Expected Fertility, Labor Market Contracts, and the Gender Wage Gap

Job Market Paper

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Abstract

This paper examines how employers' expectations about women's future fertility enlarge the gender wage gap in markets with labor contracts—standard settings in many occupations that require long-horizon, complex tasks. In such environments, salaries are preset based on expected match productivity, not solely on contemporaneous output; if employers anticipate workers' future fertility interruptions, they post lower wages today. Exploiting China's relaxation of the One-Child Policy as a quasi-experiment, a difference-in-differences design reveals that women's wages declined by 15.3% immediately after the reform, despite no short-term increase in actual births. To interpret these facts, I build a search and matching model with a joint household decision problem, in which fertility is given but uncontractable fertility-driven efforts are chosen. I further incorporate learning-by-doing human capital accumulation that depends on effort, which amplifies and propagates the impact of effort changes over time. Quantitatively, gender differences in expected productivity—rooted in the unbalanced division of household labor—account for nearly all of the pre-reform wage gap and about 80% of the post-reform widening. The policy implication is stark: women-protective rules that merely keep women employed through legislative contract provisions may not reduce the gap; by strengthening employers' present-value pricing, they can be offset by ex-ante wage markdowns applied to all women.

Keywords: gender wage gap; fertility; search and match

JEL codes: J16; J31; J64

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1 Introduction

Despite convergence in educational attainment and rising female labor force participation, sizable gender wage gaps persist in many countries, with little evidence of convergence in recent decades (Olivetti and Petrongolo, 2016; Blau and Kahn, 2017; Kunze, 2017). One central mechanism for the remaining gap is the child penalty: Childbirth leads many women to exit the labor market, depressing cumulative earnings and delaying advancement into managerial roles (Adda, Dustmann and Stevens, 2017; Kleven, Landais and Søgaard, 2019; Goldin, 2021; Cortés and Pan, 2023; Kleven, Landais and Leite-Mariante, 2024). Beyond this direct effect on mothers, statistical discrimination can bring down the wages of women without children: anticipating higher replacement and adjustment costs, employers embed expected motherhood risks into lower starting offers for young women (Amano-Patiño, Baron and Xiao, 2020). This paper complements that view with a new mechanism: even when separation risk is low, employers would still discount women's wages because they expect lower *post–birth productivity*, not merely higher replacement costs.

The main goal of this paper is to study how employers' expectations about future fertility shape wages while admitting the labor market realities that in many positions, pays are largely predetermined at the time of hiring and only weakly tied to contemporaneous effort. Empirically, I use the relaxation of the One-Child Policy in China as a natural experiment, showing that firms are forward-looking in contracted labor markets: women's wages decline immediately in anticipation of more births, and the decline is larger where separation risk is lower—consistent with a present-value pricing channel rather than a replacement-cost channel. Theoretically, I develop a wage-posting model with long-term matches in which employers' beliefs about on-the-job human capital accumulation are disciplined by a household time-allocation block that links training efficiency to free time (effort), not recorded hours—allowing workers to be fully employed yet operate at different intensities. Estimating the model on Chinese data, I find that the unequal division of household labor alone accounts for nearly all of the residual gender wage gap and roughly 80% of the post-reform widening, underscoring the central role of expectation-driven pricing when salaries are predetermined. This mechanism also helps explain why "position-preserving" protections like maternity leave mandates can disappoint: rules that require firms to keep jobs for mothers lengthen contract horizons and reinforce present-value pricing, amplifying ex-ante wage markdowns and stalling promotions—thereby failing to close, and potentially widening, the gender wage gap even as employment is maintained.

The first contribution of this paper is to provide empirical evidence on how employers price fertility expectations in a labor market where contracted, long-term employment relationships are prevalent. I exploit the 2013 selective relaxation of China's One-Child Policy as a plausibly exogenous shock to firms' expectations about young women's fertility over their lifetime. This relaxation permitted a second birth without fines if either parent was an only child, and the magnitude of the relaxation is larger where enforcement of the One-Child Policy was stricter (i.e., higher fines for additional children). I therefore implement a difference-in-differences design comparing more-exposed and less-exposed regions before and after the reform. Importantly, realized births changed little over the short run after this policy shift, so observed wage responses primarily reflect employers' expectation updates rather than contemporaneous household labor-supply adjustments around childbirth. China's labor market is well-suited to studying contracted wage setting: among tertiary-educated workers, government and state-affiliated employment is widespread, offering tenure-like security and infrequent within-position wage changes. In this environment, post-reform wage dynamics are especially informative about how firms update and price workers' expected future productivity even though current productivity remains unchanged.

I use data from the China Family Panel Studies (CFPS), restricting the sample to employed workers aged 18–35 with at least a high school education, i.e., workers most likely to hold contracted, long-horizon formal jobs and face fertility decisions during the childbearing years. Empirical results show that the wages of young women fell by 15.3% after the policy shift. The decline does not materially differ by marital status or stated fertility intentions. Applying the same specification to comparable young men yields no significant effect on their wages. Using men as a same-region, same-period control in a triple-difference specification confirms a wage decline for women of 14.8%.

I interpret the sizable post-reform wage decline as evidence of a long-horizon, expected-productivity channel of statistical discrimination: employers mark down young women's wages in anticipation of a lower future effective-productivity path—driven by greater childcare responsibilities—even when employment is continuous. This contrasts with Amano-Patiño, Baron and Xiao (2020), where statistical discrimination arises from higher separation risk and replacement costs. Here, discrimination persists *without* separations, and—because wages price the present value of productivity—lower separation risk actually amplifies the markdown by lengthening the expected match horizon. Sectoral heterogeneity reinforces this mechanism: in government-related sectors, where separation risk is minimal, young women experience a wage decline more than twice the full-sample estimate, whereas women in non-government sectors show no clear change.

To formalize this channel, I develop a simple unitary household model with discrete working hours (non-participation vs. a fixed full-time schedule) and continuous effort. Salaries are irrelevant to effort level as long as the worker is fully employed, but effort governs both on-the-job human capital accumulation and the effective utilization of accumulated human capital[1]. In this environment, additional births raise childcare needs that fall disproportionately on women; holding full-time on-record working hours fixed, greater childcare reduces free time, lowers work effort, and slows human capital accumulation, thereby diminishing future productivity, especially for women. Throughout, I treat fertility as exogenously decided by the social norm or policy, but efforts need to be chosen by household members to sustain family functioning and maximize household income.

I then embed this gender-specific effort choice in a wage–posting model with on-the-job search. This model is based on Burdett and Mortensen (1998), but is extended to allow productivity to grow at the effort-determining rate. Firms post a fixed salary at hire and do not renegotiate within the position; workers stochastically receive outside offers and can move to higher-pay firms, capturing career advancement. Because future effort is not contractable and salaries are rigid within matches, at the posting stage, firms form gender-specific beliefs about effort, which should be the same as the household's optimal choice given the number of children and thereby the future productivity level. A firm's expected profit equals the

[1]This assumption mirrors Albanesi and Olivetti (2009), who posit that the utility cost of market effort rises with home hours, making household time demands a direct wedge on effective effort. Additionally, time-use evidence shows that childcare and home production compress sleep and discretionary time, the very margins that sustain focus and off-the-clock learning and working, and "greedy jobs" reward long hours and continuous availability while penalizing flexibility (Goldin, 2014).

present value of the effective productivity path minus the posted salary, integrated over the tenure distribution implied by offer-arrival and separation hazards. In this fixed-salary / rising-productivity environment, downward revisions to women's expected growth following the policy translate immediately into lower posted wages.

Introducing within-match productivity growth yields several implications. First, human capital accumulation becomes a key driver of the upper tail and of average wage levels: by widening the support of equilibrium wages, it allows the model to match high wages that would otherwise be dismissed as measurement error (e.g., Van den Berg and Ridder (1998)). Second, higher posted wages lower poaching hazards and lengthen tenure; when productivity rises with tenure, longer matches translate into higher effective productivity. In this sense, higher posted wages employ and retain more productive workers, unlike in the canonical model, where wage levels are independent of individual productivity. Third, growth with tenure amplifies standard comparative statics: differences in separations or offer-arrival rates map into larger wage gaps when productivity accumulates on the job. This amplification carries a policy warning: rules that lengthen non-adjustment horizons—such as position-preserving job protection—raise the weight employers place on expected future productivity and can intensify ex-ante wage markdowns.

To quantify the role of expected productivity in contracted labor market, I estimate the model by matching its implied wage distributions to the observed distributions before and after the 2013 relaxation of China's One-Child Policy (CFPS 2012 vs. 2014; sample as in the reduced-form analysis). The procedure is standard: guess a parameter vector, solve the firm's wage-posting problem to obtain the equilibrium offer distribution that is also consistent with workers' acceptance decisions, compare the resulting wage CDF with the empirical CDF, and iterate until the distance is minimized. Apart from the wage distributions, the only external inputs are time-use data and the fertility composition.

In the 2012 baseline, virtually all workers were either childless or had one child. The estimates imply that women's effort falls from roughly 90% when childless to about 74% with one child, while men's effort declines only modestly from about 94% to 87%. Correspondingly, annual human capital growth rate is roughly 0.21 for childless workers, but falls to about 0.18 for mothers versus 0.20 for fathers during prime working years. Crucially, this unequal division of family work by itself explains essentially the entire gender wage gap in 2012. Although the estimated market parameters also differ by gender—men face higher job-offer arrival rates while separation rates are similar—these frictions add little once the household time-allocation gap is taken into account.

Re-estimating the model on the post-reform wage distribution without imposing a mechanism ex ante, the results indicate a sizable downward revision in firms' expectations of women's future productivity. Gender gaps in offer-arrival rates also widen, consistent with firms' greater reluctance to extend offers or promotions to women. Under a simple calibration—assuming a second child adds 25% to mothers' family-work time and leaves fathers' time unchanged (consistent with the null wage response for men in the empirical results)—the estimates imply that employers expect roughly 56% of one-child parents to have a second child. Feeding this time-use shift and fertility composition into the 2012 labor-market environment while holding other parameters fixed, the model attributes nearly 80% of the simulated 15.5% widening of the gender wage gap to employers' reductions in women's expected future productivity associated with a potential second birth.

Taken together, the reduced-form and structural evidence points to a common policy lesson: position-preserving job protection is insufficient—and can even worsen—gender gaps. This class of policies includes maternity leave mandates and special protections for mothers with young children. By lengthening the effective contract horizon, such rules raise the weight employers place on expected future productivity. When women are perceived to accumulate human capital more slowly, a longer horizon translates into larger up-front wage markdowns for all women. Moreover, retaining matches that would otherwise separate lowers the average expected productivity of the protected pool, prompting broader markdowns at hiring. Policies that reduce childcare burdens and equalize effort—rather than merely prolong tenure—are therefore more likely to address the expectations-based mechanism behind the gap. Put differently, without tackling the unequal division of household labor, maternity leave–style protections alone are unlikely to close the remaining gender wage gap.

Although the empirical setting is China and the focus is the gender wage gap, the mechanism generalizes to labor markets where long-horizon contracts are prevalent. These include government and state-affiliated employment and many high-skill occupations, where long horizons arise endogenously from costly matching, firm-specific training, and expensive replacement. In such environments, wages are set against expected productivity over the match, so employer beliefs about future effort—whether tied to anticipated childcare, immigration constraints, or health risks—carry greater weight in wage setting. In short, while this paper studies employers' pricing of fertility expectations, the logic applies broadly to contract-driven pay: when salaries are predetermined, any concern about future productivity is amplified in posted wages.

After a brief discussion of the literature, the remainder of this paper is organized as follows. Section 2 presents the empirical setting, background on China's One-Child Policy, the identification strategy, data, and main results, together with placebo and mechanism tests. Section 3 develops the theoretical framework that links household time allocation, effort, and learning-by-doing to wage posting with on-the-job search, and characterizes equilibrium. Section 4 maps the model to the data: it details the numerical solution, reports the estimation for 2012 and 2014, and conducts counterfactual decompositions. Section 5 concludes and discusses policy implications.

Literature Review This paper speaks to several strands of research. First, a large literature documents persistent gender wage gaps and traces a central role to children. Surveys show that gaps have narrowed only slowly despite convergence in education and participation (Blau and Kahn, 2017; Olivetti and Petrongolo, 2016; Kunze, 2017), with childbirth generating sizable, long-lived penalties through reduced hours and interrupted human capital accumulation (Adda, Dustmann and Stevens, 2017; Kleven, Landais and Søgaard, 2019; Cortés and Pan, 2023; Goldin, 2021; Kleven, Landais and Leite-Mariante, 2024). Classic models of household specialization and effort (Becker, 1985; Albanesi and Olivetti, 2009; Albanesi, Olivetti and Petrongolo, 2023) and the "greedy jobs" perspective (Goldin, 2014) rationalize why home responsibilities compress focus, availability, and off-the-clock learning that sustain productivity growth at work. Correll, Benard and Paik (2007) applies a laboratory experiment to causally identify the penalization of mothers in hiring and wages, reinforcing the idea that employers perceive women as less productive due to their family responsibilities. However, while most of this literature focuses on post-birth employment outcomes, this paper emphasizes that gender-based disadvantages emerge even before childbirth and for all women regardless of their fertility intentions, particularly in less mobile labor markets where employers preemptively adjust wages based on expected, rather than actual, fertility behavior.

A second strand links gender gaps to statistical discrimination in frictional labor markets. Structural and reduced-form studies emphasize higher expected separations and replacement costs for mothers (Bowlus and Grogan, 2009; Bartolucci, 2013; Amano-Patiño, Baron and Xiao, 2020). The contribution of this paper is to shift the mechanism from expected separations to expected productivity while separation is negligible: when pay is set against the present value of match productivity, beliefs about future effort and learning-by-doing receive more weight, so markdowns can arise even without separations, and, in fact, are larger when separations are rarer and matches are longer.

Finally, the paper contributes to work on China's fertility policy and women's labor outcomes. The One-Child Policy (OCP) and its local enforcement have been studied extensively (Gu et al., 2007; Scharping, 2013; Ebenstein, 2010; Huang, Lei and Zhao, 2016), and recent papers exploit relaxations to study realized fertility and labor effects, including field-experimental evidence on employer discrimination (He, Li and Han, 2023), city-level impacts on gaps (Agarwal et al., 2024), and post-birth outcomes (Li, 2022). In contrast, I use nationally representative survey data to capture aggregate effects and leverage the 2013 selective relaxation as a shock to expected fertility—with little short-run change in births—to isolate forward-looking wage setting under contracts. Rather than simply documenting a wage decline, I develop a model in which effort drives human capital growth and estimate it before and after the reform, quantifying how the unequal division of household labor is translated into expectation-driven pricing and how much of the post-reform widening it can explain.

2 Empirical Evidence

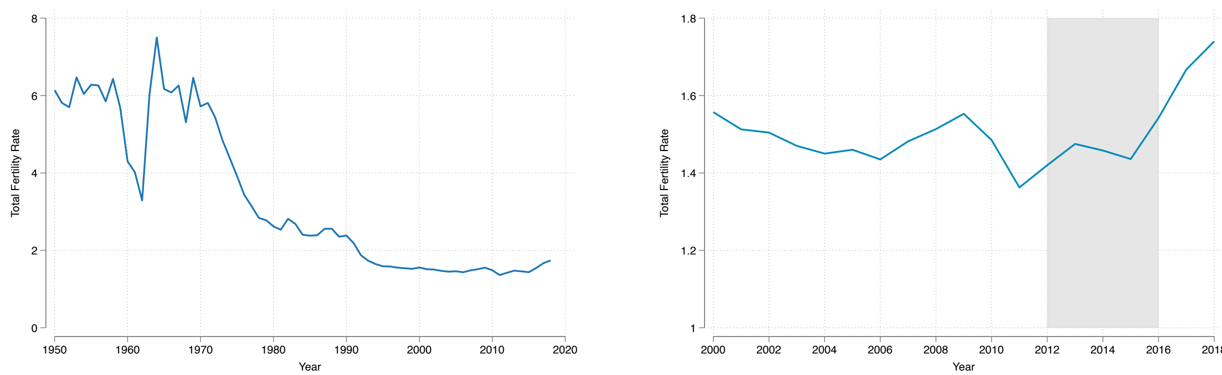
In this section, I exploit the relaxation of China's One-Child Policy (OCP) as a natural experiment to study the impact on young women's wages. The design uses three sources of variation: (i) a pre/post policy shift, (ii) cross-provincial differences in pre-relaxation enforcement intensity measured by *FinesRate*, and (iii) gender, with men serving as a natural control group in triple-differences (DDD) specifications. I estimate individual-level DiD and DDD models with province and year fixed effects (and their interactions in DDD).

The main result is a 15.3% decline in women's wages after the relaxation in provinces more exposed to the policy shift. Importantly, neither the aggregate trend in births nor province–year DiD regressions indicate a contemporaneous rise in birth rates following the first relaxation. Taken together, the evidence points to an *employer-side* response to higher *expected* lifetime fertility—rather than a *household-side* labor-supply adjustment—as the proximate driver of the wage decline.

A remaining puzzle is the magnitude of the effect in a low-mobility labor market where separations are uncommon, especially in government-related employment that offers near-lifetime job security. In such long-term relationships, pay reflects the present value of match productivity and is only weakly tied to contemporaneous effort; firing is also costly or institutionally constrained. Consistent with this long-term pricing channel, the wage decline is concentrated among young women and is *more pronounced* in government-related sectors, where employment relationships are stickier. These patterns align with firms discounting the future productivity of young women *ex ante* in response to higher expected fertility, even absent immediate changes in births.

2.1 Background: One-Child Policy in China and its Relaxation

China's family–planning policy began in 1971. Prior to 1979 it was largely voluntary. In 1979 the government enacted a stringent population–control policy—the One-Child Policy (OCP). Urban couples were generally limited to one child, while many rural couples could have a second child if the first was a girl. Authorities implemented a system of financial penalties and administrative sanctions to deter excess fertility, including wage deductions for salaried workers, reduced land allocations in rural areas, denial of certain public services, and, most commonly, fines for unauthorized births. The policy substantially lowered fertility: according to World Bank data, China's total fertility rate declined from 6.08 in 1970 to 2.51 by 1990 (Figure 1).



(a) 1950–2018 (b) 2000–2018

Figure 1: Trends in the total fertility rate in China. The shaded region in (b) denotes the sample period. *Source:* United Nations, Department of Economic and Social Affairs, Population Division (2019).

Concerns about persistent low fertility and population aging later prompted gradual relaxations of the OCP. The first major step—the "selective two-child" policy—was announced in 2013, allowing a second birth if either parent was an only child; provinces implemented the change in 2014–2015. Despite expectations of a baby boom, birth rates showed little response. In November 2015 the central government adopted a universal two-child policy, effective January 2016, permitting two children for all couples.

OCP implementation strength measurement Although the One-Child Policy was enforced nationwide, its implementation intensities varied significantly across provinces based on local preferences and social norms related to childbearing. The primary penalty for policy violation was a fine, which was substantial and varied from province to province. For example, Zhejiang, Fujian, and Guangdong—three consecutive and economically developed provinces along China's southeastern coast—imposed different fine rates for excess births. In Zhejiang, under the Regulations on Family Planning of Zhejiang Province enacted in 1990 and was still valid in 2000s, couples who exceeded birth quotas were required to pay 20%

to 50% of their household income annually for five years. In Fujian, around the 2000s, fines ranged from two to three times the household's total income from the previous year. In Guangdong, the 2002 regulations specified fines between three and six times the household's annual disposable income. The variation in punishment across provinces is a proxy for measuring the intensity of implementing the One-Child Policy, helping to identify the impacts of relaxing birth restrictions. In addition to economic penalties, urban residents working in the government or state-owned enterprises would lose their jobs and social welfare benefits for exceeding birth quotas. Moreover, coercive measures such as forced abortions were implemented in some regions.

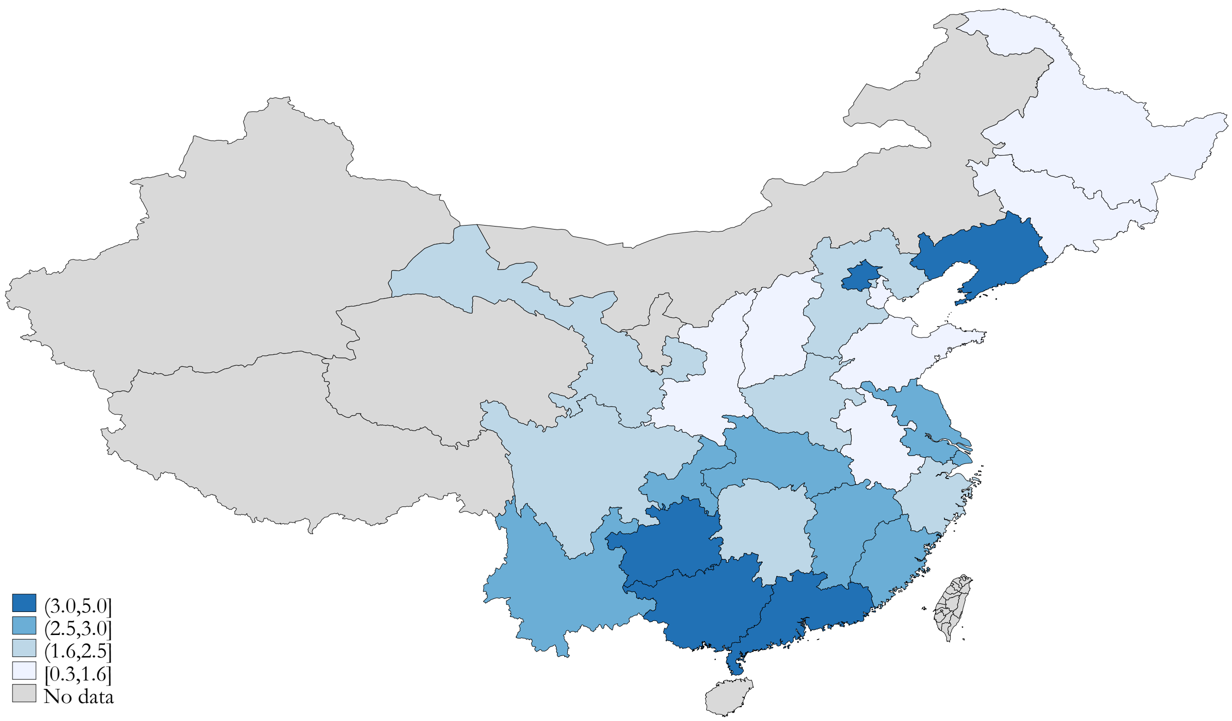


Figure 2: Harmonized OCP fine rates by province, 2000 (multiples of average annual household income). Source: Ebenstein (2010), based on Scharping (2013) and CHNS.

To obtain a unified measure of enforcement intensity, I follow the literature in using province-level "fine rates" expressed as multiples of average annual household income. Scharping (2013) compiles the official provincial schedules for 1979–2000; building on these, Ebenstein (2010) combine the schedules with China Health and Nutrition Survey data to construct annual province-specific fine rates that are widely used as an OCP enforcement proxy (see, e.g., Ebenstein (2010); Huang, Lei and Zhao (2016)). I use the 2000 cross-section of this measure—denoted *FinesRate*—as the baseline indicator of pre-relaxation enforcement intensity. Figure 2 displays the provincial distribution.

2.2 Empirical Strategy

The relaxation of the One-Child Policy provides me with a unique opportunity to understand how the policy relaxation affects young women's wages. I employ a difference-in-differences strategy to identify the policy impacts. *FinesRate*, the province-level monetary penalties to excess births before the relaxation is used to measure the change in exposure and thus help define the treatment and control group.

This empirical strategy assumes that OCP enforcement levels are unrelated to economic conditions that could influence wages. To assess plausibility, I regress *FinesRate* on provincial GDP, urban unemployment, average urban wages, and the urban employment count, using data from the National Bureau of Statistics of China.[2] Because GDP and unemployment are available in 2000, while average wages and urban employment are available from 2008 onward, I run cross-sections for 2000 (GDP, unemployment) and for 2013 (all four variables), the year when the main policy shift occurred. Results in Table 1 show small and insignificant coefficients; joint F-tests fail to reject that the economic covariates explain *FinesRate* (p-values 0.683 and 0.432). I therefore find no evidence that enforcement intensity is systematically related to provincial economic conditions, supporting the exclusion restriction underlying the design.

[2] I focus on the urban sector because the wage outcomes analyzed below are urban.

Table 1: Correlation between Fine Rates and Provincial Economic Conditions

	(1)	(2)
	Year = 2000	Year = 2013
	Dependent variable: *fine rates*	
GDP	0.000	-0.000
	(0.000)	(0.000)
Unemployment rate	-0.349	-0.585
	(0.430)	(0.509)
Number of employed		0.002
		(0.004)
Average wage		0.000
		(0.000)
Constant	3.477**	3.647
	(1.429)	(2.635)
Observations	25	25
R^2	0.034	0.166
F-stat	0.387	0.998
P-value	0.683	0.432

Standard errors in parentheses.

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

I follow a standard difference-in-differences strategy, and the estimation equation is

$$\log \text{wage}_{it} = \beta_0 + \beta_1 \text{yschool}_{it} + \beta_2 \text{age}_{it} + \beta_3 \text{age}_{it}^2 + \beta_4 \text{Treat}_{pt} + \mu_p + \gamma_t + \varepsilon_{it} \tag{1}$$

where i refers to individual worker, t calendar time, and p provinces. Using the logarithm level of nominal yearly salary income, "$\log \text{wage}_{it}$", as the dependent variable, I include years of schooling, "yschool", "age", and "squared age" as the independent variables based on the Mincer equation. Province fixed effect (μ_p) and year fixed effect (γ_t) are also included in this specification. The coefficient of interest is β_4 in front of the binary indicator Treat_{pt}, which is constructed by the time indicating variable Time_t and treatment group indicating variable FinesHigh_p:

$$\text{Treat}_{pt} = \text{FinesHigh}_p \times \text{Time}_t$$

where

$$\text{Time}_t = \begin{cases} 1, & \text{if } t > 2013 \\ 0, & \text{if } t \leq 2013 \end{cases}, \text{ and } \quad \text{FinesHigh}_p = \begin{cases} 1, & \text{if FinesRate}_p > 2.5 \\ 0, & \text{if FinesRate}_p \leq 2.5 \end{cases}$$

where 2.5 is the median of fines rates for all provinces in 2000 from Ebenstein (2010), in terms of *years of household income*. In other words, I assign provinces that had a higher fine rate for excess births in 2000 to the treatment group and others to the control group. $\text{Treat}_{pt} = 1$ implies workers in province p are under treated in year t, in comparison with the same-gender workers that are either never treated or not yet treated. As the fines rates are continuously measured, I create a new variable to capture treatment intensity, denoted as $\text{Con\_Treat}_{pt} = \text{FinesRate}_p \times \text{Time}_t$, and substitute for the binary Treat_{pt} variable in regression model described by Equation 1.

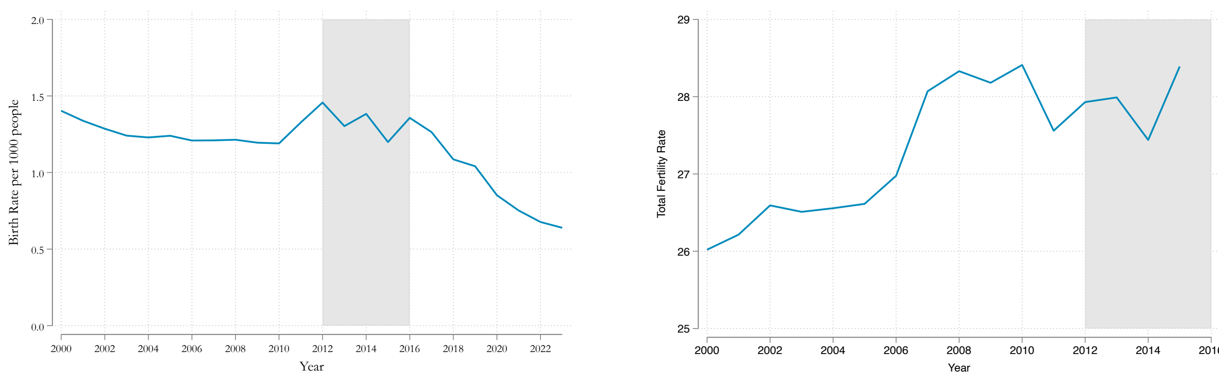
I applied Equation 1 to female workers and male workers to assess differential post-policy effects by gender. As a complementary approach, I also use men's wages as the natural control group for possible wage-determining factors that are systematically different across the treated and untreated province-year combinations, and following Jayachandran and Lleras-Muney (2009) the estimating model for Triple DiDs is

$$\begin{aligned} \log \text{wage}_{it} &= \beta_0 + \beta_1 \text{yschool}_{it} + \beta_2 \text{age}_{it} + \beta_3 \text{age}_{it}^2 \\ &\quad + \beta_4 \text{Treat}_{pt} \times \text{female}_i + \mu_p \times \gamma_t + \gamma_t \times \text{female}_i + \mu_p \times \text{female}_i + \varepsilon_{it} \end{aligned} \tag{2}$$

The model incorporates a full set of double interactions of fixed effects, including the province-year, gender-year, and province-gender fixed effects. This enables us to interpret the coefficient of interest, β_4, as the impact of an increase in expected fertility rates on female wages, accounting for all variations in province, year, and gender. I use the continuously measured treatment intensity Con_Treat as a second treatment specification in the estimation as well.

2.3 Data and Sample Selection

The primary dataset utilized in this paper is the China Family Panel Survey (CFPS), a representative family survey that started in 2010. Subsequent surveys were conducted biennially from 2012 onwards. I have included data from three survey rounds, spanning from 2012 to 2016. The exclusion of the 2010 survey is based on a change in employment criteria in 2012, leading to a low and incomparable fraction of employed individuals in 2010[3]. Furthermore, the One-Child Policy was relaxed to a universal Second Child Policy in 2016, and people realized that actual birth rates did not experience significant growth after the policy relaxation, as shown in Figure 3a. All add complexity to the interpretation of results, making the inclusion of samples post-2016 potentially challenging. The CFPS survey covers 25 provinces, excluding minority-concentrated areas such as Xinjiang, Tibet, Qinghai, Inner Mongolia, Ningxia, and Hainan. In subsequent surveys, a small number of households report living in those provinces because of migration. I exclude them due to the small size and the fact that the One-Child Policy did not initially apply to minority populations.



(a) Birth rate in China

(b) Mean age at childbearing in China

Figure 3: Fertility-related statistics in China since 2000. The shaded region denotes the sample period. *Source:* United Nations, Department of Economic and Social Affairs, Population Division (2019).

Within each survey, I refine the sample to include individuals aged 18 to 35 years, residing in urban areas, and possessing 12 or more years of schooling. This age range aligns with the prime reproductive years for females. The educational attainment requirement is based on

[3]The fraction of employed people in 2010 is only 62.96 percent, whereas in the later three rounds of the survey, this number increases to 83.29, 84.82, and 85.06, respectively

the premise that the expectation channel can only function effectively when employees can enter into long-term contracts with employers. In situations where employers can hire and terminate employment at their discretion, there would not adjust their job offer because of future concerns. Consequently, young individuals without a high-school diploma (with fewer than 12 years of schooling) are excluded from the main analysis.

Table 2 presents the summary statistics for all individuals meeting the sample selection criteria and having a valid wage record in the year they were surveyed. The sample size remains relatively consistent across the three survey rounds. The average age is approximately 28. Given that the application of the One-Child Policy varies based on people's *hukou* and ethnicity status, I provide a summary of the number of individuals with urban *hukou* status and those belonging to minority groups. It appears that, across the three rounds, the majority of individuals in the sample possess urban hukou status, less than 5% belong to minority groups, and these proportions remain relatively stable. The average years of schooling and the logarithm level of average nominal wages exhibit a slightly upward time trend. However, the gender wage ratio, defined as the wages of females divided by the wages of males, declines from 78.67% to 71.89%, indicating an widening gender wage gap.

Table 2: Summary Statistics by Year and Gender

	2012		**2014**		**2016**	
	Female	Male	Female	Male	Female	Male
n	466	486	540	563	506	556
Age	28.34	28.74	27.81	28.45	28.68	29.15
Urban *hukou* (%)	85.00	84.77	71.48	74.42	71.34	73.92
Minority (%)	2.80	2.06	4.63	3.20	4.35	3.24
Years of schooling	14.37	14.26	14.59	14.55	14.72	14.68
	(1.748)	(1.806)	(1.829)	(1.877)	(1.763)	(1.947)
ln(wage)	9.99	10.23	10.12	10.43	10.20	10.53
	(0.916)	(0.921)	(0.915)	(0.980)	(1.270)	(1.114)
Total n	952		1,103		1,062	
Gender wage ratio (F/M)	0.7867		0.7334		0.7189	

Notes: Entries are means; standard deviations in parentheses. Urban *hukou* and minority are in percent.

2.4 Empirical Results

The results for the two main regression Equation 1 and Equation 2 are presented in Table 3, and all the standard errors are clustered at the province level, which is the level where the variation happens.

Table 3: Effect of OCP Relaxation on Female/Male Workers' Wages

	(1)	(2)	(3)	(4)	(5)	(6)
	Female	Male	Female	Male	Triple-DiD	
	Dependent variable: *ln_wage*					
Years of schooling	0.089***	0.067***	0.089***	0.067***	0.078***	0.078***
	(0.016)	(0.022)	(0.016)	(0.022)	(0.009)	(0.010)
Age	0.432***	0.501***	0.431***	0.501***	0.455***	0.455***
	(0.101)	(0.066)	(0.101)	(0.066)	(0.072)	(0.072)
Age^2	-0.007***	-0.008***	-0.007***	-0.008***	-0.007***	-0.007***
	(0.002)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
Treat	-0.153*	-0.031				
	(0.083)	(0.084)				
Continuous Treat			-0.059*	0.006		
			(0.031)	(0.025)		
Treat × female					-0.148*	
					(0.078)	
Cont. Treat × female						-0.075**
						(0.030)
Observations	1,467	1,519	1,467	1,519	2,986	2,986
R^2	0.200	0.216	0.200	0.215	0.239	0.239
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Clustered SE: Province	Y	Y	Y	Y	Y	Y

Notes: Standard errors in parentheses. "Treat" is a binary indicator for exposure to the OCP relaxation; "Continuous Treat" is a continuous exposure measure. Triple-DiD columns report the interaction with the female indicator.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

The results for Equation 1 with binary indicator variable, Treat_{pt}, are presented in the first two columns. The estimated coefficient for Treat_{pt} is negative for female workers, indicating a statistically significant decrease of 15.3 percent in wages for young women in provinces experiencing more exposure to the policy change, as compared to unaffected females. However, for male workers, the estimated coefficient is not statistically significant from 0, suggesting that the change in policy does not impact their wages.

Columns (3) and (4) exhibit the results using continuously measured treatment variable Con_Treat_{pt}. Notably, the alteration in treatment measurement does not alter the primary finding: when assessed based on fines rates, female workers in provinces that previously imposed fines equivalent to 1 more year of household income for excess births experience a 5.9 percent decrease in wages after OCP relaxation, while if we time this number with the median 2.5, the results is 14.75 percent, quite comparable to what we got from binary treatment indicator. Conversely, the strictness of OCP implementation in the past did not have a significant effect on the wages of male workers.

In addition to the key coefficients of variable Treat_{pt} and Con_Treat_{pt}, other estimated results fall within the range found in the literature. Specifically, an additional year of schooling is associated with an 8.9 percent increase in wages for women and a 6.7 percent increase for men. Furthermore, wages exhibit a positive correlation with age (experience), although with a significantly diminishing return.

Triple difference-in-differences results specified in Equation 2 are presented in the last two columns. The effects of increased expected fertility rates on young females' wages is negative and significant. Under a counterfactual scenario without policy relaxation, the wage level for the same female worker would have risen by 14.8 percent. Taking the average real wage level before the policy shift for females in the treatment group as a reference point, this implies an annual wage income loss of approximately $34,650.87 \times 0.148 = 5,128.33$ RMB, which is approximately 833.87 USD using the 2013 exchange rate. If we interpret the results by the continuously-measured treatment intensity, it implies that if the province one female lives set its fine up 1 more years of household income for the excess birth, the female's annual wage income would decrease by 7.5 percent, which equal to 2598.82 RMB and 422.57 USD.

Response in Extensive Employment Margin When it comes to wage change, the naturally-related question is whether the employment rates change simultaneously. If the possibilities of finding jobs for young females increase after the policy relaxation, the wage decrease could happen because more low-skilled people are in the labor market pool and the average wage level is lower. To test this hypothesis, I run the same regression equations as in Section 2.2 but now use *"have a waged job or not"* as the dependent variable and, therefore, apply a Probit model.

The results are shown in Table 4. None of the key coefficients are significant, implying that the employment rates for either males or females do not respond to the policy relaxation. The selection criteria for the labor market do not change for young females, but the payments for working decrease.

Table 4: Effect of OCP Relaxation on Female Workers' Employment Status

	Panel A: Binary			**Panel B: Continuous**		
	Female	Male	Triple-DiD	Female	Male	Triple-DiD
Years of schooling	0.173***	0.122***	0.150***	0.174***	0.122***	0.150***
	(0.018)	(0.017)	(0.012)	(0.018)	(0.017)	(0.012)
Age	-0.032	0.152*	0.058	-0.034	0.157*	0.057
	(0.057)	(0.086)	(0.060)	(0.058)	(0.086)	(0.060)
Age^2	0.001	-0.002	-0.001	0.001	-0.002	-0.001
	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)	(0.001)
Treat	0.115	-0.030				
	(0.125)	(0.140)				
Treat × female			0.147			
			(0.220)			
Treat (cont.)				0.026	0.063	
				(0.049)	(0.042)	
Treat (cont.) × female						-0.035
						(0.075)
Observations	2,241	2,068	4,296	2,241	2,068	4,296
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Clustered SE: Province	Y	Y	Y	Y	Y	Y

Notes: Standard errors in parentheses. "Treat" is an indicator for exposure to the OCP relaxation; "Treat (cont.)" is a continuous exposure measure. "Triple-DiD" includes the interaction with the female indicator. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

2.5 Results Interpretation: Employers' Response to Expectation Changes

2.5.1 More Babies as Expected?

One potential explanation for the widening gender wage gap in treated provinces after the OCP relaxation is a differential increase in births ("motherhood penalty" channel). Prior work (e.g., Correll, Benard and Paik (2007)) documents that mothers may earn less due to reduced hours, acceptance of lower pay for schedule flexibility, or job changes that shorten commutes but lower wages. To assess this channel, I use province–year birth rates from the National Bureau of Statistics of China (live births per 1,000 population), covering 2011–2019 for the same 25 provinces as in the main analysis. Figure 3 plots the aggregate trend: after 2000, the birth rate peaks in 2012, fluctuates for several years, and then declines sharply after 2016. The sample window in this paper (2012–2016) straddles the 2013 relaxation, with the peak occurring in the pre-shock year. In other words, relative to 2012, subsequent years do not exhibit a higher number of births, which is inconsistent with a contemporaneous motherhood-penalty mechanism driving the observed wage gap widening.

I then estimate province–year DiD models analogous to Section 2.2. Let Time_t indicate the post-2013 period, FinesHigh_p equal one for provinces with above-median 2000 fine rates, and FinesRate_p denote the continuous fine measure. The binary treatment is $\text{Treat}_{p,t} = \text{Time}_t \times \text{FinesHigh}_p$; the continuous-intensity treatment is $\text{ConTreat}_{p,t} = \text{Time}_t \times \text{FinesRate}_p$. The two-way fixed effect specification is

$$\text{birth\_rate}_{p,t} = \beta_0 + \beta_1 \, \text{Treat}_{p,t} + \mu_p + \gamma_t + \varepsilon_{p,t},$$

with an analogous regression using $\text{ConTreat}_{p,t}$ in place of $\text{Treat}_{p,t}$. Columns (1)–(2) of Table 5 report pooled DiD without fixed effects; Columns (3)–(4) include province and year fixed effects. Standard errors are robust.

Across all specifications, the interaction terms are small and statistically indistinguishable from zero, indicating no detectable differential birth-rate response between treated and control provinces following the first OCP relaxation.

Table 5: Impact of OCP Relaxation on Birth Rates

	(1)	(2)	(3)	(4)
	Dependent variable: *birth_rate* (per 1,000)			
Time	0.010	0.016		
	(0.051)	(0.094)		
FinesHigh	0.027			
	(0.056)			
Time × FinesHigh	0.030		0.030	
	(0.072)		(0.021)	
FinesRate		0.022		
		(0.027)		
Time × FinesRate		0.003		0.003
		(0.034)		(0.008)
Observations	216	216	216	216
R^2	0.011	0.018	0.928	0.928
Year fixed effects			Y	Y
Province fixed effects			Y	Y
Robust SE	Y	Y	Y	Y

Notes: Province–year panel for 25 provinces, 2011–2019. Standard errors in parentheses. *Time*= 1 for post-2013 years; *FinesHigh*= 1 for provinces with above-median 2000 OCP fine rates; *FinesRate* is the continuous fine measure. Standard errors are in parentheses.
$^{*}\, p < 0.10$, $^{**}\, p < 0.05$, $^{***}\, p < 0.01$.

2.5.2 Employer–side Interpretation and an Older-Worker Placebo

Since neither the aggregate series nor the province–year DiD estimates show a birth-rate response, contemporaneous fertility changes are unlikely to drive the observed wage effects. Instead, the patterns are consistent with employer-side pricing based on long-horizon expectations: following the OCP relaxation, firms anticipate higher lifetime fertility for young women and discount wages ex ante even in the absence of immediate births.

To probe this "long-term expectation" channel, I re-estimate (1) and (2) on workers aged 35–54 who remain employed. These workers have largely passed prime childbearing ages; under the hypothesis above, their wages should not respond to the relaxation. Table 6 shows that the coefficients on *Treat* and *Treat (cont.)* in the female and male samples, as well as the triple-difference terms *Treat* × *Female* and *Treat (cont.)* × *Female*, are small and not statistically different from zero. This placebo supports the interpretation that the post-relaxation wage decline is specific to younger women through the expectation channel, rather than a contemporaneous shock affecting all women.

Table 6: Effect of OCP Relaxation on Older Workers' Wages (Ages 35–54)

	Panel A: Binary treatment			**Panel B: Continuous treatment**		
	Female	Male	Triple-DiD	Female	Male	Triple-DiD
Years of schooling	0.173***	0.122***	0.150***	0.174***	0.122***	0.150***
	(0.018)	(0.017)	(0.012)	(0.018)	(0.017)	(0.012)
Age	-0.032	0.152*	0.058	-0.034	0.157*	0.057
	(0.057)	(0.086)	(0.060)	(0.058)	(0.086)	(0.060)
Age^2	0.001	-0.002	-0.001	0.001	-0.002	-0.001
	(0.001)	(0.002)	(0.001)	(0.001)	(0.002)	(0.001)
Treat	0.115	-0.030				
	(0.125)	(0.140)				
Treat × Female			0.147			
			(0.220)			
Treat (cont.)				0.026	0.063	
				(0.049)	(0.042)	
Treat (cont.) × Female						-0.035
						(0.075)
Observations	2,241	2,068	4,296	2,241	2,068	4,296
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Clustered SE: Province	Y	Y	Y	Y	Y	Y

Notes: Province–year panel; standard errors in parentheses. "Treat" is the post–OCP relaxation indicator interacted with above-median pre-policy fine rates; "Treat (cont.)" uses the continuous fine measure. "Triple-DiD" includes interactions with the female indicator and the full set of province×year, year×gender, and province×gender fixed effects.
$^{*}\ p < 0.10$, $^{**}\ p < 0.05$, $^{***}\ p < 0.01$.

2.6 Why is the Wage Decline so Large? A Puzzle

A natural question is why an expectation-based response by employers would generate such a *large* fall in young women's wages. The canonical statistical–discrimination story emphasizes higher expected separation (replacement) costs after childbirth, which depresses starting wages. That mechanism is weak in China's high-skill labor market, where employment relationships are unusually long-lived and separations are infrequent.

Table 7 documents this low mobility: between 2010 and 2016, on average 67.9% of workers remained with the *same* employer two years later (versus roughly 40.5% in the United States over a comparable horizon). Household survey evidence from 2010 is consistent with this: over half of respondents report that their current job is their first job, with the pattern most pronounced among the tertiary-educated. A key driver is the prevalence of government-related employment (civil service and Party organs; public institutions such as schools, hospitals, research institutes; state-owned or state-controlled enterprises), which offers near-tenure and continues to attract highly educated workers—even among younger cohorts (ages 18–35).

Table 7: Labor market transition matrix (two-year horizon), employed individuals

	Number employed	Unemployed	Employed same employer	Employed changed jobs	Out of labor force
2010–2012	2,367	1.48%	74.31%	13.48%	10.73%
2012–2014	3,327	1.41%	62.79%	23.47%	12.32%
2014–2016	2,925	1.26%	66.56%	20.24%	11.93%
Average	2,873	1.39%	67.89%	19.06%	11.66%

Notes: The data are from China Family Panel Survey. Sample are restricted to all persons older than 16 years old, who lives in urban area, appeared in two consecutive rounds and have valid employment status records in non-farm sector in both rounds.

Taken together, China's labor market for high-skilled workers is low-dynamism and prevalent with long-term employment relationship . In such environments, pay is tied to the *present value* of match productivity, contemporaneous effort is only weakly coupled to current wages, and firing is costly or institutionally constrained. Employers therefore discount young women's wages ex ante if they expect postpartum productivity to be lower, even when separations do not rise.

A sectoral test of the long-term pricing channel. If long-term contracting and firing frictions are the conduit, the effect should be strongest where such frictions are most salient: government-related employment. I split the sample into government-related sectors (Government/Party/people's organizations; public institutions/research institutes; SOEs/state-controlled firms) and other sectors, and re-estimate the female DiD in (1) (reporting female coefficients) alongside the triple-differences specification (2). I report results using both the binary treatment and the continuous intensity.

Table 8: Wage Regressions by Government-Related Sector

	Panel A: Government-related = Y				**Panel B: Government-related = N**			
	Female		Triple-DiD		Female		Triple-DiD	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Dependent variable: *ln_wage*							
Years of schooling	0.036	0.037	0.052*	0.053*	0.092***	0.093***	0.091***	0.091***
	(0.030)	(0.031)	(0.028)	(0.028)	(0.021)	(0.021)	(0.010)	(0.010)
Age	0.119	0.111	0.132	0.128	0.524***	0.523***	0.517***	0.517***
	(0.101)	(0.104)	(0.113)	(0.110)	(0.121)	(0.121)	(0.076)	(0.076)
Age^2	-0.001	-0.001	-0.002	-0.002	-0.009***	-0.009***	-0.008***	-0.008***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.001)	(0.001)
Treat	-0.194				-0.145			
	(0.150)				(0.100)			
Treat (cont.)		-0.136***				-0.049		
		(0.045)				(0.040)		
Treat × Female			-0.260				-0.197	
			(0.215)				(0.123)	
Treat (cont.) × Female				-0.200***				-0.070
				(0.063)				(0.042)
Observations	429	429	873	873	1,038	1,038	2,113	2,113
R^2	0.252	0.256	0.312	0.314	0.222	0.222	0.267	0.267
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Province FE	Y	Y	Y	Y	Y	Y	Y	Y
Industry FE	Y	Y	Y	Y	Y	Y	Y	Y
Clustered SE: Province	Y	Y	Y	Y	Y	Y	Y	Y
Gov-related	Y	Y	Y	Y	N	N	N	N

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

As shown in Table 8, despite smaller samples, government-related wages for young women fall significantly under the *continuous* exposure measure, with magnitudes larger than in the full sample (cf. Table 3). In contrast, wage changes in non-government sectors are statistically indistinguishable from zero. This pattern is consistent with a long-term pricing (expectations) mechanism operating most strongly where employment relationships are the stickiest.

2.7 Mechanism: Worse new jobs *vs.* Stalled Promotions

Because employment on the extensive margin does not change after the OCP relaxation, the wage decline must operate along the intensive margin. Two natural channels are: (i) new entrants accepting lower–quality jobs ("worse new jobs"), and (ii) slower within–job advancement ("stalled promotion"). I test the promotion channel using CFPS questions on supervisory responsibility—*Has subordinates?* and, if yes, *How many?*. The estimation specification is analogous to (2):

$$\begin{aligned}
Prob.\left(\text{has subordinates}\right)_{it} &= \beta_0 + \beta_4 \text{Treat}_{pt} \times \text{female}_g + \mu_p \times \gamma_t + \gamma_t \times \text{female}_i + \mu_p \times \text{female}_i + \varepsilon_{it} \\
\log\left(\text{num of subordinates}\right)_{it} &= \beta_0 + \beta_4 \text{Treat}_{pt} \times \text{female}_g + +\mu_p \times \gamma_t + \gamma_t \times \text{female}_i + \mu_p \times \text{female}_i + \varepsilon_{it}
\end{aligned}$$

Table 9 shows no evidence of a promotion channel: neither the probability of supervising others nor the (log) number of direct reports changes for young women relative to men. Given the limited sample for counts, these results are suggestive but consistent with the "worse new jobs" interpretation rather than stalled promotions.

Table 9: Effect of OCP Relaxation on Female Workers' Promotions

	has_subordinates		*ln_subordinates_num*	
	(1)	(2)	(3)	(4)
Female × Treat	0.163		-0.017	
	(0.264)		(0.400)	
Female × Treat (cont.)		0.096		-0.058
		(0.142)		(0.152)
Sample mean	0.200	0.200	1.700	1.700
Observations	2,003	2,003	514	514

Notes: Standard errors in parentheses.
has_subordinates is an indicator for holding a supervisory position. *ln_subordinates_num* is the log number of direct reports.
Treat denotes exposure to the OCP relaxation (indicator); *Treat (cont.)* is the continuous exposure measure.
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Finally, age heterogeneity adds corroborating evidence for the "worse new jobs" channel: younger women—who are more likely to be new entrants—experience significantly larger wage declines, while women ages 29–35 do not. I apply Equation 2 to young females in different age groups and present the results in Table 10, where the first and fourth columns are the same as the last two columns in Table 3 and serve as benchmark. Since the mean age of childbirth age in China is round 28 in 2012 [4], I separate the sample by the age of 28. Among all treated young females, the annual wage income for females aged 18 to 28 decreases by an additional 21.5 percentage points compared to the average after the policy relaxation, while females aged 29 to 35 show no significant effect.

[4] United Nations, Department of Economic and Social Affairs, World Fertility Data (2019)

Table 10: Effect of OCP Relaxation on Female Wages by Age Group

	Panel A: Binary treatment			Panel B: Continuous treatment		
	≤ 35	> 28	≤ 28	≤ 35	> 28	≤ 28
	Dependent variable: *ln_wage*					
Years of schooling	0.078***	0.097***	0.060***	0.078***	0.097***	0.060***
	(0.009)	(0.018)	(0.012)	(0.010)	(0.017)	(0.012)
Age	0.455***	0.312	0.690***	0.455***	0.311	0.686***
	(0.072)	(0.348)	(0.228)	(0.072)	(0.350)	(0.229)
Age^2	-0.007***	-0.005	-0.012**	-0.007***	-0.005	-0.012**
	(0.001)	(0.006)	(0.005)	(0.001)	(0.006)	(0.005)
Treat $\times$ Female	-0.148*	0.052	-0.363**			
	(0.078)	(0.108)	(0.162)			
ConTreat $\times$ Female				-0.075**	0.008	-0.162*
				(0.030)	(0.032)	(0.084)
Observations	2,986	1,503	1,483	2,986	1,503	1,483
R^2	0.239	0.278	0.232	0.239	0.278	0.232
Year fixed effects	Y	Y	Y	Y	Y	Y
Province fixed effects	Y	Y	Y	Y	Y	Y
Clustered SE: Province	Y	Y	Y	Y	Y	Y

Standard errors in parentheses.

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

3 Theoretical Model

In this section I develop a tractable framework that links household time allocation to wage setting in frictional labor markets with long-horizon matches. The household block delivers effort as a function of free time; effort governs learning-by-doing human capital accumulation and the effective use of accumulated capital. I then embed the difference in effort-driven productivity in a wage-posting model with on-the-job search: firms post fixed salaries at hire, matches separate either exogenously or when better offers arrive, and wages are priced off the present value of expected productivity over the match. I characterize the firm's profit function, the law of motion and cross-sectional stock of wages, prove continuity of the offer distribution, and state the equilibrium conditions. The key implication is that, even holding employment fixed, differences in expected effort and accumulation translate into systematic wage differences through present-value pricing.

3.1 Household Collective Decision Model

Household problem. A household consists of one man (M) and one woman (F). Each individual has a unit time endowment. The household chooses time allocated to childcare (t_h^j) and market work (t_w^j) for $j \in \{M, F\}$ to maximize

$$\begin{aligned}
\max_{\{t_h^j, t_w^j\}_{j\in\{M,F\}}} \quad & \mathbf{1}\{\text{emp}^M = 1\}\, w^M + \mathbf{1}\{\text{emp}^F = 1\}\, w^F + \phi_H \log n + \phi_L \log\bigl(f^M + f^F\bigr) \\
\text{s.t.} \quad & \text{emp}^j = \mathbf{1}\{t_w^j \geq \underline{t}_w\}, \qquad j \in \{M, F\}, \\
& \gamma^F t_h^F + \gamma^M t_h^M \;\geq\; g(n), \\
& f^j = 1 - t_h^j - t_w^j, \qquad j \in \{M, F\}, \\
& t_h^j \geq 0,\;\; t_w^j \geq 0,\;\; t_h^j + t_w^j \leq 1, \qquad j \in \{M, F\}.
\end{aligned}$$

Here, w^j is the full-time wage for spouse j, and employment is a discrete outcome determined by the full-time threshold $\underline{t}_w$. The household derives utility from the number of children n (with weight $\phi_H > 0$) and from the *sum of partners' free time* $f^M + f^F$ (with weight $\phi_L > 0$). Childcare requires at least $g(n)$ units of effective time, where γ^j denotes spouse j's productivity in childcare.

Assumption 3.1 (Childcare productivity)**.** Women are, on average, more productive in childcare than men: $\gamma^F > \gamma^M$.

n is taken as exogenous in this section (so $\phi_H \log n$ is constant for the time-allocation problem). The function $g(n)$ is increasing in n with $g(0) = 0$, and all parameters are assumed positive unless stated otherwise.

Human capital accumulation rate. Because employment is discrete, optimal market hours satisfy $t_w^j \in \{0, \underline{t}_w\}$. Hence, among employed individuals ($t_w^j = \underline{t}_w$), variation in human capital accumulation is driven by effort ε^j, which depends on available *free time* $f^j \equiv 1 - t_h^j - \underline{t}_w$. Free time matters for at least two reasons: (i) adequate rest and recovery improve focus during fixed working hours; and (ii) many job-related investments occur off the clock (self-study, professional reading, business travel, client-maintenance activities). More free time implies more effort at work, and more effort at work accelerates the learning-by-doing process. Since childcare requirements $g(n)$ increase with the number of children n, available free time $f^j(n)$ is decreasing in n. I therefore model the accumulation rate for an individual with n children as

$$\alpha^j(n) = A\big(\varepsilon^j(n)\big), \qquad \varepsilon^j(n) = B\big(f^j(n)\big),$$

where $A : [0,1] \to \mathbb{R}_+$ and $B : [0,1] \to [0,1]$ are increasing functions. For simplicity, I assign the function forms of $\rho(\cdot)$ and $A(\cdot)$ as

$$\begin{aligned} \varepsilon\big(f^j(n)\big) &= e^{-\exp(\psi f^j(n))} \quad \in (0,1) \\ \alpha^j(n) &= \rho_0 + \rho_1 \varepsilon^j(n). \end{aligned}$$

Implications. This simple household model yields several comparative-statics:

(i) $n = 0$**.** With no children, the childcare constraint is slack. If both spouses work full time ($t_w^j = \underline{t}_w$), then $f^M = f^F = 1 - \underline{t}_w$, and women and men will have the exactly the same human capital accumulation rates.

(ii) **Small** n**.** If $g(n)$ is low enough that childcare can be met while both spouses work full time—e.g.,

$$g(n) \leq (\gamma^F + \gamma^M)\big(1 - \underline{t}_w\big),$$

then both participate in the labor market. By Assumption 3.1, the efficient allocation tilts childcare toward the woman, so $f^M > f^F$ and, since $A(\cdot)$ is increasing, men accumulate human capital faster on average.

(iii) **Large** n**.** As n rises, the childcare requirement tightens. When it becomes infeasible (or too costly in terms of free time) to satisfy $g(n)$ with both spouses employed at $\underline{t}_w$, the household chooses a corner solution with one full-time caregiver. Because $\gamma^F > \gamma^M$, reallocating the woman to childcare yields more effective units per hour, so she is the marginal worker more likely to exit employment.

Thus, even conditional on continued employment, increases in n reduce available free time $f^j(n)$ and therefore lower human capital accumulation $\alpha^j(n) = A\big(f^j(n)\big)$—with larger effects on women when childcare is disproportionately assigned to them.

Employers' aggregate expectations about accumulation. With the household block in mind, a firm posting wage w to a worker of gender j faces uncertainty about the worker's future fertility $n \in \{0, 1, 2, \dots\}$ and, therefore, about the match–specific accumulation rate $\alpha^j(n) = A\big(f^j(n)\big)$. Let p^j_n denote the firm's predictive probabilities. Expected profit at w integrates over n:

$$\mathbb{E}\big[\Pi^j(w)\big] \; = \; \sum_{n \geq 0} \Pi^j\big(w; \alpha^j(n)\big)\, p^j_n.$$

Because employers form aggregate expectations and Assumption 3.1 holds, the gender-specific accumulation rates generally satisfy $\alpha^M \geq \alpha^F$.

Proposition 3.2 (Endogenous gender gap in accumulation)**.** *Under Assumption 3.1 ($\gamma^F > \gamma^M$), in any period in which both spouses are employed at $t^j_w = \underline{t}_w$,*

$$f^M \geq f^F \quad \Rightarrow \quad \alpha^M = A(f^M) \geq A(f^F) = \alpha^F,$$

with strict inequality whenever $n > 0$ and A is strictly increasing.

3.2 Labor Market Environment

Time is continuous and I focus on steady-state analysis. The economy features two segmented labor markets indexed by gender $j \in \{F, M\}$. In each submarket there is a continuum of infinitely lived, ex ante identical workers of measure m^j and a unit measure of firms. Differences in human capital accumulation rates α^j (stemming from the household problem above) imply potentially different wage-posting equilibria across gender j.

Each worker is either employed (1) or unemployed (0). Unemployed workers receive a benefit b^j, while employed workers earn a posted wage w^j. Regardless of employment status, job offers arrive as independent Poisson processes with rates λ^j_0 for the unemployed and λ^j_1 for the employed. Offers take the form of take-it-or-leave-it contracts with wage draws w^j

from a continuous cdf F^j. The parameters $(b^j, \lambda_0^j, \lambda_1^j)$ may differ by gender.

Firms post wages and commit to them; there is no on-the-spot renegotiation or counteroffering after an outside offer arrives.[5] Separations occur either exogenously at rate δ^j or endogenously when the worker receives a strictly higher outside offer. Thus, the separation hazard for a worker employed at wage w is

$$h^j(w) \;=\; \delta^j \;+\; \lambda_1^j\big(1 - F^j(w)\big),$$

where the second term is the arrival rate of offers exceeding w. Ties have measure zero by continuity of F^j. I interpret an outside offer as a new position on the wage ladder—either an internal promotion or a move to another firm; the model abstracts from this distinction.

3.3 Workers' Human Capital Accumulation Process

Production uses only workers' human capital: one unit of *effective* human capital produces one unit of output, so a firm's output equals the sum of its employees' effective human capital. For a worker of gender $j \in \{F, M\}$ with tenure $\tau \geq 0$ and n children, I write

$$\text{effective human capital} \;=\; \varepsilon^j(n)\, z^j(\tau),$$

where the effort (or intensity) at work $\varepsilon^j(n) \in (0, 1)$ multiplies the accumulated stock $z^j(\tau)$. Effort depends on available free time—hence on n—as discussed in Section 3.1; fewer free–time resources imply lower $\varepsilon^j(N)$. Lower efforts slows down the accumulation of human capital via learning-by-doing, and thus match-specific accumulation is also affected by the number of children, n:

$$z^j(\tau) = z_0 \exp\big(\alpha^j(n)\,\tau\big), \qquad z_0 > 0, \quad \alpha^j(n) > 0, \tag{3}$$

and accrues only while employed; upon separation, tenure resets and z^j reverts to z_0. Thus a worker's instantaneous productivity is $\varepsilon^j(n)\, z_0 \exp\big(\alpha^j(n)\,\tau\big)$.

[5] This "no counteroffer" assumption is standard in on-the-job search models (e.g., Burdett–Mortensen, 1998) and is also consistent with environments featuring rigid pay scales and limited within-firm bargaining (e.g., civil-service settings).

The accumulation rate $\alpha^j(n)$ is determined in the household block as a function of efforts and therefore, free time of gender j (Section 3.1). By Proposition 3.2, $\alpha^M(n) \geq \alpha^F(n)$ whenever both spouses work. This pattern aligns with "greedy jobs" where long hours and continuous availability are rewarded while flexibility is penalized Goldin (2014): free time provides the slack for off–the–clock investment. Because family responsibilities compress women's free time and therefore, on average, their implied accumulation rates—and hence wages—are lower. Equilibrium wage schedules in segmented markets inherit these gender–specific $\varepsilon(n)\alpha^j(n)$.

3.4 Firms' Wage–Posting Problem

Firms post wages in segmented submarkets indexed by gender $j \in \{F, M\}$. Let F^j denote the cdf of posted offers with support $\left[R^j, \overline{w}^j\right]$, and let $(\lambda_0^j, \lambda_1^j, \delta^j)$ be, respectively, the offer arrival rates for unemployed and employed workers and the exogenous separation rate. Given a posted wage w, an employed worker faces a constant separation hazard

$$h^j(w) \;\equiv\; \delta^j + \lambda_1^j\big[1 - F^j(w)\big], \qquad \text{and } S^j(\tau \mid w) = \exp\big(-h^j(w)\tau\big),$$

where $S^j(\tau \mid w)$ is the survival probability in the current job up to tenure τ.

Per–worker value to the firm. A worker with tenure τ supplies match-specific effective human capital $\varepsilon^j(n)z^j(\tau) = \varepsilon^j(n)z_0 \exp(\alpha^j(n)\tau)$ (Section 3.3). The firm's expected profit *per hire* at posted wage w equals the expected integral of the contemporaneous profit margin up to the (random) separation time:

$$\mathbb{E}\big[\pi^j(w)\big] = \int_0^\infty \big[\sum_{n=0}^{n=\overline{n}} z_0\varepsilon^j(n)e^{\alpha^j(n)\tau}p_n^j - w\big]\; S^j(\tau \mid w)\, d\tau \tag{4}$$

$$= \sum_{n=0}^{n=\overline{n}} \frac{z_0\varepsilon^j(n)p_n^j}{h^j(w) - \alpha^j(n)} - \frac{w}{h^j(w)} \tag{5}$$

which is finite whenever $\alpha^j(n) < h^j(w)$ for all n. A sufficient, wage–uniform condition is $\alpha^j(n) < \delta^j$ since $h^j(w) \geq \delta^j$ for all w.

Employment stock at wage w. Let m^j be the measure of type–j workers and u^j the steady–state unemployment mass. Denote by $G^j(w)$ the cross–sectional wage cdf among the employed. The law of motion for $G^j(w,t)$ is

$$\frac{\partial G^j(w,t)}{\partial t} = \underbrace{\lambda_0^j\big[F^j(w) - F^j(R^j)\big]\, u^j(t)}_{\text{inflow from unemployment}} - \underbrace{\big[\lambda_1^j\big(1 - F^j(w)\big) + \delta^j\big]\, G^j(w,t)\,\big(m^j - u^j(t)\big)}_{\text{outflow to higher offers or separations}},$$

for $w \geq R^j$ (the inflow term is zero for $w < R^j$). In a stationary equilibrium,

$$\frac{\partial G^j(w,t)}{\partial t} = 0 \quad \text{and} \quad \lambda_0^j\big[1 - F^j(R^j)\big]u^j = \delta^j\big(m^j - u^j\big),$$

so

$$G^j(w) = \frac{\delta^j\big[F^j(w) - F^j(R^j)\big]}{h^j(w)\,\big[1 - F^j(R^j)\big]}, \qquad h^j(w) \equiv \delta^j + \lambda_1^j\big(1 - F^j(w)\big).$$

Let $l^j(w)$ denote the equilibrium stock of employees (per unit mass of firms) earning exactly w. Using G^j and F^j,

$$\begin{aligned}
l^j(w) &= \lim_{\varepsilon\to 0} \frac{G^j(w) - G^j(w-\varepsilon)}{F^j(w) - F^j(w-\varepsilon)}\,(m^j - u^j) \\
&= \lim_{\varepsilon\to 0} \frac{\dfrac{\big[F^j(w) - F^j(R^j)\big]\delta^j}{h^j(w)\big[1 - F^j(R^j)\big]} - \dfrac{\big[F^j(w-\varepsilon) - F^j(R^j)\big]\delta^j}{h^j(w-\varepsilon)\big[1 - F^j(R^j)\big]}}{F^j(w) - F^j(w-\varepsilon)}\,(m^j - u^j) \\
&= \lim_{\varepsilon\to 0} \frac{F^j(w)\,h^j(w-\varepsilon) - F^j(w-\varepsilon)\,h^j(w) + F^j(R^j)\big[h^j(w) - h^j(w-\varepsilon)\big]}{\big[h^j(w)\,h^j(w-\varepsilon)\big]\big[1 - F^j(R^j)\big]\big[F^j(w) - F^j(w-\varepsilon)\big]}\,(m^j - u^j)\,\delta^j.
\end{aligned}$$

Define $h^j(w) = h^j(w-\varepsilon) + \Delta_\varepsilon$, then

$$\Delta_\varepsilon = \lambda_1^j\big[F^j(w-\varepsilon) - F^j(w)\big].$$

Hence,

$$
\begin{aligned}
l^j(w) &= \lim_{\varepsilon\to 0} \frac{\big[F^j(w) - F^j(w-\varepsilon)\big]\big[h^j(w-\varepsilon) + \lambda_1^j F^j(w-\varepsilon)\big] + F^j(R^j)\Delta_\varepsilon}{\big[h^j(w)\,h^j(w-\varepsilon)\big]\big[1 - F^j(R^j)\big]\big[F^j(w) - F^j(w-\varepsilon)\big]}\,(m^j - u^j)\,\delta^j \\
&= \lim_{\varepsilon\to 0} \frac{h^j(w-\varepsilon) + \lambda_1^j\big[F^j(w-\varepsilon) - F^j(R^j)\big]}{\big[h^j(w)\,h^j(w-\varepsilon)\big]\big[1 - F^j(R^j)\big]}\,(m^j - u^j)\,\delta^j \\
&= \lim_{\varepsilon\to 0} \frac{h^j(w-\varepsilon) + \lambda_1^j\big[F^j(w-\varepsilon) - F^j(R^j)\big]}{\big[h^j(w)\,h^j(w-\varepsilon)\big]\big[1 - F^j(R^j)\big]}\,\lambda_0^j\big[1 - F^j(R^j)\big]u^j \\
&= \lim_{\varepsilon\to 0} \frac{\delta^j + \lambda_1^j\big(1 - F^j(w-\varepsilon)\big) + \lambda_1^j\big[F^j(w-\varepsilon) - F^j(R^j)\big]}{h^j(w)\,h^j(w-\varepsilon)}\,\lambda_0^j u^j \\
&= \lim_{\varepsilon\to 0} \frac{\delta^j + \lambda_1^j\big(1 - F^j(R^j)\big)}{h^j(w)\,h^j(w-\varepsilon)}\,\lambda_0^j u^j.
\end{aligned}
$$

In equilibrium, no firm posts below R^j so $F^j(R^j) = 0$, and if F^j is continuous, which I am going to prove in the next part,

$$
l^j(w) = \frac{\big(\delta^j + \lambda_1^j\big)\lambda_0^j u^j}{\big[h^j(w)\big]^2} = \frac{\big(\delta^j + \lambda_1^j\big)\lambda_0^j}{\big[h^j(w)\big]^2}\cdot\frac{\delta^j m^j}{\delta^j + \lambda_0^j},
$$

No mass points in F^j. In a wage-posting equilibrium, F^j is continuous on its support. If there were an atom at some interior w, a firm that deviates to $w+\varepsilon$ (for arbitrarily small $\varepsilon > 0$) would induce a discrete fall in $h^j\big(w+\varepsilon\big) = \delta^j + \lambda_1^j\big(1 - F^j(w+\varepsilon)\big)$, which strictly raises the employment stock $l^j(w)$ and lengthens expected tenure per hire. The only first–order cost is the infinitesimal increase in the posted wage, which reduces the per-hire margin by ε. By (7), the discrete gains dominate, yielding a profitable deviation—contradicting equilibrium profit equalization. Hence F^j has no atoms on its support. A full proof follows.

Lemma 3.3 (No mass points in F^j on the interior of its support)**.** *Fix $j \in \{F, M\}$. In any stationary wage–posting equilibrium, the offer cdf F^j is continuous on $(R^j, \overline{w}^j)$.*

Proof. Suppose, toward a contradiction, that F has an atom at some interior $w \in (R, \overline{w})$; i.e., for arbitrarily small $\varepsilon > 0$,

$$
F(w+\varepsilon) = F(w) + v(w) \quad \text{with } v(w) > 0.
$$

Since $h(w) = \delta + \lambda_1\big(1 - F(w)\big)$, we have

$$h(w+\varepsilon) = \delta + \lambda_1\big(1 - F(w) - v(w)\big) = h(w) - \lambda_1 v(w) > 0.$$

(i.) In-firm stock strictly rises. Using the steady-state stock

$$l(w) = \frac{(\delta+\lambda_1)\,\lambda_0\,u}{[h(w)]^2},$$

it follows that

$$l(w+\varepsilon) = \frac{(\delta+\lambda_1)\,\lambda_0\,u}{[h(w)-\lambda_1 v(w)]^2} = l(w) + \frac{\lambda_1 v(w)\,(\delta+\lambda_1)\,\lambda_0\,u}{h(w)^2\,[h(w)-\lambda_1 v(w)]} \; > \; l(w).$$

(ii.) Per-hire expected profit strictly rises for small ε. The per-hire value is

$$\pi(w) = \int_0^\infty (ze^{\tilde{\alpha}\tau} - w)\,h(w)\,e^{-h(w)\tau}d\tau = \frac{z\,h(w)}{h(w)-\tilde{\alpha}} - w \qquad (\tilde{\alpha} < h(w)).$$

where, for simplicity, $ze^{\tilde{\alpha}\tau} \equiv \sum_n z_0\varepsilon^j(n)e^{\alpha^j(n)\tau}p_n^j$ Hence

$$\pi(w+\varepsilon) - \pi(w) = z\left(\frac{h(w)-\lambda_1 v(w)}{h(w)-\lambda_1 v(w)-\tilde{\alpha}} - \frac{h(w)}{h(w)-\tilde{\alpha}}\right) \; - \; \varepsilon.$$

Because $v(w) > 0$ is fixed while $\varepsilon \to 0$, the bracketed term is a strictly positive constant (it reflects the discrete drop in h), whereas the wage increment contributes only $-\varepsilon = o(1)$. Thus, for all sufficiently small $\varepsilon > 0$,

$$\pi(w+\varepsilon) > \pi(w).$$

(iii) Total expected profit strictly rises. Total profit is

$$\mathbb{E}[\Pi(w)] = \pi(w)\,l(w) = \left(\frac{z\,h(w)}{h(w)-\tilde{\alpha}} - w\right)\frac{(\delta+\lambda_1)\,\lambda_0\,u}{[h(w)]^2}.$$

Combining (i)–(ii),

$$\mathbb{E}[\Pi(w+\varepsilon)] - \mathbb{E}[\Pi(w)] = \big(\pi(w+\varepsilon) - \pi(w)\big)\, l(w+\varepsilon) + \pi(w)\big(l(w+\varepsilon) - l(w)\big) \;>\; 0,$$

for all sufficiently small $\varepsilon > 0$. This yields a profitable deviation, contradicting equilibrium profit equalization. Therefore F cannot have a mass point at any interior w, and F is continuous on $(R, \overline{w})$. $\square$

Expected profit per posted wage (flow). Standard steady-state flows and continuity of F^j imply

$$\lambda_0^j\big[1 - F^j(R^j)\big]u^j \;=\; \delta^j\big(m^j - u^j\big), \qquad G^j(w) \;=\; \frac{\delta^j\big[F^j(w) - F^j(R^j)\big]}{h^j(w)\big[1 - F^j(R^j)\big]},$$

and

$$l^j(w) \;=\; \frac{\big(\delta^j + \lambda_1^j\big)\,\lambda_0^j\,u^j}{\big[h^j(w)\big]^2},$$

where we have used $F^j(R^j) = 0$ in equilibrium and the steady-state relation to substitute out $(m^j - u^j)$.

Consider a firm that posts wage w in submarket $j \in \{F, M\}$. Let $n^j(w)$ denote the *per–firm inflow* (per unit time) of new hires at wage w. A cohort hired at rate $n^j(w)$ has survival $S^j(\tau \mid w) = \exp[-h^j(w)\tau]$, so after tenure τ the expected number of remaining employees from that cohort is $n^j(w)\, S^j(\tau \mid w)$. In a stationary equilibrium $n^j(w)$ is time–invariant and depends only on w.

The expected profit generated by that cohort is

$$\begin{aligned}
\mathbb{E}\big[\Pi^j_{\text{cohort}}(w)\big] &= \int_0^\infty \Big[\sum_n z_0\varepsilon^j(n)e^{\alpha^j(n)\tau}p_n^j - w\Big]\; n^j(w)\, S^j(\tau \mid w)\, d\tau \\
&= n^j(w)\int_0^\infty \Big[\sum_n z_0\varepsilon^j(n)e^{\alpha^j(n)\tau}p_n^j - w\Big]e^{-h^j(w)\tau}d\tau \\
&= n^j(w)\left(\sum_{n=0} \frac{z_0\varepsilon^j(n)p_n^j}{h^j(w) - \alpha^j(n)} - \frac{w}{h^j(w)}\right),
\end{aligned}$$

which is finite whenever $\alpha^j(n) < h^j(w)$. The equilibrium stock of employees at wage w (per firm), $l^j(w)$, is the integrated surviving cohort:

$$l^j(w) = \int_0^\infty n^j(w)\, S^j(\tau \mid w)\, d\tau = \frac{n^j(w)}{h^j(w)}.$$

Hence,

$$\mathbb{E}\big[\Pi^j(w)\big] = \left(\sum_{n=0} \frac{z_0 \varepsilon^j(n) p_n^j h^j(w)}{h^j(w) - \alpha^j(n)} - w \right) l^j(w). \tag{6}$$

Substituting the equilibrium stock into (6) yields the expected profit per posted wage [6]:

$$\mathbb{E}\big[\Pi^j(w)\big] = \frac{\big(\delta^j + \lambda_1^j\big)\lambda_0^j u^j}{\big[h^j(w)\big]^2} \left(\sum_{n=0} \frac{z_0 \varepsilon^j(n) p_n^j h^j(w)}{h^j(w) - \alpha^j(n)} - w \right). \tag{7}$$

Expression (7) makes the trade-off transparent: raising w lowers the separation hazard $h^j(w)$, which (i) enlarges the employment stock $l^j(w) \propto [h^j(w)]^{-2}$ (attraction/retention) and (ii) increases the value of long tenures via $z_0 h^j(w)/\big(h^j(w) - \alpha^j\big)$. At the same time, a higher w compresses the per-hire margin through the $(-\,w)$ term. The net effect on profits is therefore ambiguous and pins down the equilibrium wage distribution.

3.5 Workers' Lifetime Utility Maximization Problem

The worker's side problem is standard as in the search and match literature. Workers maximize their life-time utility on either employed or unemployed status. I use subscripts (0) and (1) to denote them respectively. When unemployed, workers' the Bellman equation can be expressed as solution to the following asset pricing problem: the interest rate on being unemployed is equal to unemployed benefits, b^j, plus the expected gain from searching for an acceptable job that the wage level is above R^j.

$$r\, V_0^j = b^j + \lambda_0^j \int_{R^j}^{\overline{w}^j} \big[\, V_1^j(x) - V_0^j \,\big]\, dF^j(x).$$

[6]When $\alpha^j(n) = 0$ and $\varepsilon^j(n) = 1$ the expressions collapse to the standard Burdett–Mortensen case.

Similarly, when employed at wage w^j, $V_1^j(w^j)$ follows:

$$r\,V_1^j(w) = w + \lambda_1^j \int_w^{\overline{w}^j} \big[\, V_1^j(x) - V_1^j(w) \,\big]\, dF^j(x) + \delta^j \big[\, V_0^j - V_1^j(w) \,\big], \qquad w \in [R^j, \overline{w}^j].$$

As argued in Burdett and Mortensen (1998), $V_1(\cdot)$ is increasing in w^j and V_0 is independent of w^j, and thus, we could identify a reservation wage R^j such that

$$V_0^j = V_1^j(R^j), \qquad V_1^j(w) \geq V_1^j(R^j) \;\; \forall\, w \geq R^j.$$

and derive an equation in a format similar to that in Burdett and Mortensen (1998):

$$R^j - b^j = \big(\lambda_0^j - \lambda_1^j\big) \int_{R^j}^{\overline{w}^j} \big[\, V_1^j(x) - V_0^j \,\big]\, dF^j(x) \tag{8}$$

$$= \big(\lambda_0^j - \lambda_1^j\big) \int_{R^j}^{\overline{w}^j} \frac{1 - F^j(x)}{r + \delta^j + \lambda_1^j\big(1 - F^j(x)\big)}\, dx \tag{9}$$

3.6 Equilibrium

Definition 3.4 (Stationary mixed–strategy wage–posting equilibrium)**.** Fix $j \in \{F, M\}$ and parameters $(m^j, r, b^j, \lambda_0^j, \lambda_1^j, \delta^j, z_0, \alpha^j)$. A stationary mixed–strategy wage–posting equilibrium is a tuple

$$\big(F^j, G^j, u^j, R^j, \overline{w}^j\big),$$

where F^j is the cdf of posted wage offers in market j, G^j is the cross–sectional cdf of accepted wages among the employed, u^j is the mass of unemployed workers, R^j is the reservation wage, and $\overline{w}^j$ is the upper endpoint of the offer support, such that:

(i) **Worker optimality.** Unemployed workers follow a reservation rule with threshold R^j (accept any $w \geq R^j$); employed workers accept any strictly higher offer $w' > w$. Ties have measure zero by continuity. The reservation rule satisfies Equation 9.

(ii) **Firm optimality and no profitable deviation.** A measure–zero firm takes F^j (and thus $h^j(w) = \delta^j + \lambda_1^j[1 - F^j(w)]$) as given and chooses a (possibly mixed) strategy σ—a probability measure on $[R^j, \overline{w}^j]$—to maximize

$$\max_{\sigma} \int \mathbb{E}\big[\Pi^j(w; F^j)\big] \; d\sigma(w)$$

where $\mathbb{E}[\Pi^j(w; F^j)]$ is the expected profit the firm can get when posting a wage w, and

$$\mathbb{E}\big[\Pi^j(w; F^j)\big] = \frac{\big(\delta^j + \lambda_1^j\big)\lambda_0^j u^j}{\big[h^j(w)\big]^2} \left(\sum_{n=0} \frac{z_0 \varepsilon^j(n) p_n^j h^j(w)}{h^j(w) - \alpha^j(n)} - w \right) \qquad \text{(cf. (7))}.$$

There exist a nonempty set $W^j \subseteq [R^j, \overline{w}^j]$ and a constant $\bar{\pi}^j$ such that

$$\mathbb{E}[\Pi^j(w; F^j)] = \bar{\pi}^j \text{ for all } w \in W^j, \qquad \mathbb{E}[\Pi^j(w; F^j)] \leq \bar{\pi}^j \text{ for all } w \notin W^j,$$

and any equilibrium mixed strategy satisfies $\operatorname{supp}(\sigma) \subseteq W^j$.

(iii) **Support and regularity.** F^j is a valid cdf with support $[R^j, \overline{w}^j]$, $F^j(R^j) = 0$, and (if finite) $F^j(\overline{w}^j) = 1$. There are no mass points on $(R^j, \overline{w}^j)$ (Lemma 3.3).

(iv) **Flow balance and cross–sectional consistency.** Given F^j and R^j,

$$\lambda_0^j \big[1 - F^j(R^j)\big] \, u^j = \delta^j \big(m^j - u^j\big),$$

and the employed wage distribution satisfies

$$G^j(w) = \frac{\delta^j \big(F^j(w) - F^j(R^j)\big)}{h^j(w) \big(1 - F^j(R^j)\big)}, \qquad w \in [R^j, \overline{w}^j].$$

(v) **Stationarity.** F^j, G^j, u^j, R^j, and $\overline{w}^j$ are time–invariant and consistent with the hazards $h^j(w)$ induced by F^j.

4 Solution Method and Estimation Results

This section brings the model to the data and quantifies the role of expected productivity in labor markets with contracted labor. I recast the equilibrium conditions in a form suitable for computation, describe the algorithm that recovers the offer distribution from parameters, and estimate these parameters by matching the model-implied wage CDFs to the CFPS distributions (2012 baseline and 2014 post-reform). I show the baseline fit, re-estimate the model after the policy relaxation, and conduct counterfactuals that (i) equalize family-work time, and (ii) update only fertility expectations to quantify how expectation-driven pricing accounts for both the baseline gap and its post-reform widening.

4.1 Numerical Solution

Fix a gender submarket j and suppress the superscript j. In equilibrium, firms' and workers' behaviors are optimal and flows are time-invariant. I rewrite the equilibrium conditions, which are equivalent to what have been introduced in Section 3 but now more suitable for deriving a numerical solution.

First, for all w in the support of the posted-offer cdf F,

$$\mathbb{E}[\Pi(w)] = \frac{\big(\delta+\lambda_1\big)\lambda_0 u}{\big[h(w)\big]^2}\left(\sum_{n=0}\frac{z_0\varepsilon(n)p_n h(w)}{h(w)-\alpha(n)} - w\right) = \bar{\pi}. \tag{10}$$

where $h(w) \equiv \delta + \lambda_1\big[1-F(w)\big]$ and thus, $h(R) = \delta+\lambda_1$, and $h(\overline{w}) = \delta$. Equation 10 evaluated at the endpoints gives

$$\begin{aligned}\mathbb{E}\big[\Pi(R)\big] &= \frac{\lambda_0 u}{\delta+\lambda_1}\left(\sum_{n=0}\frac{z_0\varepsilon(n)p_n(\delta+\lambda_1)}{\delta+\lambda_1-\alpha(n)} - R\right),\\ \mathbb{E}\big[\Pi(\overline{w})\big] &= \frac{\big(\delta+\lambda_1\big)\lambda_0 u}{\delta^2}\left(\sum_{n=0}\frac{z_0\varepsilon(n)p_n\delta}{\delta-\alpha(n)} - \overline{w}\right).\end{aligned}$$

Equalizing profits at R and $\overline{w}$ pins down the upper endpoint:

$$\overline{w} = \sum_{n=0}\frac{z_0\varepsilon(n)p_n\delta}{\delta-\alpha(n)} - \frac{\delta^2}{(\delta+\lambda_1)^2}\left(\sum_{n=0}\frac{z_0\varepsilon(n)p_n(\delta+\lambda_1)}{\delta+\lambda_1-\alpha(n)} - R\right) \tag{$*$}$$

which is well defined under $\alpha(n) < \delta$.

Equivalently, rearranging (10) yields the convenient implicit equation (constant in w on the support)

$$\sum_{n=0} \frac{z_0 \varepsilon(n) p_n}{(h(w)-\alpha(n))\, h(w)} - \frac{w}{[h(w)]^2} = \sum_{n=0} \frac{z_0 \varepsilon(n) p_n}{(\delta+\lambda_1-\alpha(n))\,(\delta+\lambda_1)} - \frac{R}{(\delta+\lambda_1)^2} \qquad (**)$$

On the worker side, remember R is the reservation wage and $\overline{w}$ the upper support. The reservation condition implies

$$R-b = (\lambda_0-\lambda_1)\int_R^{\overline{w}} \frac{1-F(x)}{r+\delta+\lambda_1\big(1-F(x)\big)}\, dx. \qquad (***)$$

Numerical solution (computing $F(w)$ given parameters). The equilibrium is characterized by three equations, $(*)$, $(**)$, and $(***)$. Our goal is to solve for the endogenous offer cdf $F(w)$.

Fix a gender submarket j and suppress the superscript. Given $\theta = (r, z, b, \delta, \lambda_0, \lambda_1, \alpha)$, recover the stationary F on $[R, \overline{w}]$ as follows. For a trial reservation wage $R \in [R_{\min}, R_{\max}]$, compute the upper endpoint $\overline{w}$ from $(*)$. On a fine grid $\{w_k\}_{k=1}^K \subset [R, \overline{w}]$, solve the implicit equal–profit condition $(**)$ for the hazard $h(w_k)$ at each node. Then back out the offer cdf pointwise via

$$F(w_k) = 1 - \frac{h(w_k)-\delta}{\lambda_1},$$

and impose $F(R) = 0$ and $F(\overline{w}) = 1$. A monotone interpolant of $\{(w_k, F(w_k))\}_{k=1}^K$ yields a continuous F on $[R, \overline{w}]$.

Given the implied F, evaluate the reservation equation $(***)$ and define the one–dimensional residual

$$\phi(R) \equiv \big(R-b\big) - (\lambda_0-\lambda_1)\int_R^{\overline{w}} \frac{1-F(x)}{r+\delta+\lambda_1\big(1-F(x)\big)}\, dx.$$

Choose R^* such that $\phi(R^*) = 0$ (e.g., by a bracketing root–finder), and take the associated F as the equilibrium offer distribution.

Finally, the observed cross–sectional wage cdf G is recovered from F via the one–to–one mapping

$$G(w) = \frac{\delta\, F(w)}{\delta+\lambda_1\big(1-F(w)\big)} \qquad \Longleftrightarrow \qquad F(w) = \frac{(\delta+\lambda_1)\, G(w)}{\delta+\lambda_1 G(w)}\,.$$

4.2 Estimation Results

I use Chinese wage data in 2012 to estimate the parameters governing the equilibrium wage distribution before the relaxation of the policy. I retain the sample used in the difference-in-differences analysis in Section 2 and construct gender-specific wage distributions before and after the OCP relaxation (2012 and 2014). For each gender–year cell we estimate the parameter vector

$$\theta = (\alpha^j, \lambda_0^j, \lambda_1^j, \delta^j, b^j, z_0)$$

by squared minimum distance, matching the employed-wage CDF implied by the model, $G(\cdot;\theta)$, to the empirical CDF $\widehat{G}(\cdot)$. The model assumes a common initial human capital level z_0, so I work with residual wages net of year, province, education, and industry fixed effects to align the data with the model's z_0 normalization.

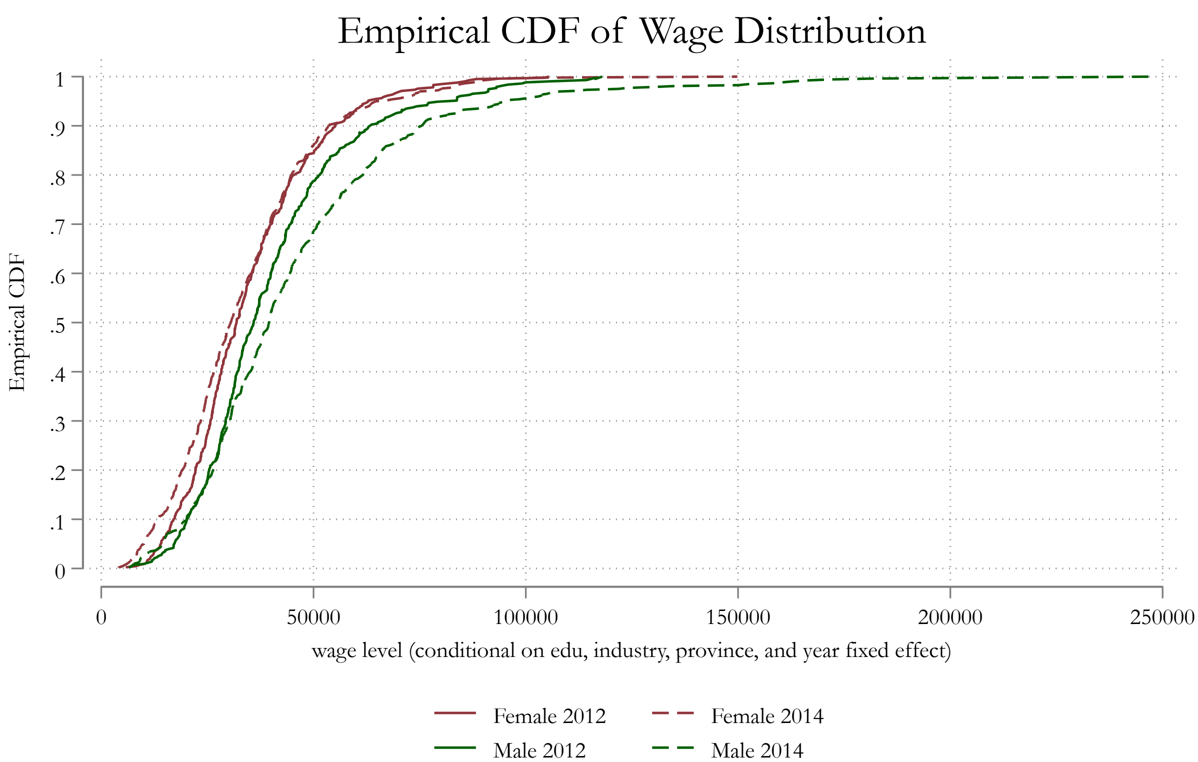


Figure 4: Employed-wage CDFs before and after the OCP relaxation

Figure 4 plots the empirical CDFs: solid lines are pre-OCP (2012) and dashed lines are post-OCP (2014). Two features stand out. First, the female CDF moves little overall, with a modest leftward shift in the lower quantile. An intuitive interpretation would be that new labor-market entrants find it harder to get a good job than before and thus compensate for a lower-paying one. Second, the male CDF shifts notably to the right, especially in the upper quantile, implying men's wages seem to increase, especially if they are more likely to get promoted.

An advantage of the structural approach is that it localizes where in the distribution the gender gap widens and maps those movements to primitives $\theta = (\alpha^j(n), \lambda_0^j, \lambda_1^j, \delta^j, b^j, z_0)$, rather than summarizing changes with a single mean effect.

There parameters I need to pin down include : ψ,ρ_0 and ρ_1 in human capital accumulation function:

$$\varepsilon\big(f^j(n)\big) = e^{-\exp(\psi f^j(n))} \quad \in (0,1) \qquad \alpha^j(n) = \rho_0 + \rho_1 \rho\big(f^j(n)\big)$$

unemployment benefits b^j, discount rate r in workers' decision problem,

$$R^j - b^j = (\lambda_0^j - \lambda_1^j) \int_{R^j}^{\overline{w}^j} \frac{1 - F^j(x)}{r + \delta^j + \lambda_1^j\big(1 - F^j(x)\big)}\, dx;$$

and a bunch of labor market parameters that govern both the workers' problem and the firm's profit maximization behavior such as exogenous separation rates δ^j, job offer arrival rates for unemployed and employed workers λ_0^j and λ_1^j.

$$\mathbb{E}\big[\Pi^j(w)\big] = \frac{\big(\delta^j + \lambda_1^j\big)\lambda_0^j u^j}{\big[h^j(w)\big]^2} \left(\sum_{n=0} \frac{z_0 \varepsilon^j(n) p_n^j h^j(w)}{h^j(w) - \alpha^j(n)} - w \right).$$

In addition, to fully capture the dynamics, the probabilities that the workers will be an n-child j-gender worker $p^j(n)$ are required, as well as how much free time they are able to enjoy $f^j(n)$. If we write all those estimators in a vector, $\Theta = \{\psi, \rho_0, \rho_1, b^j, r, \delta^j, \lambda_0^j, \lambda_1^j, \sum_{n=0}^{\overline{n}-1} p^j(n), f^j(n)\}$, where parameters with superscript j are gender-specific.

I use the China Family Panel Studies (CFPS)—the same source as in the empirical analysis—to externally calibrate several parameters. The fertility–state probabilities $p^j(n)$

are defined as the probabilities employers assign, at hiring, to a worker of gender j having n children over the expected match horizon. In the 2010 CFPS, by age 35 (the terminal age in my sample) more than 90% of workers have at least one child, while only a small fraction (3–7%) have two. The mean age at first birth is about 28 for women and roughly 30 for men. Because age is not a state variable in the model, I assume employers do not observe a recruit's exact age and instead use a flat prior over the hiring-age window $[a_L, a_H] = [18, 35]$. Let $\bar{a}^F = 28$ and $\bar{a}^M = 30$ denote mean first-birth ages, and let $q^j \in (0, 1)$ be the baseline probability of having at least one child by age 35 (set to 0.90 for women in the baseline). Then the calibration for gender j is

$$p^j(0) = \frac{\bar{a}^j - a_L}{a_H - a_L} + \frac{a_H - \bar{a}^j}{a_H - a_L}(1 - q^j), \qquad p^j(1) = \frac{a_H - \bar{a}^j}{a_H - a_L}\, q^j,$$

and, pre–relaxation, $p^j(2) = 0$. For women, this implies

$$p^F(0) = \frac{28 - 18}{35 - 18} + \frac{35 - 28}{35 - 18} \times 0.10 \approx 0.629, \qquad p^F(1) = \frac{35 - 28}{35 - 18} \times 0.90 \approx 0.371,$$

with $p^F(2) = 0$ prior to the policy relaxation. (Analogous expressions apply for men using $\bar{a}^M = 30$ and the chosen q^M.)

The free time available to a gender–j worker with n children, $f^j(n)$, is not directly observed. I proxy it with the complement of time spent on household chores and family care, $t^j(n)$, from the CFPS time–use module, and use the functional form as

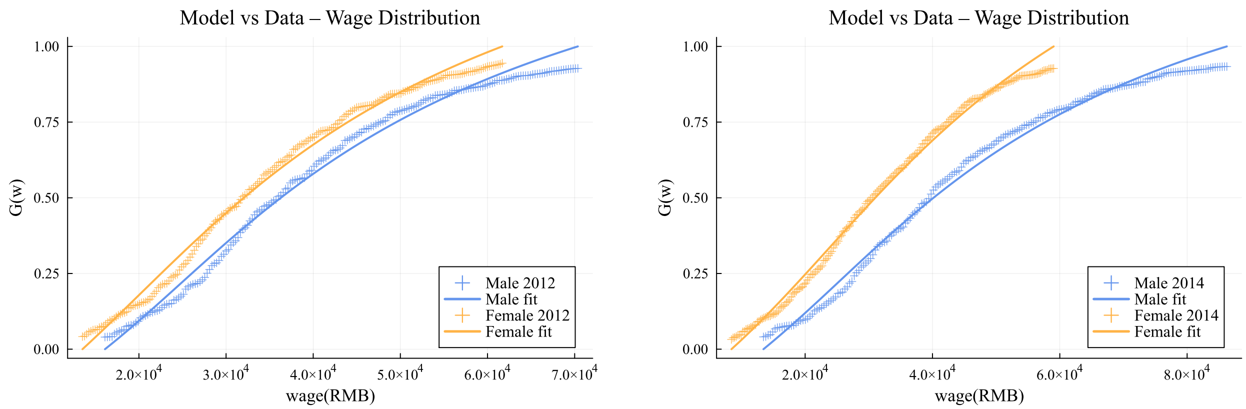
$$\varepsilon^j(n) = e^{-\exp(\varphi t^j(n))}, \quad \text{where } \varphi < 0$$

so that effort falls with family–care time (equivalently, rises with free time). CFPS weekday time–use data imply: women without children spend about 0.85 hours/day on family work (men: 0.54); with one child, women spend 2.44 hours/day (men: 1.18). These moments discipline $t^j(n)$ in the calibration and, through the mapping above, pin down $\varepsilon^j(n)$. In addition to survey-based inputs, I set the discount rate to $r = 0.05$, following standard practice.

As a short summary, the parameters I externally calibrate are: $r = 0.05$, $p^F(0) = 0.63$,

$p^F(1) = 0.37$, $p^M(0) = 0.73$, $p^M(1) = 0.27$, $t^F(0) = 0.85$, $t^F(1) = 2.44$, $t^M(0) = 0.54$, $t^M(1) = 1.18$, and the rest of parameters need to estimated by minimum squared distance are $\widehat{\Theta} = \{\varphi, \rho_0, \rho_1, z_0, b^j, \delta^j, \lambda_0^j, \lambda_1^j\}$.

The estimation results using 2012 gender-specific wage distribution are shown in Table 11 and the model fit is shown in Figure 5a.



(a) 2012: empirical vs. model CDFs by gender

(b) 2014: empirical vs. model CDFs by gender

Figure 5: Model fit for employed wage distributions

Table 11: Results for Estimated Parameters

Panel A: Human capital accumulation

Effort/accumulation mapping parameters: $\varphi = -2.11$, $\rho_0 = 0.04$, $\rho_1 = 0.18$

	$e(0)$	$e(1)$	$e(2)$	$\alpha(0)$	$\alpha(1)$	$\alpha(2)$
Female	0.90	0.74	0.71	0.21	0.18	0.17
Male	0.94	0.87	0.87	0.21	0.20	0.20

Panel B: Labor market parameters

Year-specific common productivity level: $z_0^{2012} = 25600$, $z_0^{2014} = 27400$

Year–Gender	b	λ_0	λ_1	δ	$P = (p(0), p(1), p(2))$
2012–Female	5560	0.58	0.31	0.29	$[0.63, 0.17, 0.00]$
2012–Male	5370	0.87	0.47	0.30	$[0.73, 0.27, 0.00]$
2014–Female	–	0.43	0.31	0.29	$[0.63, 0.16, 0.20]$
2014–Male	–	0.89	0.54	0.29	$[0.73, 0.12, 0.15]$

Notes: In Panel A, $e(n)$ denotes effort/intensity for n children (bounded in $[0, 1]$); $\alpha(n)$ is the human capital accumulation rate. In Panel B, P reports fertility-state probabilities by gender–year; the 2012 P vectors are taken from the data, whereas the 2014 P vectors are estimated. Dashes "–" indicate unchanged relative to the previous year. The initial human capital level z_0 and unemployment benefit b are measured in RMB per year. Standard errors for estimated parameters are forthcoming.

Panel A of Table 11 reports the parameters linking family-work time to effort (φ) and effort to human-capital accumulation (ρ_0, ρ_1). The estimate $\varphi < 0$ indicates that more family work lowers effort, and $\rho_1 = 0.18 > 0$ implies that a worker exerting full effort ($\varepsilon^j(n) = 1$) accumulates human capital at a rate 0.18 above the baseline $\rho_0 = 0.04$. Given these parameters, a woman without children supplies about 90% of her potential effort; having one child raises family-work time from 0.85 to 2.44 hours per weekday and reduces effort to 74%. By comparison, a childless man's effort is slightly higher than a childless woman's, and one child lowers paternal effort only to 87%. Consequently, women with children experience a substantially lower accumulation rate than childless women, whereas the decline in men's accumulation with fatherhood is moderate.

The effort and accumulation levels for individuals with two children are harder to pin down because $t^j(n = 2)$ is scarcely observed in the data. Prior to the policy relaxation, the share with two children was negligible; after the relaxation, realized second births remained low and thus, highly selective. To impute time use at $n = 2$, I propose the following assumption :

Assumption 4.1 (Time use with a second child)**.** For women, the increase in family–work time from one to two children equals 25% of the increase from zero to one child; for men, family–work time is unchanged when moving from one to two children. Formally,

$$t^F(2) \;=\; t^F(1) \;+\; 0.25\left[t^F(1) - t^F(0)\right], \qquad t^M(2) \;=\; t^M(1).$$

This assumption is consistent with diminishing marginal time costs of additional children documented in time–use data. Using the American Time Use Survey, Price (2008) report that mothers' weekly family–work time rises from 2093.5 minutes (one child) to 2662.2 minutes (two children), a 27.2% increase, placing the 25% increment well within the empirical range. By contrast, holding fathers' family–work time fixed at the second child is motivated by the empirical results in Section 2: women's wages respond negatively to higher expected fertility, while men's wages do not—suggesting that employers do not expect men's productivity (and thus their family–work time in our mapping) to change appreciably with a second child. Under Assumption 4.1, using the baseline time–use moments $t^F(0) = 0.85$, $t^F(1) = 2.44$,

$t^M(0) = 0.54$, and $t^M(1) = 1.18$ (hours/weekday), we obtain

$$t^F(2) = t^F(1) + 0.25\,[t^F(1) - t^F(0)] \approx 2.84, \qquad t^M(2) = t^M(1) \approx 1.18.$$

Mapping time use into effort and accumulation yields $\varepsilon^F(2) \approx 0.71 \;\Rightarrow\; \alpha^F(2) \approx 0.17$, and $\varepsilon^M(2) = \varepsilon^M(1) \approx 0.87 \;\Rightarrow\; \alpha^M(2) = \alpha^M(1) \approx 0.20$. Thus, with two children, women's effort falls to about 71% and their accumulation rate to roughly 0.17, whereas men with two children remain at the same effort and accumulation levels as men with one child.

The first two rows of Panel B in Table 11 report the 2012 estimates entering workers' and firms' problems. The initial human capital level z_0 is common across genders by assumption. Women enjoy slightly higher on-the-home benefits than men ($b^F > b^M$). On labor-market dynamics, men have higher offer-arrival rates than women—both when unemployed (λ_0) and on-the-job (λ_1)—in 2012. Exogenous separation rates are similar across genders ($\delta^F \approx \delta^M$).

Estimation for year 2014. With the 2012 estimates as a baseline, I re-estimate the model using the 2014 wage distributions with four clarifications. First, the human capital accumulation parameters (φ, ρ_0, and ρ_1) are held fixed. Second, the parameters decided in the household sector don't change over the two years because there is no evidence of a short-run household response in terms of having more babies after the relaxation of OCP. It includes the unemployment benefits b^j and family work time $t^j(n)$. However, third, in employers' minds, the policy changed expected fertility rates for the workers in the long run. There would be enough people who would have a second child so that $p^j(n = 2) > 0$, and women with two children will spend more time on family work, thus reducing their expected productivity. This change in expected productivity will also change other labor market dynamics, such as job offer arrival rates. Even though we do not know how those labor market dynamics parameters are affected by the human capital accumulation process, we can directly estimate them by minimizing the squared distance between the empirical wage distribution and the model-simulated wage distribution. Lastly, to reduce dimensionality, I keep $p^j(n = 0)$ at the same level in 2014 as in 2012. The underlying assumption is that people do not change their decision about the first birth as stated in Assumption 4.2. There will still be 90% of people who would have children, and they would have their first child at 28 for women and at 30 for men. The only thing that changes is for those who have their first kid, how many of them will have the second.

Assumption 4.2 (First-birth margin fixed)**.** The relaxation of the OCP does not alter the first-birth decision over the sample window; only the conditional probability of a second birth changes.

Figure 5b displays the 2014 model fit to the empirical cdf. The corresponding estimates in Panel B of Table 11 are informative. First, as anticipated, $p^F(2) > 0$ and is sizable. Interpreting $p^j(n)$ as the probabilities employers assign at hiring to a worker of gender j having n children over the expected match horizon, the estimates $p^F(1) = 0.16$ and $p^F(2) = 0.20$ imply that, among women expected to have any children, the conditional second–birth share is

$$\frac{p^F(2)}{p^F(1) + p^F(2)} \approx \frac{0.20}{0.16 + 0.20} \approx 0.56,$$

abstracting from birth spacing. For men, the P–vector is not identified under our maintained assumption that a second birth does not affect male productivity; the entries reported in Table 11 are therefore assigned to mirror the female composition for comparability. This imputation has no impact on the equilibrium because, by assumption, a second child does not shift male effort or accumulation.

Second, the shifts in labor-market dynamics are revealing. Following the OCP relaxation, the job–offer arrival rate for unemployed women, λ_0^F, falls markedly. This implies it became harder for women to receive offers while unemployed (or, equivalently, for new entrants to get offers without accepting lower pay). Interpreting λ_0^F as the ease of job finding for recent graduates, its decline indicates that younger women increasingly had to trade down on wages to obtain employment—consistent with the empirical result that the policy penalized younger cohorts more strongly by a larger wage decrease. By contrast, men's on-the-job offer arrival rate, λ_1^M, rises—indicating a greater likelihood of receiving higher-pay offers (i.e., promotions or upward job-to-job moves along the wage ladder).

As a further check, I overlay the simulated wage CDFs for 2012 and 2014 to compare relative movements with the empirical CDFs in Figure 4. The simulation overlay (Figure 6) mirrors the data: the female CDF shifts left in the lower quantiles, while the male CDF shifts right in the upper quantiles.

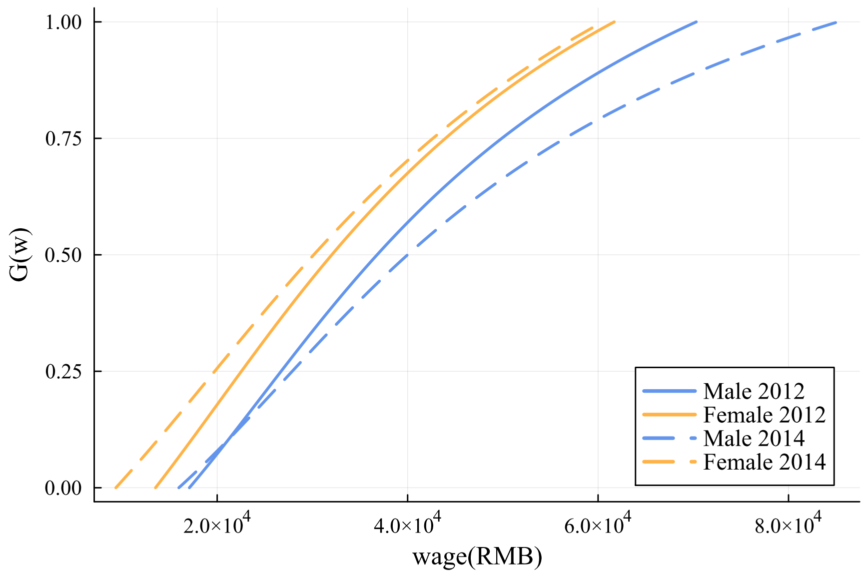


Figure 6: Simulated CDFs (2012 vs. 2014) by Gender: Overlay

4.3 Counterfactual Analysis

In this subsection, I will use the parameter estimation results to conduct several counterfactual analyses.

Gender wage gap decomposition Using the 2012 parameters, I assess how much of the residual gender wage gap (13.3% after controlling for education, industry, etc.) is attributable to differential family–work time by gender. I conduct a counterfactual in which men and women are assigned the same daily family–work time in 2012, holding all other parameters fixed, solve the model, and compute the implied average wages. Panel A of Table 12 shows that, under equalized family–work time, the gender wage ratio is essentially one. In other words, the remaining 2012 gap can be accounted for almost entirely by the unbalanced division of household work.

Empirically (Section 2), young women's wages fell by 15.3% after the policy relaxation. The model results echo this pattern: the simulated gender wage ratio declines from 0.867 in 2012 to 0.732 in 2014 (Table 12), a 13.5 percentage–point drop—about 15.5% if men's wages are held fixed. To gauge the contribution of employers' updated fertility expectations, I run a second counterfactual that changes only the expected fertility distribution $p^j(n)$ to its 2014 values, keeping all other labor–market parameters at their 2012 levels. The implied wage ratio is 0.761, a 12.2% decline. Thus, almost 80% of the observed widening of the gender wage gap is explained by lower predicted productivity for women induced by higher expected fertility—the expectation channel emphasized in this paper.

Table 12: Gender Wage Ratio Decomposition and 2012–2014 Comparison

Scenario	**Female/Male ratio**	**Gender wage gap (pp)**
Panel A: 2012 Baseline Counterfactuals		
Baseline (2012)	0.867	13.3
Female with male family–work hours	1.021	-0.02
Male with female family–work hours	0.987	0.01
Panel B: 2012 vs. 2014 (productivity/distribution shifts)		
Baseline (2014)	0.732	26.8
2012 parameters on 2014 productivity	0.761	23.9

Notes: Gender wage gap is defined as 1−ratio and reported in percentage points (pp). From 2012 to 2014, the ratio changes by −13.5 pp. Holding parameters at 2012 values and changing only expected productivity implies a −10.6 pp change in the ratio.

Propagation via human capital accumulation. In the model, family work matters through two channels. First, more family work reduces on-the-job effort. Second, lower effort slows human capital accumulation, thereby dampening future productivity. The first channel is standard; the novelty here is that household responsibilities depress on-the-job productivity even when full-time employment status is unchanged. To isolate the role of accumulation, consider a counterfactual in which human capital growth does not depend on effort—i.e., α^j is fixed and independent of n (for example, $\alpha^j = \rho_0 + \rho_1 \varepsilon^j(0)$). Fertility then affects only the effort level, not the accumulation rate. Because both genders' expected human capital growth rates increase, the wage distributions move rightward, but the gender wage gap narrows. Statistical discrimination still operates—effort differs across n and cannot be contracted upon—but heterogeneity appears only in flow productivity, not in its change in growth. Without n-variation in α^j, there is no dynamic propagation via tenure accumulation, and the present-value wage gap is much smaller (Figure 7). Quantitatively, the imputed gender wage ratio rises to 0.923, compared with 0.878 in the baseline with α^j varying by n.

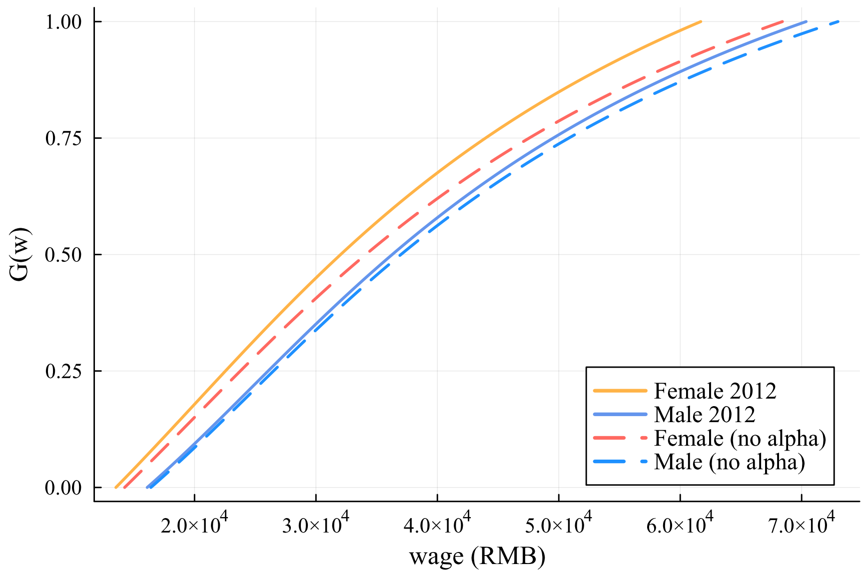


Figure 7: Simulated CDFs (2012 vs. 2014) by gender: overlay

5 Conclusion

This paper demonstrates that when pay is contracted in a long-horizon employment relationship, employers' expectations about workers' future caregiving burdens are reflected in wages in ways that can materially widen gender gaps. Exploiting China's 2013 selective relaxation of the One-Child Policy as a quasi-experiment, I document a 15.3 percent decline in young women's wages with no short-run increase in births, consistent with forward-looking wage setting rather than contemporaneous labor-supply changes. I then develop a search and match model, but extend it to allow for increasing productivity over the match and augment it with a household time-allocation block: effort depends on free time, human capital accumulates with effort, and wages are posted against the present value of an increasing productivity path over the match. Estimating the model on Chinese data, I find that the unequal division of family work alone explains essentially the entire baseline (2012) gender wage gap, and that more than 80 percent of the post-reform widening can be attributed to employers' downward revisions of women's expected productivity associated with anticipated second births.

Two broader implications follow. First, in contract-driven pay settings—government and state-affiliated employment, and many high-skill jobs where matching is costly, training is firm-specific, and replacement is expensive—wages tilt toward expectations about future productivity rather than current output. Any belief that women will accumulate human capital more slowly is amplified by the longer wage-contract horizon and shows up as larger up-front markdowns. Second, policies that merely preserve positions (e.g., job-protection mandates) lengthen the horizon over which present-value pricing applies and can therefore intensify ex-ante markdowns and stall promotions, even as employment is maintained. By contrast, policies that reduce childcare burdens—expanding affordable childcare, designing leave that equalizes caregiving across parents like Sweden's "second dad paternity leave"—directly raise expected effort and accumulation and are more likely to narrow gaps.

The analysis has limitations. Time-use inputs for family work are observed, but the second-child increment is calibrated due to the absence of a clear within-window birth response; employer expectations about second births therefore rest on an assumption and may be imprecise. On the modeling side, fertility is treated as exogenous rather than the outcome of a household decision. A natural next step is to link the household and labor-market blocks more tightly to study feedback: if employers statistically discriminate against women, lower expected returns to market work may induce higher fertility or reduced effort, which in turn validates firms' priors. Such a reinforcing cycle could help explain why, under rising childcare burdens, women disproportionately cut effort while men do not.

Overall, the evidence and the model point to a common conclusion: when salaries are predetermined and productivity grows on the job, expectation-driven pricing is an important force behind gender wage gaps. Closing those gaps requires moving the lower expectations that firms legitimately price for women, rather than only lengthening their tenure within jobs.

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