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Aligning School and Work: How After-School Care Shapes Maternal Labor Supply

Guanfu Fang

Shanghai University of International
Business and Economics

Benyu Qian

Jinan University

Zhong Zhao

Renmin University of China
and IZA@LISER

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Aligning School and Work: How After-School Care Shapes Maternal Labor Supply^{*}

Abstract

This paper examines the labor market effects of China's After-School Care policy—commonly known as the After-School 3:30 PM Service reform—on maternal employment and earnings. Introduced in 2017, the policy addresses the longstanding mismatch between early school dismissal times and standard work schedules by extending on-campus supervision and enrichment activities into the late afternoon. Exploiting the staggered rollout of the program across counties and over time, we show that the policy significantly increases mothers' labor supply and earnings. We document null effects for fathers. The policy operates by reducing maternal housework time, enabling longer commutes, and facilitating access to formal employment, while also inducing firms to increase job postings for female workers. Effects are concentrated among mothers with egalitarian gender attitudes, with minimal impact on those holding traditional views.

JEL classification

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Keywords

after-school care, maternal labor supply, gender norms

Corresponding author

Zhong Zhao

mr.zhong.zhao@gmail.com

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1 Introduction

Across developed and developing economies alike, the structure of the school day is misaligned with the standard work schedule, creating a persistent constraint for working parents—particularly mothers. In more than 70 percent of countries worldwide, primary schools end before 4:00 p.m., forcing parents to choose between employment opportunities and childcare responsibilities. This misalignment often results in reduced work hours, lower earnings, and continued gender inequality in caregiving roles. While extended school days and public childcare are frequently proposed as solutions, empirical evidence regarding their labor market impacts remains mixed, with heterogeneous effects across different institutional and cultural contexts (Dehos & Paul, 2023; Felfe et al., 2016; Martínez & Perticarà, 2017).¹ Understanding the conditions under which such policies effectively increase maternal labor supply has become increasingly important as dual-earner households grow more prevalent in modern economies.

This paper investigates these issues by examining a large-scale, nationwide after-school care (ASC) policy in China, commonly known as the After-School 3:30 PM Service reform. Compulsory schooling in China typically ends around 3:30 p.m., creating a substantial gap until the end of the standard workday. Beginning in 2017, the ASC policy mandated public primary and middle schools to offer subsidized, supervised care during afternoon hours. The policy's staggered implementation across counties provides quasi-experimental variation suitable for causal inference. China's context—marked by rapid economic change alongside deeply rooted Confucian gender norms—offers a unique opportunity to study not only the average effects of such reforms but also how their effectiveness is shaped by prevailing sociocultural beliefs.

¹ A related literature examines the transition from half-day to full-day schooling and generally finds positive effects on maternal labor supply (Berthelon et al., 2023; Contreras & Sepúlveda, 2017; Padilla-Romo & Cabrera-Hernández, 2019). However, full-day schooling extends instructional hours but often still ends before standard work hours conclude, whereas ASC directly targets the post-dismissal gap by providing supervised care into late afternoon.

We ask three primary research questions: (1) Can the provision of subsidized ASC generate substantial improvements in maternal labor market outcomes? (2) What are the underlying mechanisms? (3) To what extent do individual-level gender role attitudes moderate these effects? Using data from the 2014–2022 waves of the China Family Panel Studies (CFPS) and a difference-in-differences (DID) design, we find that the reform significantly increases maternal employment rates by 4.6 percentage points, weekly working hours by 2.5 hours, and earnings by 18 percent. In contrast, we find no statistically significant effects for fathers. These results remain robust across multiple specifications addressing potential concerns regarding parallel trends, confounding policies, selective migration, and heterogeneous treatment effects across implementation waves.

We further investigate the underlying mechanisms. Consistent with a reduction in childcare constraints, the reform significantly decreases mothers' weekday housework time, with no corresponding change on weekends. We also document increases in commuting time and a higher likelihood of formal employment, suggesting improvements in job access and employment quality. To capture demand-side responses, we assemble large-scale online recruitment data and show that the reform is associated with a significant increase in job vacancies explicitly targeting female candidates. Vacancies targeting men decline modestly but insignificantly. These findings indicate that ASC policies can reshape employer expectations and reduce gender-based hiring discrimination.

Moreover, we document substantial heterogeneity in policy effectiveness by gender role attitudes. Labor supply responses are concentrated among mothers with more egalitarian gender beliefs, while effects are negligible for those with traditional views. In contrast, husbands' gender attitudes do not significantly moderate the treatment effects.

This paper makes three contributions to the literature. First, it provides novel causal evidence on the labor market effects of ASC programs in a major emerging economy. Prior studies have yielded mixed results, with ASC shown to raise maternal employment in Chile (Martínez & Perticará, 2017) and Canada (He et al., 2024) but generating limited effects in Switzerland (Felfe et al., 2016) and Germany (Dehos & Paul, 2023). In contrast to prior work examining localized or voluntary program expansions, we study a nationwide, mandatory

reform in China with rapid rollout and near-universal school participation. This scale allows us to examine the labor market effects of a large institutional change, including potential firm-side recruitment responses. Importantly, we demonstrate that individual-level gender norms critically mediate childcare policy effectiveness. This heterogeneity provides a novel potential explanation for the mixed results in prior literature and highlights the role of cultural context in shaping policy effectiveness.

Second, we provide a more comprehensive account of the market-level effects of childcare policy by examining mechanisms on both the labor supply and labor demand sides. Existing research has focused primarily on mothers' labor supply responses, particularly reductions in unpaid household labor (for reviews, see Akgunduz & Plantenga, 2018; Halim et al., 2023; Morrissey, 2017), with limited attention to how childcare policies affect job search constraints or employer behavior. We show that ASC enables mothers to accept jobs with longer commuting time and improves employment quality, consistent with a relaxation of spatial job search constraints (Borghorst et al., 2024; Bütikofer et al., 2025; Le Barbanchon et al., 2021). Moreover, using firm-level hiring outcomes, we document adjustments in employers' recruitment behavior, providing novel evidence that childcare reforms can influence labor demand and reduce gender-based hiring discrimination.

Third, we contribute to the literature on family policy in China. While prior work has largely emphasized early childhood education and childcare (Du & Dong, 2013; Li, 2017; Lin & Wang, 2019; Ren, 2024), we highlight the importance of childcare for families with school-aged children—an increasingly salient yet understudied margin. To our knowledge, this is the first causal evaluation of China's ASC policy. Our findings suggest that extending supervised coverage beyond the regular school day can be an effective tool for improving maternal labor market outcomes and reducing gender inequality, with implications for the design of family policies in China and beyond.

The remainder of the paper is as follows. Section 2 introduces the Chinese institutional background and the conceptual framework. Section 3 describes the data and variables. Section 4 discusses the empirical strategy. Section 5 presents the empirical results on the effect of the ASC policy on maternal labor supply and explores mechanisms. Section 6 concludes.

2 Background and Conceptual Framework

2.1 Background

Rapid socioeconomic change and urbanization in China have intensified a long-standing mismatch between school dismissal times and standard work schedules. During compulsory education — particularly at the primary school level — students are typically dismissed around 3:30 p.m., well before the end of the standard workday. This mismatch has become known as the "3:30 p.m. problem." The timing gap is especially salient for dual-earner households, many of which lack adult supervision for children during late-afternoon hours. Parents, particularly mothers, may therefore face binding constraints on work schedules, commuting choices, and labor supply.

To address this childcare gap, the Chinese government launched the nationwide After-School 3:30 PM Service in 2017. The policy encouraged, and later required, public primary and lower-secondary schools to provide on-campus supervision after regular school hours, generally until 5:30 or 6:00 p.m. The central objective was to reduce parents' after-school childcare burden by aligning school supervision more closely with local work schedules.

The program combines custodial supervision, homework support, and nonremedial enrichment activities. Schools supervise students while they complete written assignments, with the policy goal that primary school students complete most homework on campus and lower-secondary students complete a substantial share of their homework. Schools may also offer activities such as sports, arts, science education, reading, and student clubs. At the same time, the policy prohibits schools from using ASC time to teach new curricular content or provide exam-oriented remedial instruction. The program is therefore designed to provide supervision and homework support rather than to extend formal academic instruction.

The program follows a "5+2" model: schools provide services on all five school days each week for at least two hours per day, with ending times coordinated with local working hours. Participation is voluntary for families, but schools are encouraged—and after 2021 generally required—to make the service available to students with after-school care needs. Importantly, ASC is a broadly accessible school-based service rather than a targeted program restricted to selected groups. Local implementation documents sometimes use priority language for

vulnerable students, including left-behind children, children of migrant workers, and economically disadvantaged students. We interpret this language as referring to efforts to ensure access and affordability, including fee reductions or waivers for disadvantaged students, rather than as indicating that eligibility was limited to these groups.² Left-behind children are children whose parent or parents migrate for work and who live apart from at least one parent for an extended period.

Financing arrangements vary across regions but are regulated by local governments. Following a 2018 State Council directive, local governments may finance ASC through public education budgets, regulated service fees, or administrative charges. Some major cities and provinces fully fund the program through public budgets, while others permit modest family fees subject to nonprofit accounting, voluntary participation, and fee reductions or waivers for disadvantaged students. Fee standards are set jointly by local development, education, and finance authorities based on local income levels and service content. Schools use these funds to cover operating costs and to compensate participating teachers or qualified external providers.

Education authorities monitor implementation through school evaluation systems, routine inspections, and complaint-handling procedures. When schools fail to provide required services, charge irregular fees, or receive substantial public complaints, local education bureaus may require correction or rectification. These monitoring arrangements helped expand program availability. By October 2021, official monitoring data indicate that more than 99 percent of monitored schools offered after-school care, and participation rates approached 90 percent of enrolled students.³

The policy rollout occurred gradually, generating quasi-experimental variation in exposure. Initial guidance issued in 2017 was advisory, encouraging local experimentation and phased adoption. Early implementation was concentrated in economically developed eastern cities beginning in 2017, while many central and western regions adopted the program later. In

² For examples of local implementation rules, see Jiangxi Provincial Department of Education, Implementation Measures for After-School Services in Compulsory Education Schools in Jiangxi Province, available at http://jyt.jiangxi.gov.cn/jxjyw/gfxwj100/pc/content/content_2032365966447280128.html; and Jiaohe Municipal Education Bureau, Policy Interpretation of the Notice on Fees for After-School Services in Compulsory Education Schools, available at http://xxgk.jiaohe.gov.cn/gzbm/swj_1/xxgkml/202211/t20221115_1086386.html.

³ See http://www.moe.gov.cn/jyb_xxgk/xxgk_jyta/jyta_jijiaosi/202301/t20230118_1039652.html.

2021, the "Double Reduction" reform upgraded ASC from a discretionary initiative to a mandatory national requirement starting in the fall semester, although local implementation timelines continued to vary.

There is no authoritative national dataset documenting precise implementation dates at the local level. We compile county-level information on the rollout of the ASC policy from official government implementation plans, education bureau announcements, and local statistical yearbooks. Figure 1 presents the cumulative distribution of adoption years across the 162 counties covered by the CFPS sample. Implementation began in 2017 and continued through 2022, with substantial variation in timing across counties. This staggered rollout underlies our DID identification strategy.

To examine whether ASC adoption timing is systematically related to pre-reform county characteristics, we estimate a Cox proportional hazards model using the county-level implementation year as the event timing. The covariates are drawn from the China County Statistical Yearbook and job-posting data, supplemented by provincial and municipal statistical yearbooks where necessary. Table A1 shows that most pre-reform county characteristics are not systematically associated with adoption timing. GDP per capita and local government revenue are significantly associated with adoption timing, suggesting that local economic and fiscal capacity may have shaped the rollout. By contrast, pre-reform labor-demand measures, including total vacancies and female-oriented vacancies, do not significantly predict implementation timing.

These pretreatment differences raise concerns that treated and control counties may have followed differential trends even absent the reform. To address this potential threat to identification, our robustness checks allow for flexible differential trends by interacting baseline county characteristics with year fixed effects, permitting counties with different observable attributes to evolve along distinct trajectories.

2.2 Conceptual Framework

2.2.1 Overview and Motivation

We develop a simple conceptual framework to examine how after-school care (ASC) affects maternal labor supply and the allocation of time between market work and home

production. The framework combines a time-specific childcare constraint with a minimum-hours requirement in formal employment. By extending school supervision, ASC can make formal jobs feasible and induce a discrete increase in market hours. When this increase exceeds the reduction in maternal childcare, less time remains for housework and leisure. Preferences over home production determine which mothers make this transition. Appendix A provides the derivations.

2.2.2 Setup

A mother allocates her time endowment T among market work m , housework h , childcare c , and leisure l :

$$m + h + c + l = T.$$

Preferences are

$$U = u(x) + \phi[v(h) + \psi(c + \gamma A)] + g(l),$$

where $A \in \{0,1\}$ denotes access to ASC and $\gamma > 0$ measures the effective childcare supplied by the program. All utility components are increasing and twice continuously differentiable. Consumption utility $u(\cdot)$ is concave, and $v(\cdot)$, $\psi(\cdot)$, and $g(\cdot)$ are strictly concave, with negative second derivatives. In particular, $g'(l) > 0$ and $g''(l) < 0$. The parameter $\phi > 0$ captures the relative weight on housework and childcare. Higher values represent stronger preferences for home production relative to consumption and leisure, which we interpret as more traditional gender norms.

Consumption is $x = w_j m + y$, where y is exogenous nonlabor income and $j \in \{0,1\}$ indexes employment type. Informal employment permits flexible hours, including nonemployment at $m = 0$. Formal employment pays a higher wage, $w_1 > w_0 > 0$, but imposes a minimum-hours requirement $\bar{m} \in (0, T)$. Work must be supplied in a continuous block, and school supervision imposes an upper bound $\tau(A)$ on feasible hours:

$$m \leq \tau(A), \quad m \geq \bar{m} \quad \text{in formal employment.}$$

ASC extends supervision, so $\tau(1) > \tau(0)$. If

$$\tau(0) < \bar{m} < \tau(1),$$

formal employment is infeasible before ASC and becomes feasible afterward, provided sufficient time remains for other activities. We focus on interior allocations of housework,

childcare, and leisure; Appendix A discusses corner solutions.

2.2.3 Childcare and Housework

Let λ denote the shadow value of time. The interior first-order conditions for nonmarket activities are

$$\phi v'(h) = \phi \psi'(c + \gamma A) = g'(l) = \lambda.$$

Strictly concave leisure utility allows the shadow value of time to vary with the allocation of time. Holding λ fixed, the childcare condition implies one-for-one displacement of maternal childcare in effective units:

$$\frac{\partial c}{\partial A} = -\gamma, \quad \text{holding } \lambda \text{ fixed.}$$

This is a fixed-shadow-value benchmark and an approximation when the policy changes λ little, rather than an unconditional prediction. A rise in λ reinforces the reduction in maternal childcare; a fall in λ offsets it. ASC has no direct technological effect on housework. Its effect operates through the shadow value of time: when time becomes more valuable, optimal housework falls.

2.2.4 Labor Supply, Gender Norms, and Participation Threshold

Once formal employment is feasible, the mother compares its indirect utility with that of the flexible alternative. Define the net advantage of formal employment as

$$D(A; \phi) = V_1(A; \phi) - V_0(A; \phi).$$

We focus on mothers for whom formal employment entails more market hours than the preferred informal allocation. At a given level of ASC provision, these longer hours reduce optimal housework and childcare. The envelope theorem then implies that $D(A; \phi)$ decreases with ϕ : mothers with stronger preferences for home production incur a larger utility cost from the transition. If this net advantage changes sign over the relevant range of preferences, a unique threshold $\phi^*(A)$ separates mothers who choose formal employment from those who prefer the flexible alternative.

ASC expands access to formal jobs. Where both sectors are feasible, it also raises the participation threshold under the conditions in Appendix A, including an unconstrained informal work schedule and a higher shadow value of time in formal employment. When formal jobs are initially infeasible, the policy instead opens this option to mothers for whom

its net advantage is positive; no pretreatment indifference threshold is required. These results predict stronger transitions into formal work among mothers with less traditional gender norms.

2.2.5 Implications for Time Allocation

For a mother who switches into formal employment, let Δ denote the change from the pre-ASC allocation to the post-ASC allocation. The time constraint implies

$$\Delta h + \Delta l = -\Delta m - \Delta c.$$

If maternal childcare falls and the increase in market hours exceeds this time saving, $\Delta m > -\Delta c$, the time available for housework and leisure contracts. Under the stated preferences, both activities increase with this residual time budget. Its contraction therefore implies

$$\Delta h < 0, \quad \Delta l < 0.$$

Appendix A shows that $\Delta m > \gamma$ is a sufficient condition for this pattern when both allocations are interior. For mothers whose market hours remain unchanged, some of the childcare time released by ASC is instead allocated to housework and leisure. The aggregate housework response therefore depends on the relative size of these adjustments.

Because ASC extends weekday supervision, the direct scheduling mechanism operates on weekdays. It predicts smaller weekend responses provided mothers do not substantially shift market work or home production across days. These are conditional predictions about time allocation, rather than a requirement that all mothers reduce housework or that weekend effects are exactly zero.

3 Data and Variables

3.1 Data Sources and Variables

Our primary data source is the CFPS, a nationally representative longitudinal household survey administered by Peking University's Institute of Social Science Survey. Launched in 2010, the CFPS follows a biennial schedule and tracks individuals even after they separate from their original households, achieving an average reinterview rate of approximately 80 percent. The baseline wave in 2010 surveyed 14,960 households comprising 42,590 individuals across 162 counties in 25 provinces. The survey collects comprehensive information on demographics, employment, income, time use, education, health, and migration history. As

documented by Xie et al. (2017), key demographic and socioeconomic indicators in the CFPS—including age, gender, ethnicity, education, income, and health status—closely align with those from other nationally representative datasets such as the China General Social Survey and the Chinese Household Income Project, supporting the external validity of our findings.

We examine several dimensions of parental labor market performance, including employment status, weekly hours worked, and annual income. Employment status is coded as a binary indicator equal to one if the respondent reports being currently employed and zero otherwise, capturing changes along the extensive margin of labor supply. For respondents who are not working, the CFPS further records the primary reason for nonemployment. We use this information to construct an indicator equal to one if the respondent reports not working due to family caregiving responsibilities—such as caring for children or elderly family members or performing household chores—and zero otherwise. Weekly hours worked measures the total time devoted to market work activities and captures changes along the intensive margin of labor supply. Annual income, measured in Chinese yuan, reflects total labor income from all sources, including wages, salaries, and self-employment income.

We also construct an indicator for formal employment, defined as holding a job that provides the full set of four types of social insurance plus the housing provident fund. Finally, the CFPS records respondents' one-way commuting time to their primary workplace, which we use to examine potential changes in job matching and spatial accessibility.

To measure individual attitudes toward traditional gender norms, we construct a Gender Stereotype Index (GSI) using responses from the 2014 CFPS wave. Respondents rated their agreement with four normative statements regarding gender roles on a five-point Likert scale: (i) "Men should prioritize their careers while women should prioritize the family," (ii) "For women, marrying well is more important than performing well," (iii) "Women should have at least one child," and (iv) "Men should share household responsibilities equally with women." Following Kleven (2022) and Zhang et al. (2024), we standardize each response to have mean zero and unit variance, then average these standardized values across all four items to obtain an individual-level index. Higher GSI values indicate stronger adherence to traditional gender

norms. For heterogeneity analysis, we classify female respondents and their spouses into two groups based on whether their GSI score falls above (traditional views) or below (egalitarian views) zero. This binary classification allows us to examine whether labor market responses to the ASC policy vary systematically with prevailing gender attitudes within the household.

To examine employer demand-side responses to the ASC policy, we use job-posting data from six major Chinese online recruitment platforms: 51job, BOSS Zhipin, Zhaopin, Liepin, Lagou, and Kanzhun. The data cover vacancies posted by publicly listed firms and their affiliates, including parent companies, subsidiaries, and joint ventures. We aggregate vacancies to the firm-year level and construct two primary outcomes: the number of male-oriented vacancies and the number of female-oriented vacancies posted by a firm in a given year. The job-posting panel is separate from the CFPS household sample, so observations in the vacancy analysis refer to firm-year observations rather than individual respondents. To align the descriptive statistics with the fixed-effects regressions, we exclude firms that appear only once in the vacancy panel.

A key measurement challenge is that China prohibited explicit gender preferences in job advertisements beginning on March 1, 2019. Before the ban, employers could directly state gender requirements in job postings, allowing us to identify male- and female-oriented vacancies from explicit gender labels. We therefore use the 2014–2018 postings with explicit gender labels as the training sample for a text classification procedure. Specifically, we train a BERT-based model to predict the gender orientation of vacancies using vacancy-level textual and categorical information, including job descriptions, job titles, recruitment categories, and primary occupation classifications. In a held-out validation sample from the pre-2019 period, the model achieves a prediction accuracy of 98 percent.

We then apply the trained model to postings from 2019–2022, when explicit gender labels are no longer available, and combine the predicted post-ban classifications with the directly observed pre-ban classifications. This procedure yields a consistent firm-year panel of male- and female-oriented vacancies covering 2014–2022, which spans both the pre-reform period and the full implementation of the ASC policy.

3.2 Sample Definition and Descriptive Statistics

Our primary analysis sample consists of mothers with at least one child aged 6–15, as the ASC policy targets children enrolled in compulsory education. In China, compulsory education consists of six years of primary school followed by three years of lower secondary school. Children typically enter primary school at age six and complete compulsory schooling by age fifteen absent grade retention. This age range thus corresponds closely to the population directly affected by the ASC reform.

For the father analyses, we define fathers as the husbands of mothers in the primary analysis sample who can be matched through the CFPS household roster. The father sample is smaller than the mother sample because some mothers are not matched to a coresident husband in the survey, including mothers who are unmarried, divorced, widowed, or whose husbands are absent from the household roster. The difference in sample size therefore reflects the household-based sample construction rather than differential treatment of mothers and fathers in the empirical analysis.

Several outcomes have smaller sample sizes because of CFPS skip patterns and item-specific non-response. For employment status, respondents who are students, retired, or otherwise not working are coded as nonemployed when their labor-force status can be inferred from the survey skip pattern. Similarly, the indicator for nonemployment due to caregiving is defined for the full employment-status sample. Observations with genuinely missing employment status are excluded. Weekly working hours, annual income, formal employment, and commuting time have smaller samples because these variables are either conditionally reported or subject to item non-response. We therefore report outcome-specific sample sizes in all tables and assess the robustness of our results to item non-response using inverse-probability weighting (Wooldridge, 2007).

Table 1 presents descriptive statistics for the main maternal variables. The average mother in our sample is 37.3 years old. Maternal employment is relatively high: 81.2 percent of mothers are employed, working an average of 34.2 hours per week and earning 16,352 yuan annually. Mothers also perform substantial unpaid labor, spending 2.5 hours per weekday and 2.8 hours per weekend day on housework. Notably, 15.6 percent of mothers report

nonemployment due to caregiving or housework responsibilities, underscoring the importance of childcare-related constraints in this population. Table A2 reports corresponding descriptive statistics for the father sample.

4 Empirical Strategy

4.1 Baseline Specification

We employ a DID framework to estimate the causal effects of the public ASC program on parental labor market outcomes. Identification exploits the staggered rollout of the program across counties between 2017 and 2022. Our approach compares changes in labor market outcomes for parents in counties that have implemented the program with changes in counties that have not yet adopted it, conditional on county and year fixed effects.

Our baseline specification is:

$$Y_{ict} = \beta_1 ASC_{ct} + X_{ict}\gamma + \alpha_c + \lambda_t + \varepsilon_{ict} \quad (1)$$

where Y_{ict} denotes labor market outcomes for individual i residing in county c in year t . We examine three primary outcomes: employment status, weekly hours worked, and annual labor earnings. The key explanatory variable, ASC_{ct} , is an indicator equal to one if county c has implemented the ASC program by year t , and zero otherwise.

The specification includes county fixed effects (α_c) to absorb time-invariant differences across counties, such as persistent economic conditions or cultural norms, and year fixed effects (λ_t) to control for common macroeconomic shocks and secular trends. The vector X_{ict} includes individual-level covariates: age, educational attainment, marital status, number of children and ethnicity. We cluster standard errors at the county level to account for serial correlation and within-county dependence in the error term ε_{ict} (Bertrand et al., 2004).

The coefficient of interest, β_1 , captures the average change in parental labor market outcomes associated with ASC implementation, relative to changes in counties that have not yet adopted the program. Under the identifying assumptions discussed below, β_1 can be interpreted as the average effect of ASC access on parental labor market outcomes.

4.2 Identification Assumptions and Validity Checks

Identification in this staggered-adoption setting requires a series of cohort-by-period

parallel-trends assumptions, rather than a single aggregate parallel-trends condition (Baker et al., 2026). Specifically, for each adoption cohort and period, parental labor market outcomes in treated counties would have followed the same trajectory as those in appropriate not-yet-treated comparison counties absent ASC implementation. The design also requires no anticipation: parents should not adjust their labor supply before the program is implemented in response to future ASC availability.

We assess these assumptions in several ways. First, we estimate event-study specifications to examine whether treated and comparison counties exhibit differential pre-treatment trends. Second, we conduct placebo tests using parents with children younger than six, who are not eligible for ASC. Third, we augment the baseline specification by interacting pre-treatment county characteristics with year fixed effects, allowing counties with different observable attributes to follow distinct time trends (Gentzkow, 2006; Li et al., 2025).

Given the strongly gendered allocation of childcare responsibilities in China, we estimate the baseline specification separately for mothers and fathers. We view fathers as a context-specific placebo group. The ASC policy should primarily relax after-school childcare constraints for mothers, while fathers' labor supply should be less directly affected by school dismissal times. Evidence from the 2016 CFPS wave, which predates the reform, supports this interpretation. As shown in Table A3, among children aged 6–15, mothers are the primary daytime caregiver for 38.7 percent of children, the primary nighttime caregiver for 59.0 percent, and the primary provider of homework support for 55.0 percent. The corresponding shares for fathers are only 3.4 percent, 4.8 percent, and 14.8 percent, respectively. These patterns are consistent with prior evidence on gendered caregiving roles in China (Chen, 2024; Fang & Miao, 2024; Qi & Dong, 2016). Null effects for fathers would therefore support the interpretation that the estimated effects operate through relaxed maternal childcare constraints rather than through broader county-level labor market shocks.

Finally, because conventional two-way fixed effects estimates may be sensitive to heterogeneous treatment effects and non-transparent weighting in staggered designs, we report additional estimators designed for staggered-adoption settings in the robustness section (Baker et al., 2026). These include the estimators proposed by Callaway and Sant'Anna (2021),

Borusyak et al. (2024), and Wooldridge (2025), as well as stacked DID specifications.

To examine the validity of the parallel-trends and no-anticipation assumptions and to trace the dynamic effects of the reform, we estimate the following event-study specification:

$$Y_{ict} = \sum_{k=-4}^4 1(t - \text{PolicyYear}_c = k) \cdot \text{Treatment}_c \delta_k + X_{ict} \gamma + \alpha_c + \lambda_t + \varepsilon_{ict} \quad (2)$$

where PolicyYear_c denotes the implementation year in county c and $1(t - \text{PolicyYear}_c = k)$ indicates that year t is k periods relative to implementation. Treatment_c is an indicator for counties that ever adopt the policy. We bin observations occurring four or more years before implementation into $k=-4$ and those occurring four or more years after into $k=4$ to address sparse support in the tails. The omitted category is $k=-1$, the year immediately preceding adoption.

The coefficients δ_k describe the evolution of parental labor market outcomes before and after ASC implementation. Statistically insignificant coefficients for the pre-treatment periods, $k < -1$, provide evidence against differential pre-trends and anticipatory responses. The post-treatment coefficients, $k \geq 0$, show the timing and persistence of the treatment effects.

Because the CFPS is conducted biennially, individuals are observed at two-year intervals rather than annually, and treatment status may change between survey waves. Accordingly, the event-study coefficients should be interpreted as changes over approximately two-year intervals, rather than as precise annual dynamic effects. The estimates therefore capture medium-run responses to ASC implementation, rather than immediate short-run adjustments.

4.3 Heterogeneity by Gender Norms

To investigate whether the program's labor market effects vary with preexisting gender norms, we extend our baseline specification to include an interaction term:

$$Y_{ict} = \rho_1 \text{ASC}_{ct} + \rho_2 \text{ASC}_{ct} \times \text{Egalitarian}_i + \rho_3 \text{Egalitarian}_i + X_{ict} \gamma + \alpha_c + \lambda_t + \varepsilon_{ict} \quad (3)$$

where Egalitarian_i is an indicator equal to one if individual i holds relatively egalitarian views on gender roles, and zero otherwise. We measure gender attitudes using responses from the 2014 CFPS wave, which predates policy implementation in all sample counties. This timing ensures that our measure reflects pretreatment beliefs rather than attitudes potentially influenced by the reform itself.

The coefficient ρ_1 identifies the average treatment effect for individuals with traditional

gender attitudes, while ρ_2 captures the differential effect for those with egalitarian views. The level term Egalitarian_i controls for baseline differences in labor market outcomes between the two groups. All other specification elements remain as in equation (1), including county and year fixed effects, individual-level controls, and standard errors clustered at the county level.

5 Empirical Results

5.1 Baseline Results

Panel A of Table 2 reports the baseline DID estimates of the ASC policy's impact on maternal labor market outcomes. The estimates indicate statistically significant and economically meaningful increases along multiple dimensions of labor supply. Exposure to the policy raises maternal employment by 4.6 percentage points (5.6 percent relative to the sample mean), increases weekly working hours by 2.5 hours (7.2 percent), and raises annual income by approximately 2,946 yuan (18.0 percent), relative to mothers in counties that had not yet implemented the program.

The magnitude of these effects is broadly comparable to, though toward the upper end of, estimates reported in existing literature. For example, Martínez and Perticarà (2017) find that access to ASC in Chile increases female employment by 4.7 percent and hourly earnings by 16.8 percent. He et al. (2024) find that a subsidized before- and after-school care reform in Quebec, Canada increases maternal employment income by 11 percent ten years after implementation. In contrast, studies in Switzerland (Felfe et al., 2016) and Germany (Dehos & Paul, 2023) report more limited labor supply responses. One potential explanation for the relatively large estimates in our setting is the design of China's program, which explicitly mandates on-campus homework completion. By shifting a time-intensive academic task from households to schools, the policy may generate larger time savings for mothers than enrichment-oriented or voluntary ASC programs studied elsewhere.

Panel B of Table 2 reports the corresponding estimates for fathers using the same empirical specification. In contrast to the results for mothers, we find no statistically significant effects on fathers' employment, working hours, or earnings, and the point estimates are small. This pattern is consistent with the context-specific placebo interpretation discussed in Section 4:

after-school childcare and homework supervision in China are borne primarily by mothers, while fathers' labor supply is less directly constrained by school dismissal times. It is also consistent with recent evidence that child penalties in China fall mainly on mothers rather than fathers (Huang et al., 2025; Meng et al., 2023; Zhang et al., 2024). The null effects for fathers therefore reinforce the childcare-constraint mechanism and help distinguish our findings from broader local labor-market shocks, which would likely affect both mothers and fathers.

Figure 2 provides complementary evidence supporting our identification strategy. The event-study estimates trace the evolution of parental labor market outcomes before and after ASC implementation, with 95 percent confidence intervals reported throughout. Because the CFPS is conducted biennially, these coefficients should be interpreted as medium-run responses over roughly two-year intervals rather than as precise annual dynamics. For mothers, the pre-implementation coefficients are small and statistically indistinguishable from zero, and joint tests fail to reject the null that these coefficients are jointly zero. These patterns provide evidence against differential pre-trends and anticipatory labor-supply responses before ASC implementation. After implementation, the coefficients become positive and statistically significant for employment, working hours, and earnings, with effects that increase over time. Consistent with the baseline estimates, we find no comparable changes for fathers. This pattern supports the interpretation that ASC primarily affects maternal labor supply by relaxing after-school childcare constraints, rather than by capturing broader county-level labor market trends.

5.2 Robustness Checks

In this subsection, we perform a range of robustness tests to assess the validity of our main findings regarding the impact of the ASC program on maternal labor market outcomes. These include checks for the parallel trends assumption crucial for our DID strategy, placebo tests using alternative populations, controlling for potentially confounding concurrent policies, addressing concerns about sample selection, and employing alternative DID estimators robust to staggered adoption and heterogeneous treatment effects. Our main results remain robust and consistent across these specifications.

5.2.1 Parallel Trends Assumption

The credibility of our DID design hinges on the parallel trends assumption—that, absent the reform, labor market outcomes in early-adopting counties would have evolved similarly to those in later-adopting or non-adopting counties. The event-study evidence presented above offers initial support for this assumption, showing no systematic pretreatment differences between treated and untreated counties. The absence of detectable effects for fathers—who bear substantially lower childcare responsibilities—provides further reassurance that our estimates are not driven by spurious county-level trends.

We strengthen this evidence through three additional robustness exercises. First, we examine whether the reform influenced individual characteristics that should be orthogonal to the policy, including years of schooling, marital status, minority ethnicity, religious affiliation, and urban hukou status. As reported in Table A4, the estimated coefficients are uniformly small and statistically indistinguishable from zero, suggesting that our findings reflect the causal impact of the ASC reform rather than confounding demographic changes.

Second, we conduct placebo tests using an ineligible population. Because the ASC program targets families with school-aged children, mothers whose youngest child is under age six are unlikely to be directly affected. Consistent with this expectation, applying our main specification to this group yields no significant effects on employment, hours worked, or earnings (Panel A of Table A5).

Third, to address concerns about heterogeneous trends correlated with adoption timing, we augment the baseline specification by interacting year fixed effects with a set of pre-reform county characteristics measured in 2016. As shown in Panel B of Table A5, the estimated treatment effects remain stable in magnitude and statistically significant, indicating that our results are not driven by differential trends associated with observable county attributes.

5.2.2 Staggered DID Issues

Beyond the assessment of pre-treatment trends, staggered treatment adoption raises additional concerns about the interpretation of conventional TWFE estimates. Recent econometric work shows that conventional two-way fixed effects (TWFE) DID estimators may be sensitive to treatment-effect heterogeneity and non-transparent weighting when treatment

timing varies across units (Baker et al., 2026; De Chaisemartin & d'Haultfoeuille, 2023). We therefore reestimate our main effects using several alternative estimators designed for staggered-adoption settings.

First, we implement the group-time average treatment effect estimator proposed by (Callaway & Sant'Anna, 2021). This estimator compares each treated cohort with appropriate comparison units, such as not-yet-treated counties, and then aggregates cohort-specific treatment effects. Second, we apply the imputation-based estimator proposed by Borusyak et al. (2024), which constructs counterfactual outcomes for treated units using information from counties that have not yet implemented the ASC policy in a given year. Third, we implement the stacked DID approach of Cengiz et al. (2019), reorganizing the data by policy adoption year. For each implementation year, we construct an event-specific sample that compares counties newly adopting the policy in that year to counties remaining untreated throughout the corresponding pre- and post-adoption window, thereby isolating variation driven by differences in adoption timing. Fourth, we adopt the extended TWFE estimator developed by Wooldridge (2025), which saturates the regression with interactions between implementation-year indicators and calendar-year indicators, allowing treatment effects to vary flexibly across adoption years and over time.

As reported in Panels C–F of Table A5, the estimates obtained from these alternative approaches are closely aligned with our baseline results in both magnitude and statistical significance. This consistency suggests that our findings are robust to concerns about staggered adoption and treatment effect heterogeneity, and that potential biases associated with conventional TWFE estimation are unlikely to drive our conclusions.

5.2.3 Other Confounding Factors

Our identification strategy requires that the staggered rollout of the ASC program is not confounded by other contemporaneous reforms or shocks that also affected maternal labor supply. We therefore examine three potentially relevant sources of confounding: maternity-leave extensions, long-term care insurance, and COVID-19 disruptions.

First, provincial maternity-leave extensions may have affected mothers' labor market outcomes during our sample period. These reforms were implemented unevenly across

provinces, with two main waves: one following the universal two-child policy in 2016 and another following the three-child policy and related family-support measures in 2021–2022. Provinces differed in both the timing of reform and the number of additional leave days provided. To account for these changes, we compile province-year information on statutory maternity-leave duration and include the number of maternity-leave days as a control in the baseline specification. As reported in Panel G of Table A5, the estimated effects of ASC remain very similar in magnitude and statistical significance after adding this control.

Second, the rollout of long-term care insurance may also affect maternal labor supply by reducing household eldercare responsibilities. China introduced long-term care insurance through phased city-level pilots, beginning with the first national pilot program in 2016 and expanding to additional areas in 2020. Because these programs provide formal care services or financial support to eligible individuals with functional limitations, they may relax eldercare constraints within households. We therefore construct a county-level indicator for the implementation of long-term care insurance and include it as an additional control. As shown in Panel H of Table A5, the ASC estimates remain stable after controlling for the rollout of this policy.

Third, the ASC rollout partly overlaps with the COVID-19 period, during which school closures, workplace restrictions, and local lockdowns may have affected parental labor supply (He et al., 2025; Qiu et al., 2020). We address this concern in two ways. First, we re-estimate the baseline specification using only the pre-COVID period from 2014 to 2019, thereby excluding observations potentially affected by pandemic-era disruptions.⁴ The estimates reported in Panel I of Table A5 remain similar to the baseline results in sign and magnitude. Second, using the full sample, we control for local pandemic exposure with the logarithm of one plus cumulative confirmed COVID-19 cases in the mother's city during 2020–2022. The estimates reported in Panel J of Table A5 remain stable after including this control.

⁴ We define the pre-COVID period as ending in 2019 because major COVID-related disruptions in China began in January 2020, when Wuhan was placed under lockdown and strict travel and public-health restrictions were introduced. Although the first cases were reported in late 2019, these developments occurred after the main 2019 survey period and are unlikely to affect the 2019 CFPS outcomes.

5.2.4 Sample Selection, Migration, and Item Non-Response

A potential concern is that ASC availability may induce selective migration. For example, mothers seeking better labor market opportunities may move into counties that implement the program, which would bias the estimates upward. Conversely, measurement error may arise if mothers move out of treated counties after implementation but are still coded according to their baseline county of residence.

To assess the extent of such concerns, we first examine changes in population composition. As reported in Table A4, we find no evidence of significant shifts in observable demographic characteristics following program implementation, suggesting limited influence of migration on our main results. To further probe this issue, we re-estimate the baseline specification among mothers who remain in the same county throughout the observed sample period. The results, presented in Panel K of Table A5, remain highly consistent with our baseline findings, lending additional support to the robustness of our identification strategy.

We also address item non-response for outcomes with smaller samples. Weekly working hours and annual income are either conditionally reported or subject to item non-response. Following (Wooldridge, 2007), we implement an inverse-probability-weighted robustness check. For each outcome, we first estimate the probability that the outcome is observed using baseline demographic characteristics, household characteristics, county fixed effects, and year fixed effects. We then re-estimate the baseline specification on the nonmissing sample, weighting each observation by the inverse of its predicted response probability. The results, reported in Table A6, remain similar to the baseline estimates in sign, magnitude, and statistical significance. This pattern suggests that item non-response is unlikely to drive our main findings.

5.2.5 Additional Estimates and Specification Checks

We conduct additional analyses to examine whether the estimated effects depend on specific features of the ASC reform. First, we estimate separate effects across the three institutional phases of the reform. The ASC policy evolved from an advisory phase in 2017–2018 to a non-mandatory expansion phase in 2019–2020, and then to a mandatory phase in 2021–2022. Because these phases differed in policy intensity and implementation incentives, the pooled baseline estimate may mask heterogeneity across stages of the reform.

To examine this issue, we estimate phase-specific effects using an approach conceptually similar to a stacked DID design. For each treatment cohort, we exclude counties treated in earlier phases and censor later-treated comparison counties when they adopt ASC. This procedure avoids using already-treated counties as controls. Table A7 reports positive employment and working-hour estimates across all three phases, with statistically significant increases in the non-mandatory expansion phase. The income estimates, however, differ substantially: 11,369 yuan in the advisory phase, 3,072 yuan in the non-mandatory expansion phase, and a statistically insignificant 393 yuan in the mandatory phase.

Two findings help contextualize the large advisory-phase income estimate. First, mean annual maternal income over the sample period is higher in advisory-phase counties: 41,440 yuan, compared with 18,600 yuan in non-mandatory-phase counties and 14,014 yuan in mandatory-phase counties. The advisory-phase estimate thus corresponds to approximately 27.4 percent of its cohort-specific mean income. Second, Table A8 shows that the formal-employment response is largest in advisory-phase counties, at 19.6 percentage points, compared with 8.7 percentage points in non-mandatory-phase counties and a small, statistically insignificant estimate in mandatory-phase counties. This pattern is consistent with changes in employment composition contributing to earnings gains. A shift toward formal jobs could raise annual income without comparably large increases in overall employment or weekly working hours.

The advisory-phase income estimate is also sensitive to the tails of the income distribution. As Table A9 shows, winsorizing income at the 1st and 99th percentiles reduces the estimate to 10,280 yuan. Winsorizing at the 5th and 95th percentiles reduces it further to 5,076 yuan, approximately 55 percent below the original estimate, although it remains statistically significant at the 5 percent level. These findings support a positive income response but indicate substantial sensitivity in its magnitude. The phase-specific regressions also rely on more limited identifying variation than the baseline analysis. We therefore interpret the size of the advisory-phase income effect and comparisons across phases cautiously.

Second, we assess whether the baseline results are disproportionately driven by any single phase of the reform. We sequentially exclude each phase and re-estimate the baseline

specification. Specifically, we re-estimate the model after excluding the advisory phase, the non-mandatory phase, and the mandatory phase, respectively. The results, reported in Table A10, remain qualitatively similar across these alternative samples. The estimated effects continue to point to increases in maternal employment, working hours, and annual income, although statistical significance varies across outcomes and samples. These results suggest that the main findings are robust to excluding any single stage of policy implementation.

Finally, we examine whether the effects differ by children's age. The ASC program should be more consequential for mothers of younger school-age children, who are more likely to require adult supervision during after-school hours. We therefore distinguish mothers with children aged 6–10 from those with children aged 11–15 and estimate whether the treatment effects are larger for the younger-child group. The results, reported in Table A11, show that the effects of ASC are larger for mothers with children aged 6–10 than for mothers with children aged 11–15. This pattern is consistent with the interpretation that ASC increases maternal labor supply primarily by relaxing after-school childcare constraints.

5.3 Mechanism Analysis

5.3.1 Supply-Side Responses

To explore the channels underlying the positive maternal labor supply responses documented above, we examine how the ASC policy affects mothers' time allocation and employment characteristics. A central hypothesis is that the policy relaxes binding time constraints created by the misalignment between school dismissal times and standard work schedules, thereby lowering the effective cost of maternal employment.

We first assess this hypothesis by analyzing mothers' time spent on housework, separately for weekdays and weekends. Columns (1) and (2) of Table 3 show that exposure to the ASC program significantly reduces weekday housework, with mothers spending 0.18 fewer hours per day on weekdays. In contrast, we find no statistically significant effect on weekend housework. This pattern aligns with the institutional design of the policy, which provides childcare only on weekday afternoons, and suggests that the time savings are concentrated in periods directly affected by the reform rather than reflecting a general reallocation of household

labor. Consistent with this interpretation, Column (3) of Table 3 shows a 4.8 percentage point reduction in nonemployment due to caregiving responsibilities.⁵

We next examine whether the relaxation of childcare constraints is associated with changes in mothers' access to employment opportunities. Prior work shows that caregiving responsibilities can restrict mothers' job search radius and spatial access to work (Borghorst et al., 2024; Bütikofer et al., 2025; Le Barbanchon et al., 2021). We use one-way commuting time as a proxy for spatial labor market access. Column (4) of Table 3 shows that the ASC policy increases mothers' one-way commuting time by 2.24 minutes, or 13.5 percent relative to the baseline mean of 16.59 minutes.⁶

We further examine whether the reform is associated with changes in employment arrangements. Column (5) of Table 3 shows that exposure to the ASC policy increases the probability that mothers hold formal employment—defined as jobs providing the full set of four social insurance programs and the housing provident fund—by 9.7 percentage points. In the Chinese labor market, access to this full bundle of social benefits is commonly viewed as an important dimension of job quality, reflecting more stable employment relationships and greater social protection. Together with the increase in commuting time, this evidence suggests that the policy facilitates mothers' access to a broader range of employment opportunities that were previously difficult to reconcile with childcare responsibilities.⁷

5.3.2 Demand-Side Responses

The preceding results suggest that the ASC policy relaxes mothers' childcare constraints and increases their access to jobs requiring regular schedules, longer commutes, and more stable employment arrangements. We next examine whether firms also adjust their recruitment

⁵ We also examine whether ASC changes the source of after-school supervision. Table A12 shows that the ASC policy reduces maternal involvement in homework support, while father-only and relatives-only arrangements do not change significantly. This pattern suggests that ASC primarily substitutes for mothers' after-school educational supervision rather than shifting responsibilities to fathers or other relatives. Consistent with this interpretation, supplementary estimates using CFPS parental-time measures from the 2020 and 2022 waves show that the ASC policy reduces mothers' childcare time, while fathers' childcare time does not increase significantly (Table A13). Because these parental-time measures are available only in later waves, we interpret them cautiously.

⁶ Although the average increase in commuting time is modest in absolute terms, it masks heterogeneity across the commuting-time distribution. Unconditional quantile regressions reported in Table A14 show little change at lower percentiles but larger effects at higher percentiles. This pattern suggests that the ASC policy does not uniformly increase commuting time for all mothers. Rather, it appears to expand the feasible job-search radius for mothers whose employment opportunities require longer commutes, consistent with the childcare-constraint mechanism.

⁷ The mechanism outcomes also vary in sample size because some variables are conditionally reported or subject to item non-response. As a robustness check, Table A15 reports inverse-probability-weighted estimates for weekday housework, weekend housework, commuting time, and formal employment. The results remain similar to the baseline estimates.

behavior in response to the reform. This analysis complements the household-level evidence by asking whether expanded after-school care affects not only mothers' labor supply decisions, but also employers' demand for female workers.

The ASC policy may affect firm hiring through several channels. Before the reform, employers may perceive mothers of school-age children as facing binding late-afternoon childcare constraints, increasing expected absenteeism, schedule instability, or turnover. By extending supervised care into the late afternoon, the ASC policy may make mothers' labor supply more predictable and reduce the perceived cost of hiring women for jobs requiring fixed schedules, regular workplace presence, or late-afternoon availability. If firms update their expectations in response to improved childcare availability, the ASC policy may increase female-oriented vacancies.

To examine this possibility, we use the job-posting data described in Section 3.1. The analysis uses two samples: a 2014–2018 sample in which gender orientation is directly observed from explicit gender labels, and an extended 2014–2022 sample in which post-2018 gender orientation is predicted using the text-classification procedure described in the data section. We estimate firm-level regressions of the following form:

$$Y_{fct} = \beta_1 ASC_{ct} + \alpha_c + \lambda_t + \varepsilon_{fct} \quad (4)$$

where Y_{fct} denotes the number of gender-oriented vacancies posted by firm f , located in county c , in year t . The key explanatory variable, ASC_{ct} , equals one if county c has implemented the ASC policy by year t . The specification includes firm fixed effects, which absorb time-invariant differences in firms' recruitment practices, and year fixed effects, which control for aggregate labor-market conditions and national regulatory changes. Standard errors are clustered at the county level.

Table 4 reports the baseline firm-level estimates. Across both the directly observed and extended samples, ASC implementation increases female-oriented vacancies, while the effects on male-oriented vacancies are small and statistically insignificant. The estimates are similar across the two samples, suggesting that the results are not driven by the shorter pre-2019 period. This pattern is consistent with firms expanding recruitment toward female workers after the relaxation of after-school childcare constraints.

We conduct several additional analyses to assess whether these patterns reflect demand-side adjustment rather than pre-existing local labor-demand shocks. As discussed in the background section, pre-reform labor-demand measures do not significantly predict ASC adoption timing in the survival analysis reported in Table A1.⁸ We also estimate an event-study specification for female-oriented vacancies. Figure A1 shows that the pre-treatment coefficients are small and statistically indistinguishable from zero, and the joint test fails to reject parallel pre-trends. These results reduce concerns that the increase in female-oriented vacancies reflects differential pre-existing labor-demand trends.

We next examine whether the vacancy response is stronger in jobs where mothers' improved late-afternoon availability is likely to matter more. Columns (1) and (2) of Table 5 classify vacancies by occupation. The increase in female-oriented vacancies is concentrated in presence-intensive service and administrative occupations, including retail, catering, sales, customer service, and clerical positions. By contrast, the effect is smaller and statistically insignificant for professional or less directly school-pickup-constrained occupations, such as IT/software, online work, design, and consulting.

Columns (3) and (4) classify vacancies using schedule-related keywords in the job-posting text. The ASC policy increases female-oriented vacancies among postings that explicitly mention standard daytime schedules or regular workplace presence, while the effect for flexible or remote postings is smaller and statistically insignificant. Although occupation categories and schedule-related keywords are imperfect proxies for actual work schedules, the consistency between the occupation-based and text-based results supports the interpretation that firms respond more strongly in jobs where mothers' increased late-afternoon availability is economically relevant.

Finally, we examine whether the demand-side response extends to women who are unlikely to have school-age children. Because the job-posting data do not report applicants' fertility status, we use vacancies targeting female candidates under age 26 as a proxy for this group. This cutoff is motivated by China's legal minimum marriage age for women of 20 and

⁸ The reform was initiated and guided by higher-level education authorities, while local implementation was organized primarily through local education bureaus and public schools. Because adoption decisions were made within the education system rather than through employer participation, firms had limited scope to directly influence implementation timing through labor demand or lobbying.

by the fact that the children directly affected by ASC are aged 6–15. Women under age 26 are therefore unlikely to have children in the age range covered by the policy. Table A16 shows no statistically significant effect of ASC on vacancies targeting women under 26. This result suggests that the increase in female-oriented vacancies is unlikely to reflect a general expansion in demand for all female workers. Instead, it is more consistent with firms responding to the improved availability of women likely to have school-age children.

Taken together, the evidence suggests that the ASC policy affects firms' recruitment behavior in addition to mothers' labor supply. We interpret the vacancy results as suggestive evidence of demand-side adjustment rather than as a full general-equilibrium estimate. The household-level results further support this interpretation: the ASC policy increases mothers' employment, hours, and earnings, but has little effect on fathers' labor-market outcomes. This gender asymmetry is difficult to reconcile with a broad local labor-demand shock alone and is more consistent with a childcare-constraint mechanism operating through the gendered allocation of after-school care.

5.4 Role of Gender Norms

We next examine whether the labor market effects of the ASC policy vary with gender role attitudes within the household. We interpret the GSI as a proxy for the relative weight placed on home production versus market work. This interpretation is consistent with the conceptual framework in Section 2, where mothers with more traditional gender attitudes place greater value on non-market time, including childcare, housework, and family responsibilities, while mothers with more egalitarian attitudes are more willing to reallocate time toward market work when childcare constraints are relaxed. The GSI may also be viewed as a proxy for the intrahousehold sharing rule that governs how mothers' time is allocated between home production and paid employment.

Under this framework, the effect of the ASC policy should be larger for mothers with more egalitarian attitudes. The policy relaxes a binding after-school childcare constraint and makes formal or more stable employment more feasible. Mothers with egalitarian views are more likely to use this newly available time to increase labor supply, accept jobs requiring longer or more regular hours, and shift from home production to market work. By contrast, mothers with

more traditional views may continue to prioritize domestic responsibilities even after the childcare constraint is relaxed, leading to smaller labor market responses.

Panel A of Table 6 supports this prediction. The interaction terms between ASC exposure and egalitarian maternal attitudes are positive and statistically significant for employment, weekly working hours, and annual income. In contrast, the estimated effects for mothers with traditional attitudes are small and statistically insignificant. These results suggest that childcare provision relaxes an important time constraint, but the extent to which mothers translate this relaxation into market work depends on their own preferences and gender-role attitudes.

Panel B examines whether husbands' gender attitudes moderate mothers' labor market responses. The interaction terms involving husbands' egalitarian attitudes are small and statistically insignificant across outcomes. This pattern suggests that mothers' own attitudes and constraints are more directly related to the marginal allocation of their time after ASC becomes available. We do not interpret this as ruling out household bargaining, but rather as indicating that mothers' own preferences appear to be the more relevant margin in this setting.

We further examine whether the mechanisms vary by maternal gender norms. Table 7 shows that the reduction in weekday housework and the increase in formal employment are significantly larger among mothers with more egalitarian attitudes, while there is no corresponding differential effect on weekend housework. This pattern is consistent with the household production interpretation: when the ASC policy relaxes a weekday-specific childcare constraint, egalitarian mothers are more likely to shift time away from home production and toward higher-quality market work. The absence of a differential effect on weekend housework further supports the interpretation that the policy operates through after-school time constraints rather than through a general change in household production preferences.

6 Conclusions

Work–family conflict remains a central challenge for working parents worldwide, driven in part by institutional misalignments between school schedules and standard work hours (Stier et al., 2012). These frictions disproportionately affect mothers, who continue to bear primary

responsibility for childcare in many settings. This paper examines whether expanding access to ASC can mitigate these constraints and improve maternal labor market outcomes.

Exploiting the staggered rollout of China's ASC policy, we find that local ASC availability increases maternal employment, weekly work hours, and annual income, with no corresponding effects on fathers' labor market outcomes. Consistent with a time-constraint mechanism, the policy reduces mothers' weekday housework, increases commuting time, and raises the likelihood of formal employment. We also document an increase in job vacancies targeting female workers, suggesting adjustments on the labor demand side.

Importantly, policy effects are highly heterogeneous with respect to gender role attitudes within the household. Maternal labor supply responses are substantially larger among women with more egalitarian gender views, while mothers with traditional attitudes exhibit negligible responses. In contrast, husbands' gender attitudes do not significantly moderate the effects. These findings indicate that while childcare provision relaxes an important structural constraint, the extent to which mothers translate time savings into market work depends critically on individual beliefs and social norms.

Our results suggest that expanding childcare access can effectively support maternal employment where families face binding time constraints. However, policy effectiveness may be limited where traditional gender norms remain strong. In such contexts, childcare policies may require complementary efforts to shift social norms around caregiving and paid work. Understanding how formal policy interventions interact with informal normative constraints remains an important avenue for future research.

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Figures and Tables

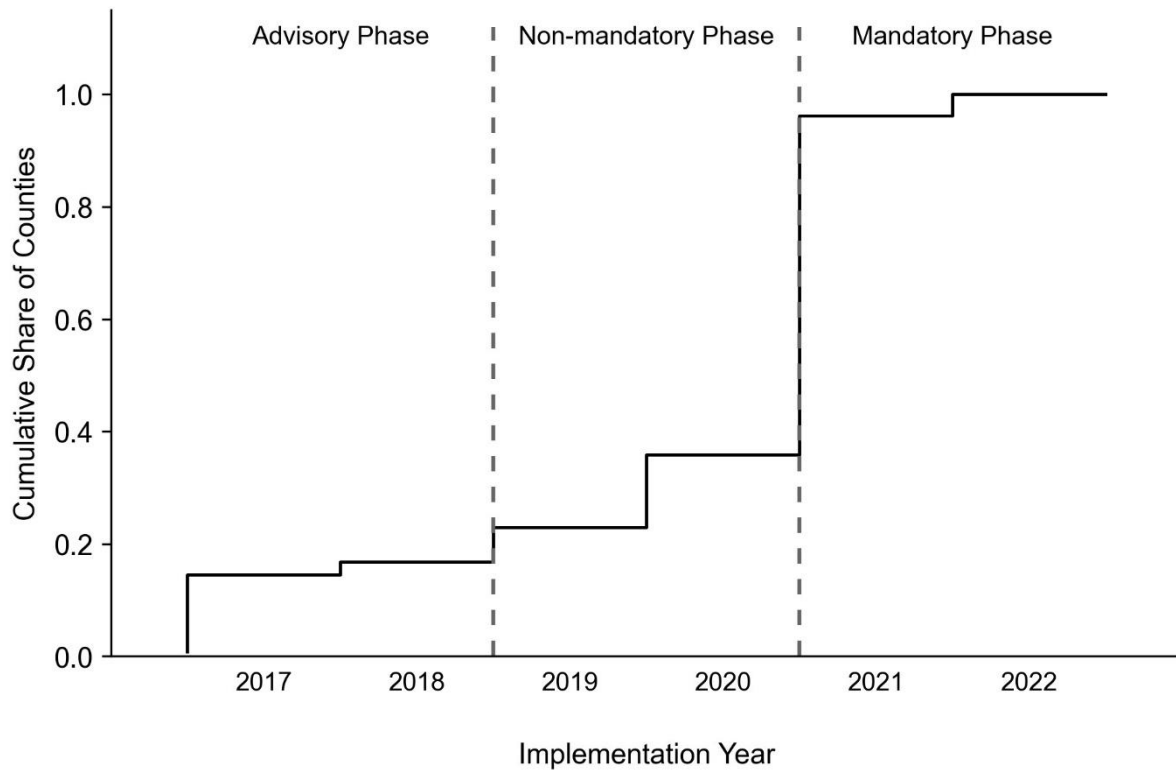


Figure 1 Cumulative Adoption of ASC across Counties

Notes: This figure plots the cumulative share of CFPS sample counties that had adopted ASC by implementation year. Dashed vertical lines separate the advisory phase, the non-mandatory expansion phase, and the mandatory phase of the reform.

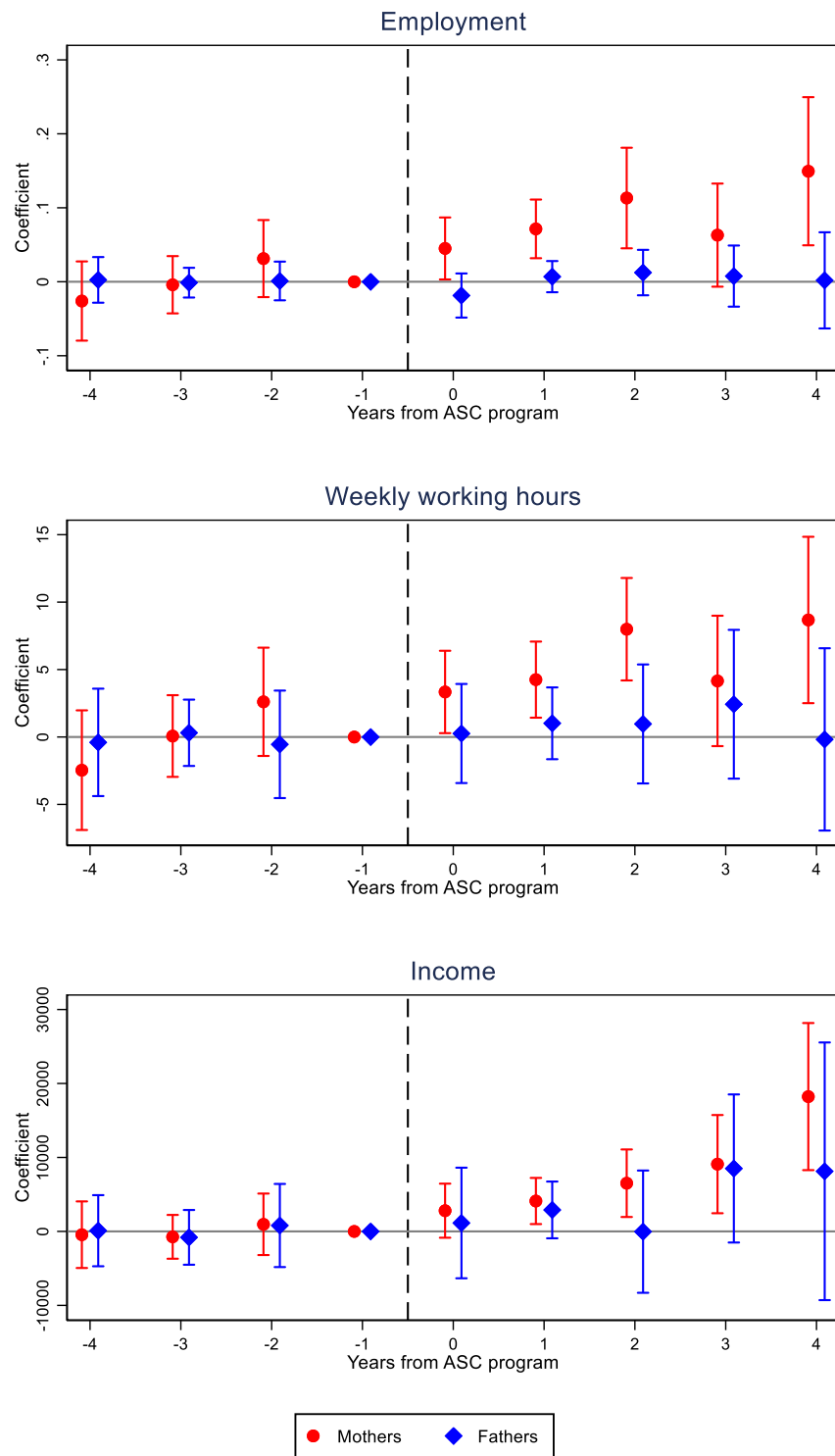


Figure 2 Event-Study Estimates for Baseline Labor-Market Outcomes

Notes: This figure depicts the coefficient estimates and associated 95% confidence intervals from the event study model outlined in Equation (2). The dependent variables are employment, weekly working hours, and annual income. Annual income is measured in yuan. The omitted category is the year preceding ASC program implementation. Control variables include maternal age, years of education, number of children, marital status, and ethnicity. All specifications include county fixed effects and year fixed effects. Standard errors are clustered at the county level.

Table 1 Summary Statistics

VARIABLES	(1) Observations	(2) Mean	(3) SD	(4) Minimum	(5) Maximum
Panel A CFPS maternal variables					
Employment	12,462	0.812	0.391	0	1
Annual income	8,127	16,352	23,828	0	190,000
Weekly working hours	10,514	34.19	26.58	0	98
Formal employment	10,116	0.110	0.313	0	1
Housework hours per day on weekday	12,317	2.515	1.825	0	10
Housework hours per day on weekend	12,310	2.791	1.814	0	10
Nonemployment due to caregiving	12,462	0.156	0.363	0	1
One-way commuting time in minutes	6,577	16.59	17.77	0	180
Egalitarian gender views	10,454	0.445	0.497	0	1
ASC	12,462	0.177	0.381	0	1
Survey year	12,462	2018	2.756	2014	2022
Age	12,462	37.34	6.172	20	76
Years of education	12,462	8.057	4.456	0	22
Married	12,462	0.971	0.167	0	1
Han ethnicity	12,462	0.677	0.468	0	1
Number of children	12,462	1.558	0.708	1	7
Panel B Job vacancies variables					
Male job vacancy count	14,078	3.868	14.09	0	475
Female job vacancy count	14,078	2.761	9.989	0	262

Notes: Panel A reports individual-level data from the 2014–2022 waves of the CFPS for mothers with children aged 6–15.

Panel B reports firm-year observations from the job-posting panel, which covers vacancies posted during 2014–2018 by publicly listed firms and their affiliates on six major Chinese online recruitment platforms: 51job, BOSS Zhipin, Zhaopin, Liepin, Lagou, and Kanzhun. Vacancy outcomes are measured as counts of male- or female-oriented job postings by a firm in a given year. Because the job-posting panel is separate from the CFPS household sample, the number of firm-year observations in Panel B is not comparable to the number of mother-year observations in Panel A. Firms appearing only once in the vacancy panel are excluded to align the descriptive statistics with the fixed-effects regression sample.

Table 2 Parental Labor Market Outcomes

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
Panel A Mothers			
ASC	0.046*** (0.016)	2.466** (1.110)	2,946.075** (1,289.632)
Observations	12,462	10,514	8,127
R-squared	0.103	0.089	0.381
Mean of Y	0.812	34.19	16,352
Panel B Fathers			
ASC	0.001 (0.008)	0.362 (1.207)	2,556.797 (1,902.807)
Observations	9,904	7,931	6,338
R-squared	0.033	0.074	0.321
Mean of Y	0.961	49.14	31,480
Year FE	YES	YES	YES
County FE	YES	YES	YES

Notes: This table reports the effects of the ASC policy on parental labor market outcomes. Panel A presents results for mothers with children aged 6–15, and Panel B presents corresponding results for their husbands. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level.

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

Table 3 Time Allocation and Employment Characteristics

VARIABLES	(1) Weekday housework	(2) Weekend housework	(3) Non-employment for caregiving	(4) One-way commuting time	(5) Formal employment
ASC	-0.182** (0.086)	-0.036 (0.090)	-0.048*** (0.014)	2.240** (0.952)	0.097*** (0.025)
Observations	12,317	12,310	12,462	6,577	10,116
R-squared	0.130	0.091	0.113	0.164	0.293
Mean of Y	2.515	2.791	0.156	16.59	0.110
Year FE	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES

Notes: This table reports the effects of the ASC policy on time allocation and employment characteristics of mothers with children aged 6–15. The dependent variables in Columns (1)–(5) are weekday housework hours per day, weekend housework hours per day, an indicator for nonemployment due to caregiving, one-way commuting time in minutes, and an indicator for formal employment, respectively. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 4 Gender-Specific Job Vacancies

Sample	2014-2018		2014-2022	
	(1)	(2)	(3)	(4)
VARIABLES	Male vacancies	Female vacancies	Male vacancies	Female vacancies
ASC	-0.219 (0.231)	0.451** (0.207)	-0.052 (0.283)	0.336** (0.151)
Observations	14,078	14,078	32,228	32,228
R-squared	0.805	0.756	0.507	0.453
Year FE	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Mean of Y	3.868	2.761	4.258	1.534

Notes: This table reports the effects of the ASC policy on gender-oriented job vacancies posted by publicly listed firms and their affiliates. Vacancy data are collected from major Chinese online recruitment platforms and aggregated to the firm-year level. Vacancy outcomes are measured as counts of job postings by a firm in a given year. Columns (1) and (2) use the 2014–2018 sample, for which gender orientation is directly observed from explicit gender labels in job postings. Columns (3) and (4) use the extended 2014–2022 sample. Because explicit gender preferences in job advertisements were prohibited nationwide after 2018, gender orientation for vacancies posted during 2019–2022 is predicted using a BERT-based text-classification model trained on 2014–2018 postings with explicit gender labels. The dependent variables are male-oriented vacancies in Columns (1) and (3) and female-oriented vacancies in Columns (2) and (4). All specifications include firm and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 5 Female-Oriented Vacancy Responses by Job Type and Schedule Requirements

VARIABLES	(1) Presence-intensive service/admin	(2) Professional or less constrained	(3) Standard-hours / presence-coded	(4) Flexible / remote-coded
ASC	0.518*** (0.145)	-0.624 (1.017)	0.113*** (0.043)	0.014 (0.015)
Observations	32,228	32,228	32,228	32,228
R-squared	0.328	0.144	0.321	0.294
Year FE	YES	YES	YES	YES
Firm FE	YES	YES	YES	YES
Mean of Y	0.837	1.798	0.162	0.0184

Notes: This table reports the effects of the ASC policy on female-oriented vacancies by job type and schedule requirements. The unit of observation is the firm-year, and vacancy outcomes are measured as counts of female-oriented job postings. The analysis uses the extended 2014–2022 job-posting panel; gender orientation for 2019–2022 vacancies is predicted using the text-classification procedure described in the data section. Columns (1) and (2) classify vacancies by occupation. Presence-intensive service and administrative occupations include retail, catering, sales, customer service, and clerical positions. Professional or less directly school-pickup-constrained occupations include IT/software, online work, design, and consulting. Columns (3) and (4) classify vacancies using schedule-related keywords in the job-posting text. Standard-hours / late-afternoon presence vacancies are postings that mention standard daytime schedules or regular workplace presence, including "9:00 a.m.–6:00 p.m.," "9-to-6 schedule," "day shift," "regular office schedule," and "fixed working hours." Flexible / remote vacancies are postings that mention flexible, remote, or output-based arrangements, including "flexible work," "flexible working hours," "remote work," "work from home," "flexible schedule," and "project-based work." All specifications include firm and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. ***p<0.01, **p<0.05, *p<0.10.

Table 6 Labor Market Outcomes and Gender Role Attitudes

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
Panel A Mother's Gender Attitudes			
ASC * Egalitarian mother	0.040** (0.019)	2.917** (1.363)	3,501.733** (1,574.438)
Egalitarian Mother	0.010 (0.010)	0.225 (0.766)	880.766 (561.324)
ASC	0.021 (0.019)	0.330 (1.431)	1,289.716 (1,406.203)
Observations	10,453	8,676	6,618
R-squared	0.105	0.099	0.401
Panel B Father's Gender Attitudes			
ASC * Egalitarian father	0.001 (0.011)	-2.146 (1.510)	-1,029.341 (2,677.196)
Egalitarian Father	-0.002 (0.005)	0.086 (0.821)	-958.302 (796.000)
ASC	0.001 (0.011)	0.705 (1.539)	5,050.948* (2,747.079)
Observations	8,012	6,254	4,945
R-squared	0.042	0.090	0.332
Year FE	YES	YES	YES
County FE	YES	YES	YES

Notes: This table reports heterogeneous treatment effects of the ASC policy by household gender role attitudes for mothers with at least one child aged 6–15. Panel A examines whether mothers' own gender attitudes moderate the policy's effects on maternal labor market outcomes. Panel B examines whether husbands' gender attitudes moderate the policy's effects on their wives' labor market outcomes. Egalitarian Mother (Panel A) is an indicator equal to one if the mother's Gender Stereotype Index (GSI) is below zero, reflecting more egalitarian views, and zero otherwise. Egalitarian Father (Panel B) is defined analogously based on the husband's GSI. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. ***p<0.01, **p<0.05, *p<0.10.

Table 7 Time Allocation, Employment Characteristics, and Gender Role Attitudes

VARIABLES	(1) Weekday housework	(2) Weekend housework	(3) Non-employment for caregiving	(4) One-way commuting time	(5) Formal employment
ASC * Egalitarian mother	-0.161** (0.076)	-0.033 (0.092)	-0.033** (0.016)	2.596** (1.144)	0.047** (0.022)
ASC	-0.090 (0.094)	0.024 (0.100)	-0.025 (0.018)	1.412 (1.188)	0.070*** (0.024)
Egalitarian mother	-0.081 (0.051)	-0.083* (0.048)	-0.008 (0.008)	0.779 (0.762)	0.004 (0.007)
Observations	10,382	10,377	10,453	5,411	8,624
R-squared	0.137	0.096	0.113	0.189	0.285
Year FE	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES

Notes: This table reports heterogeneous treatment effects of the ASC policy by mothers' gender role attitudes for mothers with at least one child aged 6–15. The dependent variables in Columns (1) to (5) are weekday housework hours per day, weekend housework hours per day, an indicator for nonemployment due to caregiving, one-way commuting time in minutes, and an indicator for formal employment status, respectively. Egalitarian Mother is an indicator equal to one if the mother's GSI is below zero, reflecting more egalitarian views, and zero otherwise. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. ***p<0.01, **p<0.05, *p<0.10.

Appendix A: Theoretical Derivations

A1 Optimization Problem

For each feasible sector $j \in \{0,1\}$, the mother solves

$$V_j(A; \phi) = \max_{m,h,c,l} \{u(w_j m + y) + \phi[v(h) + \psi(c + \gamma A)] + g(l)\},$$

subject to $m + h + c + l = T$, nonnegativity of all time inputs, and

$$\mathcal{M}_0(A) = [0, \tau(A)], \quad \mathcal{M}_1(A) = [\bar{m}, \tau(A)].$$

The formal sector is infeasible when $\tau(A) < \bar{m}$. Within each feasible sector, the objective is concave and the constraints are linear. We maintain the curvature assumptions in Section 2.2 and use interior nonmarket allocations for the strict comparative statics below. For local derivatives, we extend A continuously over $[0,1]$ and take $\tau(A)$ to be differentiable and nondecreasing. Discrete changes between $A = 0$ and $A = 1$ are considered in A6.

A2 First-Order Conditions

Let λ_j be the multiplier on the time constraint. Let α_j and μ_j be the nonnegative multipliers on the lower and upper bounds on market hours, respectively, with $\underline{m}_0 = 0$ and $\underline{m}_1 = \bar{m}$. At an interior nonmarket allocation,

$$\phi v'(h_j^*) = \phi \psi'(c_j^* + \gamma A) = g'(l_j^*) = \lambda_j.$$

The condition for market hours and the complementary-slackness conditions are

$$w_j u'(x_j^*) - \lambda_j + \alpha_j - \mu_j = 0,$$

$$\alpha_j(m_j^* - \underline{m}_j) = 0, \quad \mu_j[\tau(A) - m_j^*] = 0.$$

When neither hours bound binds, $w_j u'(x_j^*) = \lambda_j$. A binding hours constraint introduces the corresponding multiplier. The leisure condition makes the shadow value of time endogenous: a reduction in leisure raises λ_j because $g''(l) < 0$.

A3 Childcare Comparative Statics

Suppressing sector subscripts and holding ϕ fixed, differentiation of the childcare condition gives

$$\phi \psi''(c + \gamma A)(dc + \gamma dA) = d\lambda.$$

Along a differentiable allocation within a sector,

$$\frac{\partial c}{\partial A} = -\gamma + \frac{1}{\phi\psi''(c + \gamma A)} \frac{\partial \lambda}{\partial A}.$$

Thus, one-for-one crowd-out is exact when the shadow value of time is held fixed and is an approximation when the second term is small. Since $\psi'' < 0$, an increase in λ produces a larger decline in maternal childcare, whereas a decrease in λ offsets the direct displacement. Remaining in the same sector does not, by itself, imply that λ is constant. These local derivatives do not describe a discrete sector switch; A6 examines that change directly.

A4 Housework Response

Differentiating the housework and leisure conditions at a fixed ϕ yields

$$\phi v''(h) dh = d\lambda, \quad g''(l) dl = d\lambda.$$

Both activities decline when the shadow value of time rises. To establish the residual-time argument, define $R = T - m - c = h + l$. Conditional on market work and childcare, optimal housework satisfies $\phi v'(h) = g'(R - h)$. Differentiation gives

$$\begin{aligned} \frac{\partial h}{\partial R} &= \frac{g''(l)}{\phi v''(h) + g''(l)} \in (0,1), \\ \frac{\partial l}{\partial R} &= \frac{\phi v''(h)}{\phi v''(h) + g''(l)} \in (0,1). \end{aligned}$$

Housework and leisure are therefore both normal with respect to the residual time budget under the stated preferences. This property follows from strict concavity; it does not require an additional normal-goods assumption. Because the utility components are separable, changes in consumption and effective childcare do not directly alter this conditional allocation. A fall in R reduces both housework and leisure, including for a finite change between interior optima.

A5 Participation Threshold

Consider a level of ASC provision at which both sectors are feasible, and define $D(A; \phi) = V_1(A; \phi) - V_0(A; \phi)$. At differentiable optima, the envelope theorem gives

$$D_\phi = v(h_1^*) + \psi(c_1^* + \gamma A) - v(h_0^*) - \psi(c_0^* + \gamma A).$$

We focus on the relevant range of preferences in which formal employment entails more hours, $m_1^* > m_0^*$, and both nonmarket allocations are interior. At a fixed A and ϕ , the allocation of housework, childcare, and leisure solves the same strictly concave problem in each sector, but with less nonmarket time in formal employment. The first-order conditions imply

$$\lambda_1 > \lambda_0, \quad h_1^* < h_0^*, \quad c_1^* < c_0^*.$$

Consequently, $D_\phi < 0$. If D is continuous and changes sign over this preference range, an indifference threshold exists and is unique:

$$D(A; \phi^*(A)) = 0.$$

Mothers below the threshold prefer formal employment; those above it prefer the flexible alternative. The mother is indifferent at the threshold itself. Monotonicity ensures uniqueness, while the sign change is required for an interior threshold to exist.

To examine how the threshold changes with ASC, apply the constrained envelope theorem within a region where both sectors remain feasible. Because $\phi\psi'(c_j^* + \gamma A) = \lambda_j$,

$$\frac{\partial V_j}{\partial A} = \gamma\lambda_j + \mu_j\tau'(A).$$

It follows that

$$D_A = \gamma(\lambda_1 - \lambda_0) + (\mu_1 - \mu_0)\tau'(A).$$

If informal hours are below the supervision ceiling, $\mu_0 = 0$. Together with $\lambda_1 > \lambda_0$, $\mu_1 \geq 0$, and $\tau'(A) \geq 0$, this condition implies $D_A > 0$. In fact, with a common ceiling, $m_1^* > m_0^*$ implies that informal hours are below it. The implicit-function theorem therefore yields

$$\frac{d\phi^*(A)}{dA} = -\frac{D_A}{D_\phi} > 0.$$

If these conditions hold throughout an interval, greater ASC provision raises the threshold over that interval. If formal employment is infeasible before ASC, its pretreatment value is not a feasible choice and a pretreatment indifference threshold need not exist. The appropriate comparison is then between no access to formal work before the policy and the set of mothers with $D(1; \phi) > 0$ afterward.

A6 Switching and Time Reallocation

For the same mother, let $\Delta m = m_1^*(1; \phi) - m_0^*(0; \phi) > 0$ denote the discrete increase in market hours when she switches from informal work before ASC to formal work afterward. Define changes in other time uses analogously. The time constraint gives

$$\Delta h + \Delta l = -\Delta m - \Delta c.$$

Therefore, if childcare falls and $\Delta m > -\Delta c$, residual time declines. The monotonicity established in A4 implies $\Delta h < 0$ and $\Delta l < 0$. This conclusion uses the actual change in maternal childcare, without imposing one-for-one crowd-out.

A sufficient condition can be stated directly in terms of the program's childcare provision. Define effective childcare as $z = c + \gamma A$ and the effective time endowment for nonmarket activities as $S = T - m + \gamma A$. Then

$$h + z + l = S.$$

At interior allocations, the conditional objective is $\phi[v(h) + \psi(z)] + g(l)$ and the lower bound $z \geq \gamma A$ is slack. The conditional allocation therefore depends on S and ϕ . Differentiating its first-order conditions and resource constraint gives

$$\frac{\partial \lambda}{\partial S} = \frac{1}{\frac{1}{\phi v''(h)} + \frac{1}{\phi \psi''(z)} + \frac{1}{g''(l)}} < 0.$$

Each of h , z , and l increases with S . Since ASC changes A from zero to one,

$$\Delta S = \gamma - \Delta m.$$

If $\Delta m > \gamma$, the effective time endowment contracts, so housework, effective childcare, and leisure all fall, while the shadow value of time rises. Maternal childcare then satisfies $\Delta c = \Delta z - \gamma < -\gamma$. Because housework and leisure also decline, the increase in market hours exceeds the total reduction in maternal childcare. Thus, this sufficient condition establishes the time-compression mechanism without treating childcare crowd-out as exactly one-for-one.

For a nonswitcher whose hours remain unchanged, $\Delta S = \gamma > 0$. At interior allocations, housework and leisure increase, and maternal childcare falls by less than the program's effective provision, $-\gamma < \Delta c < 0$. Accordingly, a negative average housework response depends on the relative magnitude and prevalence of the responses among switchers and nonswitchers. The weekday interpretation assumes that changes in work schedules and home production do not generate substantial substitution across days; the framework does not impose a zero weekend effect.

A7 Corner Solutions

At a zero allocation of a nonmarket activity, the corresponding first-order equality is replaced by a Kuhn–Tucker inequality. When the relevant one-sided derivatives are finite,

$$h = 0 \Rightarrow \lambda \geq \phi v'(0), \quad c = 0 \Rightarrow \lambda \geq \phi \psi'(\gamma A),$$

$$l = 0 \Rightarrow \lambda \geq g'(0).$$

If marginal utility becomes infinite at zero, the corresponding corner is excluded whenever an interior allocation is feasible. At admissible corners, the strict responses derived above can become weak, and maternal childcare cannot fall below zero. The strict comparative statics and threshold derivatives therefore apply to the interior regions specified above; boundary allocations require the appropriate complementary-slackness conditions.

Appendix B: Supplementary Figures and Tables

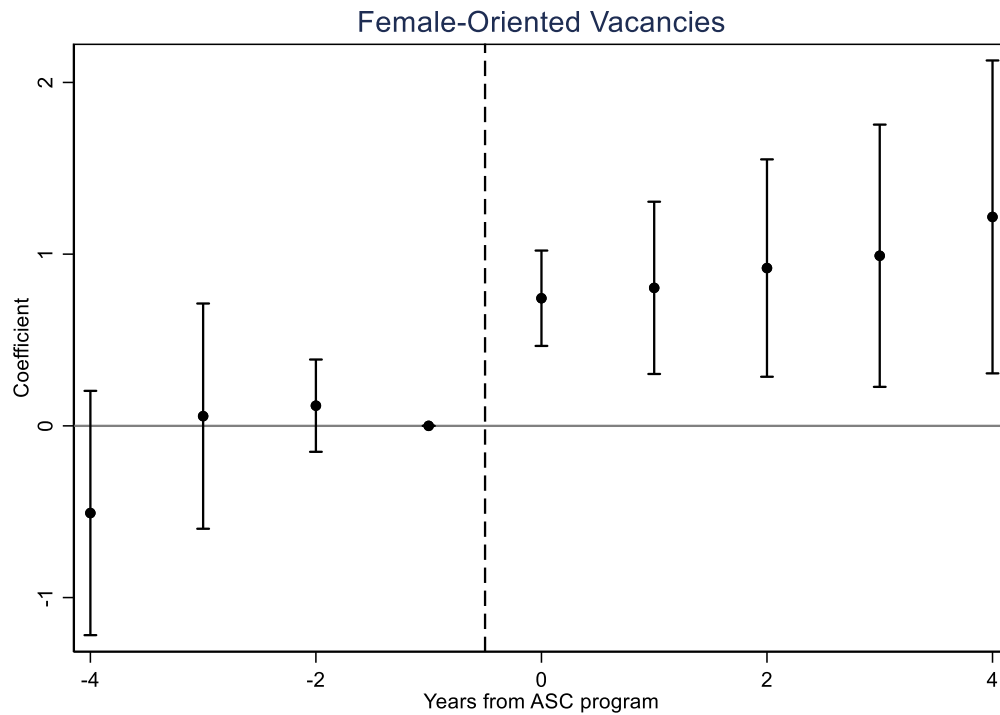


Figure A1 Event-Study Estimates for Female-Oriented Vacancies

Notes: This figure plots event-study estimates of the effect of ASC implementation on female-oriented vacancies, with 95 percent confidence intervals. The sample uses the extended 2014–2022 job-posting panel. For vacancies posted after 2018, when explicit gender preferences in job advertisements were restricted, gender orientation is predicted using the text-classification procedure described in the data section. The omitted category is the year immediately preceding ASC implementation. All specifications include firm and year fixed effects. Standard errors are clustered at the county level.

Table A1 Determinants of ASC Adoption Timing

	(1) ASC adoption timing
Total population	0.644 (0.309)
Employment	0.978 (0.090)
GDP per capita	0.816* (0.089)
Local government revenue	1.946** (0.627)
Local government expenditure	1.315 (0.652)
Primary school enrollment	1.061 (0.175)
Lower-secondary school enrollment	0.672 (0.163)
Urbanization rate	1.002 (0.009)
Secondary industry share	1.008 (0.022)
Tertiary industry share	0.994 (0.027)
Total job vacancies	0.883 (0.275)
Female-oriented job vacancies	1.699 (0.609)
Observations	742

Notes: This table reports hazard ratios from a Cox proportional hazards model of county-level ASC adoption timing. Observations are county-year risk-set observations. Values greater than one indicate a higher hazard of adoption, corresponding to earlier implementation; values below one indicate a lower hazard of adoption, corresponding to slower implementation. Explanatory variables are drawn from the China County Statistical Yearbook and job-posting data, supplemented by provincial and municipal statistical yearbooks where necessary. Robust standard errors, clustered at the county level, are reported in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A2 Summary Statistics for Paternal outcomes

VARIABLES	(1) Observations	(2) Mean	(3) SD	(4) Minimum	(5) Maximum
Employment	9,904	0.961	0.195	0	1
Annual income	6,338	31,480	30,666	0	200,000
Weekly working hours	7,931	49.14	23.32	0	120
Egalitarian gender views	8,012	0.507	0.500	0	1
ASC	9,904	0.168	0.374	0	1
Survey year	9,904	2017	2.753	2014	2022
Age	9,904	39.35	6.240	22	77
Years of education	9,904	8.906	3.966	0	22
Married	9,904	0.990	0.100	0	1
Han ethnicity	9,904	0.635	0.481	0	1
Number of children	9,904	1.559	0.709	1	7

Notes: This table reports summary statistics for the father sample using the 2014–2022 waves of the CFPS. Fathers are defined as husbands of mothers with children aged 6–15 who can be matched through the CFPS household roster.

Table A3 Pre-Reform Care Arrangements and Homework Support

	Share
Panel A. Primary Daytime Caregiver	
Childcare Center	7.79%
Paternal Grandparents	15.99%
Maternal Grandparents	2.40%
Father	3.40%
Mother	38.73%
Child Self-Care	16.87%
Other Caregivers	14.83%
Panel B. Primary Nighttime Caregiver	
Childcare Center	0.36%
Paternal Grandparents	13.95%
Maternal Grandparents	1.68%
Father	4.76%
Mother	58.95%
Child Self-Care	14.15%
Other Caregivers	6.16%
Panel C. Homework Support Arrangement	
Mother Only	54.95%
Father Only	14.80%
Both Parents (No Other Helpers)	10.83%
Other Relatives Only (Without Parents)	12.22%
Parents and Other Relatives Together	7.20%

Notes: This table reports the distribution of care arrangements and homework-support providers for children aged 6–15 before ASC implementation, using the 2016 CFPS wave. Panels A and B report the child's primary daytime and nighttime caregiver, respectively. Panel C reports homework-support arrangements.

Table A4 Placebo Tests Using Predetermined Characteristics

VARIABLES	(1) Years of education	(2) Married	(3) Minority ethnicity	(4) Religious affiliation	(5) Urban hukou
ASC	0.126 (0.162)	0.004 (0.006)	0.007 (0.020)	-0.003 (0.015)	-0.004 (0.014)
Observations	12,462	12,462	12,462	10,486	12,388
R-squared	0.348	0.036	0.231	0.201	0.330
Mean of Y	8.057	0.971	0.323	0.290	0.204
Year FE	YES	YES	YES	YES	YES
County FE	YES	YES	YES	YES	YES

Notes: This table reports the effects of the ASC policy on predetermined socioeconomic characteristics of mothers with children aged 6–15. The dependent variables in Columns (1)–(5) are years of schooling, marital status, minority ethnicity, religious affiliation, and urban hukou status, respectively. All specifications include county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A5 Robustness Checks

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
Panel A Mothers with children under 6			
ASC	0.029 (0.039)	0.650 (2.205)	133.258 (2,229.024)
Observations	4,983	4,649	4,191
Panel B Heterogeneous Trends			
ASC	0.042** (0.017)	2.368** (1.146)	2,444.487** (1,181.614)
Observations	12,462	10,514	8,127
Panel C Group-Time Estimator			
ASC	0.060** (0.024)	3.134** (1.553)	3,639.560** (1,768.878)
Observations	9,365	7,078	4,510
Panel D Imputation-Based estimator			
ASC	0.054** (0.022)	3.223** (1.415)	3,260.704** (1,649.003)
Observations	12,458	10,506	8,118
Panel E Stacked DID			
ASC	0.055*** (0.016)	3.592*** (1.089)	3,895.648*** (1,422.600)
Observations	28,037	23,660	17,633
Panel F Extended-TWFE estimator			
ASC	0.054*** (0.019)	3.226** (1.242)	3,276.539** (1,424.208)
Observations	12,457	10,505	8,117
Panel G Maternity Leave			
ASC	0.044*** (0.016)	2.665** (1.105)	2,848.713** (1,269.600)
Observations	12,462	10,514	8,127
Panel H Long-term Care Insurance			
ASC	0.049*** (0.017)	2.867** (1.127)	2,667.773** (1,311.802)
Observations	12,462	10,514	8,127
Panel I Pre-COVID Sample			
ASC	0.048** (0.021)	2.494* (1.403)	3,206.762* (1,898.651)
Observations	8,466	6,636	5,444
Panel J Full Sample with COVID-19 Controls			
ASC	0.046*** (0.016)	2.420** (1.135)	2,917.337** (1,349.686)
Observations	12,462	10,514	8,127
Panel K Non-Migrants			
ASC	0.041** (0.016)	2.361** (1.104)	2,232.722* (1,308.076)
Observations	12,339	10,405	8,020

Notes: This table reports robustness and validity checks for the estimated effects of the ASC policy on maternal labor market outcomes among mothers with children aged 6–15. Panel A presents a placebo analysis using mothers whose children are all younger than age six and are therefore ineligible for ASC. Panel B allows for differential trends by interacting year fixed effects with county-level GDP per capita and local government revenue measured in 2016. Panel C reports estimates using the group-time average treatment effect estimator of Callaway and Sant'Anna (2021). Panel D uses the imputation-based estimator of Borusyak, Jaravel, and Spiess (2024). Panel E uses the stacked DID estimator of Cengiz et al. (2019). Panel F uses the extended two-way fixed effects estimator of Wooldridge (2025). Panel G additionally controls for the number of statutory maternity-leave days, and Panel H additionally controls for the rollout of long-term care insurance. Panel I restricts the sample to the pre-COVID period, 2014–2019. Panel J uses the full 2014–2022 sample and controls for the log of cumulative confirmed COVID-19 cases in the mother's city during 2020–2022. Panel K restricts the sample to mothers who remain in the same county throughout the sample period. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A6 Inverse-Probability-Weighted Estimates for Baseline Labor-Market Outcomes

VARIABLES	(1) Weekly working hours	(2) Annual income
ASC	2.513** (1.116)	2,966.680** (1,249.207)
Observations	10,514	8,127
R-squared	0.089	0.371
Year FE	YES	YES
County FE	YES	YES

Notes: This table reports inverse-probability-weighted estimates addressing item non-response and conditional reporting in the CFPS. The dependent variables in Columns (1) and (2) are weekly working hours and annual income, respectively. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** p<0.01, ** p<0.05, * p<0.10.

Table A7 Phase-Specific Effects of ASC on Mothers' Labor Market Outcomes

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
Panel A. Advisory Phase, 2017–2018			
ASC	0.035 (0.031)	2.003 (2.354)	11368.825*** (3454.746)
Observations	20614	16960	13177
Panel B. Non-Mandatory Expansion Phase, 2019–2020			
ASC	0.069*** (0.022)	3.853** (1.573)	3072.201* (1839.767)
Observations	20172	16691	12831
Panel C. Mandatory Phase, 2021–2022			
ASC	0.044 (0.030)	2.764 (1.975)	393.449 (2247.591)
Observations	12482	10515	7854

Notes: This table reports phase-specific estimates of the effects of the ASC policy on labor market outcomes for mothers with children aged 6–15. The reform is divided into three implementation phases: the advisory phase (2017–2018), when local governments were encouraged to explore after-school service provision; the non-mandatory expansion phase (2019–2020), when implementation expanded but was not yet nationally required; and the mandatory phase (2021–2022), when ASC provision became a nationwide policy requirement. For each phase, we estimate a cohort-specific stacked DID specification that excludes earlier-treated counties and censors later-treated counties after their own adoption. Observations in stacked DID specifications reflect the expanded panel structure resulting from stacking cohort-specific sub-samples. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A8 Effects of ASC on Formal Employment by Adoption Phase

Adoption phase	Dependent variable: Formal employment		
	Advisory 2017–2018	Non-mandatory expansion 2019–2020	Mandatory 2021–2022
	(1)	(2)	(3)
ASC	0.196*** (0.040)	0.087** (0.035)	-0.003 (0.028)
Observations	16,767	16,385	10,193
R-squared	0.358	0.330	0.339

Notes: This table reports the effects of ASC on formal employment among mothers with children aged 6–15, separately by the phase in which their county adopted the policy. The dependent variable is an indicator for formal employment. All specifications follow the phase-specific stacked DID design used in Table A7. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A9 Advisory-Phase Income Effects under Alternative Winsorization Thresholds

Winsorization percentiles	Dependent variable: Annual income (yuan)		
	None	1st–99th	5th–95th
	(1)	(2)	(3)
ASC	11,368.825*** (3,454.746)	10,280.109*** (2,984.598)	5,075.556** (2,051.049)
Observations	13,177	13,177	13,177
R-squared	0.358	0.371	0.362

Notes: This table examines the sensitivity of the advisory-phase income estimate to alternative winsorization thresholds.

Column (1) uses income without winsorization. Columns (2) and (3) winsorize income at the 1st–99th and 5th–95th percentiles, respectively, replacing values outside these thresholds with the corresponding cutoff values. All columns use the same sample and stacked DID specification as the advisory-phase income regression in Table A7. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A10 Robustness to Excluding Each Reform Phase

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
Panel A. Excluding Advisory Phase			
ASC	0.056*** (0.019)	2.842** (1.282)	1,866.886 (1,457.602)
Observations	11,754	9,922	7,567
Panel B. Excluding Non-Mandatory Expansion Phase			
ASC	0.036* (0.021)	2.251 (1.427)	3,151.097* (1,803.314)
Observations	10,538	8,887	6,806
Panel C. Excluding Mandatory Phase			
ASC	0.053** (0.021)	3.727** (1.508)	4,526.447** (1,839.401)
Observations	4,857	4,097	3,239

Notes: This table reports robustness checks that sequentially exclude each implementation phase of the ASC reform. Panel A excludes observations from the advisory phase (2017–2018), Panel B excludes observations from the non-mandatory expansion phase (2019–2020), and Panel C excludes observations from the mandatory phase (2021–2022). These checks assess whether the baseline estimates are disproportionately driven by any single stage of the reform. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level.

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

Table A11 Heterogeneous Effects of ASC by Child Age

VARIABLES	(1) Employment	(2) Weekly working hours	(3) Annual income
ASC * Younger child	0.066*** (0.020)	2.610** (1.202)	2,563.232** (1,242.605)
ASC	0.008 (0.018)	0.879 (1.332)	1,413.840 (1,490.834)
Younger child	-0.030*** (0.008)	-2.105*** (0.632)	-411.922 (583.601)
Observations	12,462	10,514	8,127
R-squared	0.105	0.090	0.382
Year FE	YES	YES	YES
County FE	YES	YES	YES

Notes: This table reports heterogeneous effects of the ASC policy on maternal labor market outcomes by child age. The sample includes mothers with children aged 6–15. Younger child is an indicator equal to one for mothers with a child aged 6–10 and zero for mothers whose children are aged 11–15. Annual income is measured in yuan. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. ***p<0.01, **p<0.05, *p<0.10.

Table A12 Homework-Support Arrangements

VARIABLES	(1) Mother involved	(2) Father only	(3) Relatives only
ASC	-0.038** (0.018)	0.001 (0.005)	0.007 (0.017)
Observations	12,462	12,462	12,462
R-squared	0.167	0.075	0.210
Year FE	YES	YES	YES
County FE	YES	YES	YES
Mean of Y	0.481	0.0447	0.250

Notes: This table reports the effects of the ASC policy on homework-support arrangements among households with children aged 6–15. The dependent variables are indicators for whether the mother is involved in homework support, whether the father is the only homework-support provider, and whether relatives are the only homework-support providers. The mother-involved measure captures any maternal involvement in homework support, either alone or with others. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A13 Parental Childcare Time

VARIABLES	Childcare Time (hours per day)	
	(1) Mothers	(2) Fathers
ASC	-0.299* (0.177)	-0.135 (0.116)
Observations	3,996	2,984
R-squared	0.244	0.170
Year FE	YES	YES
County FE	YES	YES
Mean of Y	2.650	0.925

Notes: This table reports the effects of the ASC policy on parental childcare time among households with children aged 6–15. The dependent variables measure the hours that mothers and fathers spend with their children, including time devoted to care, companionship, education, and school pickup and drop-off. These measures are available only in the 2020 and 2022 CFPS waves. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level.

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

Table A14 Unconditional Quantile Regression on Commuting Time

VARIABLES	One-way commuting time			
	(1) 20th percentile	(2) 40th percentile	(3) 60th percentile	(4) 80th percentile
ASC	0.352 (0.553)	0.352 (0.493)	1.488* (0.761)	3.176** (1.236)
Observations	6577	6577	6577	6577

Notes: This table reports the effects of the ASC policy on the commuting time of mothers with children aged 6–15 using unconditional quantile regression (RIF-regression). Commuting time is measured in minutes. Columns (1) through (4) correspond to the 20th, 40th, 60th, and 80th quantiles of the commuting time distribution, respectively. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A15 Inverse-Probability-Weighted Estimates for Mechanism Outcomes

VARIABLES	(1) Weekday housework	(2) Weekend housework	(3) One-way commuting time	(4) Formal employment
ASC	-0.182** (0.087)	-0.037 (0.090)	2.721** (1.208)	0.098*** (0.025)
Observations	12,317	12,310	6,577	10,116
R-squared	0.130	0.091	0.148	0.293
Year FE	YES	YES	YES	YES
County FE	YES	YES	YES	YES

Notes: This table reports inverse-probability-weighted estimates addressing item non-response and conditional reporting in the CFPS. The dependent variables in Columns (1)–(4) are weekday housework hours per day, weekend housework hours per day, one-way commuting time in minutes, and an indicator for formal employment status, respectively. All specifications include controls for age, years of education, number of children, marital status, and Han ethnicity, as well as county and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A16 Vacancies Targeting Young Women

VARIABLES	(1) Vacancies targeting women under 26
ASC	0.010 (0.007)
Observations	32,228
R-squared	0.256
Year FE	YES
Firm FE	YES
Mean of Y	0.0163

Notes: This table reports the effect of the ASC policy on vacancies targeting female candidates under age 26. The sample uses the extended 2014–2022 job-posting panel. The dependent variable is the number of vacancies targeting female candidates under age 26. All regressions include firm and year fixed effects. Standard errors, reported in parentheses, are clustered at the county level. *** p<0.01, ** p<0.05, * p<0.1.