

The effect of mandated benefits can be predicted in the same framework used to study the implications of payroll taxes.[10] The payroll tax is modeled as a government policy that shifts back the labor demand curve. The payroll tax increases the cost of hiring, decreases wages received by workers and reduces overall employment. On the other hand, mandated benefits, such as health insurance, are considered preferable to a payroll tax as long as the mandated benefit provides some value to workers, because they can have a smaller reduction on employment in equilibrium. The magnitude of the effects on employment depends on how much workers value the benefit provided. The model predicts that when workers value the mandated benefits at the same rate that it costs to provide the benefits, the wage received by workers will be reduced while employment will remain unchanged in the new equilibrium.

Panel (1) of Figure 3 depicts the labor market for female workers. The initial equilibrium is the interaction of D_0 and S_0. Consider the case where a government mandates health insurance during leave and the leave is only available to female workers. The demand curve for females will shift downward, from D_0 to D_1. On the other hand, the equilibrium wage and employment for male workers would stay at the D_0 and S_0 level, and this would create a difference in employment rates between genders. (Note: the labor market for male workers is not depicted.) Suppose

[10] See Summer (1989), Gruber and Krueger (1991) and Baicker and Chandra, 2006 for conceptual framework.

female workers value the insurance during leave at the same rate as the provision of insurance. This will shift the female labor supply curve to the right, and the new equilibrium in the market is the intersection of D_1 and S_1. The female wage decreases by the price of health insurance. This create a gender wage gap. However, female workers' total compensation is equal to male workers.

We have seen that the wage and/or employment of female workers may decrease after the provision of health insurance. The magnitude of the effects on employment and wages may also depend on institutional constraints in the labor market. Nondiscrimination provisions may prohibit workers with a specific demographic group being treated differently. For example, when labor law prohibits unfair dismissals and/or wage differentials, it would not be possible for a firm to dismiss female workers and lower wages of female workers in response to the cost increase due to the introduction of maternity/parental leave. Hence, after the introduction of the benefit, the wage received by female workers will remain the same level as before the introduction and female workers can receive the maternity/parental benefits that they value as well. Since it is costly for a firm to employ females, the demand for female labor would decrease, and this will shift the labor supply curve upward, which is depicted in Panel (2) of Figure 3, as a shift from S_0 to S_1. In the new equilibrium, total employment will be reduced.

Another possible effect is substitution between permanent and part-time/temporary workers. In the US, nondiscrimination provisions in the Federal Tax Code prohibit firms from discriminating on non-wage benefits. However, part-time workers are excluded from these provisions, and this has created unintended consequences in the

widening of inequality within a firm. In response to the nondiscrimination rules, firms who want to minimize the cost moved workers who need higher insurance costs into a temporary or part-time contract because part-time/temporary workers are typically not covered by the employees' benefits (See Scott et al. (1989) and Baicker and Chandra, 2006 for details.). Carrington et al. (2002) show empirical evidence that firms move workers with higher benefits into part-time and seasonal positions.

To summarize, we expect two main channels from reducing the costs of employment for women. First, we expect wages to rise. Second, we expect employment to increase, which may be manifested as increases in permanent contracts.

5 Data

To test these predictions, we require firm-level data that includes wages and contract types. To this end, I rely on a number of different datasets. The main dataset I employ is the Basic Survey on Wage Structure (BSWS). The BSWS is conducted by the Ministry of Health, Labor, and Welfare and gathers establishment-level data from randomly selected establishments. About 6 percent of establishments are surveyed each year. Establishments are asked to submit firm level records as well as a random sample of workers' payroll records from June of each year. These payroll records are accompanied by detailed demographic information, including age, gender, tenure, monthly work hours and earnings. Among these establishments, 20 percent of them will continued to be surveyed next year.[11]

[11]It's possible for an establishment to be surveyed three years in a row, but the probability is just $0.2^2 = 0.04$ percent.

Despite the fact that firms are resampled in the BSWS, there is no direct longitudinal link in the BSWS data before 2008. To create a panel, I match the BSWS to the Economic Census conducted every 2 or 3 years. The BSWS draws from the Census and the firms can be linked back to it. Once firms are linked back to the Census, a longitudinal link can then be formed. Note that only firms between Census periods can be linked. For instance, the BSWS undertaken in 1998, 1999, 2000 and 2001 used the 1996 Economic Census to draw a sample, and so firms in these years can be linked. From 2008, a link is provided that allows me to link firms across years. For this paper, I rely on the BSWS data from 1998-2001, and 2008 to 2014, since these are years overlapping with the largest reforms described above.

In cleaning the data, I make a number of sample restrictions. First, because of adjustment frictions, the margin we would expect to see immediate responses from the reforms would come from new hires. Fortunately, the tenure variable in the BSWS allows me to focus on new hires. The typical contract length is 2 years before renegotiation, so I restrict analysis to workers with tenure <2 years.

To observe how the change in the cost of employment affects labor demand, I focus on firms and establishments with 10 or more employees. This sample restriction omits self-employed firms and family businesses, that amount to approximately 10 percent of firms in the Economic Census. I also restrict my analysis to employees who are age 18 to 60 years old. 18 is the age where high school graduates can begin to enter the labor force full-time, and 60 is the age when people can start to retire.

In the analysis below, I will test for heterogeneity on a number of different dimensions. I divide the establishments into large establishments and small establishments.

Small establishments are defined as having fewer than 300 employees and the large establishments are defined as having 300 or more employees. Employment adjustment may be faster in the large firms than the small firms, because the large firms may have larger margin of adjustment. In addition, since maternity leave and parental leave are taken by mainly female workers, firms with majority female employees may benefit more from the two reforms. I also do subgroup analysis on firms with above and below median shares of female employment.

To calculate the "risk" of having a baby among working mothers, I use the Longitudinal Survey of Newborns conducted by the Ministry of Health, Labour and Welfare. This survey collects data from the universe born between Jan 10-17 and July 10-17 in 2001, linked to the national Birth Register. Using these data, I calculate the probability of giving birth by age for women on permanent contracts who would thus be eligible for parental leave.

6 Empirical Strategy

This section describes my main empirical strategy. I combine two main ingredients: (1) the probability a working woman of a particular age eligible for leave gives birth with (2) costs of employment during maternity/parental leave, which vary based on the series of reforms discussed above. My key empirical strategy will be to compare outcomes of woman of childbearing age, with older women and men of the same age.

We expect that outcomes of women of child-bearing age would be affected after the reform relative to men and older women. From the perspective of the firm, there

is no differential cost due to leave between older women and older men. All of the differential costs comes from younger women of child-bearing age. I define the "risk" faced by the firm from hiring a woman of child-bearing age as the probability that a working woman eligible for leave has a child and thus goes on leave, for each age group. I do this based on the 2001 birth registry. Therefore, risk does not time-vary (but I will control for trends by age and gender in my regression.) Denote this risk $risk_{f,a(i)}$, where the subscript f indicates that risk varies across men and women (specifically, risk of leave is 0 for men), and the subscript $a(i)$ indicates that it varies across women only based on their ages.

Next, I will define the employment costs of leave to the firm. As discussed above, before 1999, firms were required to pay approximately 13 percent of earnings in payroll taxes during the 13.5 months of mandated leave. In 2000, these costs were reduced to 4.3 percent for the last 10 months of leave, and thus the average cost fell to 6.5 percent over the 13.5 months of leave.[12] In 2001, the payroll taxes only needed to be paid during the first 3.5 months of leave.[13]. In 2014, the costs fell to zero.

To put this in perspective, average monthly earnings for permanent workers in 1999 were approximately 197,200 yen, or about 1,972 USD per month.[14] In 1999, over the course of 13.5 months of leave, the firm needed to pay approximately 197,200*0.12925*13.5 = 344 thousand yen (≈3,440 USD). To facilitate interpretation, I define the independent variable $cost_t$ as the yen cost, in 100 thousands, for the average woman's leave, i.e. 3.44 in 1999, (0.12925*3.5 +0.0425 *10)*1.972=1.73

[12]The calculation is $\left(\frac{0.13*3.5+0.043*10}{13.5}\right)*100$
[13]The calculation is $\left(\frac{0.13*3.5}{13.5}\right)*100$
[14]100 yen is ≈1 dollar on average during the 2000s.

in 2000, 0.12925*3.5*1.972=0.892 from 2001, and 0 by 2014. Therefore, a "unit" change in this variable is 100,000 yen (approximately 1,000 USD).

The key object of interest is the interaction, $cost_t \times risk_{f,a(i)}$, which varies across men and women, women of different ages, and over time. The interaction is the expected leave costs of that worker over the next year. In Figure 4, I plot the values of this interaction for female workers of different ages across years. The risk of a worker going on leave is low for the youngest and oldest female workers, and reaches its peak in the early 30s. The costs were highest before 2000, and gradually fall over the period until they reach zero by 2014.

Putting together these pieces, my main estimating equation is shown below;

$$\begin{aligned} y_{it} = \beta_1 cost_t + \beta_2 risk_{f,a(i)} + \beta_3 cost_t \times risk_{f,a(i)} \\ + \xi X_{it} + \gamma_{J(i)} + \upsilon_{it} \end{aligned} \tag{1}$$

y_{it} is the dependent variable of interest for individual i in year t. Depending on the specification, y_{it} is an indicator for a permanent contract, or earnings. X_{it} includes controls that possibly time vary, including a female indicator and a quintic in age. $\gamma_{J(i)}$ denotes the establishment fixed effect. I will also include year fixed effects, age-by-year fixed effects, female-by-year fixed effects, and female-by-age fixed effects, $\gamma_t, \gamma_{a\times t}, \gamma_{f\times t}, \gamma_{f\times a}$, respectively. With year and gender-by-age fixed effects, the main effects for $cost_t$ and $risk_{f,a(i)}$ will not be separately identified, but their interaction, which varies at $t, f, a(i)$, will be. Standard errors are clustered at the firm level to account for correlation of the error structure within firms over time.

The effect of a unit increase in the cost of leave for firms on y_{it} is given as follows:

$$\frac{\partial y_{it}}{\partial cost_t} = \beta_1 + \beta_3 risk_{a(i),f} \tag{2}$$

We would expect the employment effects to be largest for higher risk groups, and therefore we expect β_3 to have an unambiguously negative effect on outcomes like permanent employment contract and starting wages. We might expect costs to reduce employment and earnings for all workers, and thus β_1 will be negative. However, there could also be general equilibrium effects from the substitution across worker types that might show up in β_1, that would make this sign ambiguous.

7 Results

Results from running Equation 1 are described in the section. Panel (a) of Table 1 reports the regression results for being hired on a permanent employment contract. Panel (b) reports the coefficients for log starting pay.

The three coefficients of interest—on $risk_{f,a(i)}$, $cost_t$ and their interaction—are reported in the tables. The first row shows the main effect for $risk_{f,a(i)}$. The interpretation is the effect of a unit change, i.e. going from a 0 percent chance of having children (e.g. men and older women) to a probability of 1, when $cost_t$ is zero, which occurs from 2014. These effects are large and negative. $risk_{f,a(i)}$ never takes on a value of 1, however, so this is an out of sample prediction. To better interpret the coefficient, we can multiply the coefficient by the value of risk for the highest risk group—32 year old women—who have a value for $risk_{f,a(i)}$ of 0.1156.

As shown by equation (2), the coefficients on $cost_t$ show the effect of cost when risk is zero, i.e., for men and older women. Because $cost_t$ varies only at the year level, this coefficient is only identified in the Column (1) specification without year effects. We should therefore interpret this coefficient with caution. We know for instance that the share of permanent employment is trending down as shown in Figure A1.

The coefficients on the interaction, $cost_t \times risk_{f,a(i)}$ pick up how the impact of cost varies with risk; this is the main coefficient of interest of my study. Row (4) shows the prediction for this highest risk group, multiplying the coefficient on the interaction by 0.1156, the value of $risk_{f,a(i)}$ for the highest risk group, as described above.

Table (1) of Column (1) reports results without year fixed effects, allowing the coefficient on $cost_t$ to be identified. In this regression, variation comes from 1) across women over time 2) across men and women, and 3) across women of different ages. The coefficient on $risk_{f,a(i)}$ in column (1) of Table 1(a) implies a 32 year old women is -0.522*0.1156*100=6 percentage points less likely to be hired on a permanent contract, and the comparable coefficient on log wages in Table 2(b) implies -0.264*0.1156*100=3 percent lower wages. The coefficient on $cost_t$ is positive for both permanent contracts and wages, suggesting that the control groups of men and older women tend to do better when costs are high. The coefficients imply that when costs increase by 100 thousand yen (approximately 1,000 USD), the probability of being hired on a permanent contract increases by 2.44 percentage points and wages increase by 4.77 percent for those with 0 risk of giving birth (i.e. men and older women). One interpretation of this finding is that firms substitute towards these

groups relative to women of child-bearing age when costs are high. Moving to the interaction, the prediction for the highest risk group shows that for a 100 thousand yen (approximately 1,000 USD) increase in cost, the probability of starting on a permanent contract decreases by (0.0244-0.271*0.1156)*100=0.7 percentage points and starting pay decreases by (0.0477-0.404*0.1156)*100=0.1 percent.

Column (2) adds year fixed effects. The variation here now comes from 1) across men and women, and 2) across women of different ages, within a year. The prediction shows that for a 100 thousand yen increase in cost, permanent employment decreases by 3.1 percentage points and starting pay decreases by 4.6 percentage points. Column (3) adds age-by-year fixed effects to control for trends across different age groups. Adding in this fixed effect means the variation now comes from across men and women of the same age within a year. The prediction changes only slightly from the column (2). Column (4) includes a female-by-year fixed effect, which means variation comes from across women of different age groups within a year. This fixed effect lowers the estimates slightly: the prediction shows that a 100 thousand yen increase in cost decreased permanent employment by 2.2 percentage points and decreased starting pay by 3.2 percentage points. Column (5) adds female-by-age fixed effects, which means variation comes from comparing women of the same age across years, and is similar to estimates in columns (1)-(3). Finally, Column (6) adds all the age, year, and female interactions. This is the most conservative specification taking into account the trends in the share of permanent employment and the average earnings by year, age and gender seen in Figure A1 and Figure A2. The prediction falls somewhat to -1.6 percentage points for being hired on a permanent employment

contract and -3.3 percent reduction for log starting pay for a 100,000 yen increase in costs, but the sign and the significance remain unchanged from column (5).

I next examine heterogeneity on three key dimensions: firm size, female employment shares, and time period of the reform. Table 2 shows the heterogeneity results by firm size. Columns (1) and (2) show the results for permanent employment for the small firms and large firms, respectively, while Columns (3) and (4) show the coefficients for starting pay. The specification used is the same as the column (6) of Table 1, my most conservative specification including all pair-wise interactions of age, year, and female, controlling for trends in outcomes by age, year and gender. The third row of the table below the interaction reports the mean of the dependent variable, which differs significantly across small and large firms. For instance, permanent contracts are less common in small firms, where only 62% of new hires are on permanent contracts, compared with 71% at large firms. Starting wages are also higher at larger firms.

The prediction shows that a 100 thousand yen increase in cost decreased permanent employment by 0.7 percentage points (or -0.0577/0.621=-9.3% change) for small firms and -3.3 percentage points (or -0.282/0.711=-39.7%) for large firms. Similarly, a 100 thousand yen increase in the cost decreased the starting pay by 2.2 percent for small firms and 5.8 percent for large firms. One reason why the magnitude of the coefficients may be bigger at large firms than small firms is because large firms may have more discrete tasks that are easily substitutable across workers.

Table 3 shows the results by above and below 50 percentile female employment shares. The specification is again the same as column (6) of Table 1, my most con-

servative specification including all the age, year, and female pair-wise interactions. Columns (1) and (2) show the results for starting on a permanent contract for firms with low female shares and high female shares, respectively. Column (3) and (4) show the coefficients for starting pay. The third row below the results again reports the mean of the dependent variable. We see that firms that employ the most female workers are much less likely to have permanent contracts: just 46.3 percent do, compared with 67.8 percent for the other firms. Starting wages are also much lower at female firms. The results imply a 100 thousand yen increase in the cost decreased permanent employment by 2.8 percentage points (or -0.242/0.784=-30.9%) for firms with low female shares and 1.2 percentage points (or -0.103/0.462=-22.3%) for firms with high female shares. Similarly, a 100 thousand yen increase in the cost decreased the starting pay by 3.8 percent for firms with low female shares and 5.4 percent for firms with high female shares. One reason why the magnitude of the coefficients on permanent employment regressions may be bigger at firms with lower female employment rates is because these firms have a higher permanent employment rate, which may indicate more tasks for permanent workers. On the other hand, the magnitude of the coefficients on starting pay regressions is slightly bigger at firms with high female firms because these firms saved more simply because they employ more female workers.

Table 4 shows the results for different sample periods. The specification is the same as the column (4) of Table 1 instead of the column (6) to compare the coefficient on risk across periods. Column (1) and (2) show the results for the 1998 to 2001 period, and column (3) and (4) show the results for 2013-2014. The coeffi-

cient on $risk_{f,a(i)}$ in column (1) implies a 32 year old women is -0.786*0.1156*100=9 percent less likely to be hired on a permanent contract when cost is 0, and the comparable coefficient in column (2) implies -0.67*0.1156*100=7.75 percent lower wages. Similarly, the coefficient on $risk_{f,a(i)}$ in column (3) implies a 32 year old women is -0.917*0.1156*100 = 10.6 percent less likely to be hired on a permanent contract, and the comparable coefficient in column (4) implies -0.693*0.1156*100=8 percent lower wages. Thus, there is no difference over time in starting contracts and pay, ceteris paribus. The coefficient on the interaction implies that a 100 thousand yen increase in costs decreased permanent employment by 1 percentage point or (-0.082/0.713=11.5%) and decreased starting pay by 1.7 percent for the 2000 and 2001 reforms. For the 2014 reform, the coefficients for starting on a permanent contract and the coefficient on pay are similar, albeit much less precise. This is not surprising because the cost reduction was much larger for the 2000 and 2001 reforms than for the 2014 reform.

8 Conclusion

In the 1990s, the Japanese government introduced maternity and parental leave programs. Although these leave programs provide many benefits to female workers, the program incurs costs for firms who hire them. For many years, firms still had to pay for the worker's social insurance payments during leave, amounting to 13 percent of earnings. A series of reforms occurring in 2000, 2001 and 2014 gradually reduced these costs to zero. This paper uses this quasi-experimental variation in the cost of

female employment to measure the labor demand response.

This study differs from the other studies on labor demand and maternity/parental leave in three key ways: First, I focus on an experiment that reduced social insurance costs, rather than extending benefits to workers. Second, since Japan's social insurance programs are universal, I am able to separate effects of changes in costs from endogenous responses by workers and firms to the menu of benefits, a problem with U.S. studies, for instance. Third, the contingent workforce in Japan is large and used alongside permanent workers on a temporary basis, therefore making it easier to observe substitution between permanent and temporary workers within firms.

Using Japan's rich firm-worker matched data, I examine the relationship between the costs of leave and labor demand for female workers of child-bearing age. I find that a decrease in costs increased permanent employment and the starting wage of female workers. The result implies that the higher costs of female employment may have discouraged firms to hire female workers, holding back progress on gender wage and employment gaps these policies were directly designed to promote. These findings have relevance for other countries implementing mandatory benefits financed through social insurance where some workers have higher probabilities of using the benefits.

References

Asai, Y., 2015. Parental leave reforms and the employment of new mothers: Quasi-experimental evidence from japan. Labour Economics 36, 72–83.

Baicker, K., Chandra, A., 2006. The labor market effects of rising health insurance premiums. Journal of Labor Economics 24 (3), 609–634.

Baker, M., Gruber, J., Milligan, K., 2008. Universal child care, maternal labor supply, and family well-being. Journal of Labor Economics 116 (4), 709–745.

Baker, M., Milligan, K., 2008. How does job-protected maternity leave affect mothers' employment? Journal of Labor Economics 26 (4), 655–691.

Baum, C. L., 2003. The effect of state maternity leave legislation and the 1993 family and medical leave act on employment and wages. Labour Economics 10 (5), 573–596.

Carrington, W., McCue, K., Pierce, B., 2002. Nondiscrimination rules and the distribution of fringe benefits. Journal of Labor Economics 20 (S2), S5–S33.

Currie, J., Madrian, B., 2000. Chapter 50 health, health insurance and the labor market. Vol. 3, Part C of Handbook of Labor Economics Edited by Ashenfelter, O. and Card, D. Elsevier, pp. 3309–3416.

Gallen, Y., 2019. The effect of maternity leave extensions on firms and coworkers. Mimeo.

Goldschmidt, D., Schmieder, J., 2017. The rise of domestic outsourcing and the evolution of the german wage structure. Quarterly Journal of Economics 132 (3), 1165–1217.

Gruber, J., 1994. The incidence of mandated maternity benefits. American Economic Review 84 (3), 622–41.

Gruber, J., Krueger, A., 1991. The incidence of mandated employer-provided insurance: Lessons from workers' compensation insurance. Tax Policy and the Economy 5, 111–144.

Heckman, J., 1980. Sample selection bias as a specification error: An application to estimation of female labor supply functions. Female Labor Supply, ed. James Smith. Princeton University Press, pp. 206–248.

Klerman, A., Leibowitz, A., 1997. Labor supply effects of state maternity leave legislation. In: Gender and Family Issues in the Workplace Edited by Blau, F. and Ehrenberg, R. Russell Sage Foundation, Ch. 3, pp. 65–85.

Lalive, R., Schlosser, A., Steinhauer, A., Josef, Z., 2013. Parental leave and mothers' careers: The relative importance of job protection and cash benefits. Review of Economic Studies 93 (1), 113–125.

Lalive, R., Zweimuller, J., 2009. How does parental leave affect fertility and return to work? evidence from two natural experiments. The Quarterly Journal of Economics 124 (3), 1363–1402.

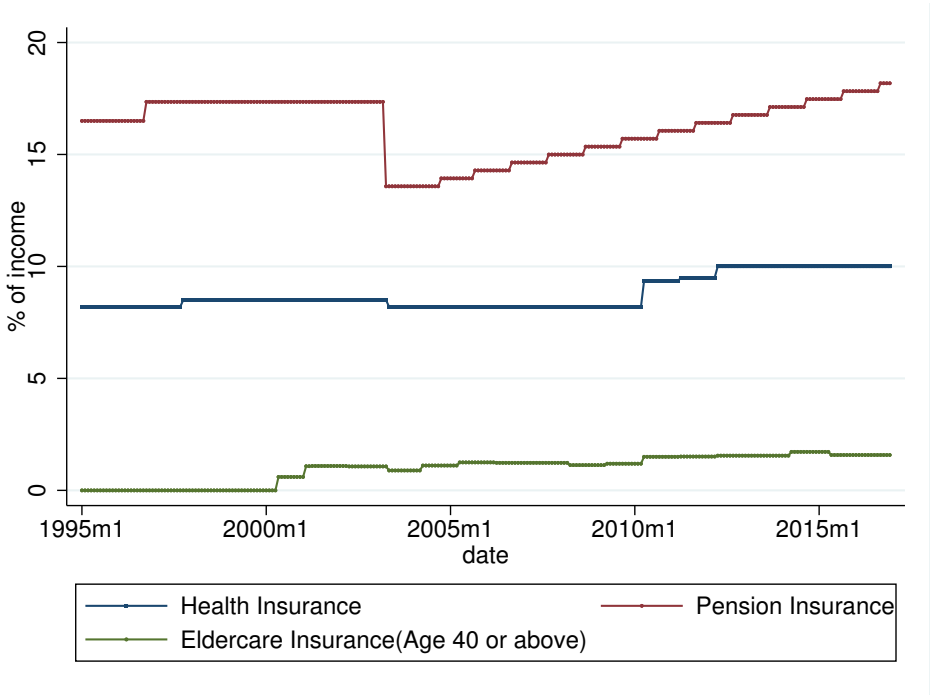
Olivetti, C., Petrongolo, B., 2017. The economic consequences of family policies: lessons from a century of legislation in high-income countries. The Journal of Economic Perspectives 31, 205230.

Schonberg, U., Ludsteck, J., 2007. Maternity leave legislation, female labor supply, and the family wage gap. IZA Discussion Papers 2699.

Scott, F., Berger, M., Black, D., 1989. Effects of the tax treatment of fringe benefits on labor market segmentation. Industrial and Labor Relations Review 42, 216–229.

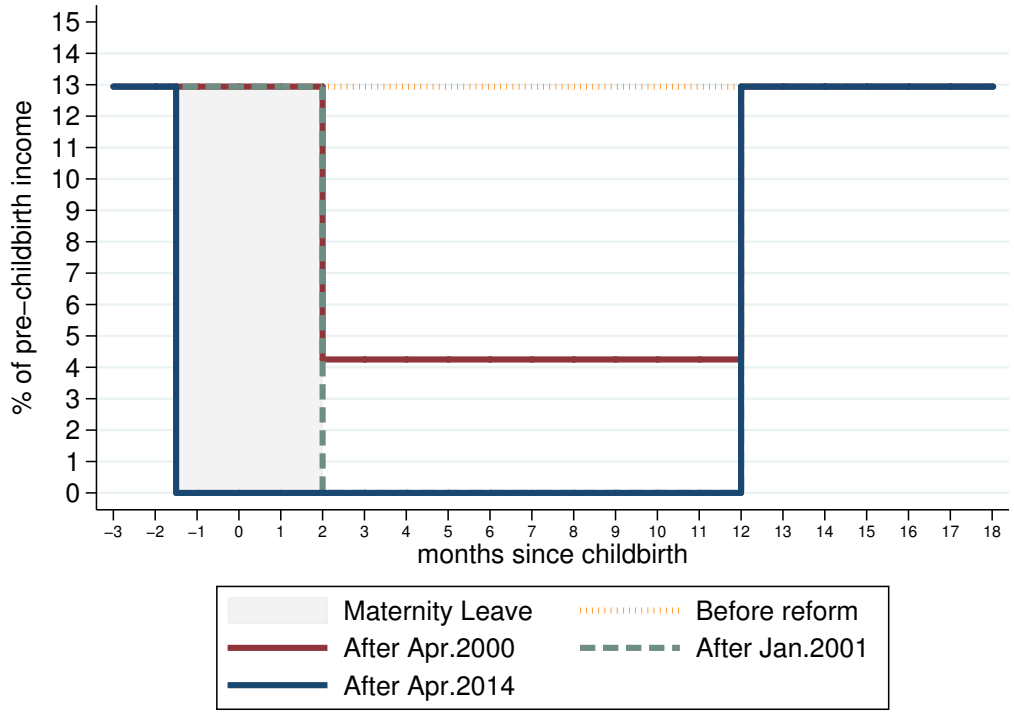
Summer, L., 1989. Some simple economics of mandated benefits. American Economic Review Papers and Proceedings 79 (2), 177–183.

Figure 1: Social Insurance Rate as % of Income



Note: The author creates the figure based on information from the Japan Health Insurance Association and Japan Pension Service. Bonus payments were excluded from income until March 2003 for the pension insurance calculation.

Figure 2: Social Insurance Costs During Leave Before and After the 2000, 2001 and 2014 Reforms



Note: The social insurance defined as the rate in 2000. The horizontal line denotes months around the childbirth, and the vertical line represents the social insurance rate. The area marked in gray is the maternity leave period, which is 1.5 months before birth and two months after birth. Mothers are allowed to take parental leave until the 12th month. The health insurance was 8.5%, and pension insurance was 17.35% in 2000, and employers are responsible for half the insurance cost during leave, i.e., 12.925 %.

Figure 3: Effects of Mandated Benefits in Female Labor Market

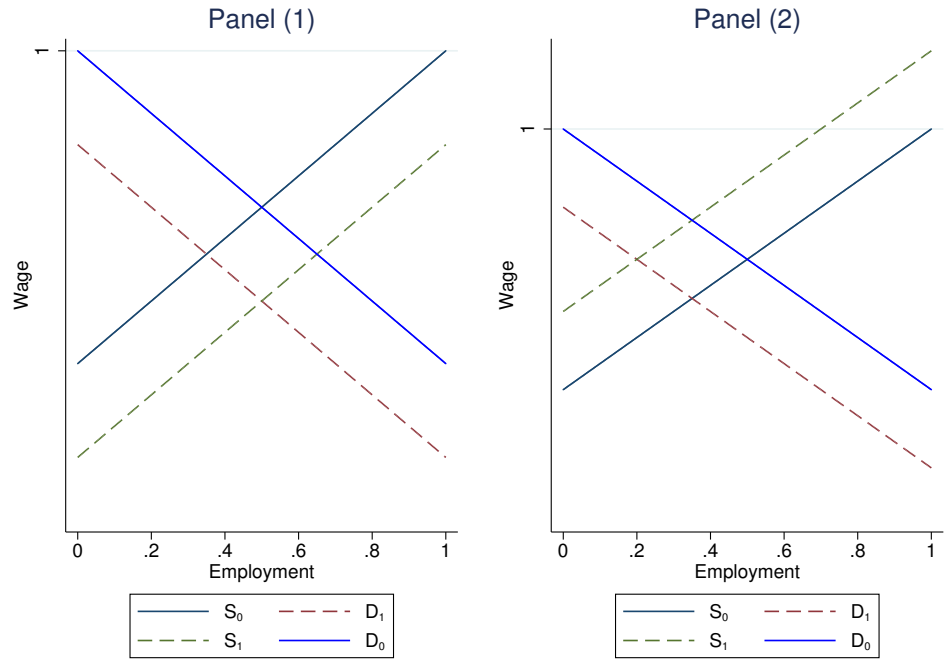


Figure 4: Cost of Employment by Age for Female Workers

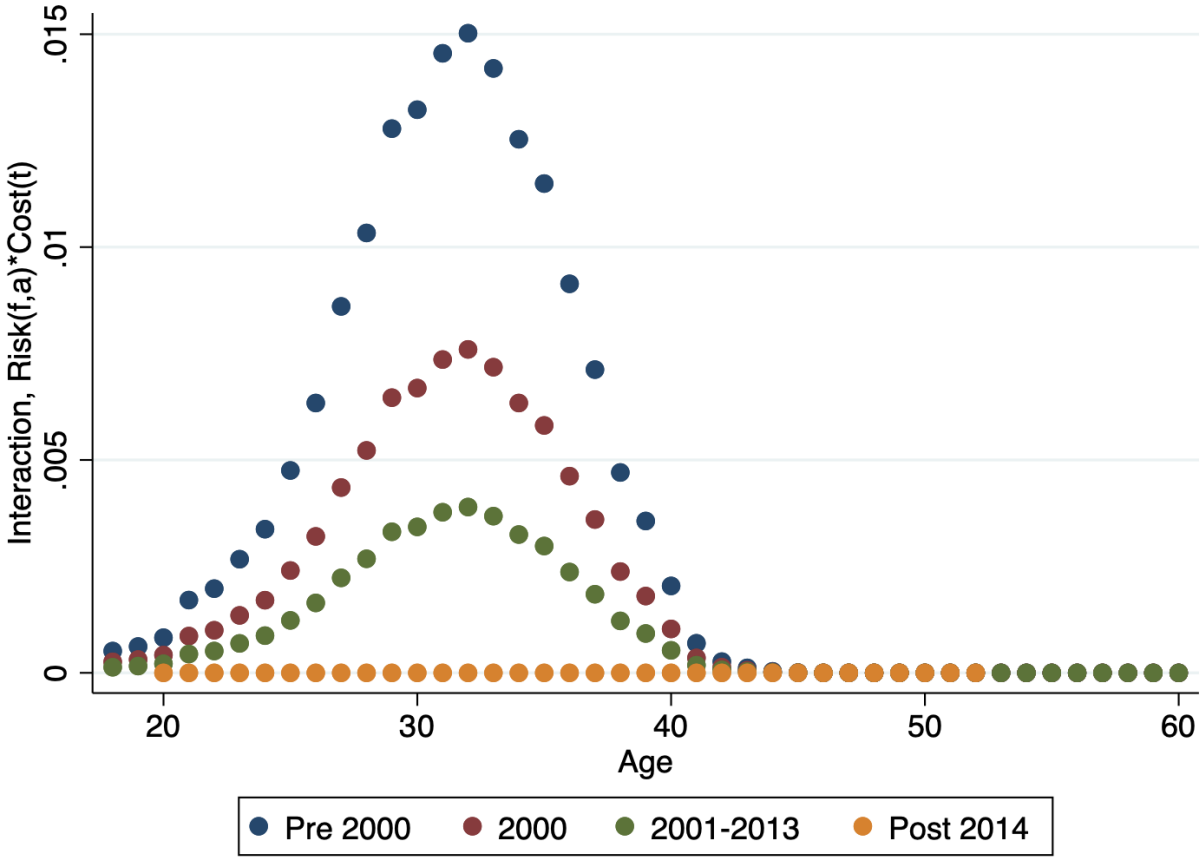


Table 1: Main Regression Results

(a) Hired on Permanent Employment Contract

	(1)	(2)	(3)	(4)	(5)	(6)
Risk	-0.522***	-0.529***	-0.670***	-0.640***		
	(0.0461)	(0.0458)	(0.0259)	(0.0470)		
Cost	0.0244***					
	(0.00144)					
Cost × Risk	-0.271***	-0.268***	-0.242***	-0.190***	-0.291***	-0.134***
	(0.0210)	(0.0210)	(0.0149)	(0.0225)	(0.0209)	(0.0318)
Prediction, Highest Risk	-.007	-.031	-.028	-.022	-.034	-.016
S.E.	.003	.002	.002	.003	.002	.004
Controls	X	X	X	X	X	X
Estab. FE	X	X	X	X	X	X
Year FE		X	X	X	X	X
AgeXYear FE			X			X
FemaleXYear FE				X		X
FemaleXAge FE					X	X

(b) Starting Pay (Log)

	(1)	(2)	(3)	(4)	(5)	(6)
Risk	-0.264***	-0.278***	-0.277***	-0.456***		
	(0.0636)	(0.0629)	(0.0372)	(0.0656)		
Cost	0.0477***					
	(0.00224)					
Cost × Risk	-0.404***	-0.400***	-0.453***	-0.275***	-0.450***	-0.284***
	(0.0291)	(0.0290)	(0.0192)	(0.0318)	(0.0287)	(0.0440)
Prediction, Highest Risk	.001	-.046	-.052	-.032	-.052	-.033
S.E.	.004	.003	.002	.004	.003	.005
Controls	X	X	X	X	X	X
Estab. FE	X	X	X	X	X	X
Year FE		X	X	X	X	X
AgeXYear FE			X			X
FemaleXYear FE				X		X
FemaleXAge FE					X	X

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2: Heterogeneity-Result by Firm Size

	(1) Perm-Small	(2) Perm-Large	(3) Pay-Small	(4) Pay-Large
Cost × Risk	-0.0577*	-0.282***	-0.194***	-0.499***
	(0.0319)	(0.0746)	(0.0460)	(0.116)
Prediction, Highest Risk	-.007	-.033	-.022	-.058
S.E.	.004	.009	.005	.013
Mean	.621	.711	7.131	7.387
Controls	X	X	X	X
Estab. FE	X	X	X	X
Year FE	X	X	X	X
AgeXYear FE	X	X	X	X
FemaleXYear FE	X	X	X	X
FemaleXAge FE	X	X	X	X

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3: Heterogeneity-Result by Female Employment Shares

	(1) Perm-Lo Fem	(2) Perm-Hi Fem	(3) Pay-Lo Fem	(4) Pay-Hi Fem
Cost × Risk	-0.242***	-0.103*	-0.330***	-0.466***
	(0.0482)	(0.0614)	(0.0622)	(0.0855)
Prediction, Highest Risk	-.028	-.012	-.038	-.054
S.E.	.006	.007	.007	.01
Mean	.784	.462	7.388	6.91
Controls	X	X	X	X
Estab. FE	X	X	X	X
Year FE	X	X	X	X
AgeXYear FE	X	X	X	X
FemaleXYear FE	X	X	X	X
FemaleXAge FE	X	X	X	X

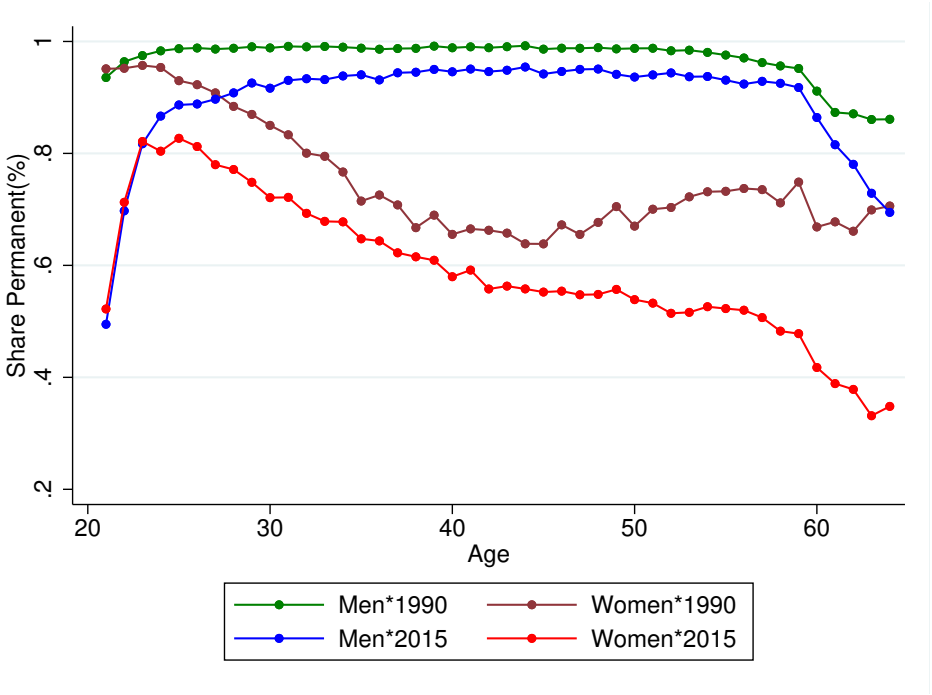
Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Heterogeneity-Different Sample Periods

	(1) perm	(2) pay	(3) perm	(4) pay
Risk	-0.786***	-0.670***	-0.917***	-0.693***
	(0.115)	(0.178)	(0.105)	(0.161)
Cost × Risk	-0.0824**	-0.147**	-0.0455	-0.123
	(0.0383)	(0.0603)	(0.128)	(0.187)
Prediction, Highest Risk	-.01	-.017	-.005	-.014
S.E.	.004	.007	.015	.022
Mean	.713	7.359	.573	7.038
Years	1998-2001	1998-2001	2013-2014	2013-2014
Controls	X	X	X	X
Estab. FE	X	X	X	X
Year FE	X	X	X	X
AgeXYear FE				
FemaleXYear FE	X	X	X	X
FemaleXAge FE				

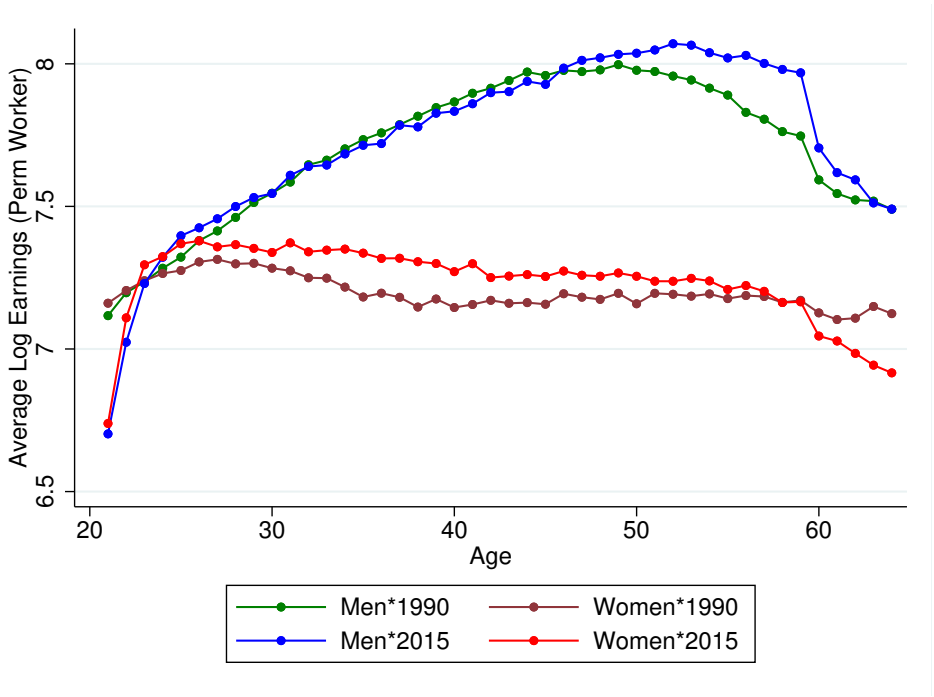
Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure A1: Permanent Share of Employment, by Age and Gender, 1990 and 2015



Note: Data are from the BSWS in 1990 and 2015. A worker not on permanent contract is defined as a temporary worker.

Figure A2: Average Earnings among Permanent Worker, by Age and Gender, 1990 and 2015



Note: Data are from the BSWS in 1990 and 2015. A worker not on permanent contract is defined as a temporary worker and excluded from the figure.