THE JAPANESE PUBLIC POLICIES ON TAX, WAGES, AND STANDARD WORK HOURS— EVIDENCE FROM MICRO DATA

by

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CHAPTER I

Introduction

This dissertation theoretically and empirically examines several topics relating to the Japanese labor market: the impact of the 2004 tax reform on the female labor supply, the possible source of wage rigidity, and the impact of Japan's reduction of weekly standard work hours on labor-market outcomes. Although Chapters II, III, and IV in this dissertation focus on different topics, they all explore new evidence regarding the Japanese labor market during the 1990s and the 2000s using Japanese micro data. Chapters II and IV examine how the reforms in the Japanese public policies, such as the spousal exemption system and standard work hours, affected labor market outcomes, employing relatively novel empirical tools, such as unconditional quantile regressions, Firpo, Fortin, and Lemieux decomposition, DiNardo, Fortin, and Lemieux decomposition, and regression discontinuity design. Chapter III focuses more on theory, and it provides a new type of efficiency wage model, exploring the impact of the performance-based layoffs on wage rigidity. The details of each chapter are presented below.

Chapter II explores how the 2004 tax reform in Japan affected the income distribution of married women: In Japan, there is a spousal exemption system whereby the income exemption amount that a husband can claim is reduced as his spouse's income increases. In 2004, part of this spousal exemption was abolished. Although it was expected that this tax reform would increase the labor supply of married women, both the average work hours and annual incomes of married women slightly decreased after the reform was introduced. To explore the real effect of the tax reform on the female labor supply, this paper theoretically and empirically examines the effect of the tax reform on the income distributions of married women. Quantile difference-in-difference estimations and FFL and DFL decompositions indicate that i) the tax reform contributed to increasing work hours (and hence the incomes) of low-income married women, and ii) the kink at 1.03 million yen in the budget line, which was made more conspicuous by the tax reform, dragged even medium- to high-income married women, who were not presumed to be affected by the tax reform, toward the center of the income distribution in response to an increase in their husbands' incomes. Consequently, the conventional distortion at 1.03 million yen in the income distribution became stronger after the reform.

Chapter III suggests a new explanation for the source of wage rigidity during recessions: The theoretical model discussed in this chapter is a new type of efficiency wage model in which performance pay is assumed, and a firm can choose how much weight to assign performance and seniority as layoff criteria. Firms' decisions follow an over-lapping generation model structure, dividing the workers into two types: new employees and experienced employees. In this setting, new employees will be able to receive wages scheduled to be paid to experienced employees in a recession only when they work hard in the first period and avoid layoffs. In contrast, wages scheduled to be paid to experienced employees in a boom do not affect incentives of new employees because they will receive the wages in the second period without being laid off. Thus, the high wages scheduled to be paid to experienced employees during a recession encourage new employees to work hard, and it becomes optimal for a firm to pay the higher fixed component of wages during a recession than during a boom. In addition, during recessions, firms are discouraged from maintaining workers' current incentives at a higher level due to the lower value of productivity, which results in the bonus that moves proportionally to the output price. Thus, the theoretical model predicts that wages during recessions become both "downwardly rigid" and "rigid" (inflexible) with respect to performance. These implications are tested using Japanese panel data from the Keio Household Panel Survey (KHPS). The empirical results support the model's predictions.

Chapter IV assesses the impact of Japan's reduction of weekly standard work hours (from 48 hours/week in 1988 to 40 hours/week by 1997) on labor-market outcomes using the regression discontinuity approach: Although the reduction of standard hours reduced total hours worked, the monthly wages of existing workers did not decrease in response to the policy change. As a result, overall labor costs per hour increased, and we do not see any evidence of job creation. Even in Japan, where wages are determined flexibly, we observe evidence similar to that in European countries such as Germany and France, where pay is largely determined through union-employer bargaining.

Chapter V concludes this dissertation.

CHAPTER II

The Impact of Tax Reform in Japan on the Income Distribution of Married Women

2.1 Introduction

There is a spousal exemption system in Japan in which the amount of income exemptions a husband can claim reduces as his spouse's income rises. This spousal exemption system has been criticized for discouraging married women from working long hours. As female labor participation has increased, criticism of the spousal exemption system has likewise escalated. As a result, part of this spousal exemption system was abolished in 2004 for taxpayers whose spouses earned annual earnings of 1.03 million yen (≈USD 10,300) or less.[1]

The original purpose of this tax reform was to mitigate the distortion it might

[1] More strictly speaking, part of the spousal exemption was abolished for taxpayers whose spouses' annual *total income* is 0.38 million yen or less, where *total income* is calculated as the total earnings minus the *employment income deduction* for employees and the total earnings minus necessary expenses for other workers. If we consider only employees, the annual *total income* of 0.38 million yen corresponds to annual earnings of 1.03 million yen. Since the amount of necessary expenses varies by individual, it is not possible to accurately relate the amount of the *total income* to the total earnings for workers other than employees without taking into consideration information about the necessary expenses applicable to each worker. In contrast, the amount of the *employment income deduction* is defined for each category of annual earnings and, hence, it is possible to relate the amount of the *total income* to the amount of annual earnings for employees. (The details will be discussed in Section 2.2.1.) For this reason, I will discuss the impact of the tax reform in the context of employees, using the amount of annual earnings, rather than the amount of the *total income* because the amount of annual earnings is more intuitively understandable. Thus, hereafter, the term "income" means annual earnings of employees.

have caused to the women's labor supply, and thus was expected to increase married women's labor supply. However, statistics show that both the average work hours for married women and their annual incomes slightly decreased after the tax reform; these trends are not observed for single women. If this is the only evidence we consider, we might be driven to conclude that the tax reform did not increase married women's labor supply. However, further considerations are required to understand the reform's real impact.

First, in the years following the 2004 tax reform, there was a trend in which the incomes of married men in Japan increased in response to the recovery from a recession that started in the late 1990s; that increase could potentially decrease the income earned by married women and, hence, conceal the reform's real impact. Therefore, it is important to estimate the impact of the tax reform on married women's labor supply after controlling for changes in exogenous variables, particularly an increase in their husbands' incomes.

Second, the impact of the tax reform should be heterogeneous across income groups among married women. Since the tax reform directly affected only low income married women[2], estimating only the mean impact among all women would be misleading and might result in missing a change in labor supply due to the tax reform. This paper explicitly takes this into account and examines the reform's impact on income distribution rather than focusing on the mean.[3]

Third, people can adjust their incomes not only by work hours/labor participation, but also by their effort at work, their job choices, and their manner of earning income (e.g., salary, dividends, or capital gains). As is often argued in the public finance

[2]More strictly speaking, individuals who claim spousal exemptions could be either husbands or wives, but here it is assumed that husbands claim the spousal exemptions. The details about the spousal exemptions rule will be discussed in Section 2.2.2.

[3]This point is very similar to that of Bitler et al. (2006), who analyzed the impact of recent welfare reforms in the United States on labor outcomes, using a random-assignment experiment. By estimating quantile treatment effects, they showed what mean impacts miss when we examine the impact of recent welfare reforms and how important it is to analyze distributional effects.

literature, people can also control their after-tax incomes by changing behaviors other than work hours in order to reduce their tax liability. "New tax responsiveness" literature holds such a general view on the concept of labor supply and uses the response of taxable income to the marginal tax rate as a summary statistic of the behavioral response to taxation (Meghir and Phillips, 2008; Feldstein, 1995). Using incomes rather than work hours to analyze the impact of the tax reform circumvents the concern that the tax reform affected people's behaviors other than their hours of work. Thus, in this paper, I mainly focus on the tax reform effect on incomes.

There have been many studies that examine the impact of the spousal exemption itself on female labor supply (Abe and Otake (1995); Higuchi (1995); Kantani (1997); Akabayashi (2006); Takahashi (2010a,b), etc.). However, only a few studies have examined the impact of the 2004 Japanese tax system reform on the female labor supply and no critical consensus has arisen concerning the real impact of the tax reform: Sakata and McKenzie (2006) found that the tax reform had no impact on the labor-participation decisions and a slightly positive effect on the number of work hours for female part-time workers who worked less than 35 hours per week. In contrast, Sakata and McKenzie (2005) found that the tax reform had a negative and statistically significant impact on both the decisions of female spouses to enter the labor market and on the number of hours they worked, conditional on their participation in the labor market.

Furthermore, these studies examine changes in labor supply at the mean, and no study has analyzed the overall distributional change due to the tax reform. However, if the tax reform had any side effects on high-income married women who were not presumed to be affected by the tax reform, restricting the sample to only female workers who were directly affected would result in missing those unexpected impacts. For example, the way high-income married women decrease their incomes in response to some exogenous shock could have been altered by the tax reform due to the fact that the tax reform changed the shape of their budget line (which could also alter their potential labor supply choice). Therefore, this paper examines what happened to the overall income distribution among married women without restricting the sample to only those female workers who were directly affected by the tax reform.

Concerning tax reforms in other countries, there have been many studies that examined the impact of tax reform on female labor supply. For example, Eissa and Liebman (1996) examined the impact of the U.S. Tax Reform Act of 1986 (TRA 86), which included an expansion of the earned income tax credit on female labor supply. They used the difference-in-difference estimation to compare the change in labor supply of single women with children with the change for single women without children, and they found that the Tax Reform Act increased labor force participation among single women with children. Furthermore, Blundell et al. (1998) examined the impact of the tax reforms in the United Kingdom on labor supply during the 1980s by comparing the labor supply responses over time for different groups defined by cohort and education level. They found positive and moderately sized wage elasticities and negative income effects for women with children.

Before moving to the detailed discussion of the 2004 tax reform, I will note important characteristics in the Japanese tax system: First, married people in Japan file separate tax returns. Second, capital income is taxed separately from labor income. Therefore, in this paper, I will examine only labor income and the tax on it, and the term "income" hereafter means only income from labor earnings, which excludes income from assets. Third, there also exist complicated income exemption schedules and policy-related customs that create kinks on the household budget line, making it nonlinear. The most important factor to note about the Japanese tax system is that 1.03 million yen (≈USD 10,300) has historically been a critical income threshold for many married women as their husbands will lose the standard spousal exemption once their annual income exceeds that threshold. Given that the income threshold for the standard spousal exemption in tax is 1.03 million yen, many firms set this amount as the critical income threshold at which the spousal allowance vanishes. Thus, in many cases, husbands will lose not only the standard spousal exemption but also the employer-provided spousal allowances once their spouses' annual (before-tax) income exceeds 1.03 million yen. Furthermore, 1.03 million yen of annual earnings is the point at which all employees begin to pay income tax. (Up to 1.03 million yen in annual earnings, employees pay zero income tax because all employees are eligible for the deduction of 1.03 million yen.) These facts give an incentive to many married women to adjust their work hours so that annual incomes do not exceed 1.03 million yen. As a result, there exists a mass point in the income distribution of married women at incomes slightly lower than 1.03 million yen.

The theoretical model in this chapter first shows that the tax reform made the kink at 1.03 million yen in the budget line more conspicuous than it had been before, which produces the following testable implications: (1) the tax reform increases work hours (and hence the incomes) for married women who earned 1.03 million yen or less but has no impact on married women who originally earned more than 1.03 million yen, when there is no change in exogenous parameters; (2) however, if an increase in a husband's income is considered, there are cases in which medium- to high-income married women greatly decrease their incomes to 1.03 million yen; (3) even in this case, the "tax reform effect" cannot be negative; (4) this "income jump" among medium- to high-income married women can be explained mostly by a negative "husband's income effect"; and (5) in some cases, this "income jump" is mitigated by a small positive "tax reform effect."

These implications are tested using Japanese panel data from the Keio Household Panel Survey (KHPS). The sample consists of 3,206 observations in the panel over the period 2003-2006. In the empirical investigation, I first conduct a difference-in-difference estimation within a framework of the quantile regressions. Both results from the conditional and unconditional quantile regressions show that the tax reform increased incomes among married women who earned 1.03 million yen or less. If we estimated the impact of the tax reform at the mean, we do not see any significant impact, implying that it is important to estimate the impact of tax reform for each quantile. These findings support the theoretical implications described above. In contrast, the tax reform effects on married women at high quantiles in the income distribution are insignificant. This is also consistent with a theoretical implication that the tax reform has no impact on high-income married women in a setting in which there is no change in exogenous parameters.[4]

Next, to introduce changes in exogenous variables, I employ a relatively new decomposition technique proposed by Firpo et al. (2007, 2010), which enables us to examine whether the "income jump" among medium- to high-income married women actually occurred in response to an increase in husbands' incomes after the tax reform. The idea in the Firpo, Fortin, and Lemieux decomposition (hereafter "FFL") is to combine the DiNardo, Fortin, and Lemieux decomposition (hereafter "DFL") proposed by DiNardo et al. (1996) with Oaxaca-Blinder decomposition to make it possible to estimate the contribution of each covariate to the structure and composition effects when analyzing distributional change. This is made possible by using the recentered influence function (hereafter "RIF") of Y as the dependent variable. In the first stage of the FFL, distributional changes in incomes are divided into a distributional change due to changes in β, i.e., a structure effect, and a distributional change due to changes in X, i.e., a composition effect, using a reweighting method proposed by DiNardo et al. (1996). In the second stage, the two components are further divided into the contribution of each explanatory variable using novel RIF re-

[4]Note that the settings of the difference-in-difference estimation corresponds to the case in which there is no change in exogenous parameters, since the husband's income and many other exogenous variables are controlled for in the estimation and the coefficients of these explanatory variables, except for the treatment dummy variable that estimates the impact of the tax reform, are also assumed to be the same before and after the tax reform.

gressions. These regressions estimate directly the impact of the explanatory variables on the distributional statistic of interest.

The results from the FFL suggest that the tax reform contributed to increasing incomes among the low-income married women. In contrast, the magnitude of the composition effect attributable to husbands' incomes is greatest among medium- to high-income married women. This indicates the possibility that the "income jump" among medium- to high-income married women in response to an increase in husbands' incomes actually occurred because this "income jump" can be explained by a large negative "husband's income effect."

Lastly, utilizing the DFL decomposition method, I visually examine how the behavioral changes confirmed in the difference-in-difference estimation and the FFL decomposition affected the overall income distribution. In this decomposition, we reconfirm the results obtained from the FFL; low-income married women and high-income married women gathered to the center of the distribution after the tax reform, given the facts that low-income married wome increased work hours because of the tax reform and that high-income married women decreased work hours in response to an increase in their husbands' incomes. Accordingly, the mass point at 1.03 million yen in the income distribution became more conspicuous following the reform. Furthermore, the DFL results also indicate that incomes would have decreased more in response to the increase in husbands' incomes if the tax reform had not occurred. These results from the FFL and DFL decompositions support theoretical implications.

The interpretation suggested for this trend is as follows: Although the kink at 1.03 million yen in the budget line had originally been an attractive option for many married women, due to the characteristics of the Japanese tax system, the tax reform made this option more attractive than before. As a result, the kink at 1.03 million yen in the budget line, which became more conspicuous after the tax reform, dragged even medium- to high-income married women, who were not presumed to be affected by the tax reform, to the mass point when there was a negative shock in an exogenous variable that could potentially discourage them from working long hours, i.e., an increase in husbands' incomes. Therefore, the increase in income among low-income married women was offset by the decrease in income among medium- to high-income married women, which resulted in the slight decrease in the average income among married women.

However, we must not forget that the tax reform still mitigated the decrease in incomes even when the "income jump" (i.e., the discontinuous income decrease from 1.41 million yen or more to 1.03 million yen among medium- to high-income married women in response to the increase in their husbands' incomes) occurred: the tax reform contributed to reducing the dispersion in the income distribution, making the mass point at 1.03 million yen in the income distribution more conspicuous than before.

This paper proceeds as follows: Section 2.2 describes the tax system in Japan, and Section 2.3 gives some information about trends in the Japanese labor market over the period 2003-2006. Section 2.4 sets forth a simple theoretical model and theoretical implications. Section 2.5 describes a strategy to test the implications of the theoretical model. Section 2.6 outlines a brief description of the data, and Section 2.7 discusses the results from the empirical analysis. Finally, Section 2.8 sets forth the conclusions.

2.2 Tax System in Japan

In this section, general information about the Japanese tax system is provided. First, married people file separate tax returns and two-earner couples are taxed as separate individuals. For one-earner couples with several dependents, the head of the household is treated as a single taxpayer. Second, income from assets is taxed separately from earnings. Thus, the incomes analyzed in this paper refer only to income from labor earnings. Third, there are income exemption systems and policy-related customs that could potentially discourage married women from working long hours. In particular, 1.03 million yen (≈USD 10,300) has historically been a critical income threshold for many married women, which is discussed in Section 2.2.3.

2.2.1 Income Tax and Residential Tax

Before moving on to the explanation of the spousal-exemption system in Japan, I will first introduce the income tax and residential tax, both of which have spousal exemption systems.

Income Tax

The income tax is a national tax, which is levied based on total earnings during the year minus certain exemptions and deductions. More specifically, the income tax is calculated as follows: {total earnings − *employment income deduction* (in Table 2.1.A) − *basic deduction* (of 0.38 million yen) − other deductions/exemptions (if applicable)[5]} × tax rate (in Table 2.1.B) −deduction in Table 2.1.B−tax credit (if applicable)[6].

The amount of the *employment income deduction* can be formulated as $aY + b$, where Y is the annual earnings, and its tax deduction schedule is listed in Table 2.1.A.[7]

[5] Other deductions/exemptions include the following: exemptions for spouse; exemption for dependents; deduction for casualty losses; deduction for medical expenses; deduction for life insurance premium; deduction for social insurance premium; deduction for fire and other casualty insurance premium; deductions for small-scale business enterprise mutual aid premiums; exemptions for handicapped person, widow, widower, or working student; exemption for aged person (abolished in 2005); and deductions for donations.

[6] If any tax credit is applicable, we can subtract the tax credit(s) directly from the income tax. The tax credits during the period 2003-2006 include tax credits for dividends; a foreign tax credit; and a special credit for acquisition of dwelling house. In 2006, a special deduction for residence seismic retrofitting was added to these tax credits (National Tax Agency, 2003, 2004, 2005, 2006).

[7] If the employed workers have made the following specified expenses during the year and if the total amount of the expenses exceeds the amount of the *employment income deduction*, the excess amount can also be deducted from the *total income*: (1) commuting expenses, (2) moving expenses upon job transfer, (3) study and training expenses, (4) expenses in acquiring qualifications, and (5) home visiting expenses for workers living away from home.

All employees are eligible for the *employment income deduction* of 0.65 million yen. If the deduction amount calculated in Table 2.1.A is less than 0.65 million yen, it is set at 0.65 million yen.

Furthermore, all workers in Japan are also eligible for the *basic deduction* of 0.38 million yen, and thus the total amount of deductions that can be applied to any employed worker is calculated as the minimum *employment income deduction* of 0.65 million yen plus the *basic deduction* of 0.38 million yen. Due to this deduction schedule, employees begin to pay income tax only after their annual earnings exceed 1.03 million yen.

Although I have focused on the discussion of employees due to the simplicity of the corresponding tax calculation, workers other than employees, of course, pay income tax. For these workers, the *employment income deduction* is replaced by the necessary expenses. To consider both types of workers, tax-exemption schedules and the limit of nontaxable income are usually defined by the amount of the *total income*, which is calculated as the total earnings minus the *employment income deduction* for employees and the total earnings minus necessary expenses for other workers. If we use the measure of *total income* instead of the total earnings, the limit of nontaxable income for the income tax is 0.38 million yen (=1.03 million yen − the minimum *employment income deduction* of 0.65 million yen).

Residential Tax

In addition to the income tax, residents in Japan are required to pay "residential tax," which consists of prefectural inhabitant tax and municipal inhabitant tax. Japanese residents pay these taxes to the municipality of residence. The residential taxes are divided into an income-based component and a uniform per-capita component. The income-based component of the residential tax is levied based on the previous year's income, and its tax rate schedule is presented in Table 2.1.C. The taxable income is calculated in the same way as the income tax, but the amount of the *basic deduction* for the residential tax is 0.33 million yen instead of 0.38 million yen. While the limit of nontaxable income for the income tax was 0.38 million yen of the *total income*, the limit of nontaxable income for the income-based component residential tax is 0.35 million yen (which corresponds to 1 million yen of annual earnings). Although the amount of the per-capita component varies among municipalities, the standard amount is about 4,000 yen (≈USD 40) per year.[8]

2.2.2 Spousal Exemptions in Japan

Both the income tax and the residential tax in Japan have spousal exemptions. Within each spousal exemption system, there are two types of spousal exemptions: the standard spousal exemption and the special spousal exemption.

The head of the household is eligible for the special spousal exemption if his/her own *total income* is 10 million yen or less per year and his/her spouse's *total income* is below certain thresholds.[9] For the sake of simplicity, in this paper I will assume that the head of household is a husband and the spouse is a wife.[10]

Table 2.2 shows the amount of spousal exemptions by the amount of spouse's annual earnings.[11] The un-parenthesized numbers represent the income-tax spousal exemptions and the parenthesized numbers represent the residential-tax spousal exemptions. The standard spousal exemption of income tax is fixed at 0.38 million yen

[8]The limit of nontaxable income for the per-capita component also varies among municipalities.

[9]There is no rule about whether the husband or the wife should claim the spousal exemption, but to be eligible for the exemption, the spouse's earnings should be lower than the threshold (1.41 million yen≈14,100 USD). Thus, in most cases, the spouse who has the highest income claims the spousal exemption, as this person is more likely to be eligible for the spousal exemption if he or she does so.

[10]According to the Keio Household Panel Survey conducted between 2004 and 2007, 96 % of married couples listed the husband as the head of the household for tax purposes.

[11]The schedule of the spousal exemption by "spouse's annual earnings" is applicable only for employees. The schedule applicable to all workers is defined by *total income*, not by the total earnings, which is calculated as the total earnings minus the *employment income deduction* for employees and the total earnings minus necessary expenses for other workers. For simplicity, here I just present the schedule for employees.

(≈USD 3,800) and ceases immediately if a wife's earnings exceed 1.03 million yen (≈USD 10,300). In contrast, the special spousal exemption declines gradually as the wife's earnings increase.

As illustrated in Table 2.2, the spousal exemption system for the residential tax is almost the same as that for the income tax, except that in the case of the residential tax, the starting amount of the spousal exemption is 0.33 million yen instead of 0.38 million yen.

Historically, the standard spousal exemption was established in 1961 to support households that have full-time housewives by offering more spousal tax exemptions to a husband whose wife worked less. However, because of the characteristic of the standard spousal exemption that it ceases at a certain threshold of the wife's income, there was a problem, in that after-tax incomes of a household in which the wife earned slightly above the threshold would be less than those of a household in which the wife earned slightly less than the threshold. This created a "dip" at the threshold on the household budget line.

To fix this "reversal" phenomenon in after-tax incomes around the threshold, the special spousal exemption was introduced in the late 1980s. As a result, the dip on the household budget line that was caused by the standard spousal exemption was removed (Abe and Otake, 1995). However, the dip on the household budget line at the threshold, currently at 1.03 million yen, actually still exists because most firms set the income threshold for losing spousal allowances at 1.03 million yen. This point will be discussed in Section 2.2.3. Note that both the standard and special spousal exemptions have characteristics that could potentially discourage married women from working long hours. This aspect of the spousal exemption system has been under a criticism for a long time, and the criticism has escalated together with the trend of an increase in the female labor participation rate.

As a result, at the end of March 2003, the Japanese Diet passed legislation to abolish part of the spousal exemption for taxpayers whose spouses had annual earnings of 1.03 million yen or less (≈USD 10,300). This legislation took effect in the 2004 tax year. (There was no change in the standard spousal exemption.) In the tax reform of 2004, the upper block of the special spousal exemption schedule in Table 2.2 was abolished. This abolition of part of the special spousal exemptions took effect from the 2004 fiscal year for the income tax and from the 2005 fiscal year for the residential tax, respectively.

2.2.3 Factors that Could Potentially Induce Labor Supply Adjustment

In Japan, there are several income- and work-hour thresholds at which people lose/gain some benefits, such as tax-waivers, exemptions, and allowances etc. These income- and work-hour thresholds potentially create incentives for people to adjust their incomes or work hours to just below or above the thresholds. In this section, I will review the major factors that could potentially cause people to adjust their labor supply and demonstrate how these factors affect income distribution.

A. Limit of Nontaxable Income

As was confirmed in Section 2.2.1, all employees in Japan begin to pay income tax only after their annual earnings exceed 1.03 million yen. This implies the possibility that some employees adjust their work hours so that their annual earnings do not exceed 1.03 million yen.[12]

[12] Although the limit of nontaxable income for the residential tax (i.e., 1 million yen) could also induce labor supply adjustment, the amount of the residential tax is smaller than that of income tax due to its smaller tax rate, and hence, if people mind a limit of nontaxable income, that for the income tax, in general, matters more. Thus, the case in which people adjust labor supply to avoid residential tax payments is considered to be a minor case, and hence it is included in the "*F. Other*" category, rather than in Factor *A*.

B. Spousal Exemption

As was confirmed in Section 2.2.2, husbands lose the standard spousal exemption once their spouses' annual earnings exceed 1.03 million yen. Additionally, since the tax reform, 1.03 million yen has been the point at which the total spousal exemption amounts start to decrease as spouses' incomes increase.

C. Social Security Payments

The following three categories comprise the social security system in Japan:

Category I: Self-employed persons, farmers, etc., age 20 or over but under 60.

Category II: Private company employees and public officers.

Category III: Dependent spouses of private company employees and public officers.

C-1. Choice Between Category I and III (Threshold: 1.3 Million Yen Ceiling)

Wives whose annual income is less than 1.3 million yen are classified into Category III and are treated as their husbands' dependents; they do not need to make payments for their own pension and health insurance. However, if their annual income equals or exceeds 1.3 million yen, then they will shift to Category I, in which they are required to pay for the national pension and the national health insurance.

The national pension as of 2003 is a lump-sum tax of 13,300 yen per month. The calculation of the national health insurance varies among municipalities. When the calculation method using the former corrected income system (Type 4) is assumed, the national health insurance is composed of an income-based component, a household property-based component, an insurance receiver per-capita component, and a household per-capita component. According to the Survey on National Health Insurance 2003 (Ministry of Health, Labor, and Welfare, Government of Japan, 2003b), as of 2003 the average premium rate of the income-based component, the household property-based component, the insurance receiver per-capita component, and the household per-capita component are 9.151 %, 9,989 yen, 24,762 yen, and 20,083 yen, respectively.

C-2. Choice Between Category II and the Other Two Categories (Threshold: 3/4 of Regular Employee's Work Hours)

In contrast, if one's work hours exceed 3/4 of a regular employee's work hours, that worker is classified into Category II, in which people are required to enroll for the employees' pension- and health-insurance systems (Welfare Pension and Health Insurance). The amounts of these premiums are proportional to their earnings and halved between employers and employees. As of 2003, the premium rates only for employees are 6.79 % for welfare pension and 4.1 % for health insurance. There might be a case that some employers limit employees' hours to avoid the enrollment in these employees' pension and health insurance systems. These cases are included in the "*F. Other*" category, rather than in Factor C–2.

D. Employer-Provided Spousal Allowances

According to a survey conducted by the Japanese Cabinet Office in 2001, the fraction of firms that have a family allowance system is 83.5 % on average, and 91.5 % among large firms with 1000 or more employees. The rule for the eligibility of employer-provided spousal allowances differs among employers: 61.5 % of firms that offer family allowances determine the eligibility for spousal allowances based on spouses' incomes.[13] Among these firms, 78.4 % of firms set 1.03 million yen as the

[13] The way employers know how much their employees' wives earned depends on the individual rules of particular companies. In many cases, each employee reports his wife's expected annual

critical income threshold at which spousal allowance vanishes because it is the income threshold for the standard spousal exemption in tax. 13.9 % of firms set the income threshold at that for the social security payment, i.e., 1.3 million yen.[14] The average amount of spousal allowances per month is 14,500 yen (≈USD 145) (Cabinet Office, Government of Japan, 2001), and it is known that this spousal allowance is also a factor that could potentially make a spike at 1.03 million yen in a wife's earning distribution (Takahashi et al., 2009).

E. Employment Insurance Premium

The employment insurance system regulates that all workers who work for 20 hours or more per week and are expected to stay employed for more than a certain length of time are covered by the employment insurance.[15]

F. Other

The remaining reasons include the following:

(a) To avoid enrollment in the employment insurance or welfare pension system due to reasons originating within the firm

(b) To avoid the reduction in the pension amount currently provided

(c) Other

income for the year during the month of March since a fiscal year starts on April 1 in Japan. If the amount that the wife has actually earned by December is significantly different from the estimate provided in March, the employee has to return the spousal allowance that he has received. When the wife's annual income unexpectedly exceeds 1.03 million yen by only a slight margin, in many cases, the employee is not required to return the spousal allowance.

[14] 5.8 % of firms have different income thresholds and the remaining 1.9 % offered no answer to the question.

[15] The expected length of employment regulated in the conditions has been revised several times. The minimum length was set at one year prior to 2009 and changed to six months in 2009. In 2010, it was further revised to 31 days.

Figure 2.1 summarizes these income/work-hour thresholds and the amount of each payment as of 2003. The Y-axis represents the amount of wives' annual earnings and the X-axis represents weekly work hours. If they choose to earn 1.03 million yen or less, their income tax becomes zero, and they also become eligible for the standard spousal exemption and employer-provided spousal allowances, although the criteria for spousal allowances differ among firms. Furthermore, they can choose among the three categories of the social security system by adjusting their incomes to below or above 1.3 million yen and their work hours to below or above 3/4 of a regular employee's work hours.

Figure 2.2 reports how each factor ($A \sim F$) affected labor supply of female workers in 2006. The data is obtained from the General Survey on Part-time Workers 2006 (Ministry of Health, Labor, and Welfare, Government of Japan, 2006). The sample is restricted to female part-time workers who answered that they adjusted their labor supply during the previous year. The height of each bin represents the percentage of those who answered each factor ($A \sim F$) affected their decisions of labor supply adjustment. According to Figure 2.2, the limit of nontaxable income (1.03 million yen) is the most influential factor affecting the labor supply because this is applicable to both single and married women. The spousal exemption schedule is the second most influential factor. Considering that the standard spousal exemption vanishes at 1.03 million yen and that the majority of firms set the income threshold for losing spousal allowances at that amount, in most cases it can be said that, in addition to Factor A, Factor B and Factor D also refer to the 1.03 million yen ceiling. From this evidence, we understand that 1.03 million yen serves as a critical income threshold for many female workers.

Figure 2.3 presents a histogram of 2003 income distribution of married women. The sample comprises married women in the Keio Household Panel Survey (KHPS). In Figure 2.3, there is a conspicuous spike at around 1 million yen, just below the

1.03 million yen threshold.[16] This reflects the fact confirmed in Figure 2.2 that there are many women who adjust their labor supply so that their annual incomes do not exceed 1.03 million yen. Furthermore, since this is the income distribution prior to the tax reform, there is another spike at the old kink in the household budget line, i.e., 0.7 million yen. This point will be discussed in Section 2.4. Throughout this paper, I will examine what happened to married women originally at the spike of 0.7 million yen after the tax reform.

In addition to these two spikes, there is also a spike, albeit somewhat smaller, at just below 1.3 million yen (Abe and Otake, 1995; Oishi, 2003). This is because married women have to make a social security payment by themselves once their income exceeds 1.3 million yen (i.e., Factor C-1). Thus, 1.3 million yen also serves as a threshold of annual income in Japan, and this will also be considered in Section 2.4.

2.3 Evidence

In this section, I will review the overall trend of the Japanese labor market from 2003 to 2006 and I will confirm how the statistics changed after the tax reform.

2.3.1 Data

The data set used in this study is Japanese panel data from Keio Household Panel Survey (KHPS) 2004-2007. The KHPS data are relatively new as the data collection for the survey began in 2004. This survey is conducted on January 31 every year and includes observations randomly chosen from almost all regions and industries in Japan. A key feature of KHPS is that it is the first nationwide longitudinal survey in Japan of individuals (4,000 households and 7,000 people) of all ages and both sexes,

[16] There are many studies that point out a considerable cluster of observations near this amount (Abe and Otake, 1995; Kantani, 1997; Higuchi, 1995; Takahashi, 2010a,b).

which also captures information regarding education, employment, income, expenses, health, and family structure. The survey has been designed to enable comparisons with major international panel surveys such as Panel Study of Income Dynamics (PSID) and European Community Household Panel (ECHP).

Since the survey asks about information concerning the previous year, information prior to the tax reform can be obtained from the 2004 survey. For example, annual incomes in the previous year are reported in this survey, and this information is used as incomes workers had originally earned before the tax reform.[17] However, only a question about the average work hours per week does not specify if it is asking about work hours in the previous year. That question simply asks: "How long do you work per week on average?" The survey is held every January and, thus, there is a possibility that some respondents provided information about their average work hours during January of the year in which the survey was held, instead of their work hours from the previous year. The concern regarding this possibility would be that work hours provided by respondents in the 2004 survey have already been affected by the tax reform. To circumvent this problem, the main focus of this paper is the analyses of annual incomes.

2.3.2 Japanese Labor Market Trends: 2003-2006

As stated in the previous section, there is a concern about using work hours reported in the January 2004 survey as the information prior to the tax reform. However, even if respondents have already started to adjust their work hours in January 2004, the average work hours in January 2004 should also include information about their 2003 work hours because it is unlikely that an individual's work hours greatly

[17] As is the case with the US Current Population Survey, the Keio Household Panel Survey collects information on the previous year's income approximately one month before income taxes are due. Thus, on the date the survey was conducted, it could be thought that most people had already calculated (or at least roughly understood) their annual income in order to be ready to submit their tax documents. Therefore, this paper proceeds assuming that the misreporting of incomes is not a serious issue.

change beginning on January 1.[18] For this reason, in this section that reviews the trend of the Japanese labor market during the period 2003-2006, I use the information about work hours provided in January 2004 as being reflective of the respondent's work hours prior to the tax reform.

During the period 2003-2006, Japan was recovering from a recession that started in the late 1990s. Table 2.3.A summarizes trends of the Japanese labor market during this period. The unemployment rate decreased significantly from 5.3 % to 4.1 %. The other statistics are calculated from the Keio Household Panel Survey (KHPS) 2004-2007. The sample is 2,896 women who participated in the Keio Household Panel Survey (KHPS) who reported information for both periods: before and after the 2004 tax reform. (The statistics from the period after the tax reform are calculated as the average among the three years immediately following the tax reform.) Thus, the differences in each of the statistics presented in the last column represent the person-specific variation of each of the statistics, which is free from the change in the composition of the sample. For example, each married women decreased weekly work hours by 1.69 on average over the periods.

As was discussed in Section 2.2.2, spousal exemptions were initially designed to support the nuclear family, in which full-time housewives specialized in domestic work and childbearing by offering more spousal tax exemptions for a husband whose wife worked less. Because this system has characteristics that could discourage married women from working long hours, it was expected that the married women's labor supply would increase after abolishing part of the special spousal exemption.

However, both average annual income and average work hours slightly decreased among married women after the tax reform, although this trend is not observed for

[18]Even if they try to adjust their work hours so that their annual income does not exceed a certain threshold, it is unlikely that people start to adjust their work hours from the very beginning of each year to control their annual income; rather, it is more natural to think that people start to adjust their labor supply in the last half of the year based on the amount they have earned by that point in time.

single women; thus, we cannot see any evidence of increased income and work hours among married women. During this period, there was little change in the average hourly wage rate for both married women and single women. In contrast, husbands' incomes increased during the same period, and the change is significantly different from zero.[19] Since this increase in husbands' incomes could be one reason for the decrease in wives' work hours, it is worth checking the possibility that this change in husbands' incomes concealed the positive impact of the tax reform on wives' work hours.

2.3.3 The Effect of the Husband's Income Increase vs. the Effect of the Spousal-Exemption Cut

In this section, I will compare the magnitude of the effect of the increase in the husbands' incomes with the effect of the spousal-exemption cut, using a simple calculation. The increase in husbands' annual before-tax earnings is about 0.13 million yen (≈USD 1,300), and for those with a husband whose marginal tax rate is 20 %, the effect of the increase in husbands' incomes on household income would be 0.104 million yen (≈USD 1,040) after tax.

Table 2.3.B summarizes the changes in the total amount of spousal exemptions for both income tax and residential tax. The sample consists of 2,135 married women from the Keio Household Panel Survey (KHPS) whose husbands' spousal exemption amount was greater than zero in 2003. The amount is calculated based on their 2003 annual income. The amount of spousal exemption after the tax reform is a simulated

[19] During the period from 2003 to 2006, husbands' annual incomes increased, while their work hours stayed almost the same. Because their hourly wage rate also increased during the same period, it is natural to assume that the increase in husbands' annual incomes was caused by the exogenous shock to their hourly wage in response to the recovering market conditions. In contrast, the increase in wives' hourly wage was very small: 1.6 %. This difference between males and females might reflect the fact that most hourly paid female workers are paid the minimum wage or slightly above the minimum wage. Thus, the market's recovery may not have been fully reflected in the hourly wages of female workers. Although the increase in hourly wage rates among female workers was small enough to be negligible, I also conducted theoretical and empirical analyses that allow for fluctuations in hourly wages, which did not change the main results.

amount of spousal exemption each woman would receive when her earnings did not change over the period. Thus, the difference of the spousal exemption between the periods reflects only the impact of the tax reform. The average loss of the total spousal exemption for income tax is 0.27 million yen (≈USD 2,700), and hence the effect of the spousal exemption cut on household income is 20 %×0.27 million yen=0.054 million yen (≈USD 540) on average and 0.076 million yen (≈USD 760) maximum. If we consider the spousal exemption for residential tax, the average loss of the spousal exemption is 0.24 million yen (≈USD 2,400) on average and 0.33 million yen (≈USD 3,300) maximum. The tax rate for the residential tax applicable to the average amount of the husband's income is 10 %. Thus, the effect of the reduction in the spousal exemption of the residential tax on household income is 10 %×0.24 million yen=0.024 million yen (≈USD 240) on average and 0.033 million yen (≈USD 330) maximum. Combining the effects for income tax and residential tax, the total effect of the spousal exemption cut amounts to 0.078 million yen (≈USD 780) on average and 0.109 million yen (≈USD 1,090) maximum.

Therefore, if we compare the effect of the spousal exemption cut and the effect of an increase in husbands' incomes, we notice that they almost offset each other. This simple calculation implies that looking only at the average income and work hours without controlling for the increase in the husbands' incomes might result in missing the increase in labor supply due to the tax reform, because the husbands' income effects might offset the tax reform effect.

In this paper, I will look more closely at what happened to the overall income distribution after the tax reform, taking into account the changes in exogenous variables.

2.4 Theory

In this section, I will consider a utility maximization problem of a wife who chooses the optimal combination of leisure and consumption given her husband's income, and I will simulate how the change in the spousal exemption system affects the wife's income choice.

2.4.1 The Tax Reform's Effect on the Optimal Income

It is assumed that a wife chooses work hours that maximize her utility taking her husband's income as given.[20][21] The *employment income deduction* can be formulated as: $aY + b$, where Y is annual income. Let w be a wife's hourly wage, h be a wife's work hours, t_W be a wife's income tax rate, and t_H be a husband's income tax rate. Then total household income, I_{HH}, can be written as:

$$I_{HH} = wh - max\{0, [wh-(awh+b)-0.38]t_W\} + I_H - (I_H - D - SE(h))t_H \quad (2.1)$$

where I_H represents the husband's annual income, and D represents income deductions that the husband can claim, other than the spousal exemption. SE is the

[20] In the literature on female labor supply, it has been common to assume that the husband's decision is predetermined or exogenous (Mincer, 1962; Killingsworth, 1983; Hyslop, 1999; Takahashi et al., 2009; Apps et al., 2012). In contrast, many recent studies analyzing households' decisions consider collective models that treat a family as a collection of agents with specific preferences (Chiappori (1988) and Chiappori (1992) etc.). However, Donni and Moreau (2007) emphasize that in most developed countries, the male labor supply is rigid and largely determined by exogenous constraints and that in this case, use of the collective model could be misleading. This is the case for the Japanese labor market as well: it is known that the male labor supply in Japan has historically been much more inelastic than the female labor supply. Kuroda and Yamamoto (2008), who estimated Frisch labor supply elasticity by gender, found that elasticity was 0.7 for males and 1.3 to 1.5 for females during the period from 1992 to 2002. Reflecting this fact, it is still common to assume the husband's exogenous labor supply in the context of a model of the female labor supply in Japan (Takahashi et al., 2009). Thus, in this paper, I assume that imposing the assumption of the husband's exogenous labor supply is not a serious issue for consideration.

[21] The unilateral influence on labor decision (from a husband to a wife) is less likely to be true in more modernized double-income couples in which the wife also earns a high income. However, in this paper, wives of interest are those with low incomes because they were the group affected by the tax reform; these women are more likely to belong to traditional households in which the husband is the main earner and the wife is the second earner. It is thought that in these households, the assumption above is likely to be true.

amount of spousal exemption. Note that the wife's choice of work hours will affect the amount of the spousal exemption; therefore, SE can be written as a function of a wife's work hours, h. In Equation (2.1), the household income is calculated as the summation of a wife's annual after-tax income (=a wife's annual earnings − income tax (if applicable)) and a husband's after-tax income (=a husband's annual earnings − income tax). The term $(awh + b)$ represents the *employment income deduction*, and 0.38 represents the amount of the *basic deduction* (million yen/ year). Since all employees are eligible for the basic deduction of 0.38 million yen and the *employment income deduction* of 0.65 million yen, the income tax becomes greater than zero when annual earnings exceed 1.03 million yen. In Equation (2.1), it is assumed that the tax credit is zero.[22][23]

According to Table 2.2, the spousal exemption (SE) can be calculated as:

- Before the Tax Reform (∼2003)

$$SE = \begin{cases} 0.76 & for\ 0 \leq wh < 0.7 \\ 1.41 - wh & for\ 0.7 \leq wh \leq 1.03 \\ 1.41 - wh & for\ 1.03 < wh < 1.41 \\ 0 & for\ 1.41 \leq wh \end{cases} \tag{2.2}$$

- After the Tax Reform (2004∼)

$$SE = \begin{cases} 0.38 & for\ 0 \leq wh < 0.7 \\ 0.38 & for\ 0.7 \leq wh \leq 1.03 \\ 1.41 - wh & for\ 1.03 < wh < 1.41 \\ 0 & for\ 1.41 \leq wh \end{cases} \tag{2.3}$$

[22] Since there is no tax credit that depends on the wife's work hours, this assumption does not change the result.

[23] Although residential tax, spousal allowances, and social security payments are not included in Equation (2.1), these will be considered in simulations performed in the latter part of this section.

where wh is the wife's annual income, and the unit is million yen (≈USD 10,000).[24] Substituting Equations (2.2) and (2.3) into Equation (2.1), household income before and after the tax reform can be calculated as:

- Household Income Before the Tax Reform (∼2003)

$$=\begin{cases} wh+0.76\,t_H+I_H(1-t_H)+Dt_H & \textit{for } 0\leq wh<0.7 \\ w(1-t_H)h+1.41\,t_H+I_H(1-t_H)+Dt_H & \textit{for } 0.7\leq wh\leq 1.03 \\ w[1-(1-a)t_W-t_H]h+(0.38+b)t_W+1.41t_H+I_H(1-t_H)+Dt_H & \textit{for } 1.03<wh<1.41 \\ w[1-(1-a)t_W]h+(0.38+b)t_W+I_H(1-t_H)+Dt_H & \textit{for } 1.41\leq wh \end{cases}$$

- Household Income After the Tax Reform (2004∼)

$$=\begin{cases} wh+0.38\,t_H+I_H(1-t_H)+Dt_H & \textit{for } 0\leq wh<0.7 \\ wh+0.38\,t_H+I_H(1-t_H)+Dt_H & \textit{for } 0.7\leq wh\leq 1.03 \\ w[1-(1-a)t_W-t_H]h+(0.38+b)t_W+1.41t_H+I_H(1-t_H)+Dt_H & \textit{for } 1.03<wh<1.41 \\ w[1-(1-a)t_W]h+(0.38+b)t_W+I_H(1-t_H)+Dt_H & \textit{for } 1.41\leq wh \end{cases}$$

Figure 2.4a illustrates the household budget lines before and after the 2004 tax reform. Since the income tax becomes greater than zero at annual earnings of 1.03 million yen, this creates a kink at 1.03 million yen in both of the budget lines. Comparing the two budget lines, we can clearly see the impact of the tax reform on the household budget line. Since wives who had a net annual income of 1.03 million yen or less lost the special spousal exemption after the tax reform, the intercept of the budget line for these people shifted down. Furthermore, since the special spousal exemption, where the amount had been reduced in stepwise fashion as the spouse's income rose before the tax reform, was abolished for those who earned between 0.7

[24]Here, I adopt the model structure assumed in Takahashi et al. (2009), that examines the effect of the spousal exemption itself (on the female labor supply), not the effect of the reform in the spousal exemption system. Thus, in Takahashi et al. (2009), only the household budget line prior to the tax reform is considered, while in this paper, both household budget lines before and after the tax reform are considered.

and 1.03 million yen in 2004, the budget line for this income group became steeper after the tax reform. This moved the kink in the household budget lines, caused by the spousal exemption system, from 0.7 million yen to 1.03 million. As a result, the kink at 1.03 million yen that had originally existed due to the limit of nontaxable income became more conspicuous after the tax reform.

To consider a more realistic budget line, in Figure 2.4b I also include a social security payment and spousal allowance. In Figure 2.4b, the dip at income equal to 1.03 million yen is caused by the loss of spousal allowance that the husband receives from his employer at this point. Furthermore, married women have to make a social security payment once their income exceeds 1.3 million yen, which creates a dip at income equal to 1.3 million yen.[25]

Next, I will confirm the impact of the tax reform on labor supply for each income group.

Proposition 2.1

Without a change in exogenous variables, such as husbands' incomes and the price of goods, the tax reform has no impact on those who originally earned more than 1.03 million yen but increases work hours for married women who originally earned 1.03 million yen or less.

Proof

Figure 2.5a depicts the impact of the tax reform on the labor supply of married women who earn 1.03 million yen or less annually, in the setting in which there is no change in the exogenous variables. If leisure is a normal good, married women who annually earned less than 0.7 million yen prior to 2004 work more after the tax reform because of the positive income effect (i.e., people at point A will move to point B after the tax reform in Figure 2.5a). Married women who annually earned between 0.7 and 1.03 million yen prior to 2004 work more after the tax reform because of

[25] The details about these dips are discussed in Section 2.2.3.

the positive substitution effect (movement from point C to point D) and the positive income effect (movement from point D to point E).

In contrast, as Figure 2.5b shows, for married women who earn more than 1.03 million yen annually, there is no change in their household budget line after the tax reform. Thus, the tax reform does not affect labor supply of these women. Q.E.D.

In Proposition 2.1, it is assumed there is no change in exogenous variables. However, Table 2.3.A confirmed that married women experienced an increase in their husbands' incomes as well as the tax reform during the period 2003-2006. If such a change in an exogenous variable is considered, there might be a case in which the tax reform affects the labor supply of married women who earned more than 1.03 million yen before the tax reform. The next proposition states what could potentially happen to the labor supply of married women who were not presumed to be affected by the tax reform when they experience an increase in their husbands' income as well as the tax reform.

Proposition 2.2

(a) After the tax reform, when an increase in a husband's income is considered, there are cases in which married women who earned 1.41 million yen or more greatly decrease their work hours to earn only 1.03 million yen.

(b) These cases can be realized only when the increase in a husband's income is sufficiently large and a wife's preference for consumption (or a household's preference for the wife's consumption) is not "too strong."

Proof of (a)

Figure 2.6a explains Proposition 2.2a graphically. The increase in a husband's income makes households richer and shifts household budget lines to the right.[26] As long as

[26] At the same time, the husband's marginal tax rate might also increase for some households that experience an income increase. This will make the budget line flatter for the wives, whose annual

leisure is a normal good, a wife decreases the number of work hours when the husband's income increases. Given an increase in a husband's income, wife A −originally at point A− will decrease the number of work hours to point A' by the income effect. In contrast, wife B −originally at point B− will jump to point B' if her husband's income increases after the tax reform. When comparing these two wives, wife A has the stronger preference for consumption (i.e., her indifference curve is flatter at each value of leisure than wife B's curve).[27]

I will confirm that the case illustrated in Figure 2.6a is possible by simulating the wives' optimal income choices with the following four representative utility functions:

1. Cobb-Douglas $\quad u(c,l) = c^{\alpha} l^{(1-\alpha)} \quad 0 < \alpha < 1$

2. CES $\quad u(c,l) = \{\alpha c^{\rho} + (1-\alpha) l^{\rho}\}^{1/\rho} \quad \rho \neq 0,\ \rho < 1,\ 0 < \alpha < 1$

3. Power function (CRRA) $\quad u(c,l) = \frac{c^{1-\gamma}}{1-\gamma} + \frac{l^{1-\gamma}}{1-\gamma},\ \ \gamma > 0$

4. Stone-Geary $\quad u(c,l) = \gamma\, ln(c-A) + (1-\gamma)\, ln(l-B) \quad c > A,\ l > B$

I calculated the optimal income for attaining the highest utility for 100 different parameter values for each functional specification.[28] To simulate what happens to the optimal income when a husband's income increases, the husband's income in the simulation increases from 5.2 million yen to 5.4 million yen, which are the average amounts of husband's income of each period. The hourly wage and the price of goods

incomes are between 1.03 million yen and 1.41 million yen, making the kink at 1.03 million yen in the budget line more conspicuous. Thus, it is more likely that the "income jump" in Proposition 2.2 will be observed when we also consider a change in the husband's tax rate.

[27]Here, a wife's preference for consumption is considered, since it is assumed that a wife solves her utility maximization problem by taking her husband's income as given. However, considering this as the household's preference for the wife's consumption will not change the result, i.e., even in the case that the way a wife works reflects her husband's preference about this matter as well, the result will not change.

[28]α used in the Cobb-Douglas and Power function (CRRA) functions, and γ in the Stone-Geary function are uniformly spaced on [0.1]. For the CES function, α is fixed at 0.5, and ρ is uniformly spaced on [-0.5, 0.5] (a case with ρ=0 was omitted). 1.95 million yen and 1 are used for the subsistence levels of consumption and leisure, respectively in the Stone-Geary. 1.95 million yen was the amount of livelihood assistance for a standard household consisting of three members living in Tokyo as of 2003.

are fixed at 1,000 yen and 100 yen, respectively. In the calculation, social security payments, which include decision making among the three categories of the social security system, residential tax, employer-provided spousal allowance, and employment insurance premium, are also considered. Concerning these payments/allowances, the amounts specified in Figure 2.1 are assumed. Table 2.4.A presents the percentages of the 100 simulated optimal incomes that belong to each income group before and after the tax reform. For example, when using a Cobb-Douglas utility function, a wife choose to earn 1.41 million yen or more before the tax reform and 1.03 million yen after the tax reform in one out of a 100 cases.

According to Table 2.4.A, the majority of the income changes for the highest income group are made within the group (i.e., changes from point A to point A' in Figure 2.6a). However, for all of the utility functions, we can observe a case in which women who earned 1.41 million yen or more before the tax reform will decrease their income to 1.03 million yen after the tax reform (i.e., the changes from point B to point B' in Figure 2.6a). Q.E.D.

Proof of (b)

Table 2.4.B reports the probability that married women in the highest income group will decrease their incomes to 1.03 million yen after the tax reform, conditional on having been in the highest income group prior to the tax reform. Thus, the number of parameter values with which the "income jump" occurs is divided by the total number of parameter values that attain annual incomes of 1.41 million yen or more, prior to the tax reform. For example, as in the Cobb-Douglas case, when there is an increase in the husband's income of 0.2 million yen, the probability of the "income jump" among high-income married women is calculated as 1/(1+40) using the numbers in Table 2.4.A. This calculation is useful in predicting how often the "income jump" occurs among people who originally earned 1.41 million yen or more prior to the tax reform. This is because, in the actual data, there are people who already earned 1.41

million yen or more prior to the tax reform, and the numbers presented in Table 2.4.B predict how the incomes of these people were affected by the tax reform.

As can be confirmed in Table 2.4.B, the greater the husband's income increase, the higher the probability of the "income jump" becomes. When the increase in a husband's annual income is only 0.1 million yen ($\approx$USD 1000), the "income jump" will never occur for almost all of the utility functions, but if the increase in a husband's income is greater than 0.1 million yen, the "income jump" occurs at a positive probability for all of the utility functions. Thus, we can say that the "income jump" stated in Proposition 2.2a can be realized only when the increase in a husband's income is sufficiently large.

Next, I will show that these cases are possible only when the preference for consumption is not "too strong." Figure 2.6b shows how wives' optimal incomes will change if the tax reform occurs and their husbands' incomes increase. To make the result as obvious as possible, Figure 2.6b reports the results with a sufficiently large increase in husbands' incomes, i.e., increase of 2 million yen. Except for a husband's income after the tax reform, all parameters are set at the same as those in Table 2.4.A. The dots in Figure 2.6b show the amount of optimal income that attains the highest utility for each value of α (i.e., the preference for consumption) when we use a Cobb-Douglas utility function. It is obvious that the higher the parameter value, α, the higher are the incomes that people choose. In Figure 2.6b, only individuals with α of 0.6 or higher earned 1.41 million yen or more before the tax reform. Thus, for an individual to be in the high income group originally, her preference for consumption has to be sufficiently strong.

Comparing optimal incomes before the tax reform with optimal incomes after the reform, we notice that individuals with α of 0.6-0.65, who originally earned 1.41 million yen or more a year, decrease their annual incomes to 1.03 million yen after the tax reform. These individuals correspond to "wife B" in Figure 2.6a. Thus, it can be said that there exists a region in the domain of α in which the "income jump" occurs. In contrast, if the preference for consumption is "too strong" (with α more than 0.65), the decrease in income is small enough that the income changes for the high-income group are made within the group (i.e., changes from point A to point A' in Figure 2.6a). Thus, it can be said that the "income jump" is possible only when their preference for consumption is sufficiently strong, but not "too strong." Q.E.D.

Proposition 2.2b provides an important implication about the levels of income and the impact of the tax reform. For a wife to be originally in the high-income group, her preference for consumption should be sufficiently strong, but if it is too strong, the decrease in income is small enough that the income changes are made within the group.

Individuals with very strong preferences for consumption are those who originally earn extremely high incomes (e.g., more than 10 million yen per year), and they are less likely to decrease their incomes to 1.03 million yen, even in response to an increase in their husbands' incomes and the tax reform. In this sense, we can say that the "income jump" is not true for those who earned extremely high incomes but is likely to be true for medium- to high-income married women.

In the next proposition, I will decompose the jump from 1.41 million yen or more to 1.03 million into a "husband's income effect" and the "tax reform effect."

Proposition 2.3

After the tax reform, when an increase in a husband's income is considered,

(a) the "income jump" in Proposition 2.2 is explained mainly by a large negative "husband's income effect";

(b) there are cases in which a decrease in the wife's income due to an increase in the husband's income is mitigated by a small positive "tax reform effect"; and

(c) there is no case in which the tax reform would have caused a further decrease in the wife's work hours. In other words, the "tax reform effect" cannot be negative.

Proof

Figure 2.7a illustrates a "husband's income effect" and a "tax reform effect." The long-dashed line represents the counterfactual budget line without the tax reform in 2004. In Figure 2.7a, a wife would have decreased her income from 1.41 million yen or more at point B, to 0.7 million yen at point C, if there had not been a tax reform in 2004. This movement from point B to point C can be thought of as a "husband's income effect," because this change was caused only by an increase in a husband's income. However, the tax reform actually occurred and not the long dashed budget line, but the bold budget line was realized in 2004. Thus, a wife that should have moved to point C will move to point B' when the tax reform occurs, since point C is no longer available. This movement from point C to point B' can be thought of as a "tax reform effect." Thus, a movement from point B to point B' is explained by a large negative "husband's income effect" and a small positive "tax reform effect." This "small positive tax reform effect" means that the decrease in income among medium- to high-income married women that occurs in response to an increase in the husbands' income is slightly mitigated by the tax reform.

Again, this will be confirmed using a simulation on the wives' optimal income choices with the four representative utility functions. In Table 2.4.C, the parenthesized numbers represent the "counterfactual" realizations of the optimal income (i.e., the percentages of the 100 simulated optimal incomes where the tax reform had not occurred in 2004). The settings are the same as those represented in Figure 2.6b. According to the result of the Cobb-Douglas utility function, in six cases out of 100, a wife who earned 1.41 million yen or more before the tax reform decreased her income to 1.03 million yen after the tax reform. In contrast, the counterfactual realizations of optimal income show that all six of these cases would have appeared in the category of "income jump" from 1.41 million yen or more to between 0.7-1.03 million yen without the tax reform; in other words, if the income earned by a married woman decreased from 1.41 million yen or more to 1.03 million yen after the tax reform, without the tax reform, her income would have decreased even further to 0.7 million yen, i.e., the kink that existed prior to the tax reform. We can observe similar results for the other preferences.[29]

Given these results, we would be inclined to conclude that the effect of the tax reform is always positive in any case; however, when we consider a small increase in a husband's income, there are also cases in which a negative "husband's income effect" is the only explanation for the "income jump," i.e., cases in which the "tax reform effect" is zero. In Table 2.4.C, although a sufficiently large increase in a husband's income of 2 million yen is considered, if the increase in a husband's income is as small as what is shown in Table 2.4.A, there are also cases in which the tax reform effect is zero, i.e., cases in which the counterfactual realizations appear in the same cell as the actual realizations.

However, we can at least say that the tax reform cannot be negative in any case. This will be confirmed using Figure 2.7b. Since both B' and C in Figure 2.7b are available on both the "after-reform" and counterfactual budget lines, a wife who

[29]Note that the kink at 1.03 million yen in the household budget line is more conspicuous when we assume a higher hourly wage rate. That kink is caused by the difference in the slopes of the budget line between the regions of 0.7-1.03 million yen and 1.03-1.41 million yen, and the difference of the slopes, $w\{(1-a)t_W - t_H\}$, becomes greater as the hourly wage rate, w, increases. Thus, the "income jump" (i.e., the discontinuous income decrease from 1.41 million yen or more to 1.03 million yen among medium- to high-income married women in response to the increase in their husbands' incomes) is more likely when we assume a higher hourly wage rate. In the current simulation, a rounded version of the average hourly wage rate among hourly-paid married women (1000 yen) is used for simplicity. However, even when we assume the actual average hourly wage rate (881 yen) of hourly-paid workers, the number of cases in which the "income jump" from 1.41 million yen or more to 1.03 million yen occurs does not change. When we assume the median hourly wage rate among hourly-paid workers, which is 800 yen, the number of cases representing the "income jump" decreases by one for each utility function in Table 2.4.C; however, it does not cause any other change. In contrast, when we assume a higher wage rate, e.g., 1,342 yen, which is the average hourly wage rate among all married women for all payment types, the number of the "income jump" increases by one for some of the utility functions.

would choose point C in the counterfactual budget line will never choose point B' on the actual budget line after the tax reform, because point C is more preferable to point B' for her. Thus, there is no case in which married women face a negative "tax reform effect" (a movement from C to B'). In other words, as long as the after-tax and counterfactual budget lines overlap in the region of more than 1.03 million yen, it is impossible for a wife who would choose point C on the counterfactual budget line to choose point B' in the actual budget line after the tax reform. Thus, if a married woman greatly decreased her income from 1.41 million yen or more to 1.03 million yen after the tax reform, she would have done the same thing even without the tax reform or would have decreased her income further to 0.7 million yen. Q.E.D.

In this section, it has been confirmed that the tax reform has a positive impact on work hours among low-income married women who originally earned 1.03 million yen or less prior to the tax reform. Without a change in exogenous variables, the tax reform would have no impact on work hours among married women who originally earned more than 1.03 million yen. If there is a change in exogenous variables, such as an increase in a husband's income, the tax reform could also affect work hours of medium- to high-income married women: the tax reform made the kink at 1.03 million yen in the budget line more conspicuous than before. As a result, there are cases in which medium- to high-income married women are dragged to the mass point at 1.03 million yen in the income distribution that had historically existed, when there is a negative shock in the exogenous variables that could potentially discourage them from working long hours, such as an increase in their husbands' incomes. Even in that case, the "tax reform effect" cannot be negative. Instead, the "income jump" among medium- to high-income married women can be explained mainly by a negative "husband's income effect" and, in some cases, the income drop is mitigated by a small positive "tax reform effect."

2.4.2 Introducing a Flexible Hourly Wage Rate

Up to this point, the hourly wage rate has been fixed in the simulation and the figures. Here, I show that the introduction of the flexible hourly wage rate does not change the theoretical implication. Table 2.4.D reports the probability of the "income jump" among married women who earned 1.41 million yen or more before the tax reform by husband's income increase and change in wife's hourly wage. From this table, we notice that the "income jump" can be realized as long as the increase in a wife's hourly wage is not too large. Thus, except for cases in which married women experience a large increase in their hourly wage rate, the main theoretical results are preserved. As was confirmed in Table 2.3.A, the increase in wives' hourly wage during the period from 2003 to 2006 was very small: 1.6 %. Thus, it is unlikely that ignoring such a small change in hourly wage rates matters in understanding the real impact of the tax reform.

2.4.3 The Tax Reform's Effect on Labor Participation

Next, I theoretically predict the effect of the tax reform on labor participation. As was confirmed in Figure 2.4, the tax reform did not change the starting wage rate but decreased the husband's after-tax income via the spousal exemption cut. This decrease in the husband's after-tax income can be thought of as a decrease in wives' non-labor income. Then, given that leisure is a normal good, the reservation wages of married women should decrease in response to the tax reform; it is thus expected that the probability of work will increase.

However, as was confirmed in Section 2.3.3, the effect of the spousal exemption cut is not exceptionally large and is, at most, 0.109 million yen ($\approx$USD 1,090). Thus, there is a considerable possibility that the effect of the spousal exemption cut is too small for the reservation wage to be lower than the market wage rate. Furthermore, if we consider that many married women would incur some fixed costs for working, such as the monetary costs of child care or loss of time due to taking their children to child care, this possibility would be more likely.

2.5 Empirical Model

In this section, I will estimate the impact of the tax reform on the labor supply of married women for each income group. Based on the shape of the budget line represented in Figure 2.4, married women can be divided into the following four groups based on their 2003 income: less than 0.7 million yen, 0.7 to 1.03 million yen, 1.03 to 1.41 million yen, and 1.41 million yen or more. Since the tax reform's impact on the labor supply of women should differ according to income level, it is extremely important to identify the tax reform's impact on each income group. Particularly, in order to determine whether the tax reform successfully worked, it is necessary to see whether married women in the lowest and the second-lowest income group increased labor supply after the tax reform. In addition to the tax reform effect on low-income married women, we are also interested in how the labor supply among medium- to high-income married women changed in response to changes in exogenous variables (such as husbands' incomes) as well as the tax reform.

I first test Proposition 2.1 in Section 2.4 by estimating difference-in-difference estimations for each quantile of incomes and work hours. Then I employ a relatively novel decomposition method, proposed by Firpo et al. (2007, 2010), to test Propositions 2.2 and 2.3 in Section 2.4. Lastly, I visually examine how the behavioral changes confirmed in the difference-in-difference estimation and the FFL decomposition affected the overall income distribution, utilizing the DFL decomposition. This method also enables us to test Propositions 2.2 and 2.3.

2.5.1 Model

In Sakata and McKenzie (2005, 2006), a difference-in-difference estimation was conducted to estimate the impact of the tax reform on the labor supply of married women. The estimated model in Sakata and McKenzie (2005, 2006) is:

$$Y_{it} = \beta_0 + \beta_1 After_t + \beta_2 Treatment_i + \beta_3 After_t \cdot Treatment_i + \gamma X_{it} + u_{it} \quad (2.4)$$

where $Treatment$=1 for the treatment group, and $After$=1 for samples in periods after the tax reform. A dummy variable for working women and their work hours are used for Y_{it} in Sakata and McKenzie (2005, 2006). Utilizing the fact that those who are expected to be affected by the tax reform are married women, they use married women as the treatment group and single women as the control group in this difference-in-difference estimation.

In this paper, I will estimate the model presented in Equation (2.4) by quantile in the income distribution. The same regressions are also performed for labor participation and work hours to strengthen the income regression results. This quantile analysis enables us to separately examine the impact of tax reform on the low-income married women who were directly affected by the reform and its impact on other married women who were not presumed to be affected by the reform. According to Proposition 2.1, we expect that the coefficient of the interaction term, β_3, is significantly positive for low quantiles and zero for high quantiles. If the coefficient of the interaction term, β_3, is significantly positive for low quantiles, it implies that the tax reform increased labor supply among low-income married women as expected.

Note that the settings of the difference-in-difference estimation corresponds to the case in which there is no change in exogenous parameters, since the husband's income and many other exogenous variables are controlled for in the estimation and the coefficients of these explanatory variables, except for the treatment dummy, are also assumed to be the same before and after the tax reform.

2.5.2 CQR vs. UQR

In the difference-in-differences estimation, both the conditional quantile regression (CQR) and the unconditional quantile regression (UQR) are used to examine the tax reform's impact on each quantile of incomes and work hours.

Firpo et al. (2009) showed that the UQR estimates can be represented as a weighted average of CQR estimates. For example, when estimating the effect of unions on wages, the CQR estimate is found to be negative for highly skilled workers (with high educational attainment and extensive labor market experience) and positive for other workers. Suppose that the explanatory variables other than the union status include only the workers' skill level (highly skilled or otherwise) and that the number of highly skilled workers is small. In CQR, wage distributions conditional on skill level are used. The union effects are first estimated separately for the highly skilled workers and for other workers, and then the CQR estimate of the union effect is calculated as the weighted average of the union effects (β) conditional on each skill level, i.e., $E_Z[\beta/Z]$, where Z represents a worker's skill. Therefore, as long as the number of highly skilled workers is smaller than that of other workers, the negative union effects on skilled workers are averaged away by the positive effects on the majority of workers. As a result, the union effects are estimated to be positive for all quantiles in CQR.

In contrast, in UQR, the unconditional distribution is used, and the fraction of highly skilled workers varies by quantile, while it is fixed across quantiles in CQR. Thus, in UQR, at high quantiles of (unconditional) wage distribution that mainly consist of highly skilled workers in the sample, more weight is placed on highly skilled workers. As a result, the UQR estimates of union effect on wages become negative for high quantiles of wage distribution (Firpo et al., 2009).

However, in the context of the tax reform effect in this paper, there is no possibility that the sign of the estimate for the tax-reform effect varies by workers' attributes, although the magnitude of the effect might vary with differences in worker attributes. Thus, it is hard to expect a case in which there is a significant gap between CQR estimates and UQR estimates.

Figure 2.8 illustrates what is estimated in the CQR regression. Suppose T is the treatment dummy and Z represents other covariates. The distribution at the left-hand side is the income distribution of the control group conditioning on attributes Z. In Figure 2.8, it is assumed for simplicity that there is no change in the income distribution of the control group during the period 2003-2006. The dashed and bold lines at the right-hand side represent the income distributions of the treatment group before and after the tax reform, respectively. $Y_{\tau/T=1,Z}$ represents income at the τ quantile of the income distribution for the treatment group with workers' attributes Z, while $Y_{\tau/T=0,Z}$ represents income at the τ quantile of the income distribution for the control group with Z. The slope of the straight dotted line that connects incomes at the τ quantile of each of the two distributions represents the CQR estimates for the coefficient of the treatment dummy ($\beta_2^{\tau,CQR}$), which is calculated as $\beta_2^{\tau,CQR} = Y_{\tau/T=1,Z} - Y_{\tau/T=0,Z}$. The figure shows an example in which the dotted line rotates around $Y_{0.2/T=0,Z}$ in response to an increase in income among low-income workers in the treatment group, and the slope becomes flatter after the tax reform. Then, the change in the slope, i.e., the gap in the slope between the two dotted lines, represents the magnitude of the interaction term between *treatment* and *after* dummy variables, i.e., $\beta_3^{\tau,CQR}$. In this way, it is expected that the coefficient of the treatment dummy variable, which is originally negative, becomes smaller negative or even positive at low quantiles after the tax reform, in response to the increase in incomes among low-income workers in the treatment group.

Whether we should use CQR or UQR depends upon what we want to know. As is indicated in Figure 2.8, since the CQR estimates capture the changes in distribution conditional on X, it will help us to understand the within-group effect of the tax reform. The UQR estimate captures the between-group effect as well as the within-group effect. The UQR estimate might be helpful when we estimate the counterfactual impact of expanding the eligibility of the spousal exemption because it captures the direct impact of increasing the proportion of the treatment group. In general, when an explanatory variable of interest, X, is binary, the coefficient of X from a UQR becomes equal to the effect of increasing the proportion of people with X=1 on the τ quantile of the distribution (i.e., $\beta_\tau = dq_\tau/dPr[X{=}1]\ \forall\tau$). This holds for all quantiles in a UQR, while in a CQR, it holds only for the mean (Firpo et al., 2009). In the context of the model presented above, the coefficient of the treatment dummy from the UQR is equal to the effect of increasing the proportion of the treatment group on the τ quantile in the income distribution.

Figure 2.9 is presented to depict the interpretation of the UQR estimate. In Figure 2.9, we see how the income distribution moves after increasing the proportion of the treatment group by one unit. Suppose that before the tax reform, when we increase the proportion of the treatment group by one unit, income at the 20th quantile of the distribution shifts from $Y^{0.2}_{t=0}$ to $Y^{0.2\prime}_{t=0}$. Then $Y^{0.2\prime}_{t=0} - Y^{0.2}_{t=0}$ represents β_2 for the 20th quantile before the tax reform, i.e., $\beta^{0.2}_{2,t=0}$. Similarly, if income at the 20th quantile of the distribution shifts from $Y^{0.2}_{t=1}$ to $Y^{0.2\prime}_{t=1}$ after the tax reform in response to the increase in the proportion of the treatment group, $Y^{0.2\prime}_{t=1} - Y^{0.2}_{t=1}$ represents β_2 for the 20th quantile after the tax reform, i.e.,$\beta^{0.2}_{2,t=1}$. Then, the difference in $\beta^{0.2}_2$ between before and after the tax reform corresponds to $\beta^{0.2}_3$. Details regarding unconditional quantile regression are presented in Firpo et al. (2009).[30]

[30]In Figure 2.9, the two figures are placed vertically to compare the magnitudes of $\beta^{0.2}_2$ and to show $\beta^{0.2}_3$ as the difference between the two $\beta^{0.2}_2$ values. However, this does not mean that $Y^{0.2}_{t=0} = Y^{0.2}_{t=1}$ holds.

2.5.3 Basics of FFL Decomposition

Before explaining how I will utilize the FFL decomposition technique in the context of the model presented in Section 2.5.1, I will briefly explain the FFL decomposition method. In the FFL, the total change in Y is divided into the composition and structure effects, and they are further decomposed into the contributions of each explanatory variable. In this sense, the Oaxaca-Blinder decomposition (OB) would be the counterpart of this method. However, two characteristics in the FFL decomposition differentiate it from the classic OB decomposition. First, in FFL, the recentered influence function (RIF) of Y, instead of Y, is used as the dependent variable, which enables us to examine distributional changes in Y; this was possible only for the mean in the classic OB decomposition (Firpo et al., 2007, 2010). Second, to calculate the counterfactual distributions, the reweighting method proposed in the DiNardo, Fortin, and Lemieux decomposition (DiNardo et al., 1996) is used. By performing the reweighting regressions, we can prevent the difference in the coefficients between the two periods from being contaminated by differences in the distribution of X between the two periods. (This point will be explained at the end of this section.)

Conceptually, the influence function (IF) represents the influence of increasing an individual observation on the distributional statistic, such as means, variances, quantiles, and Gini coecients. For the τth quantile, the influence function $IF(Y; q_\tau, F_Y)$ is known to be equal to $\{\tau - 1(Y \leq q_\tau)\}/f_Y(q_\tau)$. The IF gives us a way to explore how changes in the distribution of Y affect q_τ. The goal here is to compute the effect of changing X on quantiles. The process leading up to this goal is divided into two steps: (1) changes in $X \rightarrow$changes in the distribution of Y and (2) changes in the distribution of $Y \rightarrow$changes in q_τ. The IF only does step two. To connect steps one and two, we need the Law of Iterated Expectations (LIE). To utilize the LIE, we need the recentered version of the IF, i.e.,RIF, which is simply equal to $q_\tau + IF(Y;\, q_\tau,\, F_Y)$. The RIF has a very convenient feature in that its expectation is equal to its distributional statistics (in our case, q_τ):

$$q_\tau = E[RIF] = E_X(E[RIF/X]) \tag{2.5}$$

Then, using the LIE, the distributional statistic can also be expressed in terms of the expectations of the conditional RIF −i.e., a function of X. Thus, by using the RIF, step one can be connected to step one. The details about this process can be found in work by Firpo et al. (2007, 2010).

The counterfactual distribution in FFL is calculated by DiNardo et al. (1996) method using the following reweighting term:

$$\begin{aligned}\omega = \tfrac{h(X/t\geq 2004)}{h(X/t=2003)} &= \tfrac{P(t\geq 2004/X)P(X)/P(t\geq 2004)}{P(t=2003/X)P(X)/P(t=2003)} \\ &= \tfrac{P(t\geq 2004/X)P(t=2003)}{P(t=2003/X)P(t\geq 2004)}\end{aligned} \tag{2.6}$$

where the density $h(X/t = T)$ is the p.d.f. of attributes in year T. The second equation is derived from Bayes' rule. In the actual regression of ω, $P(t = T/X)$ can be calculated using propensity scores obtained from the probit model in which $P(t = T)$ is regressed on X, and $P(t = T)$ is calculated as the proportion of the observations from year T in the pooled data.

In Section 2.7, I will report both results without reweighting, i.e., the classic OB decomposition, and results with reweighting. For the latter, the following different pooled samples are used to obtain the composition and the structure effects.

- Estimation 1: Reweighted regression with a pooled sample of X_{03} and$X_{C,04}$: (Only the composition-effect part $((\overline{X}_{C,04} - \overline{X}_{03})'\beta^\tau_{C,03})$ is used)

$$\begin{aligned}&\overline{X}'_{C,04}\beta^\tau_{C,03} - \overline{X}'_{03}\beta^\tau_{03} \\ &\quad = \overline{X}'_{C,04}(\beta^\tau_{C,03} - \beta^\tau_{03}) + (\overline{X}_{C,04} - \overline{X}_{03})'\beta^\tau_{03} \\ &or \quad = \overline{X}'_{03}(\beta^\tau_{C,03} - \beta^\tau_{03}) + (\overline{X}_{C,04} - \overline{X}_{03})'\beta^\tau_{C,03}\end{aligned}$$

- Estimation 2: Reweighted regression with a pooled sample of $X_{C,04}$ andX_{04}:

(Only the structure-effect part $(\overline{X}'_{C,04}(\beta^{\tau}_{04} - \beta^{\tau}_{C,03}))$ is used)

$$
\begin{aligned}
&\overline{X}'_{04}\beta^{\tau}_{04} - \overline{X}'_{C,04}\beta^{\tau}_{C,03} \\
&\quad = \overline{X}'_{04}(\beta^{\tau}_{04} - \beta^{\tau}_{C,03}) + (\overline{X}_{04} - \overline{X}_{C,04})'\beta^{\tau}_{C,03} \\
&or \quad = \overline{X}'_{C,04}(\beta^{\tau}_{04} - \beta^{\tau}_{C,03}) + (\overline{X}_{04} - \overline{X}_{C,04})'\beta^{\tau}_{04}
\end{aligned}
$$

where $X_{C,04}$ is the counterfactual 2004-2006 distribution of X obtained from the reweighting method, and $\beta^{\tau}_{C,03}$ is the counterfactual values of the coefficients for the τth quantile. In calculating $X_{C,04}$, the 2003 sample is reweighted to mimic the 2004-2006 sample using the reweighting term defined in Equation (2.6), which means that we should have $plim(X_{C,04}) = plim(X_{04})$. The dependent variable used to obtain $\beta^{\tau}_{C,03}$ is the RIF of Y in 2003, thus the counterfactual values of the coefficients can be thought of as coefficients that would be realized if the determinant structure of Y stayed the same from 2003 to 2004-2006. The composition effect from Estimation 1 and the structure effect from Estimation 2 are used in the FFL decomposition.

Counterfactual values of the coefficients are used here because the difference in the coefficients between the two periods in the classic OB decomposition may be contaminated by differences in the distribution of X between the two years. In the FFL decomposition with reweighting, the (asymptotically) same X is used for the two periods to estimate the structure effect, and hence the difference in β^{τ} should not be contaminated by the difference in the distribution of X.

2.5.4 FFL Decomposition Model

Now, we return to the context of this paper. The purpose of adding the FFL decomposition is to test Propositions 2.2 and 2.3. In the decomposition, utilizing the advantage of the technique that enables us to further divide each of the structure and composition effects to a contribution of each explanatory variable, I examine whether medium- to high-income married women actually decreased their incomes by a large negative "husband's income effect" (and not "tax-reform effect"). This is possible by checking the magnitude of the composition effect attributable to husband's income; If it is greatest at medium- to high-quantiles of the income distribution with statistical significance, it supports Propositions 2.2 and 2.3 stating that the "income jump," which is explained mainly by a large negative "husband's income effect," is likely to occur for medium- to high-income married women. It is also important to check that the structure effect attributable to the treatment dummy variable is not significantly negative for these women to make sure that this "income jump" cannot be explained by the negative tax reform effect.

Furthermore, I reconfirm the results obtained from the difference-in-difference estimation in the FFL decomposition as well. If the coefficient of the treatment dummy variable becomes higher after the tax reform in the difference-in-difference regressions (i.e., the interaction term between the *treatment* and *after* dummy variables is significantly positive in the difference-in-difference regressions) and the structure effect attributable to the treatment dummy variable obtained from the FFL decomposition is significantly positive at low quantiles, this would imply that the tax reform contributed to increasing incomes among low-income women.

Note that the structure effect attributable to the treatment dummy variable corresponds to the coefficient of the interaction term between the *treatment* and *after* dummy variables in the difference-in-difference regression. Thus, by looking at the structure effect attributable to the treatment dummy variable, it is possible to reconfirm the results for low-income married women in the difference-in-difference regression.

Figure 2.10 illustrates the interpretation of the FFL estimates visually. The dashed line represents the income distribution before the tax reform, and the solid line represents that for after the tax reform. As can be confirmed in the figure, the quantiles are reassigned for each period in FFL, and FFL compares the RIF of income at quantile

τ between before and after the tax reform.

Again, in the FFL decomposition, we can observe whether incomes among medium- to high quantiles decreased after the tax reform; and, if so, whether this was caused by the "income jump," which is explained mainly by a large negative effect of husband's income. In addition, the results obtained from the difference-in-difference estimation are reconfirmed in the FFL decomposition by checking whether incomes at low quantiles increased after the tax reform; and, if so, whether this was caused by the tax reform.

2.5.5 DFL decomposition

Lastly, utilizing the DFL decomposition method, I visually examine how the behavioral changes confirmed in the difference-in-difference estimation and the FFL decomposition affected the overall income distribution. The merit of this method is that it visually decomposes the change in the distribution into two parts: structure effects and composition effects (DiNardo et al., 1996; DiNardo and Lemieux, 1997).

First, the income distribution in 2003 is expressed as:

$$F_{2003} = \int f_{2003}(Y/X)h(X/t = 2003)dX \tag{2.7}$$

where $f_{2003}(Y/X)$ is the income determination mechanism in 2003 that maps workers' attributes to the income distribution. Again, the density $h(X/t = 2003)$ is the p.d.f. of attributes in year 2003. Similarly, the distribution during the period 2004-2006 is expressed as:

$$F_{2004} = \int f_{2004}(Y/X)h(X/t \geq 2004)dX \tag{2.8}$$

What the distribution would be after the tax reform if the income determination mechanism were identical to its mechanism in 2003 is expressed as:

$$F_{2004}^{2003} = \int f_{2003}(Y/X)h(X/t \geq 2004)dX \tag{2.9}$$

This can be thought of as a counterfactual distribution in the period 2004-2006 without the tax reform because it consists of the same workers' attributes as the real 2004-2006 distribution of X but of β prior to the tax reform. This counterfactual distribution is calculated by DiNardo et al. (1996) method:

$$F_{2004}^{2003} = \int f_{2003}(Y/X)h(X/t \geq 2004)dX = \int \omega f_{2003}(Y/X)h(X/t = 2003)dX \tag{2.10}$$

where ω is given by Equation (2.6). This reweighting process is exactly the same as that in the FFL.

2.6 Regression Sample

Again, the data set used in this study is Japanese panel data from the Keio Household Panel Survey (KHPS) 2004-2007. Since the characteristics of this survey were already reported in Section 2.3.1, in this section, I will discuss how I select the regression samples from the data.

The details of the KHPS are as follows: respondents for the first wave were men and women aged between 20 and 69 as of January 31, 2004, from all of Japan. The first wave (2004) included 4,005 households, the second wave (2005) included 3,314 out of 4,005 in the first wave, the third wave (2006) 2,887 households, and the fourth wave (2007) 2,643 households. Additionally, 1,419 households were newly added to the samples in 2007, and this paper also utilizes these samples. The attrition rate from the first wave to the fourth wave was 34 %.[31]

[31] The sample selection issue arising from attrition (i.e., from the first to the fourth waves) is expected to be small, because the characteristics of the women who participated in the survey in all

Table 2.5 contains descriptive statistics for the sample of main analyses. The first column in Table 2.5 reports the overall sample including non-working women, which is used for the probit estimation first conducted to check the tax reform effect on labor participation decision.

As discussed in Section 2.2, the real schedule of the spousal exemption is defined with regard to the spouse's *total income*, which is calculated as the total earnings minus the *employment income deduction* for employees, and the total earnings minus necessary expenses for other workers. While necessary expenses vary among individuals, the amounts of the *employment income deduction*, which are the counterpart of necessary expenses, are fixed for each amount of employment incomes. As a result, the uniform correspondence between the amount of annual earnings and the amount of the spousal exemption can be obtained only from employment incomes. Therefore, analyzing the effect of the spousal exemption for employment income earners is more straightforward. For this reason, the sample of the main analyses concerning the tax reform effect on the income distribution is restricted to employees, which comprises 41 % of all the observations.

Furthermore, in the theoretical model, traditional households in which the husband is the main earner and the wife is a dependent were assumed, but this assumption is less unlikely to be true for married women who work for the government. Nishikawa (2002) shows that the gap in work hours between a husband and a wife in a couple in which a wife works in the public sector is much smaller than that for other types of couples. This finding implies that the assumption about the unilateral influence on labor decision (from a husband to a wife, and not the reverse) is unlikely to hold for these women. For this reason, in the main analyses, I also restrict the sample to employees who work in the private sector.

In the sample of the main analyses, 3,206 observations in the panel that are

years (2004-2007) are similar to those of the women who dropped from the sample at some point during that time. Table A.2.1 presents the descriptive statistics for these two groups.

employed in the private sector and work at least one hour each year are used.[32][33] The average age is 44.22 years. Among the four income groups (0-0.7 million, 0.7-1.03 million, 1.03-1.41 million, and 1.41 million+), the group that comprises the largest number of workers is the group with incomes of 1.41 million yen and above, which corresponds to 44% of all samples. Among married women, almost half of the sample is in the income groups that were directly affected by the tax reform.

Furthermore, although it is the husbands' after-tax incomes, not the before-tax incomes, that would actually affect the wives' incomes, there is a concern about using the husbands' after-tax incomes to control for husbands' incomes in estimating the tax reform effect. If the husband's after-tax incomes were used in the estimations, the change in spousal exemption would also be reflected in the variable. This means that part of the tax reform effect would be captured by the husbands' incomes as well, despite our desire to capture the impact only by the treatment dummy variable. To circumvent this problem, husband's after-tax incomes that ignore the spousal exemptions are used in the regressions because this variable comes closer to the economically relevant variable than do pre-tax incomes, without introducing the problem stated above. The average value of this variable among married women is 4.84 million yen.

2.7 Results

2.7.1 Difference-in-Difference Regression Results

Table 2.6 reports the results of the difference-in-difference estimations for labor participation, annual incomes, and work hours. The coefficient of interest is that pertaining to the interaction term between *treatment* and *after* dummy variables.

[32] Thus, if either of these three conditions, i) positive work hours, ii) employed, and iii) private sector, are not satisfied during year t, that person is dropped from the sample of year t.

[33] Only women who reported all the necessary variables will be included in the sample in the regression. A comparison of workers' characteristics between the raw data and the regression sample did not show any significant difference. Thus, the sample selection issue arising from using only people who reported all the necessary variables is expected to be small.

The treatment group is made up of married women whose husbands are eligible for the spousal exemptions (i.e., married women who are not the head of household and have husbands whose *total income* is 10 million yen or less per year). Single women and married women whose husbands are not eligible for the spousal exemptions are used as the control group.

Column 1 first reports the probit estimation result using the overall sample including non-working women. The coefficient of the interaction term between the *treatment* and *after* dummy variables is insignificant, and we do not see any evidence that the tax reform increased labor participation. This might reflect the noise in the regression as well as the possibility that the effect of the spousal exemption cut is too small for the reservation wage to be lower than the market wage rate.

The dependent variable in Column 2 to 6 is annual income, while the dependent variable for the UQR (Column 7 to 9) is the recentered influence function (RIF) of annual income.[34] The CQR of weekly work hours is also added to the table to strengthen the analyses of incomes. The explanatory variables include the treatment dummy, husband's annual after-tax income that ignores the spousal exemptions, number of children, age and its square, tenure and its square, unionized worker dummy variable, permanent employee dummy variable, the logarithm of firm size, educational dummy variables, and industry dummy variables.[35]

Column 2 and 3 present the results of the OLS regression and the fixed effects (FE) model, respectively. The OLS and FE estimates for a coefficient of the interaction

[34]The RIF for the τth quantile can be computed as $q_\tau + \{\tau - 1(Y \leq q_\tau)\}/f_Y(q_\tau)$ using the sample estimate of q_τ and the kernel density estimate of $f_Y(q_\tau)$ using the Epanechnikov kernel.

[35]Table 2.6 reports the results without hourly wage rate in X. Forty-three percent of the women in the sample are paid hourly, and hourly wage information is available only for this group. Fifty-one percent of the women in the sample are paid monthly, and the hourly wage for this group needs to be calculated by dividing the monthly salary by the number of hours worked per month as reported by the workers. This can create a measurement error in the explanatory variable. Thus, in the main analyses, I report results without hourly wage rate. However, as was confirmed in Section 2.3, the hourly wage rate was stable over the period for both married and single women; the sensitivity check that includes hourly wage rate in X thus shows that the main results are preserved when hourly wage rate included.

term are both insignificant, which reflects the fact that these estimates capture the mean impact of the tax reform. However, those who were directly affected by the tax reform are only married women whose annual income is 1.03 million yen or less; hence, there is a possibility that the mean impact estimated in the OLS and FE model missed the real impact of the tax reform on low-income married women.

To capture the heterogeneous impact of the tax reform across income groups, I conducted the same difference-in-difference estimation by quantile. Columns 4 in Table 2.6 reports the CQR estimates for women whose income is at the 30th quantile, and the coefficient of the interaction term is significantly positive at this quantile. These women belong to the lowest and second-lowest income group, and hence married women at this quantile are those who were directly affected by the tax reform. The CQR estimate shows that married women whose income is at the 30th quantile increased their incomes by 0.164 million yen per year.

In contrast, for the higher quantiles, the coefficients of the interaction term are close to zero and insignificant. We can observe similar results for the UQR from Column 7 to 9. The fact that we do not see any significant difference in the results for the interaction term between CQR and UQR reflects a uniform sign of the tax effect across worker characteristics as discussed in Section 2.5.2.

In Column 10 to 12, the same regressions are conducted for weekly work hours. We can observe results similar to the results for incomes: i.e., the significantly positive estimates for the coefficients of the interaction term for the low quantile and insignificant estimates for the quantiles above median. The CQR estimate of the interaction term indicates that the tax reform increased the weekly work hours of low-income married women by 3.589 hours. A simple calculation using the average hourly wage rate shows that this increase in work hours leads to an income increase of approximately 0.16 million yen per year, which is consistent with the result of the income regression. I also tried a regression including asset incomes and polynomials of husband's income, and the main results are preserved even when these variables are included.

Note that the results obtained from Table 2.6 are consistent with Proposition 2.1. In Table 2.6, the husband's income and many other exogenous variables are controlled for, and it is implicitly assumed that the coefficients of these explanatory variables, except for the treatment dummy, are the same before and after the tax reform. This corresponds to the case in which there is no change in exogenous variables. Proposition 2.1 states that in such an environment, the tax reform has no impact on those who originally earned more than 1.03 million yen but increases work hours and for married women who originally earned 1.03 million yen or less. This is exactly what Table 2.6 demonstrates.

In Figure 2.11, I plotted the coefficients for the interaction term between *treatment* and *after* dummy variables obtained from the CQR for all quantiles. For the income distribution, the fraction of married women who were directly affected by the tax reform should be approximately 1 for the quantiles lower than the 43th quantile (corresponding to 1.03 million yen) and 0 for the higher quantiles.[36] Reflecting this fact, Figure 2.11 shows that the magnitude of the tax reform's effect drops near the income distribution's 43th quantile.

The reason for the small tax reform effect at the 10th quantile is that the estimates at these extremely low quantiles are likely to be affected by special characteristics that are not captured by the explanatory variables. For example, these women might be in very conservative households and would, therefore, not have much incentive to work more, even after tax reform. Note that in general, both ends of the distribution are likely to include people who have special characteristics; thus, the standard errors

[36] Although earning less than 1.03 million yen does not directly equate with being eligible for the special spousal exemption (because the husbands' *total incomes* also need to be 10 million yen or less per year for the household to be eligible for the special spousal exemption), the fraction of spouses whose husbands' *total incomes* are more than 10 million yen is very small (only 1.5% in the sample). In the sample, it seems that the condition regarding the husbands' incomes is not very effective.

of the interaction term's coefficient at the 10th quantile and quantiles over the 70th quantile are, indeed, very large. This finding demonstrates the importance of estimating the impact of the tax reform by quantile, rather than estimating the mean impact, because by doing so, we can prevent the estimates from being contaminated by the imprecise estimates at the extremely low or high quantiles.

For the work-hour distribution, the magnitude of the coefficient peaks at around the 30th quantile and diminishes as we move toward the higher quantiles. This reflects the fact that the fraction of married women with an income of 1.03 million yen or less peaks at around the 30th quantile of the work-hour distribution and declines for the higher quantiles.

Considering the Difference in Husbands' Tax Rates

To take into account the difference in husbands' tax rates, I also conducted a regression in which "eligible" married women whose husbands' tax rate is 20 % are used as the treatment group and "eligible" married women whose husbands' tax rate is 10 % are used as the control group. Thus, only "eligible" women are included in the sample, and the treatment and control groups are defined by the husbands' tax rates. Since the impact of the spousal exemptions on the household budget line becomes greater as the husbands' tax rate increases, the theory predicts that married women whose husbands' tax rate is higher will be affected by the tax reform more than those whose husbands' tax rate is lower.

Figure A.2.1 plots the CQR estimates for the coefficient of the interaction term between *treatment* and *after* dummy variables in the income regression. The estimate is significantly positive at the 20th quantile meaning that the magnitude of the tax reform effect is greater for those whose husband's tax rate is higher. Thus, the difference-in-difference estimation results support the theoretical predictions.

2.7.2 FFL Decomposition Results

Next, I will report the FFL decomposition results. An extension of the classic OB decomposition and the reweighted-regression decomposition results for the 25th, 50th, 60th, 70th, and 95th quantiles in the income distribution are reported in Table 2.7. The odd-numbered columns report the results of the decomposition without reweighting. The even-numbered columns report the results of the reweighted-regression decomposition, in which $F(X)$ in 2003 is reweighted to the 2004-2006 period.

First, for both the decomposition without reweighting and the reweighted-regression decomposition, the structure effects attributable to the treatment dummy at the 25th quantile are significantly positive at the 10 % significance level. The results here are comparable with those from the difference-in-difference estimation; the structure effect attributable to the treatment dummy in the FFL decomposition corresponds to the coefficient of the interaction dummy between *treatment* and *after* dummy variables in the difference-in-difference regression. The FFL decomposition results from low-income women indicate that the increase in the coefficient of the treatment dummy, confirmed in Table 2.6, contributed to increasing incomes among low-income women.[37]

Second, the composition effects attributable to husbands' incomes are significantly negative among medium-to high-income women at the 50th and 60th quantiles. Proposition 2.2 in Section 2.4 states that there are cases in which medium-to high-income married women greatly decrease their incomes after the tax reform (from 1.41

[37]It is well known that the contribution of each covariate to the unexplained part can depend on arbitrary scaling decisions (Jones and Kelley, 1984; Jann, 2008). As a related issue, it is also known that the unexplained part of the decomposition arbitrarily depends on the choice of reference groups when indicator variables are included in the model (Jones, 1983; Oaxaca and Ransom, 1999; Firpo et al., 2010). However, identifying the magnitude of the structure effect attributable to the treatment dummy variable is not of interest here because the magnitude of the tax reform effect was already estimated in the difference-in-difference estimation. Rather, the main objective of using the FFL decomposition here is to compare the statistical significance and relative magnitude of the composition effect attributable to the husbands' incomes across all quantiles. Note that the potential problems stated above do not arise in estimating the contribution of each covariate to the explained part (Jones, 1983; Jones and Kelley, 1984; Oaxaca and Ransom, 1999; Jann, 2008; Firpo et al., 2010). Thus, it is still meaningful to compare the statistical significance and relative magnitude of the composition effect attributable to the husbands' incomes across all quantiles.

million yen or more to 1.03 million yen) in response to an increase in their husbands' incomes. In addition, Proposition 2.3 emphasizes that the discontinuous jump from 1.41 million yen or more to 1.03 million yen among these women is explained mainly by a large negative husband's income effect.

To compare the magnitude of the "husband's income effect" among all quantiles, Figure 2.12 plots the composition effect attributable to husbands' incomes in Table 2.7. From it, we can see that the magnitude of the negative "husband's income effect" peaks at around the 60th quantiles. This result indicates the possibility that the discontinuous "income jump" among medium- to high-income married women actually occurred in response to an increase in husbands' incomes, because this "income jump" can be explained by a large negative husband's income.

At quantiles over the 70th in Table 2.7, the composition effects attributable to husbands' income become smaller and insignificant. Figure 2.12 also indicates that the negative husband's income effect vanishes at these high quantiles. These results support the theoretical implication that the "income jump" is not possible for those whose preference for consumption is very strong.

It is also important to note that, as was suggested in Proposition 2.3, there is no quantile in which there is a significant negative structure effect attributable to the treatment dummy variable. Thus, the theoretical implication that the "tax reform effect" cannot be negative is supported.

2.7.3 DFL Decomposition Results

Lastly, I will confirm how these changes in the optimal income decision affected the overall income distribution of married women. Figure 2.13 depicts the DFL decomposition results for the income distribution of married women. Comparing the two "actual" distributions, the income distribution after the tax reform is less spread out, and low-income married women and high-income married women gathered to the center of the distribution after the tax reform. In addition, if we look at the counterfactual distribution, the income distribution would have shifted to the left in response to the increase in husbands' incomes during the "after" period without the tax reform.

Furthermore, the fractions of those who earned more than 2.5 million yen and those who earned between 1.41 million yen and 1.7 million yen decreased during the period. This change can be explained by the gap between the actual *before* distribution and the counterfactual distribution, which indicates a composition effect. Among the composition effects at the 50th and 60th quantiles in the FFL decomposition, only the husbands' income effect was significantly negative. Given that these quantiles correspond to 1.41 million yen and 1.7 million yen, respectively, the distributional change for the incomes of 1.41 million yen to 1.7 million yen can be explained by the husbands' income effect.

In contrast, for the incomes below 1.03 million yen, we observe a different trend. In response to an increase in the husbands' incomes, the actual *before* distribution shifted to the left, i.e., to where the counterfactual distribution occurs. However, low-income married women were also affected by the tax reform; hence, the decline of their income was mitigated by the tax reform effect, which is captured by the gap between the actual *after* distribution and the counterfactual distribution.

Although the smoothed lines in Figure 2.13 help us understand the overall trend of the income distribution, it is difficult to confirm the spikes in the income distributions by the figure. Thus, next, I will see how the spikes in the income distribution changed by Figure 2.14, in which histograms of income distribution of married women for both periods are presented. In Figure 2.14, the spike at the kink in the household budget line that had existed prior to the tax reform (0.7 million yen) shrank, and people who were originally at this spike moved to the kink at 1.03 million yen that was made more conspicuous by the tax reform. That movement resulted in the conventional spike at 1.03 million yen being more conspicuous after the tax reform.

As shown in the theoretical model, the kink at 1.03 million yen in the budget line, which became more conspicuous after the tax reform, drags even medium- to high-income married women, who are not presumed to be affected by the tax reform, to the mass point when there was a negative shock in the exogenous variables that could potentially discourage them from working long hours. Consequently, the decrease in income among medium- to high-income married women offset the increase in income among low-income women, which resulted in the slight decrease in the average income among married women.

2.8 Conclusion

In Japan, there is a spousal exemption system, wherein the amount of the income exemption a husband can claim is reduced as his spouse's income rises. In 2004, the special spousal exemption was abolished for taxpayers whose spouses had annual earnings of 1.03 million yen or less, and it was expected that the married women's labor supply would increase after that tax reform was enacted. However, both the average annual income and the average work hours slightly decreased among married women after the tax reform.

At the same time, married men's incomes increased in the years following the tax reform, which could potentially conceal the reform's real impact; hence, this paper estimates the reform's real impact by controlling for the trends in exogenous variables. Furthermore, as the impact of the tax reform could be heterogeneous across income groups due to the tax exemption system, this paper takes this into account and examines the income distribution rather than focusing on the mean impact.

The theoretical model shows that the tax reform made the kink at 1.03 million yen in the household budget line more conspicuous, which produces four testable implications: (1) the tax reform increases work hours (and hence the incomes) for married women who earned 1.03 million yen or less but has no impact on married women who originally earned more than 1.03 million yen, when there is no change in exogenous parameters; (2) however, if an increase in a husband's income is considered, there are cases in which medium- to high-income married women greatly decrease their work hours to earn only 1.03 million yen; (3) even in this case, the "tax reform effect" cannot be negative; (4) this "income jump" among medium- to high-income married women can be explained mostly by a negative "husband's income effect"; and (5) in some cases, this "income jump" is mitigated by a small positive "tax reform effect."

These implications were tested using data from the Keio Household Panel Survey (KHPS). In these empirical analyses, I first examined how married women in each income group responded to the tax reform using a difference-in-difference estimation within a framework of the conditional and unconditional quantile regressions. The results showed that the effects of the tax reform on annual incomes and work hours were significantly positive at low quantiles. If we estimated the mean impact of the tax reform, we did not see any significant impact, implying that it was important to estimate the impact of tax reform for each quantile. In contrast, the tax reform effect for high-income women was small and insignificant. This is consistent with a theoretical implication that the tax reform has no impact on high-income married women in a setting in which there is no change in exogenous parameters.

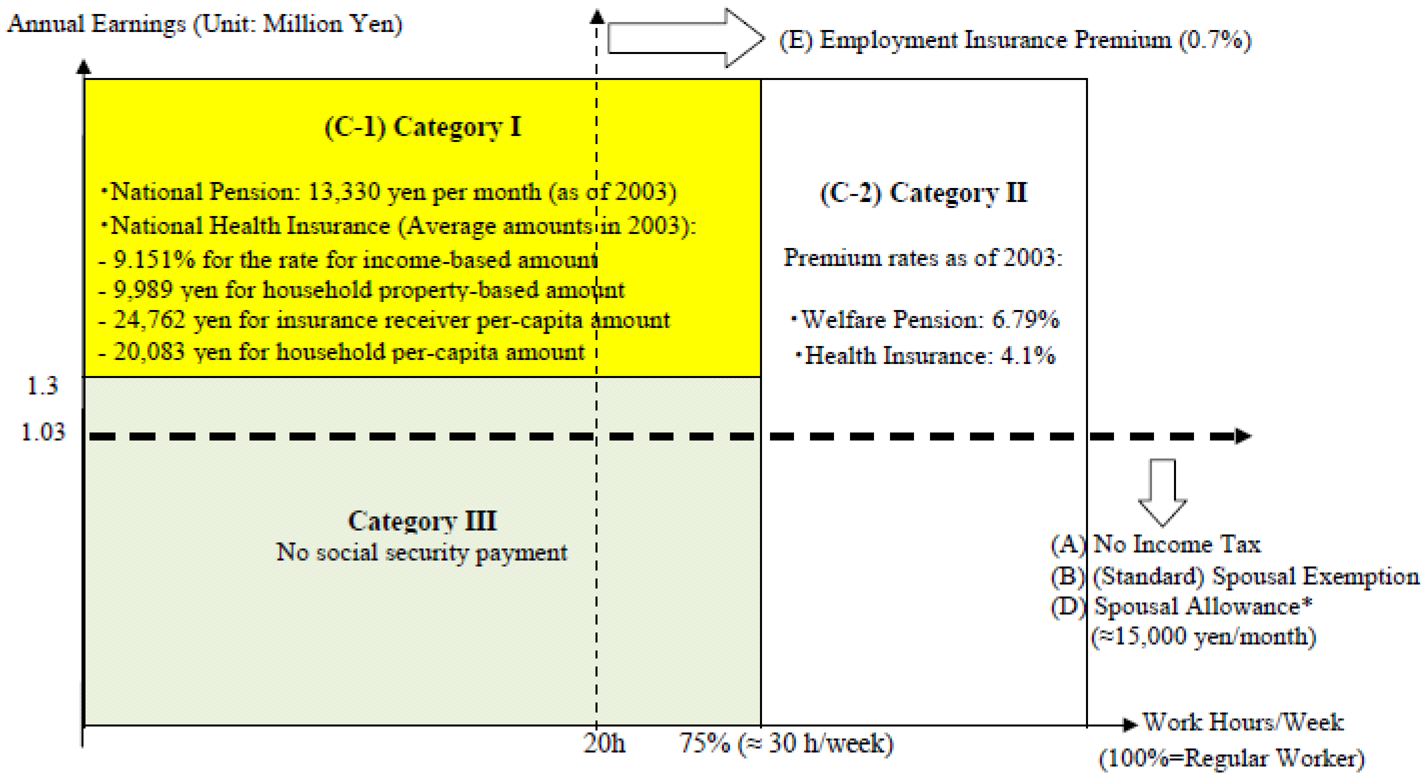
To examine whether medium- to high-income married women actually decreased their incomes by a large negative "husband's income effect" (and not "tax-reform effect"), I also adopted a relatively new decomposition technique proposed by Firpo et al. (2007, 2010). This FFL technique enables us to further decompose the structure effects and the composition effects into contributions of each individual covariate. The structure effect attributable to the treatment dummy was significantly positive for the low quantiles, implying that the tax reform contributed to increasing incomes among low-income married women. Furthermore, the magnitude of the composition effect attributable to husbands' incomes is greatest among medium- to high-income married women. This indicates the possibility that the discontinuous "income jump" among medium- to high-income married women in response to an increase in husbands' incomes actually occurred because this "income jump" can be explained mainly by a large husband's income effect.

The results of DFL decomposition also showed that low-income married women and high-income married women gathered to the center of the income distribution after the tax reform. Accordingly, the mass point at 1.03 million yen became more conspicuous after the tax reform.

Although the tax reform did contribute to increasing incomes among low-income married women, that tax reform resulted in the kink at 1.03 million yen in the household budget line becoming more conspicuous, making it a more attractive income choice than before. Due to this change in the household budget line, the annual income of some of the medium- to high-income married women decreased to 1.03 million yen when they experienced both the tax reform and an increase in their husbands' incomes. As a result, the decrease in incomes caused by a large negative husband's income effect offset the increase in income among the low-income women, which resulted in a slight decrease in the average income among married women.

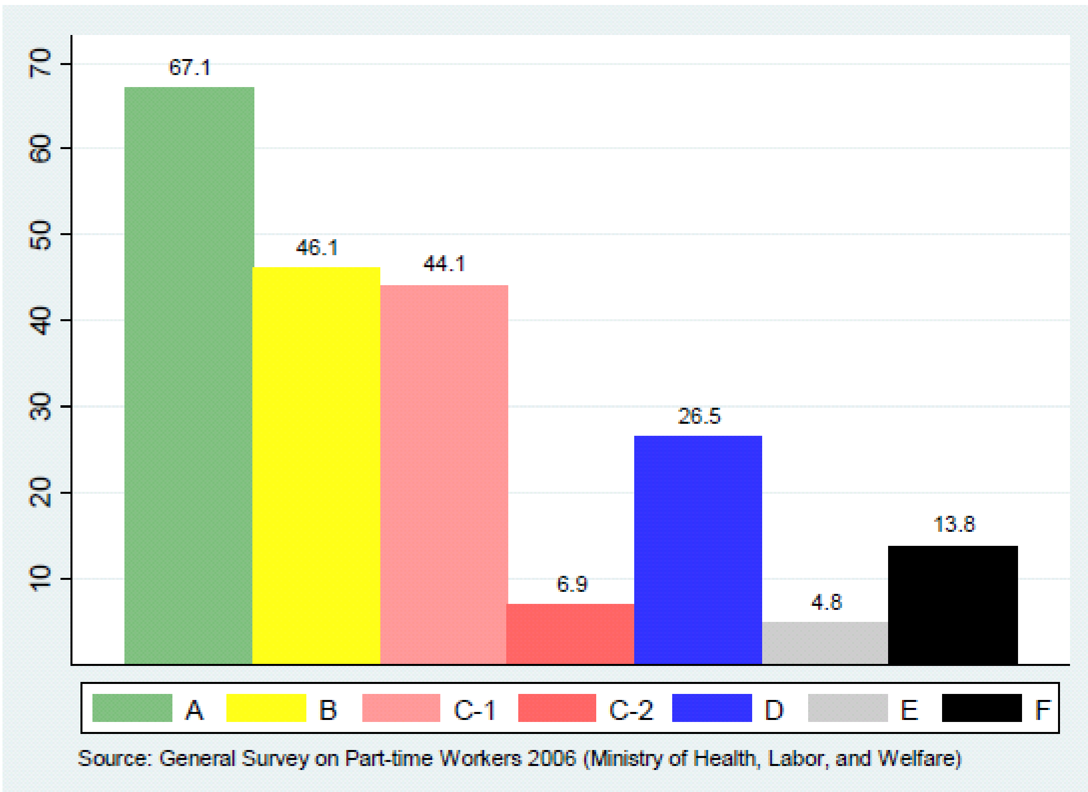
The important implications here are that although the tax reform contributed to an increase in labor supply of married women, ironically, it made the conventional distortion at 1.03 million yen in the income distribution more conspicuous and that this will be the case as long as the remaining spousal exemption exists.

Figure 2.1: Factors that Could Potentially Induce Labor Supply Adjustment



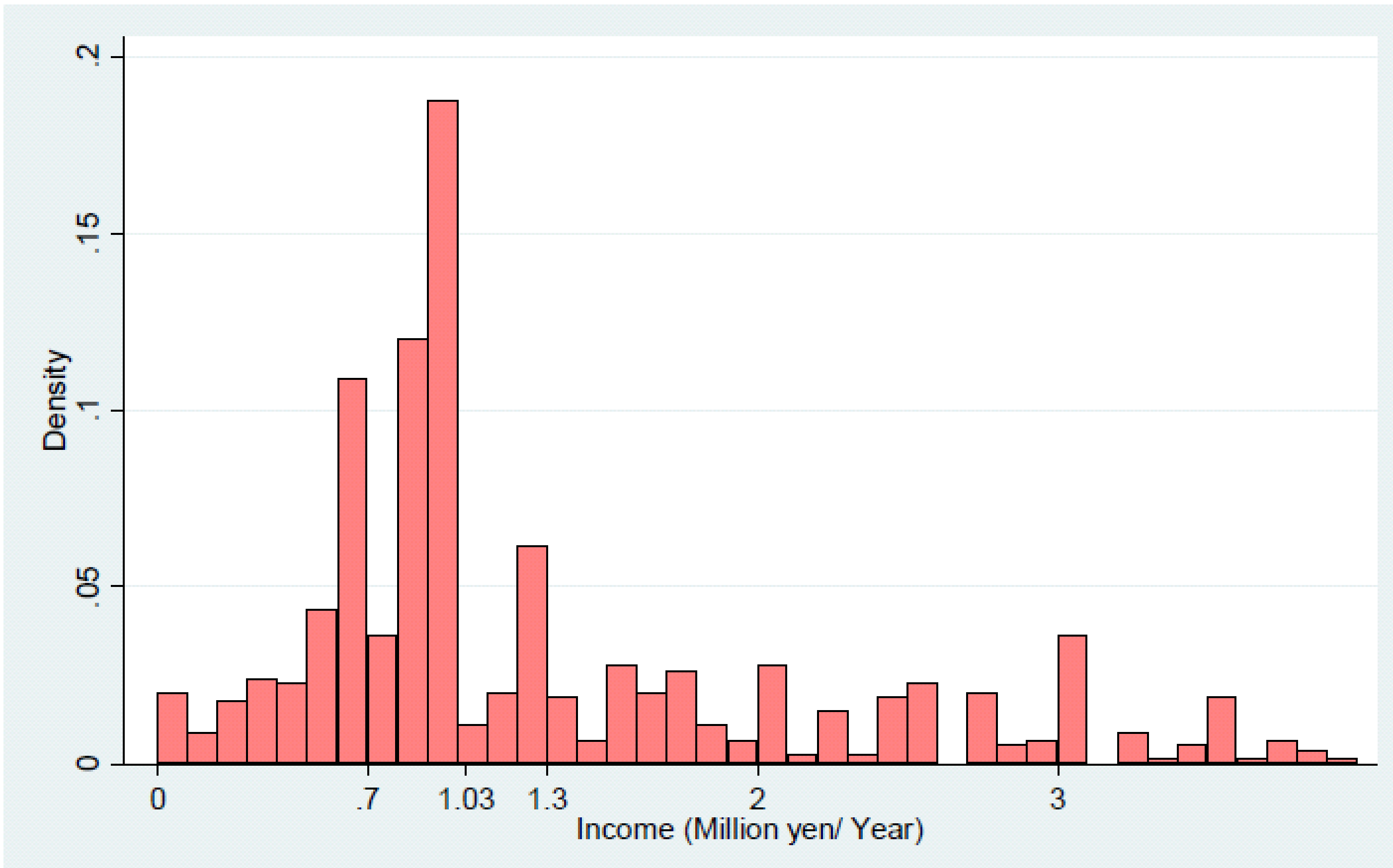
Note: This figure was created based on Table 1 of Oishi (2003). The amounts of National Health Insurance are the average amounts as of 2003 reported in the Survey on National Health Insurance 2003 (Ministry of Health, Labor, and Welfare, Government of Japan, 2003b). (*) Concerning spousal allowances, the income threshold used most frequently (i.e., 1.03 million yen) is reported.

Figure 2.2: Reasons for Labor Supply Adjustment among Female Part-time Workers



Note: The sample is restricted to female part-time workers who answered that they adjusted their labor supply. Each bin in Figure 2.2 reports the percentage of those who answered each factor ($A \sim F$) affected their decisions of labor supply adjustment. $A \sim F$ correspond to the following: A: the limit of nontaxable income (1.03 million yen), B: Spousal exemption schedule, C-1: 1.3 million yen ceiling for no social security payment, C-2: 3/4 of regular worker's work hours at which workers enroll for the employees' pension- and health- insurance systems, D: Income threshold for the employer-provided spousal allowances, E: 20 hours/week at which workers enroll for the employment insurance, and F: Other.

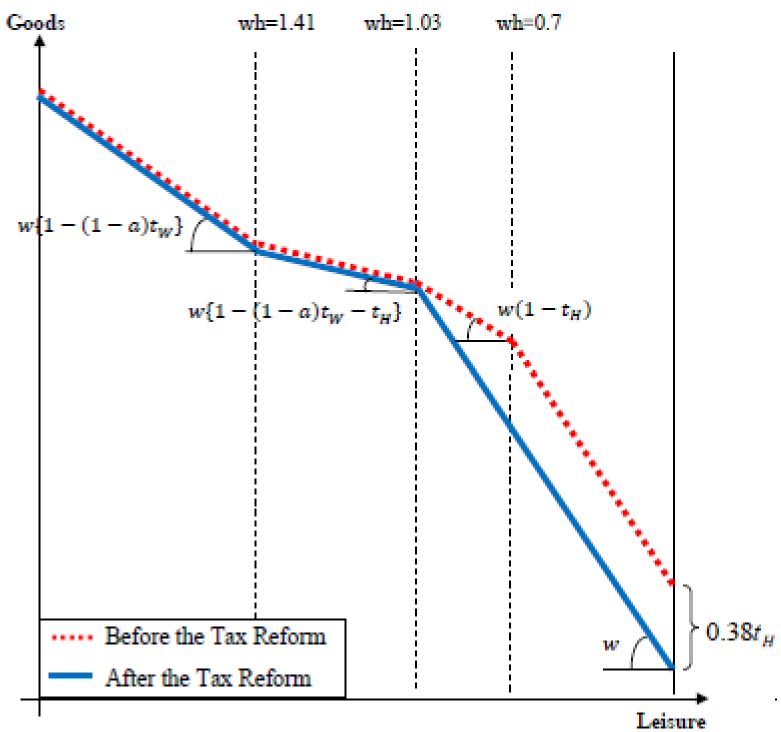
Figure 2.3: Spikes in the Annual Income Distribution of Married Women (2003)



Note: Figure 2.3 presents the 2003 income distribution of married women in the Keio Household Panel Survey (KHPS). The bin width is 0.1 million yen, but for the bin just below 1.03 million yen, people earning 0.9 to 1.03 million yen (instead of 0.9 to 1 million yen) are included in order to separate people below this threshold from those above it.

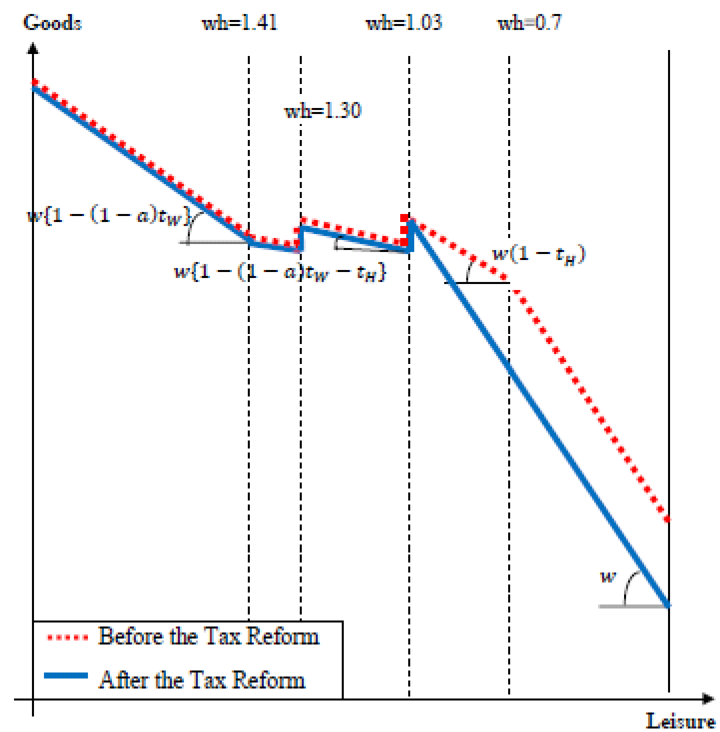
Figure 2.4: Impact of Tax reform on Household Budget Line

a. Household Budget Lines without Spousal Allowance and Social Security Payment



Note: This is the simplest version of the household budget line, which considers neither employer-provided spousal allowances nor the social security payment.

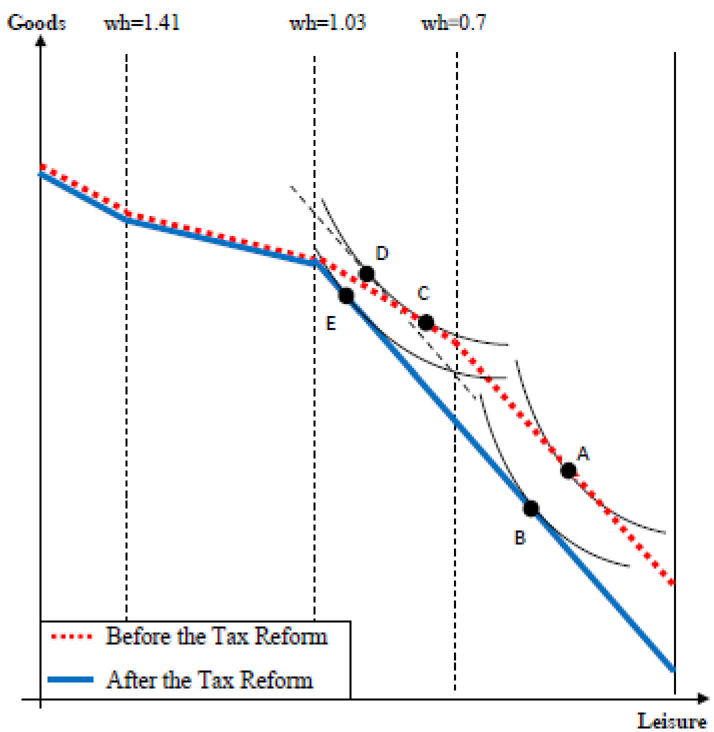
b. Household Budget Lines with Spousal Allowance and Social Security Payment



Note: In this figure, employer-provided spousal allowances that vanish at 1.03 million yen and social security payments that start from 1.3 million yen are considered.

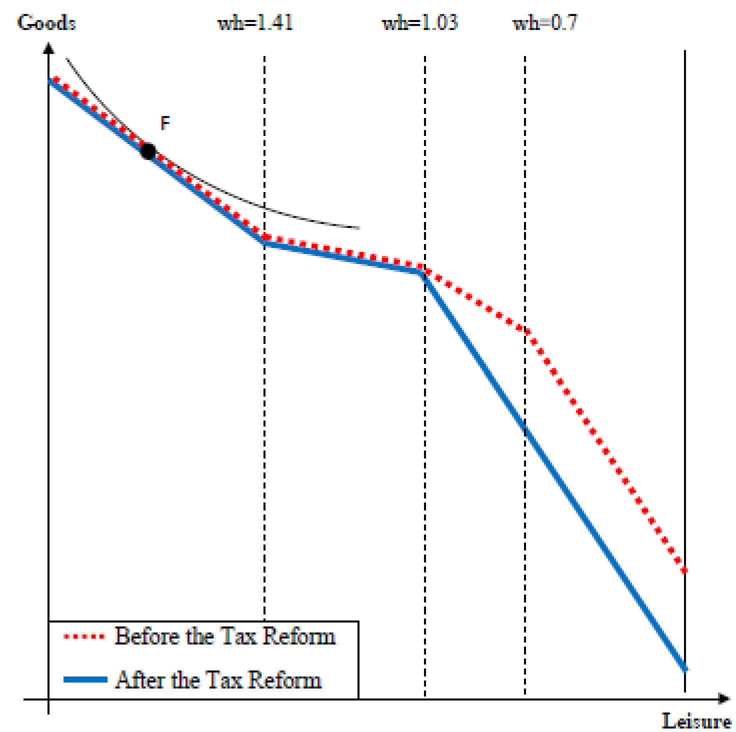
Figure 2.5: Impact of Tax reform on Labor Supply of Married Women (Without Change in Exogenous Parameter)

a. Impact on Married Women Who Earn 1.03 Million Yen or Less



Note: Married women at point A will move to point B after the tax reform because of the positive income effect. Married women at point C will move to point E because of the positive substitution effect ($C \rightarrow D$) and the positive income effect ($D \rightarrow E$).

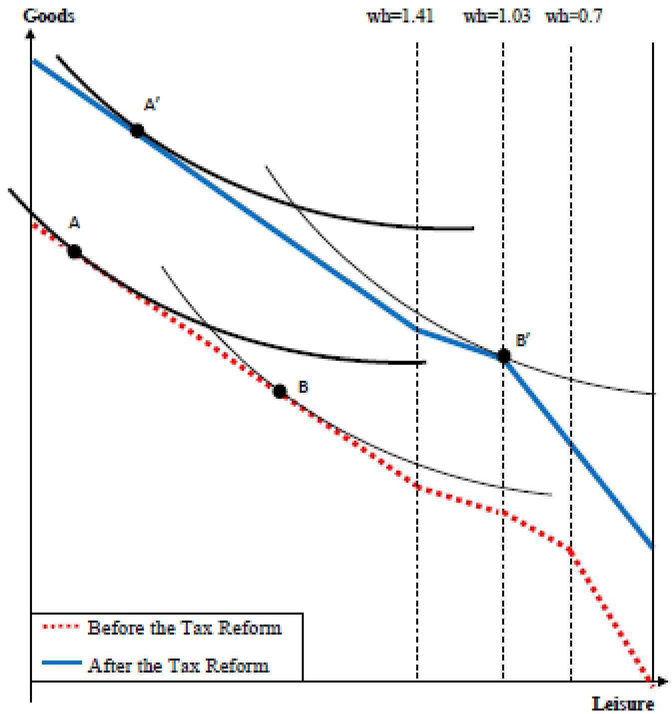
b. Impact on Married Women Who Earn More Than 1.03 Million Yen



Note: Without a change in exogenous variables, there is no impact on married women who earn more than 1.03 million yen.

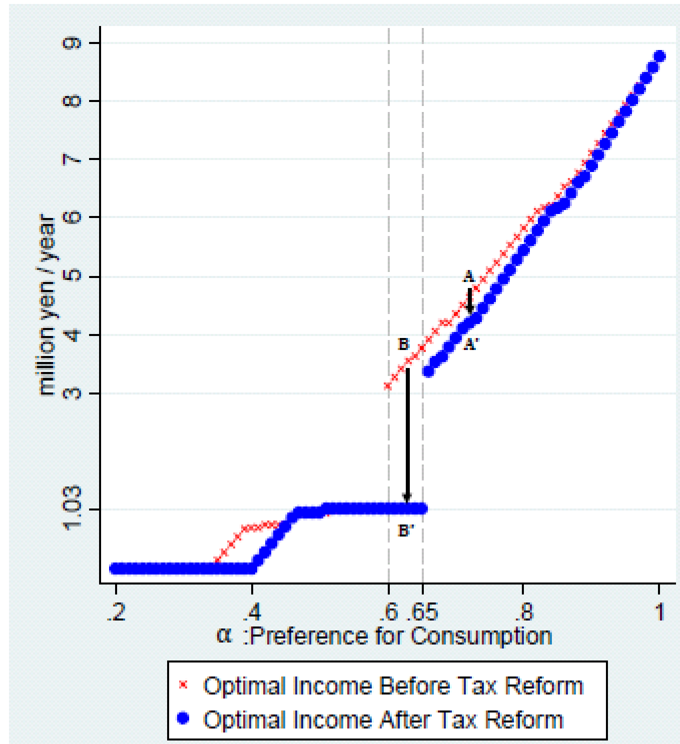
Figure 2.6: Change in Labor Supply of Medium- to High-income Married Women in Response to both Tax Reform and Increase in Husband's Income

a. "Income Jump" from 1.41 Million Yen or More to 1.03 Million Yen



Note: When an increase in a husband's income is considered, there are cases in which married women who earn 1.41 million yen or more will decrease their incomes to 1.03 million yen ($B \rightarrow B'$) after the tax reform.

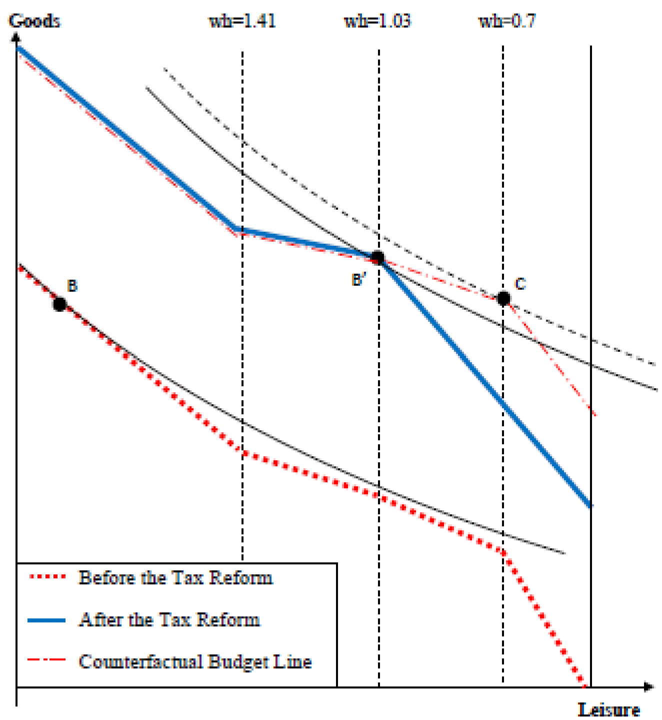
b. Wives' Optimal Income Choice by Preference for Consumption



Note: The dots in Figure 2.6b show the amount of optimal income that attains the highest utility for each value of α (i.e., the preference for consumption) when we use the Cobb-Douglas utility function. The simulation settings are the same as those in 2.4.A, except that Figure 2.6b reports the results with a sufficiently large increase in husbands' incomes (i.e., increase of 2 million yen).

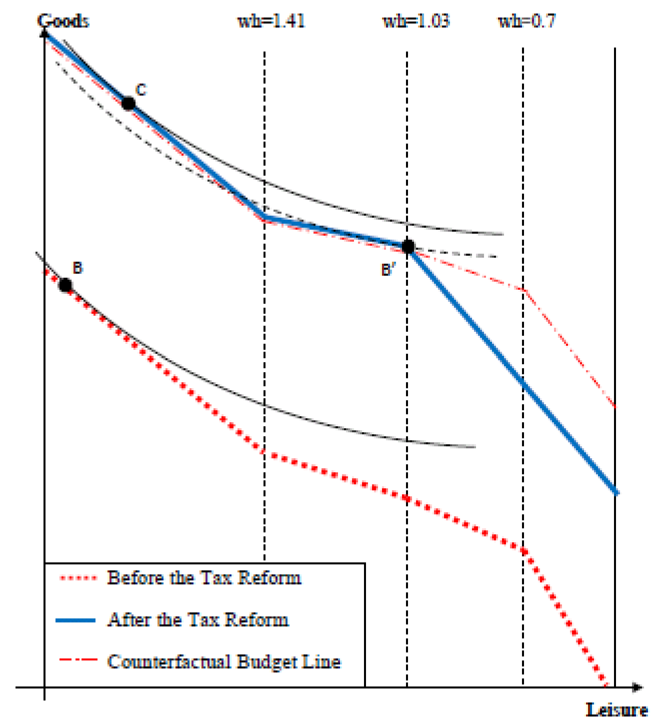
Figure 2.7: Decomposition of the “Income Jump” into “Tax Reform Effect” and “Husband’s Income Effect”

a. Decomposition of the “Income Jump”



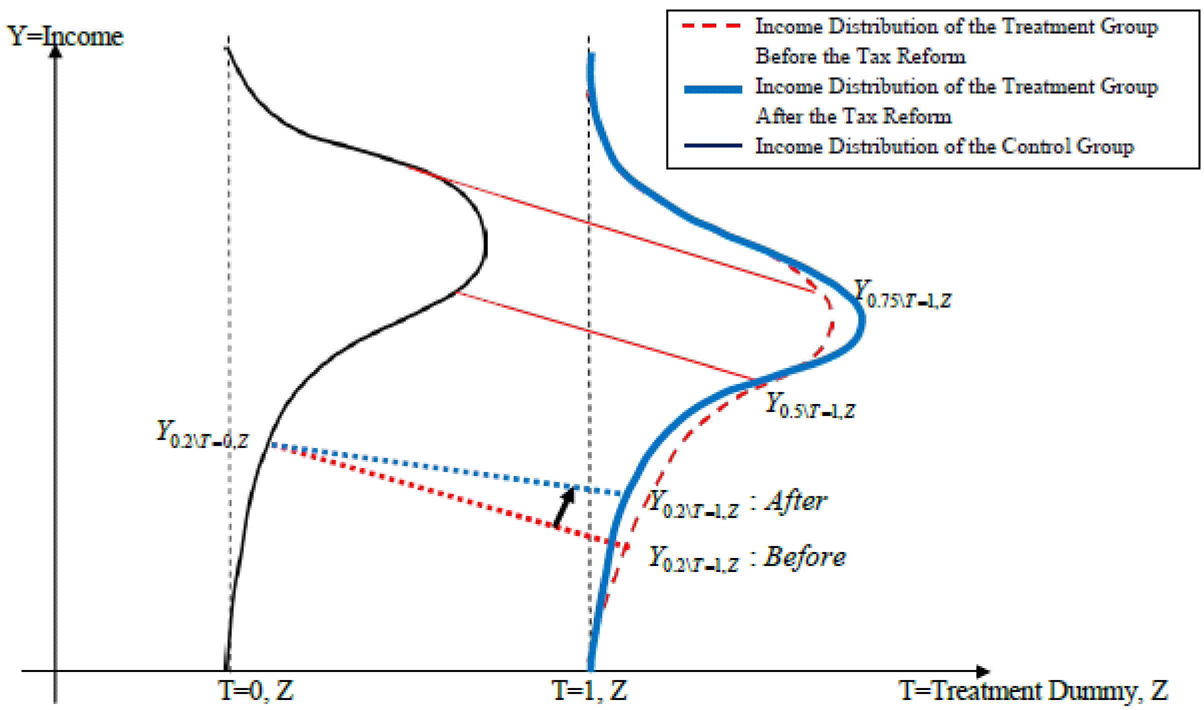
Note: The long-dashed line represents the counterfactual budget line without the tax reform in 2004. The “income jump” from B to B' is decomposed into a large negative “husband’s income effect” ($B \rightarrow C$) and a small positive “tax reform effect” ($C \rightarrow B'$).

b. “Tax Reform Effect” Cannot Be Negative



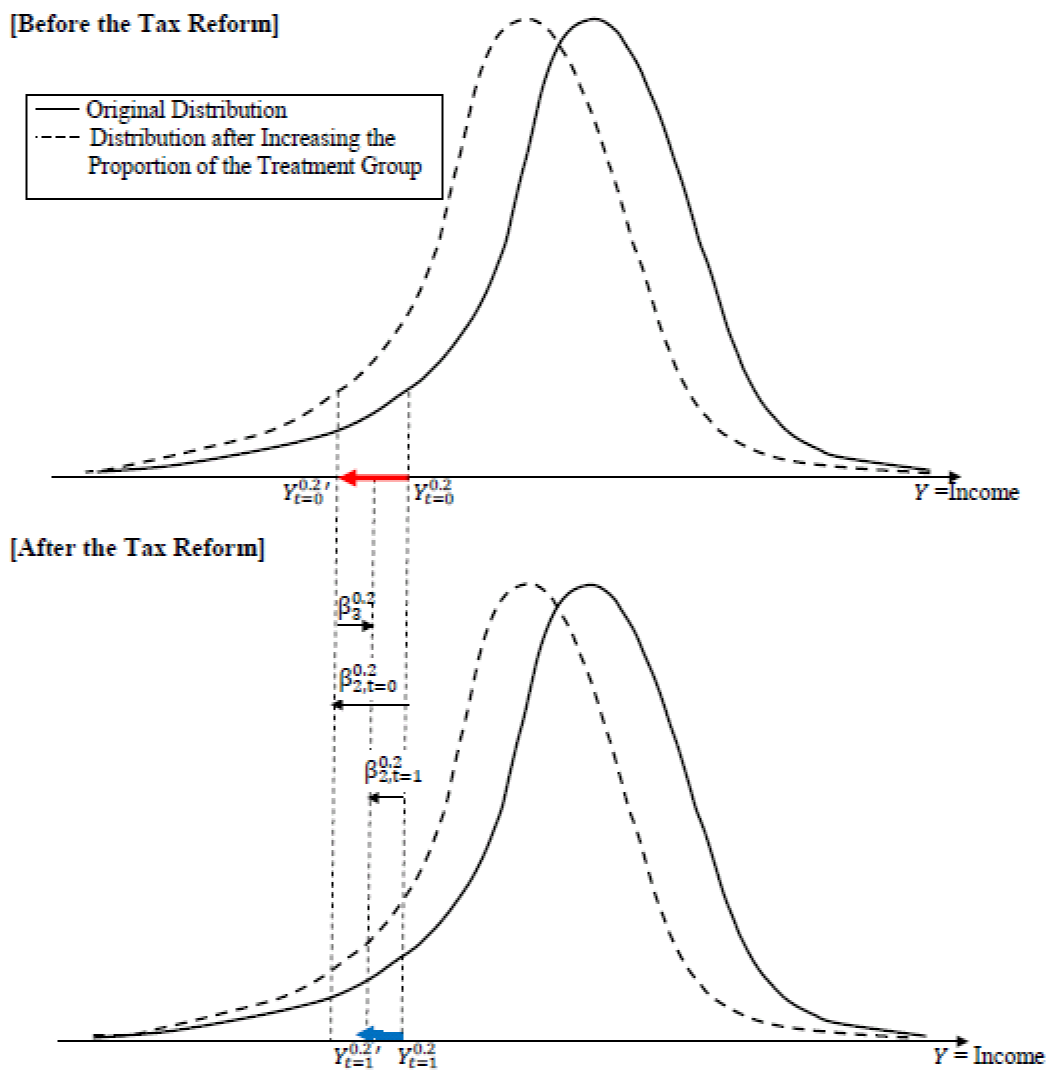
Note: This figure shows that a tax reform effect cannot be negative: both B' and C are available on the “after-reform” and counterfactual budget lines; a wife who chose point C on the counterfactual budget line would never choose point B' on the actual budget line following the tax reform, because point C is preferable to point B'.

Figure 2.8: Interpretation of CQR Estimate for β_3



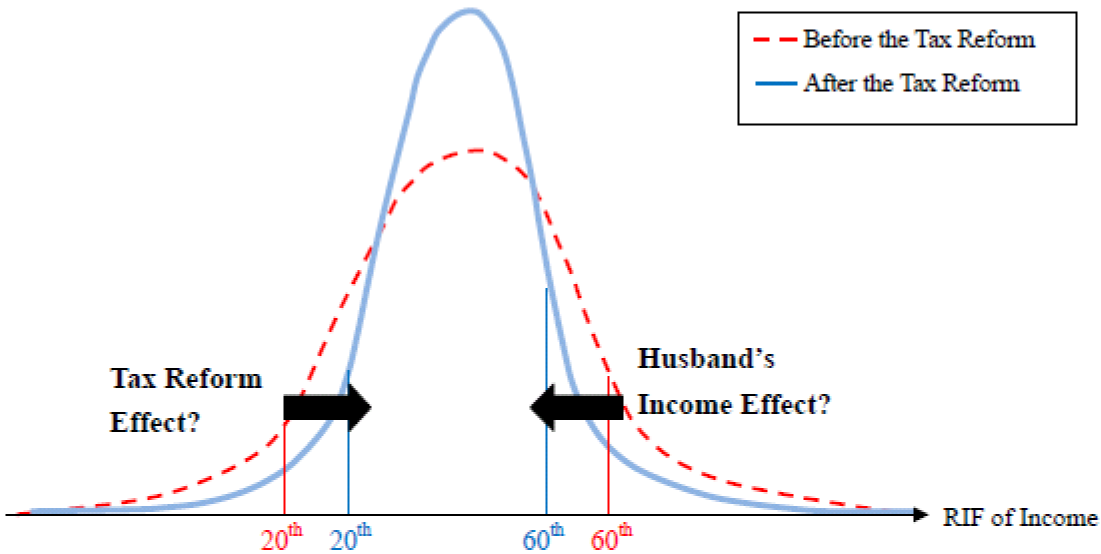
Note: Figure 2.8 illustrates the interpretation of the CQR estimate for β_3 visually. The distribution at the left-hand side is the income distribution for the control group conditioning on attributes Z. The dashed and bold lines at the right-hand side represent the income distributions before and after the tax reform, respectively. $Y_{\tau/T=1,Z}$ represents income at the τ quantile of the income distribution for the treatment group with workers' attributes Z, while $Y_{\tau/T=0,Z}$ represents income at the τ quantile of the income distribution for the control group with Z. The slope of the straight dotted line that connects incomes at the quantile of each of the two distributions represents the CQR estimates for the coefficient of the treatment dummy variable (i.e., $\beta_2^{\tau,CQR}$), which is calculated as $\beta_2^{\tau,CQR} = Y_{\tau/T=1,Z} - Y_{\tau/T=0,Z}$. The change in the slope, i.e., the gap in the slope between the two dotted lines, represents the magnitude of the interaction term between *treatment* and *after* dummy variables, i.e.,$\beta_3^{\tau,CQR}$.

Figure 2.9: Interpretation of UQR Estimate for β_3 [Change in Income Distribution When Increasing Proportion of the Treatment Group]



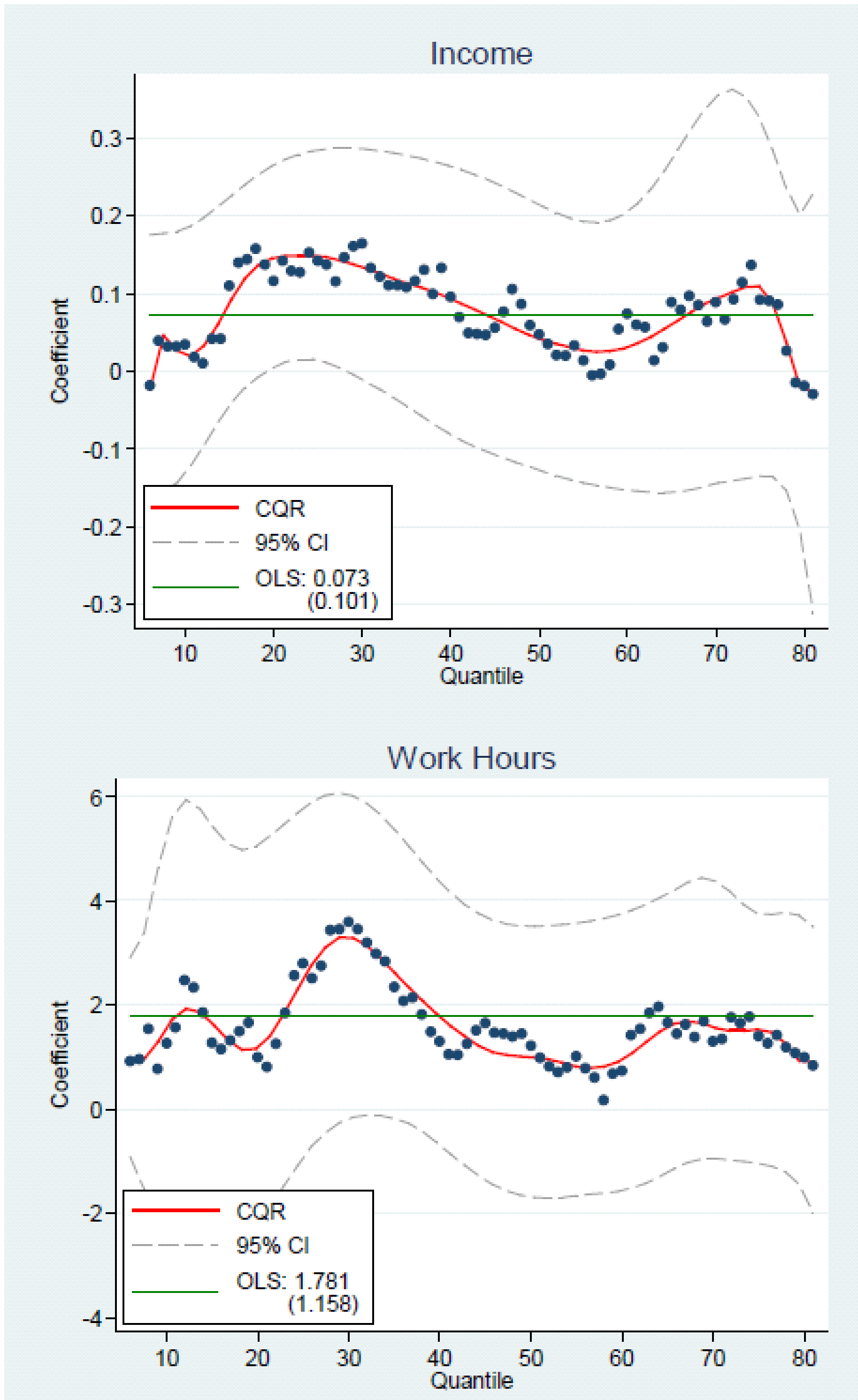
Note: Figure 2.9 illustrates the interpretation of the UQR estimate for β_3 visually. In Figure 2.9, we see how the distribution moves after increasing the proportion of the treatment group by one unit. Suppose that we start from the solid line and that the distribution shifts to the dashed distribution at the left when we increase the proportion of the treatment group by one unit. In the figure, income at the 20th quantile of the distribution shifts from $Y_{t=0}^{0.2}$ to $Y_{t=0}^{0.2\,\prime}$ when we increase the proportion of the treatment group by one unit before the tax reform. β_2 for the 20th quantile before the tax reform can be calculated as $\beta_{2,t=0}^{0.2} = Y_{t=0}^{0.2\,\prime} - Y_{t=0}^{0.2}$. Similarly, $\beta_{2,t=1}^{0.2} = Y_{t=1}^{0.2\,\prime} - Y_{t=1}^{0.2}$. Then, $\beta_3^{0.2} = \beta_{2,t=1}^{0.2} - \beta_{2,t=0}^{0.2}$. ($t$=0 and t=1 in superscript mean before and after the tax reform, respectively.) Note that in Figure 2.9, the two figures are placed vertically to compare the magnitude of $\beta_2^{0.2}$ and to show $\beta_3^{0.2}$ as the difference between the two $\beta_2^{0.2}$ values. However, this does not mean that $Y_{t=0}^{0.2} = Y_{t=1}^{0.2}$ holds.

Figure 2.10: Interpretation of FFL Results



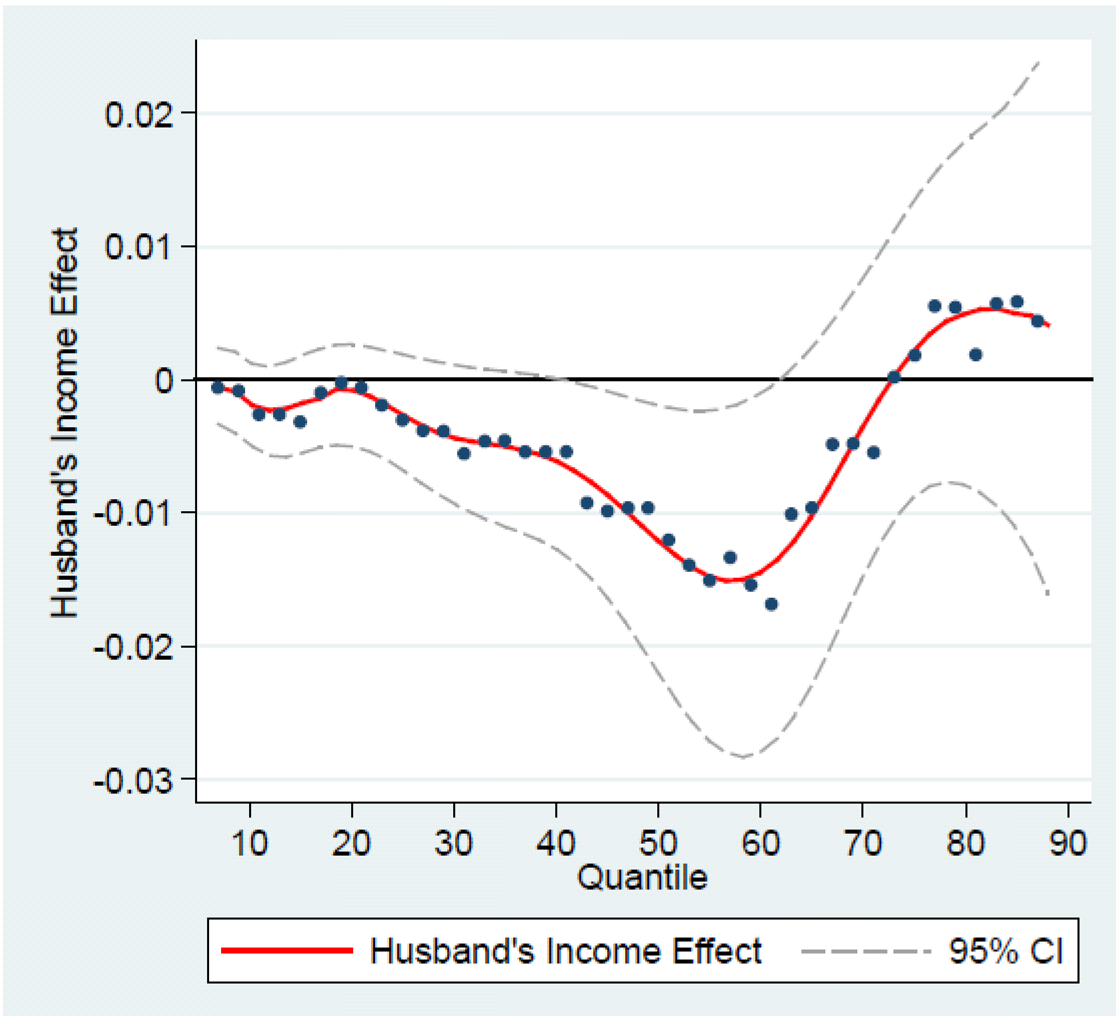
Note: Figure 2.10 illustrates the interpretation of the FFL estimates visually. The dashed line represents the income distribution before the tax reform, and the solid line represents that for after the tax reform. The quantiles are reassigned for each period in FFL, and FFL compares the RIF of income at quantile τ between before and after the tax reform. In the FFL decomposition, we can observe whether incomes among medium-to high quantiles decreased after the tax reform; and, if so, whether this was caused by the "income jump," which is explained mainly by a large negative effect of husband's income. Furthermore, the results obtained from the difference-in-difference estimation are reconfirmed in the FFL decomposition by checking whether incomes at low quantiles increased after the tax reform; and, if so, whether this was caused by the tax reform.

Figure 2.11: Coefficients for Interaction Term in the Difference-in-Difference Estimations



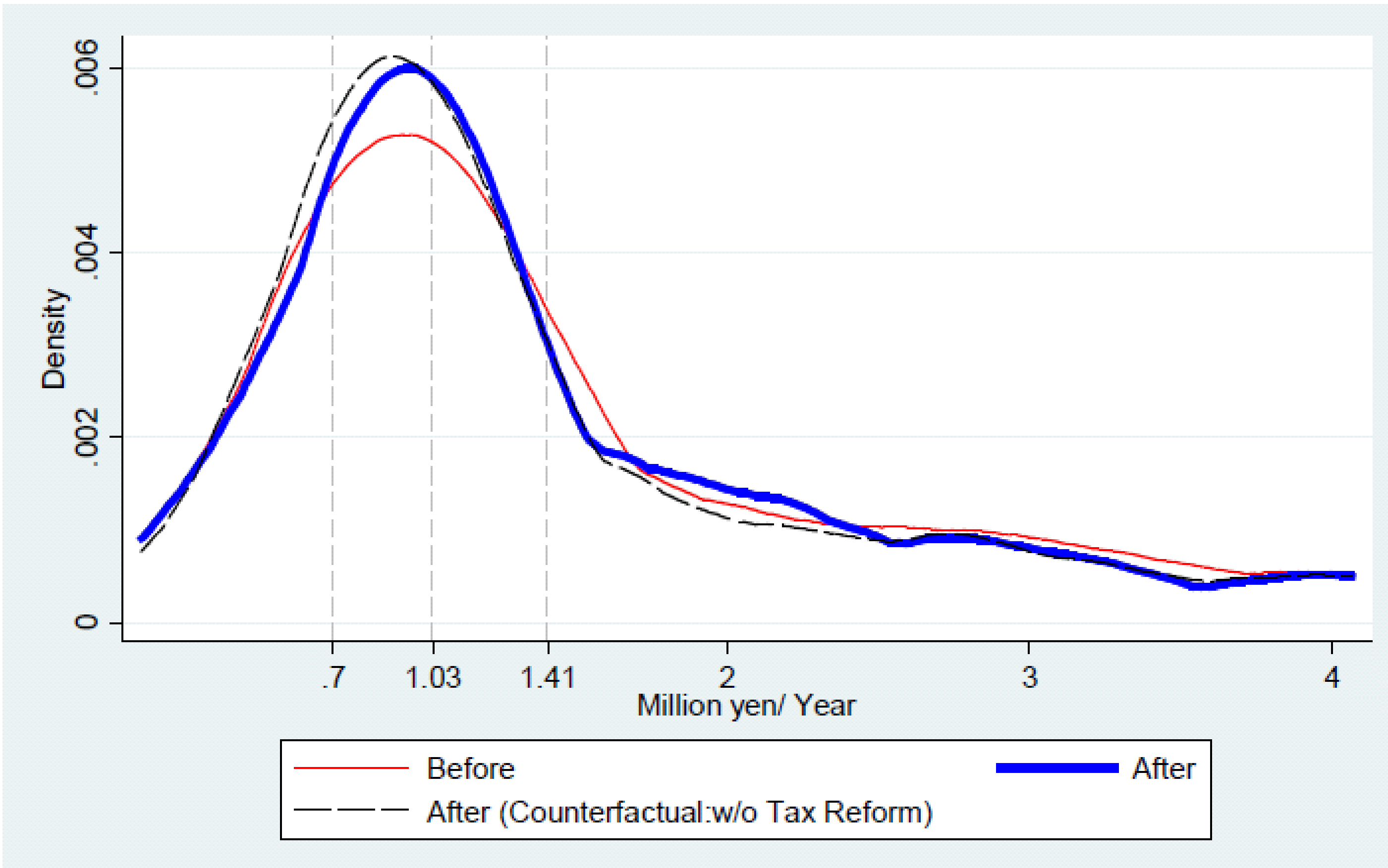
Note: The OLS and the CQR estimates for the coefficient of the ineraction term are reported. These values are obtained from Table 2.6.

Figure 2.12: Composition Effect Attributable to Husband's Income in FFL Decomposition



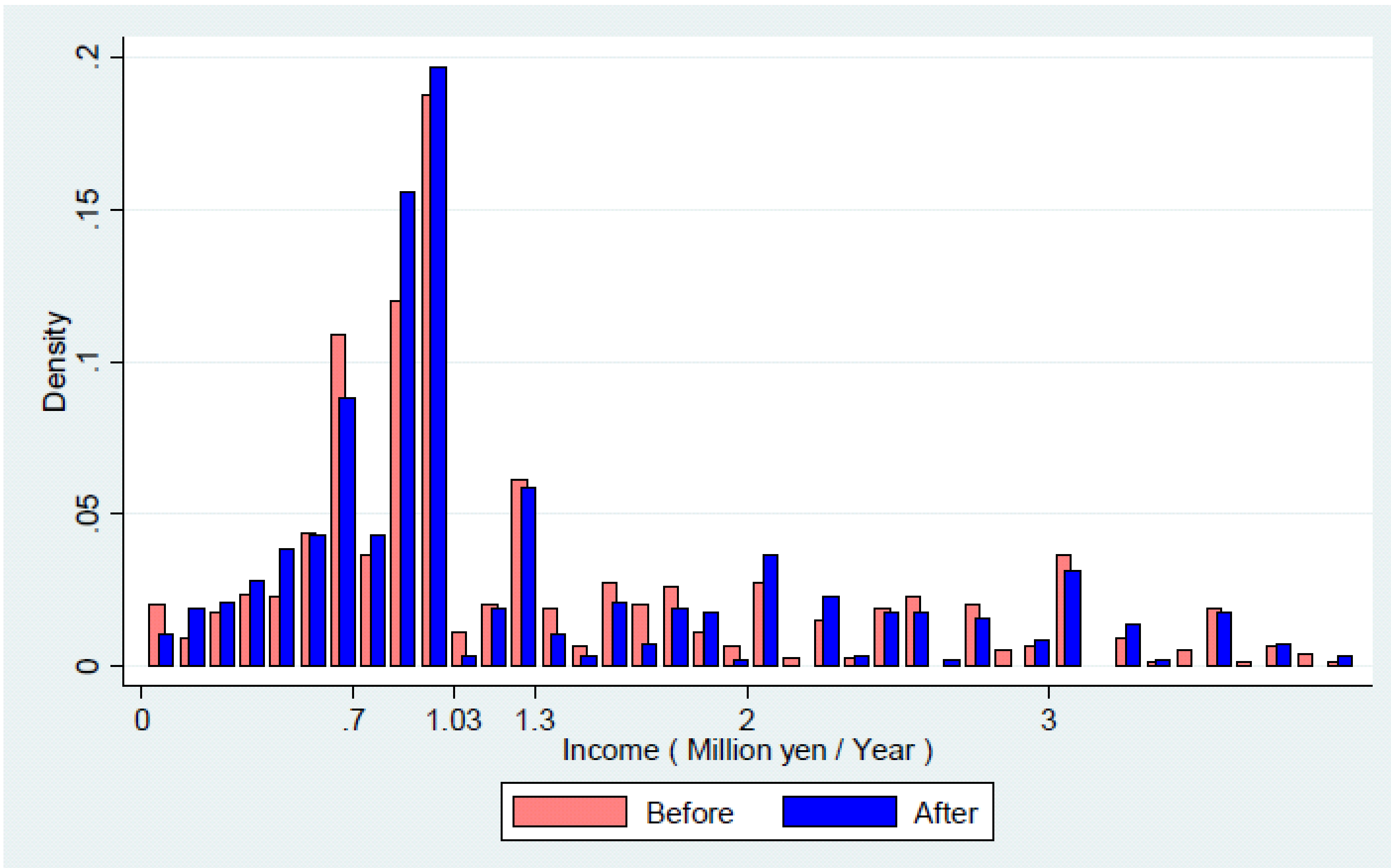
Note: These coefficients are those obtained from the FFL decomposition in Table 2.7.

Figure 2.13: Actual Income Distributions and Counterfactual Distribution (DFL Decomposition Result)



Note: The same explanatory variables as in FFL decomposition are used. Samples are married women in the Keio Household Panel Survey (KHPS), 2004-2007.

Figure 2.14: Changes in Histogram of Income among Married Women



Note: Samples are married women in the Keio Household Panel Survey (KHPS), 2004-2007. The bin width is 0.1 million yen, but for the bin just below 1.03 million yen, people earning 0.9 to 1.03 million yen (instead of 0.9 to 1 million yen) are included in order to separate people below this threshold from those above it.

Table 2.1.A: Employment Income Deduction Schedule

Parameters	Total Annual Earnings (Y)
a=0, b=0.65 million	if $Y \leq$ 1.625 million yen
a=0.4, b=0	if 1.625 million yen$<Y\leq$ 1.8 million yen
a=0.3, b=0.18 million	if 1.8 million yen$<Y\leq$ 3.6 million yen
a=0.2, b=0.54 million	if 3.6 million yen$<Y\leq$ 6.6 million yen
a=0.1, b=1.2 million	if 6.6 million yen$<Y\leq$ 10 million yen
a=0.05, b=1.7 million	if 10 million yen$<Y$

Note: The *employment income deduction* can be formulated as $aY + b$, where Y is the total annual earnings. All employees are eligible for a tax deduction of 0.65 million yen. If the amount of the tax deduction calculated from Table 2.1.A is less than 0.65 million yen, it is set at 0.65 million yen.

Table 2.1.B: Tax Rate Schedule for Income Tax (1999-2006)

Amount of Taxable Income (Unite: Million Yen≈10,000 USD)	Tax Rate (%)	Amount of Deduction (yen)
≤3.3	10 %	0
(3.3,9]	20 %	0.33 million
(9,18]	30 %	1.23 million
>18	37 %	2.49 million

Note: The schedule of the tax rate along with the deduction amount is listed in Table 2.1.B. The tax rate changed in 2007 from a 4-tier system (10 %, 20 %, 30 %, and 37 %) to a 6-tier system (5 %, 10 %, 20 %, 23 %, 33 %, and 40 %). Since incomes during 2003-2006 are analyzed in this paper, the previous 4-tier system is presented in Table 2.1.B. The tax rate is determined based on individual taxable income, which is calculated as total earnings − *employment income deduction* − *basic deduction* (of 0.38 million yen) − other deductions/exemptions (if applicable). Total income tax =Tax Rate×Taxable Income−Deduction in Table 2.1.B.

Table 2.1.C: Tax Rate Schedule for Income-Based Component of Residential Tax (1999-2006)

Amount of Taxable Income (Unite: Million Yen≈10,000 USD)	Tax Rate (%)	Amount of Deduction (yen)
≤ 2	5 %	0
(2,7]	10 %	0.1 million
>7	13 %	0.31 million

Note: Residential taxes are levied at a rate based on personal income of the previous year as well as uniform per-capita rate. The taxable income is calculated in the same way as the income tax, but the amount of the *basic deduction* for the residential tax is 0.33 million yen instead of 0.38 million yen.

Table 2.2: Spousal Exemptions of Income Tax and Residential Tax (Unit: Million Yen≈10,000 USD)

Spouse's Annual Earnings	Standard Spousal Exemption Amount	Special Spousal Exemption Amount		Total Spousal Exemption Amount	
	No Change	Before	After	Before	After
∼0.699..	0.38 (0.33)	0.38 (0.33)	0 (0)	0.76 (0.66)	0.38 (0.33)
0.70∼0.749..	0.38 (0.33)	0.33 (0.33)	0 (0)	0.71 (0.66)	0.38 (0.33)
0.75∼0.799..	0.38 (0.33)	0.28 (0.28)	0 (0)	0.66 (0.61)	0.38 (0.33)
0.80∼0.849..	0.38 (0.33)	0.23 (0.23)	0 (0)	0.61 (0.56)	0.38 (0.33)
0.85∼0.899..	0.38 (0.33)	0.18 (0.18)	0 (0)	0.56 (0.51)	0.38 (0.33)
0.90∼0.949..	0.38 (0.33)	0.13 (0.13)	0 (0)	0.51 (0.46)	0.38 (0.33)
0.95∼0.999..	0.38 (0.33)	0.08 (0.08)	0 (0)	0.46 (0.41)	0.38 (0.33)
1.00∼1.03	0.38 (0.33)	0.03 (0.03)	0 (0)	0.41 (0.36)	0.38 (0.33)
1.030..∼1.049..	0 (0)	0.38 (0.33)	0.38 (0.33)	0.38 (0.33)	0.38 (0.33)
1.05∼1.099..	0 (0)	0.36 (0.33)	0.36 (0.33)	0.36 (0.33)	0.36 (0.33)
1.10∼1.149..	0 (0)	0.31 (0.31)	0.31 (0.31)	0.31 (0.31)	0.31 (0.31)
1.15∼1.199..	0 (0)	0.26 (0.26)	0.26 (0.26)	0.26 (0.26)	0.26 (0.26)
1.20∼1.249..	0 (0)	0.21 (0.21)	0.21 (0.21)	0.21 (0.21)	0.21 (0.21)
1.25∼1.299..	0 (0)	0.16 (0.16)	0.16 (0.16)	0.16 (0.16)	0.16 (0.16)
1.30∼1.349..	0 (0)	0.11 (0.11)	0.11 (0.11)	0.11 (0.11)	0.11 (0.11)
1.35∼1.399..	0 (0)	0.06 (0.06)	0.06 (0.06)	0.06 (0.06)	0.06 (0.06)
1.40∼1.409..	0 (0)	0.03 (0.03)	0.03 (0.03)	0.03 (0.03)	0.03 (0.03)
1.41∼	0 (0)	0 (0)	0 (0)	0 (0)	0 (0)

Note: Table 2.2 shows the amount of spousal exemptions by the amount of spouse's annual earnings. The un-parenthesized numbers represent the income tax spousal exemptions and the parenthesized numbers represent the residential tax spousal exemptions. The schedule of the spousal exemption by "spouse's annual earnings" is applicable only for employees. The schedule applicable to all workers is defined by the *total income*, not by the total earnings, which is calculated as the total earnings minus the *employment income deduction* for employees and the total earnings minus necessary expenses for other workers.

Table 2.3.A: Japanese Labor Market Trends in the Period 2003-2006

	Before(2003)	After(2004-2006)	Difference
Unemployment Rate (%)	5.30 (2003)	4.10 (2006)	-1.20
Married Women			
Weekly Work Hours	30.37	28.67	-1.69
	(19.11)	(15.93)	(0.54)
Annual Income (million yen≈10,000 USD)	1.84	1.78	-0.06
	(1.92)	(1.76)	(0.04)
Husband's Annual Income (million yen≈10,000 USD)	5.23	5.36	0.13
	(3.32)	(3.23)	(0.04)
Hourly Wage Rate (100 yen≈1USD)	8.81	8.95	0.14
	(8.64)	(8.61)	(0.09)
Single Women			
Weekly Work Hours	36.18	37.46	1.28
	(18.38)	(17.56)	(1.33)
Annual Income (million yen≈10,000 USD)	2.13	2.25	0.11
	(1.40)	(1.44)	(0.06)
Hourly Wage Rate (100 yen≈1USD)	9.34	9.55	0.21
	(2.66)	(2.34)	(0.29)

Note: The unemployment rate is obtained from the Statistics Bureau. The other statistics are calculated from the Keio Household Panel Survey (KHPS) 2004-2007. Standard deviations are in parentheses. The sample is 2,896 women who participated in the Keio Household Panel Survey (KHPS) who reported information for both periods: before and after the 2004 tax reform. The statistics from the period after the tax reform are calculated as the average among the three years immediately following the tax reform. The statistics for weekly work hours, hourly wage rate, and annual income are obtained from sample participants who worked at least one hour during each period. The average husbands' annual income is obtained from all the married women in the sample. The hourly wage rate of hourly-paid workers is reported in the table due to the data limitation that hourly wage information is available only for this group.

Table 2.3.B: Changes in Total Spousal Exemption Amount (Unit: Million Yen≈10,000 USD)

	Mean	Std. Dev.	Min	Max
Income Tax				
Actual Spousal Exemption Before Tax Reform	0.640	0.184	0.030	0.760
Simulated Spousal Exemption After Tax Reform With the Same Earnings	0.369	0.050	0.030	0.380
Difference	-0.270	0.160	-0.380	0.000
Residential Tax				
Actual Spousal Exemption Before Tax Reform	0.560	0.156	0.030	0.660
Simulated Spousal Exemption After Tax Reform With the Same Earnings	0.322	0.040	0.030	0.330
Difference	-0.238	0.138	-0.330	0.000

Note: Samples are 2,135 women in the Keio Household Panel Survey (KHPS) whose husbands received spousal exemption greater than zero in 2003. The amount is calculated based on their 2003 annual income. The amount of spousal exemption after the tax reform is a simulated amount of spousal exemption each woman would receive when her earnings did not change over the period. Thus, the difference of the spousal exemption between the periods reflects only the impact of the tax reform.

Table 2.4.A: Impact of Tax Reform with an Increase in Husbands' Incomes on Wives' Optimal Income Choice

Income Group Before	Income Group After	Homothetic Cobb-Douglas	Homothetic CES	Homothetic CRRA	Non-homothetic Stone-Geary
<0.7	<0.7	39	27	0	28
	[0.7, 1.03)	0	0	0	0
	1.03	0	0	0	0
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	0	0	0	0
[0.7, 1.03)	<0.7	0	1	0	1
	[0.7, 1.03)	6	13	0	5
	1.03	8	16	7	8
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	0	0	0	0
1.03	<0.7	0	0	0	0
	[0.7, 1.03)	0	0	0	0
	1.03	6	10	10	7
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	0	0	0	0
(1.03, 1.30)	<0.7	0	0	0	0
	[0.7, 1.03)	0	0	0	0
	1.03	0	0	0	0
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	0	0	0	0
[1.30, 1.41)	<0.7	0	0	0	0
	[0.7, 1.03)	0	0	0	0
	1.03	0	0	0	0
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	0	0	0	0
≥1.41	<0.7	0	0	0	0
	[0.7, 1.03)	0	0	0	0
	1.03	1	1	1	1
	(1.03, 1.30)	0	0	0	0
	[1.30, 1.41)	0	0	0	0
	≥1.41	40	32	82	50
Observations		100	100	100	100

Note: Table 2.4.A presents the percentages of the 100 simulated optimal incomes that belong to each income group before and after the tax reform for each utility function. The parameters for a husband's income before and after the tax reform are set at the average amount of the husband's income in Table 2.3.A. The hourly wage and the price of goods are fixed at 1,000 yen and 100 yen, respectively. In the calculation, the social security payments, the decision making among the three categories of the social security system, residential tax, employer-provided spousal allowance, and employment insurance premium are also considered. Concerning these payments/allowances, the amounts specified in Figure 2.1 are assumed.

Table 2.4.B: Probability of "Income Jump" among High-Income Married Women (%) by Husband's Income Increase

Increase in Husband's Income (million yen≈10,000 USD)	Homothetic			Non-homothetic
	Cobb-Douglas	CES	CRRA	Stone-Geary
0.1	0.00	0.00	0.00	1.96
0.2	2.44	3.03	1.20	1.96
0.5	4.88	6.06	2.41	5.88
1	7.32	12.12	4.82	9.80
1.5	12.2	21.21	8.43	13.73
2	14.63	24.24	9.64	17.65

Note: Table 2.4.B reports the probability that married women in the highest income group will decrease their incomes to 1.03 million yen in response to an increase in their husbands' incomes after the tax reform, conditional on having been in the highest income group prior to the tax reform. Thus, the number of parameter values with which the "income jump" occurs is divided by the total number of parameter values that attain annual incomes of 1.41 million yen or more, prior to the 2004 tax reform. For example, in the Cobb-Douglas case with an increase in husband's income of 0.2 million yen, the probability of the "income jump" among high-income married women is calculated as 1/(1+40). The simulation settings are the same as those in Table 2.4.A.

Table 2.4.C: Impact of Tax Reform with an Increase in Husbands' Incomes on Wives' Optimal Income Choice (Counter-factual Realizations Included)

Income Group Before	After	Homothetic Cobb-Douglas	CES	CRRA	Non-homothetic Stone-Geary
< 0.7	<0.7	39 (39)	27 (27)	0 (0)	28 (28)
	[0.7, 1.03)	0 (0)	0 (0)	0 (0)	0 (0)
	1.03	0 (0)	0 (0)	0 (0)	0 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	0 (0)	0 (0)	0 (0)	0 (0)
[0.7, 1.03)	<0.7	5 (6)	14 (14)	0 (0)	8 (8)
	[0.7, 1.03)	6 (8)	9 (16)	0 (7)	6 (6)
	1.03	3 (0)	7 (0)	7 (0)	0 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	0 (0)	0 (0)	0 (0)	0 (0)
1.03	<0.7	0 (0)	0 (0)	0 (0)	0 (0)
	[0.7, 1.03)	0 (6)	0 (10)	0 (10)	0 (7)
	1.03	6 (0)	10 (0)	10 (0)	7 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	0 (0)	0 (0)	0 (0)	0 (0)
(1.03, 1.30)	<0.7	0 (0)	0 (0)	0 (0)	0 (0)
	[0.7, 1.03)	0 (0)	0 (0)	0 (0)	0 (0)
	1.03	0 (0)	0 (0)	0 (0)	0 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	0 (0)	0 (0)	0 (0)	0 (0)
[1.30, 1.41)	<0.7	0 (0)	0 (0)	0 (0)	0 (0)
	[0.7, 1.03)	0 (0)	0 (0)	0 (0)	0 (0)
	1.03	0 (0)	0 (0)	0 (0)	0 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	0 (0)	0 (0)	0 (0)	0 (0)
≥1.41	<0.7	0 (0)	0 (0)	0 (0)	0 (0)
	[0.7, 1.03)	0 (6)	0 (8)	0 (8)	0 (9)
	1.03	6 (0)	8 (0)	8 (0)	9 (0)
	(1.03, 1.30)	0 (0)	0 (0)	0 (0)	0 (0)
	[1.30, 1.41)	0 (0)	0 (0)	0 (0)	0 (0)
	≥1.41	35 (35)	25 (25)	75 (75)	42 (42)
Observations		100	100	100	100

Note: The parenthesized numbers represent the "counterfactual" realizations of the optimal income (i.e., the percentages of the 100 simulated optimal incomes where the tax reform had not occurred in 2004). The simulation settings are the same as those in 2.4.A, except that Table 2.4.C reports the results with a sufficiently large increase in husbands' incomes (i.e., increase of 2 million yen).

Table 2.4.D: Probability of "Income Jump" among High-Income Married Women (%) by Husband's Income Increase and Change in Wife's Hourly Wage

Change in Wife's Hourly Wage (yen (%))	Increase in Husband's Income (million yen≈10,000 USD) 0.2	0.5	1	1.5	2
-100 yen (-10%)	9.756	12.195	14.634	29.268	21.951
-50 yen (-5%)	4.878	7.317	9.756	29.268	17.073
-20 yen (-2%)	2.439	4.878	9.756	29.268	14.634
0 yen (0%)	2.439	4.878	7.317	26.829	14.634
+20 yen (2%)	0.000	2.439	4.878	24.39	12.195
+50 yen (5%)	0.000	0.000	2.439	12.195	9.756
+100 yen (10%)	0.000	0.000	0.000	31.707	7.317

Note: Table 2.4.D reports the probability of the "income jump" among married women who earned 1.41 million yen or more before the tax reform by husband's income increase and change in wife's hourly wage. These simulation settings are the same as those in 2.4.B. Only the result of the Cobb-Douglas utility function is reported.

Table 2.5: Descriptive Statistics

	All Women	Employees in Private Sector		
	(Sample for Probit)	All	Married Women	Single Women
Hours of Work (hours/week)	17.98	30.87	28.82	37.73
	(20.89)	(16.98)	(16.53)	(16.65)
Before-Tax Income	1.03	1.88	1.76	2.20
(million yen/year)	(1.60)	(1.62)	(1.63)	(1.40)
Husband's After-Tax Income	4.00	3.72	4.84	0.00
(million yen/year)	(3.14)	(2.99)	(2.49)	0.00
Age	48.15	44.22	46.32	37.25
	(12.86)	(11.31)	(9.78)	(13.14)
Tenure	10.04	7.03	7.50	5.44
	(10.58)	(7.62)	(7.86)	(6.56)
Number of Children	1.37	1.41	1.68	0.54
	(1.08)	(1.10)	(1.00)	(0.95)
Firm Size	149.82	233.80	229.91	246.78
	(181.84)	(197.40)	(197.18)	(197.71)
Dummy Variables				
Married	0.84	0.77	1.00	0.00
Working	0.55	1.00	1.00	1.00
Employee	0.41	1.00	1.00	1.00
Employee in Private Sector	0.37	1.00	1.00	1.00
Permanent Employee	0.29	0.59	0.57	0.68
Unionized Worker	0.07	0.14	0.12	0.17
Education Dummy Variables				
Junior High School	0.11	0.06	0.07	0.03
High School	0.55	0.57	0.60	0.46
Junior College	0.22	0.24	0.23	0.28
University	0.12	0.13	0.10	0.22
Income-group Dummy Variables				
$\text{Income}<0.7$	0.54	0.19	0.22	0.11
$0.7\leq\text{Income}\leq1.03$	0.14	0.25	0.30	0.10
$1.03<\text{Income}<1.41$	0.06	0.11	0.11	0.11
$1.41\leq\text{Income}$	0.27	0.44	0.37	0.68
Observations	10976	3206	2466	740

Note: The first column reports descriptive statistics for the probit regression, and the descriptive statistics for employees represent those of the sample used in the main analyses. Husbands' after-tax incomes that ignore the spousal exemptions are used for husbands' incomes. All the earnings data are corrected for inflation (i.e., the real earnings, calculated as nominal earnings divided by the consumer price index (CPI) of each year are reported for women's (before-tax) annual incomes and husbands' (after-tax) annual incomes that ignore the spousal exemptions.) The CPI of each year is 1, 1, 0.997, and 1 for 2003, 2004, 2005, and 2006, respectively. (The base CPI is that of 2003.) The husbands' incomes for single women are treated as zero and included in the sample. For this reason, the average value of the husbands' incomes is low for the sample that consists of both married women and single women.

Table 2.6: Difference-in-Difference Estimation: Effects of Tax Reform on Labor Supply

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Dependent Variable	Work=1	Annual Income (Million Yen)					RIF of Annual Income			Weekly Work Hours		
Quantile		Mean		30th	60th	90th	30th	60th	90th	30th	60th	90th
	Probit	OLS	FE	CQR			UQR			CQR		
Treatment	-0.334	-0.667	-0.366	-0.475	-0.464	-0.517	-0.307	-0.786	-0.537	-7.291	-4.801	-5.286
	(0.061)	(0.172)	(0.306)	(0.069)	(0.075)	(0.143)	(0.061)	(0.139)	(0.376)	(1.362)	(1.380)	(1.280)
After	0.129	-0.028	-	-0.138	-0.038	-0.082	-0.105	-0.061	0.030	-1.480	0.190	-1.130
	(0.052)	(0.090)		(0.064)	(0.069)	(0.128)	(0.046)	(0.120)	(0.257)	(1.257)	(1.267)	(1.192)
Treatment·After	-0.021	0.073	0.038	0.164	0.074	0.064	0.172	0.126	0.028	3.589	0.726	1.657
	(0.059)	(0.101)	(0.034)	(0.075)	(0.081)	(0.149)	(0.058)	(0.138)	(0.302)	(1.470)	(1.480)	(1.388)
Husband's Income	-0.053	0.036	0.198	-0.037	-0.032	-0.000	-0.031	-0.069	0.117	-0.874	-0.811	-0.537
(million yen/year)	(0.008)	(0.037)	(0.127)	(0.006)	(0.007)	(0.014)	(0.008)	(0.020)	(0.061)	(0.122)	(0.131)	(0.125)
Number of Children	-0.034	-0.152	-0.123	-0.067	-0.135	-0.171	-0.036	-0.202	-0.287	-0.873	-0.657	-0.995
	(0.020)	(0.033)	(0.070)	(0.017)	(0.018)	(0.036)	(0.020)	(0.051)	(0.112)	(0.335)	(0.332)	(0.332)
Age	0.191	0.081	-	0.072	0.100	0.100	0.024	0.098	0.318	0.335	0.153	-0.218
	(0.013)	(0.025)		(0.012)	(0.012)	(0.022)	(0.013)	(0.033)	(0.075)	(0.231)	(0.227)	(0.213)
$Age^2/100$	-0.221	-0.104	-	-0.090	-0.125	-0.125	-0.032	-0.137	-0.371	-0.525	-0.285	0.253
	(0.014)	(0.028)		(0.013)	(0.014)	(0.024)	(0.016)	(0.037)	(0.083)	(0.259)	(0.251)	(0.233)
Tenure	-	0.085	0.078	0.069	0.066	0.077	0.064	0.144	0.093	0.666	0.612	0.558
		(0.016)	(0.019)	(0.005)	(0.006)	(0.011)	(0.007)	(0.016)	(0.042)	(0.106)	(0.107)	(0.101)
$Tenure^2/100$	-	0.009	-0.229	-0.054	0.119	0.167	-0.136	-0.200	0.429	-0.790	-0.649	-0.784
		(0.065)	(0.086)	(0.018)	(0.020)	(0.039)	(0.022)	(0.054)	(0.162)	(0.357)	(0.359)	(0.336)
Unionized Worker	-	0.566	0.153	0.415	0.583	0.485	0.122	0.644	1.368	3.867	0.298	1.732
		(0.101)	(0.079)	(0.049)	(0.050)	(0.091)	(0.042)	(0.127)	(0.384)	(0.925)	(0.933)	(0.862)
Permanent Employee	-	0.614	0.133	0.302	0.424	0.654	0.073	0.648	1.564	5.319	5.660	4.787
		(0.062)	(0.057)	(0.033)	(0.034)	(0.064)	(0.036)	(0.086)	(0.180)	(0.650)	(0.635)	(0.605)
ln (Firm Size)	-	0.072	0.008	0.037	0.066	0.084	0.024	0.027	0.265	0.696	0.489	-0.037
		(0.020)	(0.017)	(0.011)	(0.012)	(0.021)	(0.012)	(0.029)	(0.068)	(0.218)	(0.213)	(0.200)
Education	Yes	Yes	No	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.083	0.409	0.137	0.160	0.294	0.355	0.190	0.349	0.270	0.151	0.163	0.100
Observations	10976	3206	3206	3206	3206	3206	3206	3206	3206	3206	3206	3206

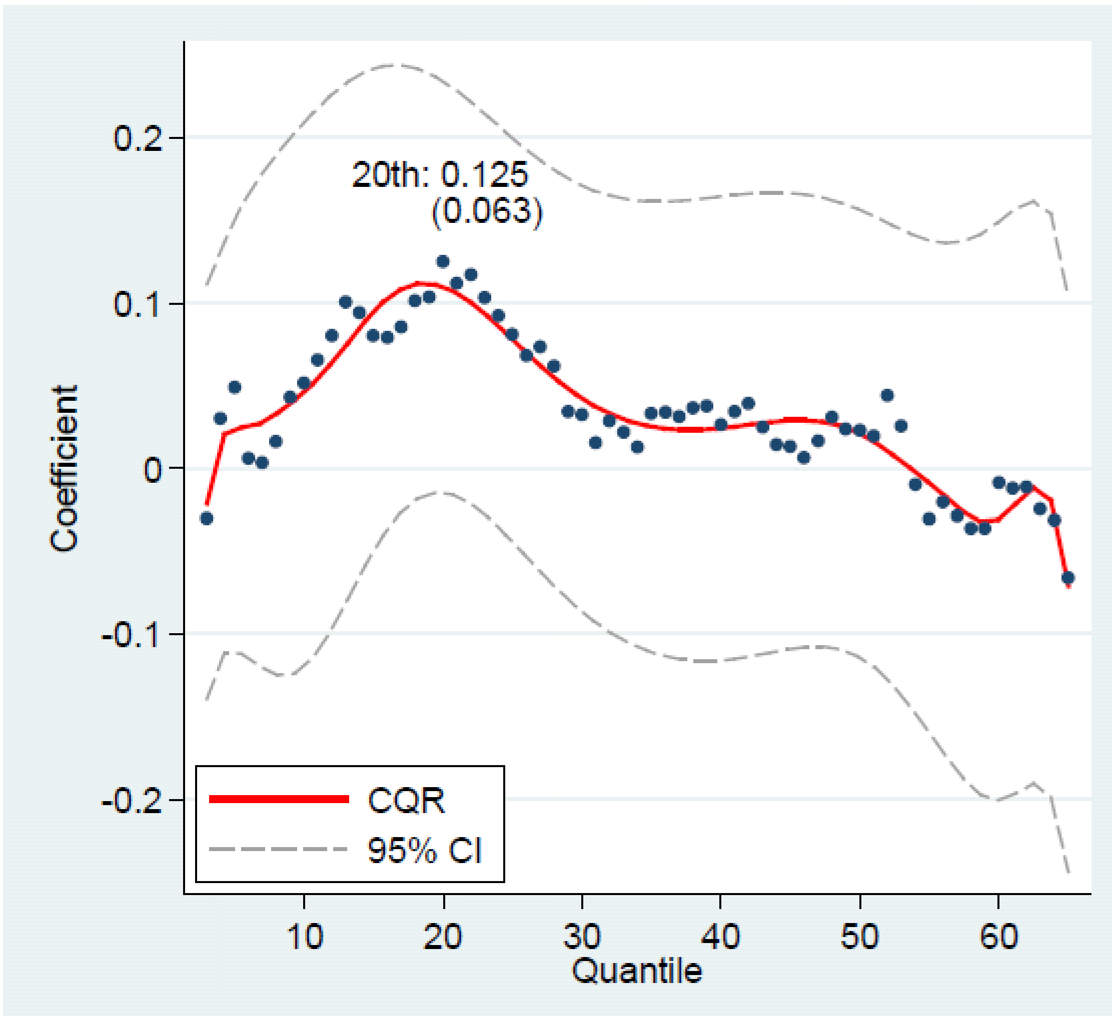
Note: Marginal effects are reported in Column 1. Standard errors are in parentheses. The Probit, OLS, and UQR standard errors are clustered at the individual level. A constant term is also included. Husbands' after-tax incomes that ignore the spousal exemptions are used for husbands' incomes. The treatment group is made up of married women whose husbands are eligible for the spousal exemptions (i.e., married women who are not the head of household and have husbands whose total income is 10 million yen or less per year). Single women and married women whose husbands are not eligible for the spousal exemptions are used as the control group.

Table 2.7: FFL Decomposition Results

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Reweighting	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes
Quantile	25th		50th		60th		70th		95th	
Total Change:	0.0002	-	-0.0106	-	-0.0072	-	-0.0616	-	-0.0544	-
	(0.0275)		(0.0354)		(0.0708)		(0.0857)		(0.1686)	
Composition Effects:										
Treatment	-0.0027	-0.0023	-0.0040	-0.0035	-0.0041	-0.0036	-0.0045	-0.0039	-0.0046	-0.0040
	(0.0033)	(0.0014)	(0.0050)	(0.0020)	(0.0055)	(0.0027)	(0.0062)	(0.0032)	(0.0096)	(0.0073)
Husband's Income	-0.0026	-0.0027	-0.0115	-0.0118	-0.0169	-0.0165	-0.0055	-0.0049	0.0202	0.0244
	(0.0026)	(0.0024)	(0.0058)	(0.0035)	(0.0093)	(0.0060)	(0.0073)	(0.0066)	(0.0200)	(0.0159)
Other	0.0047	0.0072	0.0025	0.0054	-0.0144	-0.0099	-0.0168	-0.0136	-0.0169	-0.0064
	(0.0125)	(0.0082)	(0.0207)	(0.0116)	(0.0437)	(0.0230)	(0.0542)	(0.0287)	(0.1102)	(0.0553)
Total	-0.0006	0.0021	-0.0131	-0.0098	-0.0354	-0.0300	-0.0268	-0.0224	-0.0013	0.0140
	(0.0130)	(0.0081)	(0.0225)	(0.0116)	(0.0453)	(0.0227)	(0.0544)	(0.0276)	(0.1106)	(0.0531)
Structure Effects:										
Treatment	0.1442	0.1785	0.1272	0.1189	-0.0996	-0.0063	-0.0521	0.0845	-0.0828	-0.1662
	(0.0841)	(0.0822)	(0.1070)	(0.1065)	(0.1973)	(0.1933)	(0.2348)	(0.2496)	(0.6007)	(0.5273)
Husband's Income	-0.1186	-0.1494	0.0044	-0.0090	0.0610	0.0393	-0.1029	-0.0959	0.1650	0.1349
	(0.0792)	(0.0790)	(0.0968)	(0.0963)	(0.1730)	(0.1672)	(0.2084)	(0.2164)	(0.5781)	(0.5042)
Other/Constant	-0.0248	-0.0306	-0.1291	-0.1076	0.0667	0.0416	0.1202	-0.0223	-0.1353	-0.0423
	(0.0544)	(0.0526)	(0.0774)	(0.0729)	(0.1630)	(0.1613)	(0.1952)	(0.2097)	(0.3069)	(0.2934)
Total	0.0008	-0.0015	0.0025	0.0023	0.0282	0.0746	-0.0347	-0.0337	-0.0531	-0.0735
	(0.0265)	(0.0256)	(0.0311)	(0.0292)	(0.0600)	(0.0555)	(0.0730)	(0.0719)	(0.1436)	(0.1331)

Note: Standard errors are clustered at the individual level and reported in parentheses. The dependent variable is the RIF of annual income (million yen). "*Husband's Income*" includes husband's after-tax income that ignores spousal exemptions and its square. The "*Other*" category includes age and its square, number of children, tenure and its square, unionized worker dummy variable, permanent employee dummy variable, educational dummy variables, and industry dummy variables. The standard errors of the decomposition components are computed using the delta method and take into account the variability induced by stochastic regressors (Jann, 2008). The odd-numbered columns report the results of the decomposition using RIF-regressions without reweighting. The even-numbered columns report the results of the reweighted-regression decomposition, in which $F(X)$ in 2003 is reweighted to 2004-2006.

Figure A.2.1: Tax Reform Effects for Higher Husbands' Tax Rates



Note: The sample includes "eligible" married women whose husbands' tax rates are 10% or 20%. In this regression, annual income is regressed on $I(t_H = 0.2)$, $After$, and $I(t_H = 0.2) \times After$ and other explanatory variables. Figure A.2.1 plots CQR estimates for the coefficient of $I(t_H = 0.2) \times After$. The dependent and independent variables used in this regression are the same as those in Column 4 to 6 in Table 2.6, except that the treatment group in this regression is "eligible" married women whose husbands' tax rate is 20% and the control group is "eligible" married women whose husbands' tax rate is 10%.

Table A.2.1: Check for Selection Issue due to Attrition

	All Women	Participated in All Years	Dropped
Hours of Work (hours/week)	17.60	17.41	17.98
	(21.47)	(21.30)	(21.83)
Before-Tax Income	0.98	0.98	0.98
(million yen/year)	(1.60)	(1.60)	(1.59)
Age	47.17	46.94	47.65
	(12.91)	(12.61)	(13.48)
Number of Children	1.36	1.41	1.25
	(1.07)	(1.07)	(1.08)
Married	0.84	0.85	0.82
Working	0.52	0.53	0.51
Employed	0.39	0.38	0.39
Education Dummy Variables			
Junior High School	0.13	0.11	0.15
High School	0.55	0.55	0.56
Junior College	0.21	0.21	0.19
University	0.11	0.12	0.10
Observations	3269	2189	1080

Note: The descriptive statistics represent those for 2004. The sample in the first column includes all women who participated in the 2004 survey. The sample in the second column includes only those workers who participated in all years (2004-2007), and the third column reports information about those workers who dropped from the sample at some point during the 2004-2007 period. Since the sample includes both working and non-working women, 0 hours and 0 yen are included in the calculations of the average work hours per week and the average annual income, respectively.

CHAPTER III

Why Do Wages Become More Rigid during a Recession Than during a Boom?

3.1 Introduction

When unemployed workers are available, why do firms not cut wages until the excess supply is eliminated, as is expected in the ideal market scenario depicted by conventional theory? This question has puzzled many economists, and a number of studies have attempted to solve the dilemma. Noteworthy among them are the efforts of Bewley (1999), who conducted commendable field research, and provided us with a clue to the answer. In the research, Bewley (1999) found that none of the existing theories about wage rigidity correctly explains his findings in the "real" US labor market, which implies the need for a new theoretical model.

One of the reasons why research on wage rigidity during recession is important is that wages during recessions could be a factor that would reduce high unemployment; i.e., theoretically, firms could hire more people by paying lower wages to the existing workers. In this sense, it would not matter if wages increase or do not fall during booms; however, if wages do not fall during recessions, it could be a factor that prevents new workers from being hired. Thus, if there is a special reason wages do not fall during recessions, it would be important to understand the possible source of such wage rigidity.

There are two core concepts in wage rigidity: real wage rigidity and nominal wage rigidity. Bewley (1999) defines each concept of wage rigidity as follows: "Real wages are downwardly rigid if employers feel obliged to increase pay by at least the rate of inflation in the cost of living. Nominal wages are rigid if there is resistance to cutting nominal pay but not to increasing pay by less than the rate of inflation in the cost of living."

In this paper, I will explore the mechanism in which the nominal wages become downwardly rigid during recessions (but not during booms) by showing that firms have a reason to resist cutting nominal pay only during recessions. To do so, I construct a new type of efficiency wage model in which workers' efforts could be maintained not only by high wages (under the threat of performance-based layoffs), but also by performance-based pay and promotion. In the sense that high wages induce workers' high effort via the threat of dismissal, this paper utilizes the idea of the shirking model. However, the following attributes make this model different from the existing shirking models.

First, wages are assumed to be performance-based, unlike the standard efficiency wage hypothesis, which assumes fixed wages. Recently, performance-based pay has been employed in many countries. For example, Lemieux et al. (2009) show that, in the U.S., the proportion of performance-pay jobs is increasing. Moreover, the "pay for performance" system has been widespread in many countries. This is particularly so in Japan, where most employees traditionally receive a substantial portion of their pay in the form of bonuses. This trend has been particularly widespread since the 1990s. In Japan, it is generally the case that the bonus is a payment that can fluctuate, depending on the firm's business and the worker's performance. Thus, particularly for countries such as Japan, it is important to assume performance-based pay rather than fixed wages. Given this scenario, in the model presented in this paper, a worker's total compensation is divided into two components: a fixed pay component (regular pay) and a performance-based pay component (bonus). This division enables us to understand how each component of pay contributes to wage rigidity during recessions.

Second, the layoff decision is endogenous, and is allowed to be contingent on market conditions. In Shapiro and Stiglitz (1984), the dismissal rule is exogenously given, and workers caught shirking are fired regardless of market conditions. Sparks (1986) further developed the rule of Shapiro and Stiglitz (1984) by making both workers' effort levels and the criterion for dismissal endogenous. In Sparks' model, it is assumed that workers who provide effort equal to or above the minimum effort standard are never dismissed, and a firm offers workers a labor contract that specifies a wage and the minimum effort standard. In equilibrium, workers' efforts are set equal to the minimum effort standard, which yields no dismissals regardless of market conditions, as Shapiro and Stiglitz (1984) demonstrated. However, in the real world, while unionized experienced employees' jobs are relatively secure even during recessions, it is widely observed that many non-union workers are actually laid off during recessions, depending on their performance. According to Bewley (1999), 28 % (86 %) of non-union (unionized) workers are laid off based on inverse seniority, while 57 % (7 %) of non-union (unionized) workers are laid off based on performance.[1] In order to capture these widely observed practices, I adopt a model in which firms decide how much weight they put on performance (or seniority) as a layoff criterion, rather than the "minimum effort standard."[2] In this way, it becomes possible to allow the layoff criteria to differ between unionized workers and non-union workers, and in this

[1]The remaining category is "both performance and inverse seniority," and 7 % of unionized workers and 13 % of non-union workers are categorized into this category.

[2]In most economic models, firms lay off their workers either randomly or based on seniority. For example, Baily (1977) and Macleod et al. (1994) present models in which firms lay off workers randomly, while Grossman (1983) and Reagan (1992) assume seniority-based layoffs. Nosal (1990), Strand (1991), and Strand (1992) consider both types of layoffs. Laing (1994) and Gibbons and Katz (1991) propose signaling models, in which firms may choose to lay off workers according to observed ability. Ioannides and Pissarides (1983) present a model in which a firm decides to lay off a worker based on the information of an external offer to the worker.

model, layoffs can actually occur only during recessions in equilibrium. Consequently, unlike Shapiro and Stiglitz (1984) and Sparks (1986), I arrive at an answer to the following question: Why do wages not fall during downturns in which firms lay off many workers, thereby creating high unemployment?

Third, I extend the standard shirking theory by introducing promotions. Since promotion is also an important instrument for maintaining workers' incentives, in Section 3.3, I present the model with promotions.

The main results obtained from the theoretical model are as follows: (a) Performance-based layoffs are more likely to occur when the layoff costs and the output price are low; (b) experienced employees' regular pay is likely to be downwardly rigid during periods in which performance-based layoffs occur; and (c) bonuses move proportionally to the output price. Together with the result (a), the result (b) has an implication on the downward rigidity of regular pay during recessions: without the threat of layoffs, wages scheduled to be paid in the next period do not affect current workers' efforts because workers will necessarily receive the wages without being laid off in the next period, regardless of their current effort levels. Thus, the firm cannot control the workers' efforts using their future wages, which results in a lower, at least, a less downwardly rigid regular pay without the threat of layoffs. In contrast, under the threat of layoffs, workers can receive the next period's wages only when they work hard in the current period and avoid layoffs. Therefore, the higher wages scheduled to be paid after performance-based layoffs have occurred are, the harder workers try to avoid layoffs, thus investing greater efforts. This gives the firm an incentive to raise the future regular pay to maintain workers' current efforts, which results in a downwardly rigid regular pay during recessions.

The rationale behind (c) is very simple. During recessions, firms are discouraged from maintaining workers' current incentives at a high level due to the lower value of productivity. In contrast, during booms, it is more beneficial for firms to raise bonuses so as to induce higher effort levels from workers because the value of productivity is high. As a result, the bonus moves proportionally to the output price. I also show that introducing promotions does not change these main results.

In order to test these theoretical implications obtained from the theoretical model, I conduct an empirical analysis using the Japanese panel data from the Keio Household Panel Survey (KHPS) 2004-2007. There are two steps in the estimations of the theoretical model: The first step is a layoff regression, employed to confirm that performance-based layoffs are more likely to occur for non-union workers whose performance-based layoff costs are relatively low and during recessions in which the output price is low. The second step is a wage regression, which shows that regular pay becomes downwardly rigid when performance-based layoffs are likely to occur, i.e., for non-union workers during recessions, and that bonuses just move proportionally to the output price.

This paper is organized as follows. Section 3.2 describes the framework of the basic model. Section 3.3 shows that the results obtained in the basic model remain unchanged even if promotions are introduced. Section 3.4 presents a strategy to test the implications of the theoretical model, followed by a possible solution for the selection bias problem. Section 3.5 provides a brief description of the data, and Section 3.6 discuses the results from my empirical analysis. Section 3.7 contains the conclusion to this paper.

3.2 The Basic Model

3.2.1 Model Structure

Worker Types

I assume that there are only two types of workers in each firm: new employees and experienced employees. Workers who have been recently recruited by a firm are categorized as new employees, and workers who have continued with the same employer are categorized as experienced. Here, by simplicity, workers are assumed to have a two-period time horizon. In the first period, they join a firm as new employees. In the second period, they either remain with the same employer as experienced employees or are laid off at the beginning of the second period.[3]

Output Price

Firms are assumed to be price-takers, and the output price is drawn randomly from a distribution $G(p)$ with density function $G(p) > 0$ for $p \in [p^-, p^+]$. Thus, the output price is assumed to be *i.i.d.* here, so that a higher value of p_t does not lead the firm (or others) to expect a higher value of p_{t+1}.

Wages

A worker's wage is divided into two components —a fixed pay component and a performance-based pay component— and it is expressed as follows:

$$w_t(p_t) = a_t(p_t) + b_t(p_t)\, e_t \tag{3.1}$$

where a_t represents regular (fixed) pay and b_t represents the "piece rate" paid for each unit of effort, e_t. It is assumed that both a_t and b_t are contingent on the output price, p_t, and that the contract specifies $(a_t(p_t), b_t(p_t))$ for each realization of p_t.

Layoffs

When a firm uses workers' "performance" as a criterion for layoffs, the layoff probability for a worker is assumed to be a decreasing function of the efforts offered by the same worker in the previous period. In contrast, if the firm uses "seniority" as

[3]In this model, I assume that workers who have been recently employed by a firm are not laid off. The firm thus controls the entire labor force by adjusting the number of new employees and the number of layoffs of experienced employees.

a criterion for layoffs, an experienced worker who has already worked for one period in the firm will not be laid off with a probability 1. It is assumed that the firm can choose how much weight to put on performance while choosing the layoff criterion. Let $\gamma_t \in [0, 1]$ denote the weight put on performance by the firm while deciding the layoff criterion. It is assumed that γ is also contingent on the output price, p, i.e. $\gamma(p)$.

Then, the probability of a worker being retained in the firm is expressed as:

$$\gamma_t(p_t) \cdot min\left(\frac{e_N}{\overline{e}}, 1\right) + (1 - \gamma_t(p_t)) \cdot 1 \tag{3.2}$$

where e_N is a new employee's effort. The subscripts N, E, hereafter, represent the types of workers, new employees and experienced employees, respectively. I assume that $\overline{e}$ is exogenously given, while γ_t is chosen by the firm. Note that the higher γ_t is, the higher the layoff risks that a worker faces. If $\gamma_t = 1$, the layoff decision is completely performance-based, and the fraction, $(1-e_N/\overline{e})$, of experienced employees will be laid-off. Note that a "completely performance-based" layoff decision does not mean that the firm lays off all workers whose efforts in the first period are less than $\overline{e}$: it means that an increase/decrease in e_N will be fully reflected in the layoff probability. For $\gamma_t \in [0, 1]$, changes in the effort level will be partially reflected in the layoff probability. If $\gamma_t = 0$, the layoff decision is completely seniority-based and thus, an experienced worker who has already worked for one period will be retained in the firm with a probability 1. In this case, changes in the effort level do not affect the layoff probability at all.

Timing

The timing is given by the following:

1. The output price, p_t, is observed by both the firm and its workers.

2. Each firm selects a labor contract that will be offered to the new employees.
3. New employees decide whether to accept or reject the firm's offer. Experienced employees also decide whether to stay in the same firm.
4. Both new employees and experienced employees who continue with the firm exert effort, production occurs, profits are realized, and payments are made.
5. In case of $\gamma_t > 0$, workers are laid off with a probability $\gamma_{t+1}(p_{t+1})(1 - e_N/\overline{e})$. If not laid off, workers who have finished their first period at the firm become new experienced employees. Original experienced employees who have finished their second period at the firm retire.
6. Steps from 1 to 5 are repeated.

3.2.2 Workers

All workers are assumed to be identical in that they possess the same skills and utility functions. A worker's utility is assumed to be increasing in wage income, w, and decreasing in the level of work effort, e. The posited utility function is:

$$Utility = w - e^2 \tag{3.3}$$

Let EU be the discounted expected lifetime utility of a new employee employed in period t. Assuming that a worker is paid wages at the end of a period, the next equation shows the discounted lifetime utility of a new employee hired in period t:

$$\begin{aligned} EU_t &= a_{N,t}(p_t) + b_{N,t}(p_t)\, e_{N,t} - e_{N,t}^2 + \\ \delta \int_p &\left\{ \begin{array}{l} \left(\gamma_{t+1}(p) \cdot min\left(\frac{e_{N,t}}{\overline{e}}, 1\right) + 1 - \gamma_{t+1}(p)\right) \\ \times max\left(a_{E,t+1}(p) + b_{E,t+1}(p) e_{E,t+1} - e_{E,t+1}^2, U(p)\right) \\ +\gamma_{t+1}(p) \cdot \left(1 - min\left(\frac{e_{N,t}}{\overline{e}}, 1\right)\right) U(p) \end{array} \right\} dG(p) \end{aligned} \tag{3.4}$$

where $U(p)$ is the utility of a laid off worker when the output price, p, is realized.[4] Workers who have completed the first period are allowed to quit the firm if the utility of the second period, calculated after the output price p_{t+1}, has been realized to be lower than $U(p)$. Given the contract proposed by a firm, each employed worker decides the amount of effort to invest into his/her current job with the aim of maximizing his/her expected lifetime utility.

Note that the worker never has an incentive to supply efforts beyond $\overline{e}$ because offering efforts beyond $\overline{e}$ brings disutility without lowering the probability of layoffs. In order to capture the fact that experienced employees are better protected against permanent layoffs under seniority-based layoffs than under performance-based layoffs, I assume hereafter that the exogenous variable $\overline{e}$ is large enough to ensure that the internal solution for e^*_N is always less than $\overline{e}$, which results in $e^*_N/\overline{e} < 1$. Then, solving the first order conditions, the effort supply functions can be written as:

$$e^*_{N,t} = \frac{1}{2}\Bigg(b_{N,t} + \delta E\left[max\left\{\frac{\gamma_{t+1}(p)}{\overline{e}}\left(a_{E,t+1}(p) + b_{E,t+1}(p)e_{E,t+1} - e^2_{E,t+1} - U(p)\right), 0\right\}\right]\Bigg) \tag{3.5}$$

$$e^*_{E,t+1} = \frac{1}{2}\, b_{E,t+1} \tag{3.6}$$

Equations (3.5) and (3.6) show that workers' current efforts depend on the current piece rate, b, i.e., e^*_N and e^*_E increase in b_N and b_E, respectively. It is, however, more important to note that as long as a firm pays wages that are higher than those necessary to keep workers during the second period, the efforts of a new employee, e_N, will also depend on the wages scheduled to be paid in the next period after performance-based layoffs have occurred, i.e., $\gamma_{t+1}(p) > 0$. In other words, in deciding

[4]In this paper, it is assumed that all workers have the same ability and skill, thus there is no room for the sort of inference that arises in the "career-concern" models, where the worker's outside option in the second period depends on the market's estimate of the worker's ability. Therefore, $U(p)$ does not depend on effort in the first period as it usually does in "career-concern" models.

the optimal effort level, new employees consider the future wages to be paid during periods in which performance-based layoffs occur. For example, if performance-based layoffs occur only in recessions, wages scheduled to be paid during recessions will affect new employees' efforts, while wages scheduled to be paid during booms will not.

This is because given the possibility of being laid off, new employees can receive wages scheduled to be paid in the second period only if they work hard in the first period and avoid layoffs. Then, higher wages of experienced employees during "performance-based layoff periods" encourage new employees to work hard.

In contrast, experienced employees' wages during periods in which performance-based layoffs do not occur, i.e., experienced employees' wages with $\gamma_{t+1}(p) = 0$, do not affect new employees' efforts because new employees will necessarily receive the wages without being laid off, regardless of their current effort levels. Thus, when $\gamma_{t+1}(p) = 0$, only the current piece rate can affect new employees' incentives.

In addition, since experienced workers are assumed to retire after the second period, only the current "piece rate" induces experienced employees' efforts without threat of future layoffs. Equation (3.6) explains this result.

3.2.3 Firms

All firms are assumed to produce the same output, adopt the same technology, and utilize homogeneous labor input. Output is a function of the amount of total efforts provided by new employees and experienced employees. For simplicity, I assume the following linear production function:

$$f(n_N\, e_N, n_E\, e_E) = n_N\, e_N + n_E\, e_E \tag{3.7}$$

where n_N and n_E denote the number of new employees and the number of experienced employees, respectively. In any period, the realization of the firm's profits is:

$$\Pi_t = p_t\, f(n_{N,t}\, e_{N,t}, n_{E,t}\, e_{E,t}) - n_{N,t}(a_{N,t} + b_{N,t}\, e_{N,t}) - n_{E,t}(a_{E,t} + b_{E,t}\, e_{E,t}) \quad (3.8)$$

Experienced employees are recruited from last period's new employees, so the number employed, n_E, is given by:[5]

$$n_{E,t} = \begin{cases} 0 & t = 0 \\ \left(\gamma_t \frac{e_{N,t-1}}{\bar{e}} + 1 - \gamma_t\right)\, n^*_{N,t-1} & \forall\, t \geq 1 \end{cases} \quad (3.9)$$

In period 0, when the firm is established, there is no worker who continue with the firm, and hence, there is no experienced employee at the firm in period 0 by the definition of experienced employees in this paper.

Let C be the exogenous costs associated with laying off a worker. Note that in this model, it is assumed that workers who were recently employed by a firm are not laid off. This simplification makes the "seniority-based layoffs" equivalent to "no layoff of experienced workers." Thus, layoff costs (C) are also equivalently treated as the costs of performance-based layoffs.

The firm offers a new employee hired in period t a contingent contract, $X_t =(n_{N,t}$, $a_{N,t}$, $b_{N,t}$, $\gamma_{t+1}(p_{t+1})$, $a_{E,t+1}(p_{t+1}), b_{E,t+1}(p_{t+1}); \forall p_{t+1})$ to maximize its profits subject to providing workers with a competitively determined utility level. Since a competitive firm with constant returns to scale is considered here, this utility level will be adjusted until the firm makes zero profits in equilibrium. Thus, the firm's problem can be viewed as one of maximizing a worker's utility, subject to its zero profit constraint (Arnott and Stiglitz, 1985).

[5] Here, I adopt the hiring structure assumed in Ioannides and Pissarides (1983).

Then, the firm's problem can be written as follows:

$$
\begin{aligned}
& Max_{(X_t)}\, EU_t \\
& s.t. \begin{cases}
p_t n_{N,t}\, e_{N,t} - n_{N,t}\,(a_{N,t}(p_t) + b_{N,t}(p_t)\, e_{N,t}) \\
+\delta n_{N,t} \int_p \left\{ \begin{array}{l} \left(\gamma_{t+1}(p)\frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p)\right) \\ \times\,(p_{t+1}e_{E,t+1} - a_{E,t+1}(p) - b_{E,t+1}(p)e_{E,t+1}) \\ -C\gamma_{t+1}(p)\left(1 - \frac{e_{N,t}}{\bar{e}}\right) U(p) \end{array} \right\} dG(p) = 0, \\
Equation\,(3.5), \\
Equation\,(3.6), \\
a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} \geq U(p_{t+1}) \quad \forall\, p_{t+1} \\
\gamma_{t+1}(p_{t+1}) \geq 0 \;\; \forall\, p_{t+1}, and \\
\gamma_{t+1}(p_{t+1}) \leq 1 \;\; \forall\, p_{t+1}
\end{cases}
\end{aligned}
\tag{3.10}
$$

Let λ, $\eta(p_{t+1})$, $\mu_{1t} \geq 0$, and $\mu_{2t}(p_{t+1}) \geq 0\,\forall\, p_{t+1}$ be the Lagrangian multiplier on the zero profit constraint, the Kuhn-Tucker multipliers associated with the no-quit constraint for experienced workers, $\gamma_{t+1}(p_{t+1}) \geq 0$ and $\gamma_{t+1}(p_{t+1}) \leq 1\,\forall\, p_{t+1}$, respectively.[6] From the first order condition for $a_{N,t}$, we see that:

$$\lambda = \frac{1}{n_{N,t}} \tag{3.11}$$

Substituting Equation (3.11) and $\partial e_{N,t}/\partial b_{N,t} = 1/2$ into the first conditions for $b_{N,t}$ and $a_{E,t+1}$ yields:

$$\eta(p_{t+1}) = 0 \quad \forall\, p_{t+1} \tag{3.12}$$

This means that the no-quit constraint for experienced employees is not binding. Thus, as long as the firm solves the optimization problem presented in (3.10), the no-quit constraint is automatically satisfied, and workers will never quit even without

[6]The first order conditions are solved in the Appendix.

the no-quit constraint.[7]

From the first conditions for $b_{E,t+1}$ (i.e., equation (3.4A) in the Appendix), we obtain $p_{t+1} = 2e^*_{E,t+1}$. From Equation (3.6), we also know that $2e^*_{E,t+1} = b^*_E(p_{t+1})$. Combining these two equations yields:

$$b^*_E(p_{t+1}) = p_{t+1} \quad \forall\, p_{t+1} \tag{3.13}$$

Equation (3.13) implies that the piece rate for experienced workers is set equal to the output price. Thus, in equilibrium, the marginal revenue of effort from the viewpoint of a firm (p_{t+1}), the marginal disutility of effort ($2e^*_{E,t+1}$), and the marginal cost of

[7] In the equilibrium, the no-quit constraint is automatically satisfied, so placing the no-quit constraint will not change the optimal solution. It will, however, effect, the feasible options for the firm to deviate from the solution. Here, I will show that when we do not place the no-quit constraint, the possibility that $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} = U(p_{t+1})$ holds in equilibrium is excluded, and hence in equilibrium, $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} > U(p_{t+1})$ is always satisfied—that is, workers are always paid higher wages than they would earn as a result of the outside option during the second period. In the simple principal-agent problem, maximizing profits subject to a participation constraint yields the same equilibrium as maximizing worker utility subject to a profit constraint. Therefore, in order to explain the intuition which led to the result, I assume a problem maximizing profits subject to a participation constraint. Suppose that a firm paid just the necessary level of wages to keep workers during the second period. Then, the firm could be better off by increasing $b_{N,t}$ by a small amount and decreasing $a_{E,t+1}(p_{t+1})$ for the realization of p_{t+1} with which performance-based layoffs occur by an amount that satisfies the hiring constraint. This is because if the firm paid just the necessary amount to keep workers during the second period, future wages should not affect the incentives for new employees. Therefore, decreasing $a_{E,t+1}(p_{t+1})$ during the "layoff regime" would not affect new employees' efforts but would allow the firm to increase $b_{N,t}$, and this would enable the firm to induce greater efforts from experienced employees. Furthermore, decreasing $a_{E,t+1}(p_{t+1})$ will cause the utility from the second period to drop below the utility of the outside option. Thus, in response to the decrease in $a_{E,t+1}(p_{t+1})$, workers will quit at the beginning of the second period, which would benefit the firm because the profits during periods in which performance-based layoffs occur were originally scheduled to be negative. The fact that a firm has an incentive to move from the solution derived from the maximization problem contradicts the definition of the equilibrium. Thus, as long as workers are allowed to quit, the firm pays more than necessary to prevent workers from quitting in equilibrium. (When the firm pays more than necessary to prevent workers from quitting, decreasing $a_{E,t+1}(p_{t+1})$ from the optimal level derived in this section will lower the new employees' level of effort, which results in a decrease in profits. Thus, the firm does not have an incentive to deviate from the solutions derived in the Appendix.) In contrast, if we do place the no-quit constraint, both cases, $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} > U(p_{t+1})$ and $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} = U(p_{t+1})$, can be realized in equilibrium. The reason why $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} = U(p_{t+1})$ can also be optimal in this case is that the deviation stated above—decreasing $a_{E,t+1}(p_{t+1})$ slightly and increasing $b_{N,t}$ to make up for the loss of utility—is not possible when we have no-quit constraint, because decreasing $a_{E,t+1}(p_{t+1})$ will break the no-quit constraint, taking $a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1}$ below $U(p_{t+1})$.

effort from the viewpoint of a firm ($b^*_E(p_{t+1})$) are set to be equal. Equation (3.13) also implies that firms base wages less on workers' performance during recessions. This is because, during recessions, firms are discouraged from maintaining workers' incentives at a high level since the value of productivity, p_{t+1}, is lower.

Furthermore, by solving the first order condition with respect to $\gamma_{t+1}(p_{t+1})$, we obtain the following Proposition:

Proposition 3.1

Performance-based layoffs occur when $p^2_{t+1}/4 - U(p_{t+1}) + C < 0$ is satisfied. This implies that:

1. Performance-based layoffs are more likely to occur when layoff costs, C, are low.
2. With a constant U, the above inequality is more likely to be satisfied for lower p_{t+1}.
3. Even if $U(p_{t+1})$ is allowed to vary with p_{t+1}, as long as the utility function of unemployed workers is a concave function of p_{t+1}, i.e., as long as $U'(p_{t+1}) \geq 0$ and $U''(p_{t+1}) \leq 0$ are satisfied, performance-based layoffs do not occur for sufficiently large p_{t+1}.[8]

Proof. See the Appendix.

In words, the firm's optimal layoff decision can be explained as follows: When $p_{t+1} > 2\sqrt{U(p_{t+1}) - C}$ is satisfied, $\gamma^*_{t+1}(p_{t+1}) = 1$ holds, i.e., the optimal layoff rule is the seniority-based layoff, and nobody will be laid off under the assumption that layoffs are implemented only from experienced workers.[9] Once p_{t+1} crosses the threshold value, $2\sqrt{U(p_{t+1}) - C}$, layoffs are implemented, i.e., the "no-layoff regime" will switch to the "layoff regime." When $p_{t+1} < 2\sqrt{U(p_{t+1}) - C}$ is satisfied, the layoff decision is completely performance-based, and an experienced worker will be laid off

[8] In Section 3.6, I will empirically examine with which p_{t+1} this inequality is more likely to be satisfied.

[9] The result of seniority-based layoffs means "no layoff" under this assumption.

with probability, $1 - e_N/\bar{e}$. Within the "layoff regime," a decrease in p_{t+1} will reduce e_N through the reduction in $b_{E,t+1}$, which results in an increase in layoffs.[10]

The next proposition explains how the optimal regular pay and the optimal bonuses are related to the layoff decision stated above.

Proposition 3.2

1. In periods in which performance-based layoffs occur, the regular pay of experienced employees has a higher lower-bound than the regular pay in periods in which performance-based layoffs do not occur.
2. Except for cases in which the output price greatly declines from period t to period $t+1$, workers' pay becomes more performance-based as their careers progress.
3. Experienced employees' bonuses move proportionally to the output price regardless of the possibility of performance-based layoffs.

Proof. See the Appendix.

Proposition 3.2.1 implies that it is more important for the firm to raise experienced employees' piece rate than to raise new employees' piece rate since the former can affect not only experienced employees' current efforts but also new employees' efforts in the preceding period. As a result, the piece rate paid to new employees is set below

[10] Given the linear production technology, each cohort's two-period problem is solved independently of other cohorts. With this setting, the firm's decision to lay off the existing experienced workers in period t is made independently from its decision to hire new junior worker in period t, and hence even during a period in which the firm conducts performance-based layoffs, junior workers can be hired. As Nosal (1990) also assumes, "seniority-based layoffs (layoffs by inverse-seniority)" means here that the possibility for experienced workers to be laid off is zero. It is also possible to pose a restriction that junior workers are not hired during periods in which experienced workers are laid off. However, the object of this paper is to classify periods by the layoff possibility of senior workers, rather than delving into specifics regarding the definition of each layoff type. Thus, the term "seniority-based layoffs" in this paper should be understood as a situation where the layoff probability of senior workers is automatically set at zero (regardless of their performance, market conditions, and the number of new hires).

the output price, while the piece rate paid to experienced employees is set equal to the output price. Thus, if workers encounter the same output price between period t and $t+1$, they necessarily experience an increase in the piece rate as their careers progress. Except for the case in which the output price greatly declines from period t to period $t+1$, workers experience a piece rate increase as their careers progress.

Simple calculations using the first-order condition for b_N show that b_N would become equal to p_t if the possibility of performance-based layoffs were eliminated, i.e., if $\gamma^*_{t+1}(p_{t+1}) = 0$ was satisfied with probability 1. If this could happen, the firm would not have to differentiate b_N from b_E since both piece rates would affect only workers' current efforts without a possibility of layoffs. In this way, when there is no possibility of performance-based layoffs, the two periods become independent of each other, and then the two-period problem becomes a per-period profit maximization problem.

Proposition 3.2.2 is more important because it offers an implication for wage rigidity. Future wages scheduled to be paid in the next period during which layoffs do not occur do not affect new employees' current efforts, because workers will necessarily receive the wages without being laid off in the next period, regardless of their current effort levels. Thus, the firm cannot control new employees' efforts by manipulating their future wages under the "no-layoff regime," which results in a lower, or at least a less downwardly rigid, regular pay.

In contrast, future wages scheduled to be paid in the next period during which layoffs occur affect new employees' current efforts since workers will receive the wages only when they work hard in the current period, and avoid being laid off. Thus, the higher the wages scheduled to be paid in the next period are, the harder the workers will work. This gives the firm an incentive to pay high regular pay under the "layoff regime." As a result, the regular pay of experienced workers has a lower bound, equal to the layoff costs when the output price is low enough that layoffs occur, while the regular pay can be below the layoff costs when the output price is high enough that layoffs do not occur. In this way, in this model, the downward wage rigidity during recessions occurs through a channel in which a low output price increases the probability of performance-based layoffs.

The intuition behind Proposition 3.2.3 is very simple. Firms are discouraged from maintaining workers' current incentives at a higher level due to the lower value of productivity during recessions. Thus, during recessions, firms base their wages less on workers' performance.

3.3 Promotion

So far, we have seen how a firm uses the threat of layoffs to maintain new employees' incentives and how the wage decision yields wage rigidity during periods in which performance-based layoffs occur. However, promotions can be thought of as another important instrument for maintaining new employees' incentives. Thus, in this section, I consider promotions as well as layoffs. Here, it is assumed that a new employee can be promoted at the beginning of the second period and that the promotion probability is an increasing function of the workers' efforts offered in the first period. The promotion probability is expressed as:

$$Pr(promotion) = min\left(1, \frac{e_N}{\overline{e}_P}\right) \tag{3.14}$$

This is very similar to the probability of a worker being retained in the firm, and a new employee can be promoted with probability 1 if he/she offers a level of work effort greater than or equal to $\overline{e}_P$, which is exogenously given. Promotions are generally accompanied by an increase in salary. Thus, let $R(p) \geq 0$ be the additional reward paid to the promoted workers. I assume that R is also contingent on the output price.

In these settings, Equation (3.4) in the previous section becomes:

$$
\begin{aligned}
&EU_t = a_{N,t}(p_t) + b_{N,t}(p_t)\, e_{N,t} - e_{N,t}^2 + \\
&\delta \int_p \left\{ \begin{array}{l}
\left(\gamma_{t+1}(p) \cdot min\left(\frac{e_{N,t}}{\bar{e}}, 1\right) + 1 - \gamma_{t+1}(p)\right) \times \\
max\left(a_{E,t+1}(p) + b_{E,t+1}(p)e_{E,t+1} - e_{E,t+1}^2 + min\left(1, \frac{e_{N,t}}{\bar{e}_P}\right) R(p), U(p)\right) \\
+\gamma_{t+1}(p) \cdot \left(1 - min\left(\frac{e_{N,t}}{\bar{e}}, 1\right)\right) U(p)
\end{array} \right\} dG(p)
\end{aligned}
\tag{3.15}
$$

Introducing promotions changes new employees' choice of the optimal effort, and the first order condition with respect to the new employees' efforts (the internal solution) becomes:

$$
\frac{\partial V_t^E}{\partial e_{N,t}} = -2e_{N,t} + \delta \int_p \left\{ \begin{array}{l}
\gamma_{t+1}(p)\frac{1}{\bar{e}} \left\{ \begin{array}{c} a_{E,t+1}(p) + b_{E,t+1}(p)e_{E,t+1} \\ -e_{E,t+1}^2 + \frac{e_{N,t}}{\bar{e}_P} R(p) - U(p) \end{array} \right\} \\
+ \left(\gamma_{t+1}(p)\frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p)\right) \frac{1}{\bar{e}_P} R(p)
\end{array} \right\} dG(p) = 0
\tag{3.16}
$$

Note that now R can affect new employees' efforts even in the "no-layoff regime," which can be confirmed from the fact that R remains in the equation even for the cases of $\gamma = 0$. This is because a higher R motivates workers even in the "no-layoff regime" as long as the promotion probability is based on a new employee's performance. Note that experienced employees' wages excluding R, i.e., $a_E + b_E\, e_E$, still do not affect new employees' efforts under the "no-layoff regime" because workers will necessarily receive the wages, $a_E + b_E\, e_E$, under the "no-layoff regime" regardless of whether they will be promoted or not.

Then, the firm's problem with promotions becomes:[11]

$$
\begin{aligned}
&Max_{(X_t)}\, EU_t \\
&s.t. \begin{cases}
p_t n_{N,t}\, e_{N,t} - n_{N,t}\left(a_{N,t}(p_t) + b_{N,t}(p_t)\, e_{N,t}\right) + \\
\delta n_{N,t} \int_p \left\{ \begin{array}{l}
\left(\gamma_{t+1}(p)\frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p)\right) \times \\
\left(p_{t+1}e_{E,t+1} - a_{E,t+1}(p) - b_{E,t+1}(p)e_{E,t+1} - \frac{e_{N,t}}{\bar{e}_P}R(p)\right) \\
-C\gamma_{t+1}(p)\left(1 - \frac{e_{N,t}}{\bar{e}}\right) U(p)
\end{array} \right\} dG(p) = 0, \\
Equation\,(3.16), \\
\gamma_{t+1}(p_{t+1}) \geq 0 \;\; \forall\, p_{t+1}, and \\
\gamma_{t+1}(p_{t+1}) \leq 1 \;\; \forall\, p_{t+1}
\end{cases}
\end{aligned}
\tag{3.17}
$$

Then we obtain the following proposition:

Proposition 3.3

Propositions 3.1 and 3.2 are still true even if promotions are allowed.

Proof. See the Appendix.

If we allow for promotions, the regular pay in the "layoff regime" falls by an amount that is proportional to the wage increase when promoted. Thus, there is a trade-off between a high wage increase when an employee is promoted and the high regular pay of experienced employees in the "layoff regime" because both future factors can positively affect new employees' incentives. However, even if promotions are considered, the conditional expectation of regular pay under the "layoff regime" is still set above layoff costs, C, because of the participation constraint for the experienced employees. In contrast, the regular pay under the "no-layoff regime" can be below the layoff costs, as in the previous section. As a result, in periods in which performance-based layoffs occur, the regular pay of experienced employees has a higher lower-bound than the

[11]The participation constraint for experienced employees does not change since it should be satisfied for those who are not promoted under the assumption of $R \geq 0$.

regular pay in periods in which performance-based layoffs do not occur. Therefore, it can be said that the important implications from the basic model do not change even if we allow for promotions.

3.4 Empirical Model

This section presents a strategy to test empirically the implications of the theoretical model. There are two steps in the evaluation of the theoretical model. The first is layoff regression, which tests the implications of Proposition 3.1. The second is wage regression, which tests the implications of Proposition 3.2.

3.4.1 Layoff Regression

Proposition 3.1 states that layoffs occur when the inequality, $p_{t+1}^2/4 - U(p_{t+1}) + C < 0$, is satisfied. This implies that layoffs are more likely to occur when the performance-based layoff costs, C, and the output price, p, are low.[12]

Thus, the following model is estimated:

$$Involuntary\,Leave_{it} = \gamma_0 + \gamma_1 NonUnion_{it} + \gamma_2\,Price_{it} + X_{it}\gamma + u_{it} \tag{3.18}$$

where $Involuntary\,Leave_{it}$ is an indicator function that takes a value of 1 if individual i was laid off or left his/her employer due to the reason on the firm's side during year t. $NonUnion$ is a dummy variable that takes the value of 1 if the individual is a non-union worker at the beginning of year t, and $Price_{it}$ represents the output price the firm (where individual i belongs to) faces at the beginning of year t . X_{it} represents attributes of individual i and its firm at the beginning of year t.

[12] The impact of the output price on layoffs is actually indeterminate because it depends on the functional form of $U(p_{t+1})$. However, as shown in Proposition 3.1, the negative relationship between the output price and the layoff probability is true as long as the functional form of $U(p_{t+1})$ is linear or concave, as usually assumed in economics literatures.

The union status captures the magnitude of performance-based layoff costs, C, because the costs associated with performance-based layoffs of unionized workers are expected to be higher than those of non-union workers.[13]

The expected sign for γ_1 is positive because non-union workers are expected to have a higher layoff probability than unionized workers because of their lower layoff costs. The expected sign for γ_2 is negative.

In the layoff regression, it is expected that restricting the sample to experienced workers places more focus on performance-based layoffs than on seniority-based layoffs. This is because the layoffs that happen among experienced or senior workers are less likely to be "seniority-based layoffs." Thus, Section 3.6 will present the results obtained from the sample that consists only of experienced workers. After confirming that the theoretical implication of Proposition 3.1 is true in the actual data, I will proceed to the wage regressions to test Proposition 3.2.

3.4.2 Wage Regression

Wage regressions are added to test Proposition 3.2. Proposition 3.2.2 states that the regular pay of experienced workers becomes downwardly rigid during periods in which performance-based layoffs occur. Together with Proposition 3.1, which states that performance-based layoffs are likely to occur during recessions, Proposition 3.2.2 has implications for the downward rigidity of regular pay during recessions.

Note that this type of model about the threat of performance-based layoffs is more likely to be true for non-union workers (Bewley, 1999) because there is a possibility

[13] The survey conducted by Abraham and Medoff (1984) shows that in the U.S., approximately 78 % of union groups were covered by written policies that specify seniority as the most important factor to be considered in permanent layoff decisions, while it was just 16 % for nonunion groups. The Bureau of Labor Statistics (BLS) data also reveal that in 1970-71, over 70 % of employee groups under major union agreements in the U.S. were covered by layoff provisions that specify seniority as the most important factor to be considered in permanent layoff decisions. Given these evidences, in the empirical analysis, I will use union-status as a proxy for layoff costs, C, assuming that costs associated with the performance-based layoffs of experienced/senior workers are lower for non-union workers than for unionized workers because of the contract provisions.

that the constraint under which the performance-based layoffs become optimal cannot be satisfied for unionized workers even when the output price is low because of their high "performance-based layoff costs."[14]

In the layoff regression, it was possible to roughly distinguish performance-based layoffs from seniority-based layoffs by checking whether each laid-off worker had been an experienced worker with a sufficiently long tenure with a previous employer. However, in the regular-pay regressions, which use samples of working people, it is not possible to know the types of layoff criteria (whether they are performance-based or seniority-based) assumed in their firms by just restricting the sample to experienced workers. Thus, in the regular-pay regression, I include the interaction term between the "non-union worker" dummy variable and the output price. By including this term, we can test whether wage rigidity is more likely to be observed for non-union workers whose layoffs are more likely to be based on performance.

In contrast, Proposition 3.2.3 predicts that the bonus does not depend on the types of layoffs; hence, distinguishing performance-based layoffs is not necessary for the bonus-pay regression, which means that we are not interested in a coefficient of the interaction term in the bonus regression. Instead, checking whether the coefficient of the output price is significantly positive is important because Proposition 3.2.3 suggests that the bonuses of experienced employees move proportionally to the output price.

[14]According to Bewley (1999), 28 % (86 %) of non-union (unionized) workers are laid off according to inverse seniority, while 57 % (7 %) of non-union (unionized) workers are laid off according to performance. The model explains these widely observed layoff practices as follows: The costs associated with the performance-based layoffs of unionized experienced employees are high enough that $p_{t+1}^2/4 - U(p_{t+1}) + C < 0$ is never satisfied for any p_{t+1}. Thus, the firm never lays off unionized experienced employees because of the high layoff costs. In contrast, with the relatively low costs associated with the layoffs, both cases $p_{t+1}^2/4 - U(p_{t+1}) + C > 0$ and $p_{t+1}^2/4 - U(p_{t+1}) + C < 0$, are possible for non-union workers. Then, the firm implements performance-based layoffs for non-union workers when the output price is low and no workers are laid off when the output price is high.

Thus, the estimated model is:

$$\begin{cases} lna_{it} = \alpha_0 + \alpha_1 NonUnion_{it} + \alpha_2 Price_{it} \\ \qquad\qquad\qquad\qquad +\alpha_3 NonUnion_{it} \cdot Price_{it} + X_{it}\alpha + u_{it} \\ lnb_{it} = \beta_0 + \beta_1 NonUnion_{it} + \beta_2 Price_{it} \\ \qquad\qquad\qquad\qquad +\beta_3 NonUnion_{it} \cdot Price_{it} + X_{it}\beta + e_{it} \end{cases} \tag{3.19}$$

where lna_{it} and lnb_{it} are the logarithm of regular pay and bonus pay for individual i in year t, respectively. Although the theoretical implication regarding the output prices refers to the firm-specific output price, wages are often indexed with respect to the overall price level as well, either formally or informally. In order to capture this widely observed practice, the year dummy variables are also included in X.

The implications from Proposition 3.2 can be written as follows:

$$\begin{cases} \alpha_3 < 0 \\ \beta_2 > 0 \end{cases} \tag{3.20}$$

Proposition 3.2.2 expects α_3 to be significantly negative because, in the comparison between non-union workers and unionized workers, regular pay is expected to fluctuate less in response to changes in the output price for non-union workers who are more likely to face the downward wage rigidity.

In contrast, the sign of β_2 implied by Proposition 3.2.3. is significantly positive. In order to use bonuses as the performance-based component of wages, it is necessary to ensure that bonuses are actually paid for workers' current performance. This is supported by Freeman and Weitzman (1989), Ohashi (1989), and Brunello (1991), all of whom examined the Japanese bonus system. Freeman and Weitzman (1989) state that compensating workers' efforts is one of the main purposes of bonus payment in Japan. Ohashi (1989) also found that bonuses are paid to compensate employees for their intensity of work. Brunello (1991) found no statistically significant correlation between bonuses and employment level in the car, steel, and electric-machinery industries in Japan, which implies that the profit-sharing aspect of the Japanese bonus system is not substantial. Particularly, since the 1990s, the importance of the "pay-for-performance" aspect of the Japanese bonus system has been increasing. Thus, it is thought that it is theoretically more valid to assume performance pay ($w(e) = a + be$) than fixed pay (w) for Japan.

3.4.3 Restricting the Sample to Experienced Workers

The theoretical implications of wages could be true only for experienced workers, because the crux of the model is that under the threat of layoffs, firms maintain workers' incentives by using future wages, which will be paid to "experienced workers." Therefore, in Section 3.6, I will present the empirical results for both all workers and experienced workers, and I will confirm that the expected results on wages can be obtained or become more conspicuous when the sample is restricted to experienced workers.

In contrast, the theoretical implications of layoffs would be true for new employees as well when we relax the assumption that new employees are not laid off. However, as discussed in Section 3.4.1, restricting the sample to experienced workers in the layoff regression helps us focus more on performance-based layoffs, rather than seniority-based layoffs. For this reason, in the layoff regression as well, the result with the sample that consists only of experienced workers will be reported.

3.4.4 Sample Selection Bias

Since the dependent variables in the wage regressions, regular pay and bonuses, are reported only by working people, there might be a selection bias problem. If there is a tendency that workers with specific unobserved characteristics are likely to avoid layoffs during recessions, $Price_{it}$ and the error term might be correlated. For example, if workers with very high ability are less likely to be laid off during recessions, then the lower the output price is, the more likely it is that workers included in the sample will be of high ability.[15] Since such unobservable individual characteristics might be included in the error term, $Price_{it}$ are thought to be correlated with the error term. Then, the coefficients α_2 and β_2 might be negatively biased in the OLS. Furthermore, it is thought that this selection bias is more serious for non-union workers, since their employment is thought to be less protected than that of unionized workers. This means that α_3 might also be biased if we use OLS.

Thus, it is necessary to control for workers' fixed effect by using panel data. In this paper, this problem is expected to be resolved by using the fixed effects model.

3.5 Data

The data set used in this study is Japanese panel data from Keio Household Panel Survey (KHPS), conducted annually by Keio University. The KHPS data are relatively new as the data collection for the survey began in 2004. This survey is conducted in January every year and includes observations randomly chosen from almost all regions and industries in Japan. A key feature of KHPS is that it is the first nationwide follow-up survey in Japan of individuals (4,000 households and 7,000 people) of all ages and both sexes, which also captures information regarding education, employment, income, expenses, health, and family structure. The survey has been designed to enable comparisons with major international panel surveys such as Panel Study of Income Dynamics (PSID) and European Community Household Panel (ECHP).

The details of the KHPS are as follows: Respondents for the first wave are men and women aged between 20 and 69 as of January 31, 2004, from all of Japan. The

[15]This type of selection bias can be thought of as the so-called countercyclical composition bias, although the term is usually used in the context of aggregate time-series data (Stockman (1983), Bils (1985), Solon et al. (1994), and Chang (2000)).

first wave (2004) included 4,005 households, the second wave (2005) included 3,314 out of 4,005 in the first wave, the third wave (2006) 2,887 households, and the fourth wave (2007) 2,643 households. The attrition rate from the first wave to the fourth wave is 34 %. Although 1,419 households were newly added to the samples in 2007, this paper does not utilize these samples.[16]

The industry-level CPIs are used as the measure of $Price_{it}$. Since KHPS records wage data in the previous year, the price indexes used as the independent variable in wage regressions are the annual price indexes for the previous year. Table 3.1.A reports changes in the annual CPI for each industry during the period 2003-2006. The CPI data is obtained from the Consumer Price Index data by the Statistics Bureau at the Ministry of Internal Affairs and Communications (Statistics Bureau, Ministry of Internal Affairs and Communications, 2006). To avoid a potential mismatch between the CPI data and the industry categories used in the survey, the service industry and the "financing and insurance" industry are excluded from the sample. Additionally, because, in many cases, the outputs in the mining industries are used as the intermediate goods traded between firms, the price index data in the Corporate Goods Price Index are used for the mining industry (Bank of Japan, 2006).

Table 3.1.B contains descriptive statistics for the sample of layoff regression. Samples include only individuals who were working as of January in year t. Those represented in the first column are divided into two groups: those who left/changed their employer during year t (shown in the second column) and those who continued with the same employer (shown in the third column). The layoff regression determines how the employment status and the output price at the beginning of year t affect the layoff probability during year t. Thus, Table 3.1.B reports characteristics of the workers (and the employers) at the beginning of year t and the industry-level consumer price index during January. Because the survey asks about a change in employer during

[16]The data can be extended in the future study since this survey will continue in the future.

the previous year, the information about the job change during year t can be obtained from the survey of year $t+1$. In contrast, the characteristics of the workers (and the employers) at the beginning of year t are obtained from the survey of year t. To utilize both, the sample is restricted to individuals who participated in the survey for at least two straight years.

The average age of the participants is 46.04 years and 63 % of the sample did not belong to any labor union. Individuals who left or changed their previous employer during year t represent 5 % of the sample, and approximately 20 % of these people are those who left their previous employers involuntarily. This category includes lay-offs, dismissals, and other firm-related reasons excluding bankruptcy. The criterion, "Years Needed To Be Experienced," represents the average value of the answer to the questions: *How long does it take for workers to feel they are experienced in your field?*, and the average value is calculated for each of industry $\times$ occupation cells. The sample average of the answer to that question is 2.12 years. This measure is used to classify the sample into experienced workers and others.

Table 3.1.C reports the descriptive statistics for the sample of wage regressions. In the wage regressions, the observations are restricted to individuals who earned positive values of regular pay and bonuses.[17] In the estimation sample, the average monthly regular pay from 2003-2006 is 327,301 yen (≈USD 3,273.01), while the average annual bonus for individuals who received a positive bonus amount is 960,475 yen (≈USD 9,604.75). These amounts become higher if we restrict the sample to experienced workers. In order to distinguish "experienced" worker groups from other workers, two criteria are used: tenure of three years and "Years Needed To Be Experienced." According to the general survey on working conditions conducted in 2003, three years of tenure is the most common minimum length of tenure that is necessary for an

[17] The reason why the layoff regression has more observations than the wage regressions is that only 60 % of workers in the sample receive bonuses greater than zero. The samples are unified between the regular-pay and bonus-pay regressions, thus because of the fewer samples of the bonus regression, the sample size for the regular-pay regression is also reduced.

employee to receive retirement allowances in Japan (Ministry of Health, Labor, and Welfare, Government of Japan, 2003a). The minimum tenure required for an employee to receive retirement allowances differs between layoffs and quits. For layoffs, the minimum tenure required to receive a retirement allowance is less than 1 year for 12.2 % of all firms, 1-2.9 years for 30.3 % of all firms, 2-2.9 years for 11.1 % of all firms, 3-3.9 years for 38.7 % of all firms, 4-4.9 years for 1.1 % of all firms, and more than 5 years for 6.2 % of all firms. For quits, I omit the overall distribution of firms over the criteria, but the fraction of the firms that set the minimum requirement of tenure in the 3-3.9 years category is 60 %. Thus, three years can serve as a useful threshold for classifying workers into "new" and "experienced" groups, because it becomes harder for a firm to lay workers off at a tenure of three years, at least in terms of costs.

The total annual bonus payment amounts to three-times the employee's monthly regular pay. From this data, it can be confirmed that most employed workers in performance-pay jobs in Japan receive a substantial portion of their pay in the form of a bonus. This is consistent with other previous studies that use other earning data.[18]

3.6 Empirical Results

3.6.1 Layoff Regression Results

Table 3.2 presents the regression results of the layoff equation. The dependent variable is an indicator function that takes a value of 1 if the individual was laid-off or left his/her employer due to firm-related reasons during year t. Individuals who were working at the beginning of year t are used for the sample. In this regression,

[18]It is well known that the ratio of bonuses to total pay is traditionally high in Japan compared to other countries. Nakamura and Hubler (1998) show that the ratios of bonus to regular pay in the 1980s were 0.317, 0.121, and 0.194 for Japan, Germany, and the U.S., respectively. According to Nakamura and Nakamura (1991), most employed workers in Japan are paid 25 to 33 % of their total earnings in the form of bonus payments.

layoff experience during year t is regressed on the industrial-level CPI during January, the non-union status dummy variable and other characteristics of the workers at the beginning of the year. The other characteristics of the workers include: the male dummy variable, the education dummy variables, potential experience in years (=age−6−education years) and its square, tenure and its square, the marital status dummy variable, the number of children, the logarithm of firm size, year dummy variables, payment-type dummy variables, and industry dummy variables.

The estimates in Table 3.2 show that the effect of the CPI on layoff experience is significantly negative in all columns. In other words, the data show that layoffs are more likely to occur during periods in which the output price is low, which is consistent with Propositions 3.1.2 and 3.1.3. This is also consistent with the findings of Bewley (1999) that the majority of layoffs are implemented in response to reduced demand for labor because of a decline in product demand.

Columns 3 to 6 show the results of the layoff regressions among experienced workers only. As discussed in Section 3.4, it is expected that the layoffs that occurred among experienced workers are likely to be performance-based. The estimates from these columns suggest the significant positive effect of the non-union worker dummy on the layoff experience. This is consistent with the prediction from Proposition 3.1.1, which states that performance-based layoffs are more likely to occur when the costs of performance-based layoffs are low. Thus, it is expected that the theoretical predictions concerning performance-based layoffs are true in the actual data.

3.6.2 Wage Regression Results

In the wage regressions, the same explanatory variables as those included in the layoff regressions are used with two exceptions. First, the interaction between non-union status and the industrial-level CPI is also included in the wage regression. Second, the CPI used in the wage regression is the annual CPI (by industry) instead of the CPI during January.

Table 3.3 reports the results of the regular pay regression. As discussed in Section 3.4.4, the OLS estimates might suffer from the sample selection bias. Thus, the FE estimates, not the OLS estimates, should be taken as the more reliable results. The FE estimates indicate that the coefficient of the interaction term is significantly negative, which supports Proposition 3.2.2, i.e., the regular pay of experienced employees is less responsive to changes in the output price for non-union workers whose layoffs are more likely to be based on performance. Together with the result from the layoff regression, which states that performance-based layoffs are likely to occur during recessions, this implies that downward wage rigidity occurs during recessions in the presence of performance-based layoffs.

Table 3.4 presents the results of the bonus-pay regressions. Although the OLS estimates suggest insignificant coefficients for the CPI, we confirmed in Section 3.4.4 that the coefficient of the CPI could be negatively biased in the OLS regressions due to the sample selection problem. Indeed, if we examine the FE estimates, the effect of the CPI on bonuses becomes significantly positive for experienced workers. Thus, the results of the bonus regressions are consistent with Proposition 3.2.3: *The bonuses of experienced employees move proportionally to the output price.*

These results from the wage regressions are consistent with Freeman and Weitzman (1989), who found that Japanese bonuses are much more procyclical than Japanese base wages.[19]

[19]Freeman and Weitzman (1989) also found that base wages in Japan are negatively related to employment. The results obtained from Table 3.3 are also consistent with this finding of Freeman and Weitzman (1989) because, in the current paper, base wages become downwardly rigid when employment shrinks due to layoffs. Consequently, this implies a negative relationship between base wages and employment. Moreover, Freeman and Weitzman (1989) found that bonuses are positively related to employment. Note that, in the current paper, it is assumed that bonuses are paid for the workers' efforts. As a result, they are set equal to the output price. Thus, although the model setting, itself, assumes the pay-for-performance bonus model, the results do not refute the positive correlation between the amount of bonuses and the firms' profits (or other variables that reflect the firms' profits, such as employment). Therefore, it can be said that the findings here do not contradict the findings in Freeman and Weitzman (1989), which also analyzed the Japanese labor market.

3.7 Conclusion

This paper provides a theoretical and empirical analysis of the effect of performance-based layoffs on wage rigidity in the context of performance pay. Given the findings of Bewley (1999) that performance-based layoffs frequently occur during recessions, especially for non-union workers, I constructed a theoretical model in which firms' layoff decisions can depend on workers' performance, and firms can decide how much weight they put on performance in the layoff decision. Thus, both the rules for layoffs and wages are endogenous in this model. The firms' decisions follow an over-lapping generation model structure, dividing the workers into two types: new employees and experienced employees.

The main results obtained from the theoretical model are as follows: *a) Performance-based layoffs are more likely to occur when the layoff costs and the output price are low. b) Experienced employees' regular pay is likely to be downwardly rigid during periods in which performance-based layoffs occur. c) Bonuses move proportionally to the output price.*

The reason for the result (b) is important: Without the threat of layoffs, wages scheduled to be paid in the next period do not affect current workers' efforts because workers will necessarily receive the wages without being laid off in the next period, regardless of their current effort levels. Thus, the firm cannot control new employees' efforts using their future wages, which results in a lower, at least, a less downwardly rigid regular pay without the the threat of layoffs. In contrast, under the threat of layoffs, workers can receive the next period's wages only when they work hard in the current period and avoid layoffs. Thus, the higher the wages to be paid in the next period are, the harder workers try to avoid layoffs, investing greater efforts. This gives the firm an incentive to raise the future regular pay to maintain workers' current efforts, which results in a downwardly rigid regular pay of experienced employees under the threat of layoffs.

The explanation underlying (c) is very simple. Firms are discouraged from maintaining workers' current incentives at a higher level due to the lower value of productivity during recessions, which results in the bonus that moves proportionally to the output price. This analysis also showed that introducing promotions does not change the results.

The empirical analysis in this paper uses Japanese panel data from the Keio Household Panel Survey (KHPS). All the empirical results confirmed the theoretical implication: performance-based layoffs are likely to occur for non-union workers and during recessions; and regular pay is likely to be downwardly rigid for non-union whose layoffs are more likely to be based on performance. Given that performance-based layoffs are likely to occur during recessions, the result concerning regular pay implies that downward wage rigidity occurs during recessions in the presence of performance-based layoffs. Furthermore, the bonus-pay regression confirmed that firms base wages less on workers' performance during recessions by paying lower bonuses. As a result, wages during recessions become both "downwardly rigid" and "rigid" (inflexible) with respect to performance.

This type of explanation for wage rigidity can be applied to unionized workers whose layoff decisions are likely to be performance-based in countries such as Japan where performance-based pay has been widely employed.

Appendix

The first order conditions are as follows:

$$\frac{\partial \mathcal{L}}{\partial a_{N,t}} = 1 - \lambda n_{N,t} = 0 \tag{3.1A}$$

$$\frac{\partial \mathcal{L}}{\partial b_{N,t}} = e_{N,t} - \lambda n_{N,t} e_{N,t} + \frac{\partial \mathcal{L}}{\partial e_{N,t}} \frac{\partial e_{N,t}}{\partial b_{N,t}} = 0 \tag{3.2A}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial a_{E,t+1}(p_{t+1})} &= \delta \left(\gamma_{t+1}(p_{t+1}) \frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p_{t+1}) \right) g(p_{t+1}) \\ &\quad - \lambda \delta n_{N,t} \left(\gamma_{t+1}(p_{t+1}) \frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p_{t+1}) \right) g(p_{t+1}) \\ &\quad + \frac{\partial \mathcal{L}}{\partial e_{N,t}} \frac{\partial e_{N,t}}{\partial a_{E,t+1}} + \eta(p_{t+1}) = 0 \qquad \forall\, p_{t+1} \end{aligned} \tag{3.3A}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial b_{E,t+1}(p_{t+1})} &= \delta \left\{ \begin{array}{l} \left(\gamma_{t+1}(p_{t+1}) \frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p_{t+1}) \right) \times \\ \left(e_{E,t+1} + (b_{E,t+1} - 2e_{E,t+1}) \frac{\partial e_{E,t+1}}{\partial b_{E,t+1}} \right) \end{array} \right\} g(p_{t+1}) \\ &\quad + \lambda \delta n_{N,t} \left(\gamma_{t+1}(p_{t+1}) \frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p_{t+1}) \right) \\ &\quad \times \left((p_{t+1} - b_{E,t+1}) \frac{\partial e_{E,t+1}}{\partial b_{E,t+1}} - e_{E,t+1} \right) g(p_{t+1}) + \frac{\partial \mathcal{L}}{\partial e_{N,t}} \frac{\partial e_{N,t}}{\partial b_{E,t+1}} \\ &\quad + \eta(p_{t+1}) \left(e_{E,t+1} + (b_{E,t+1} - 2e_{E,t+1}) \frac{\partial e_{E,t+1}}{\partial b_{E,t+1}} \right) = 0 \quad \forall\, p_{t+1} \end{aligned} \tag{3.4A}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial \gamma_{t+1}(p_{t+1})} &= \delta \left(\frac{e_{N,t}}{\bar{e}} - 1 \right) \left(a_{E,t+1} + b_{E,t+1} e_{E,t+1} - e_{E,t+1}^2 - U(p) \right) g(p_{t+1}) \\ &\quad + \lambda n_{N,t} \delta \left(\frac{e_{N,t}}{\bar{e}} - 1 \right) \left(p_{t+1} e_{E,t+1} - a_{E,t+1} - b_{E,t+1} e_{E,t+1} + C \right) g(p_{t+1}) \\ &\quad + \frac{\partial \mathcal{L}}{\partial e_{N,t}} \frac{\partial e_{N,t}}{\partial \gamma_{t+1}} + \mu_{1t} - \mu_{2t} = 0 \qquad \forall\, p_{t+1} \end{aligned} \tag{3.5A}$$

$$\begin{aligned} \frac{\partial \mathcal{L}}{\partial \lambda} &= p_t n_{N,t}\, e_{N,t} - n_{N,t} \left(a_{N,t}(p_t) + b_{N,t}(p_t)\, e_{N,t} \right) \\ &\quad + \delta n_{N,t} \int_p \left\{ \begin{array}{l} \left(\gamma_{t+1}(p) \frac{e_{N,t}}{\bar{e}} + 1 - \gamma_{t+1}(p) \right) \\ \times \left(p_{t+1} e_{E,t+1} - a_{E,t+1}(p) - b_{E,t+1}(p) e_{E,t+1} \right) \\ - C \gamma_{t+1}(p) \left(1 - \frac{e_{N,t}}{\bar{e}} \right) U(p) \end{array} \right\} dG(p) = 0 \end{aligned} \tag{3.6A}$$

$$\begin{aligned} \eta(p_{t+1}) \frac{\partial \mathcal{L}}{\partial \eta(p_{t+1})} &= \eta(p_{t+1}) \\ &\quad \times \left\{ a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1}) e_{E,t+1} - e_{E,t+1}^2 - U(p_{t+1}) \right\} = 0 \quad \forall\, p_{t+1} \end{aligned} \tag{3.7A}$$

Proof of Proposition 3.1

Given that $\partial e_{N,t}/\partial b_{N,t} = 1/2 > 0$ and given Equation (3.11), from Equation (3.2A), the following condition holds:

$$\frac{\partial \mathcal{L}}{\partial e_{N,t}} = 0 \tag{3.8A}$$

By substituting Equations (3.11), (3.13), and (3.8A) into Equation (3.5A), Equation (3.5A) can be written as follows:

$$\begin{aligned}\frac{\partial \mathcal{L}}{\partial \gamma_{t+1}(p_{t+1})} = \delta\left(\frac{e_{N,t}}{\overline{e}} - 1\right)\left(\frac{1}{4}p_{t+1}^2 - U(p_{t+1}) + C\right) g(p_{t+1})\\ +\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1}) = 0\end{aligned} \tag{3.9A}$$

Note that $\mu_{1t}(p_{t+1}) \geq 0$ and $\mu_{2t}(p_{t+1}) \geq 0$ are the Kuhn-Tucker multipliers associated with $\gamma_{t+1}(p_{t+1}) \geq 0$ and $\gamma_{t+1}(p_{t+1}) \leq 1$, respectively. Since $(e_{N,t}/\overline{e} - 1)$ in the first term in (3.8A) is always negative by assumption, the sign of $(\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1}))$ depends on the sign of the term, $(p_{t+1}^2/4 - U(p_{t+1}) + C)$, and can be categorized as:

$$\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1}) \begin{cases} < 0 \;\; if \;\; \frac{1}{4}p_{t+1}^2 - U(p_{t+1}) + C < 0 \\ = 0 \;\; if \;\; \frac{1}{4}p_{t+1}^2 - U(p_{t+1}) + C = 0 \\ > 0 \;\; if \;\; \frac{1}{4}p_{t+1}^2 - U(p_{t+1}) + C > 0 \end{cases} \tag{3.10A}$$

Since both $\mu_{1t}(p_{t+1})$ and $\mu_{2t}(p_{t+1})$ are always non-negative, the sign of $\{\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1})\}$ can be negative only when $\mu_{1t}(p_{t+1}) = 0$ and $\mu_{2t}(p_{t+1}) > 0$.[20] This corresponds to the case of $\gamma_{t+1}^*(p_{t+1}) = 1$, i.e., the case in which the layoff decision is completely performance-based.

The second case, $\{\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1})\} = 0$, can be true only when $\mu_{1t}(p_{t+1}) =$

[20] Note that these two multipliers cannot be positive at the same time since $\gamma_{t+1}(p_{t+1}) = 0$ and $\gamma_{t+1}(p_{t+1}) = 1$ cannot hold at the same time.

$\mu_{2t}(p_{t+1}) = 0$ holds. In this case, $\gamma^*_{t+1}(p_{t+1})$ can take any number between 0 and 1 when p_{t+1} satisfies the condition, $p^2_{t+1}/4 - U(p_{t+1}) + C = 0$, as an equality. With continuous p_{t+1}, this happens with zero probability. The last case, $(\mu_{1t}(p_{t+1}) - \mu_{2t}(p_{t+1}))>0$, can be satisfied only when $\mu_{1t}(p_{t+1}) > 0$ and $\mu_{2t}(p_{t+1}) = 0$, which corresponds to the case of $\gamma^*_{t+1}(p_{t+1}) = 0$, i.e., the case in which a firm's layoff decision is completely seniority-based.

Mathematically, the optimal layoff decision stated above can be written as:

$$\begin{cases} \gamma^*_{t+1}(p_{t+1}) = 1 \;\; if \;\; T(p_{t+1}, C) < 0 \\ \gamma^*_{t+1}(p_{t+1}) \in [0,1] \;\; if \;\; T(p_{t+1}, C) = 0 \\ \gamma^*_{t+1}(p_{t+1}) = 0 \;\; if \;\; T(p_{t+1}, C) > 0 \end{cases} \tag{3.11A}$$

where $T(p_{t+1}, C) \equiv p^2_{t+1}/4 - U(p_{t+1}) + C$

It is obvious that the function, $T(p_{t+1}, C)$, is increasing in C, and thus the case of $\gamma^*_{t+1}(p_{t+1}) = 1$ is more likely to occur when C is low. In other words, firms are more likely to implement performance-based layoffs when layoff costs are low.

If we assume that the utility of the laid off workers, U, does not depend on p_{t+1}, the value of $\partial T/\partial p_{t+1} = p_{t+1}/2 - U'(p_{t+1})$ is always positive since, in that case, the second term, $U'(p_{t+1})$, is zero. On the other hand, if we allow U to increase in p_{t+1}, the sign of $\partial T/\partial p_{t+1} = p_{t+1}/2 - U'(p_{t+1})$ becomes indeterminate and varies according to the functional form of $U(p_{t+1})$. However, even if we allow U to depend on p_{t+1}, we can still say that $T(p_{t+1}, C) > 0$ is satisfied for sufficiently large p_{t+1} as long as the law of diminishing marginal utility is satisfied for $U(p_{t+1})$ with respect to p_{t+1}, i.e., $U'(p_{t+1}) \geq 0$ and $U''(p_{t+1}) \leq 0$.

For ease of understanding, I will divide the function $T(p_{t+1}, C)$ into two parts: $p^2_{t+1}/4 + C$ and $U(p_{t+1})$. Figure A.3.1 depicts $p^2_{t+1}/4 + C$ and $U(p_{t+1})$ as functions of p_{t+1}. As shown in (3.9A), when the curve, $p^2_{t+1}/4 + C$, is above $U(p_{t+1})$, $\gamma^*_{t+1}(p_{t+1}) = 0$

is satisfied, and performance-based layoffs do not occur. Depending on the value of C, the curve, $p_{t+1}^2/4+C$, can move up or down in Figure A.3.1. It is also possible to move $U(p_{t+1})$ up and down by changing the intercept of $U(p_{t+1})$. However, as long as the law of diminishing marginal utility is satisfied for $U(p_{t+1})$ with respect to p_{t+1}, i.e., $U'(p_{t+1})>0$, and $U''(p_{t+1})<0$, $p_{t+1}^2/4+C$ finally moves above the $U(p_{t+1})$ curve as p_{t+1} increases since $p_{t+1}^2/4+C$ is a convex function of p_{t+1}. In other words, the value of $T(p_{t+1},C)$ goes to infinity as p_{t+1} increases except when $U(p_{t+1})$ is a convex function of p_{t+1}. Thus, the firm sets $\gamma_{t+1}^*(p_{t+1})=0$, i.e., the firm stops implementing layoffs for sufficiently large p_{t+1} as long as $U(p_{t+1})$ is concave. Q.E.D.

Proof of Proposition 3.2.1

Given Equations (3.11) and (3.13), Equation (3.8A) can be rewritten as:

$$\frac{\partial\mathcal{L}}{\partial e_{N,t}}=p_t-2e_{N,t}+\delta\frac{1}{\bar{e}}\int_p\gamma_{t+1}(p)\left(\frac{1}{4}p_{t+1}^2-U(p_{t+1})+C\right)dG(p)=0 \tag{3.12A}$$

As shown in (3.11A), $\gamma_{t+1}^*(p_{t+1})=1$ holds for any p_{t+1} that satisfies $p_{t+1}<2\sqrt{U(p_{t+1})-C}$ and $\gamma_{t+1}^*(p_{t+1})=0$ holds for any p_{t+1} that satisfies $p_{t+1}>2\sqrt{U(p_{t+1})-C}$. Therefore, the inside of the integral in (3.12A) becomes zero for the realizations of $p_{t+1}>2\sqrt{U(p_{t+1})-C}$, and the second term in (3.12A) remains only for the case in which $p_{t+1}<2\sqrt{U(p_{t+1})-C}$:

$$\frac{\partial\mathcal{L}}{\partial e_{N,t}}=p_t-2e_{N,t}+\delta\frac{1}{\bar{e}}\int_{p<2\sqrt{U(p)-C}}\left(\frac{1}{4}p_{t+1}^2-U(p_{t+1})+C\right)dG(p)=0 \tag{3.13A}$$

By the condition of $p_{t+1}<2\sqrt{U(p_{t+1})-C}$, the second term in (3.13A) is negative. Thus, the following inequality can be obtained:

$$p_t>2e_{Nt}^* \tag{3.14A}$$

By Equation (3.5), we also have:

$$e^*_{Nt} \geq \frac{1}{2} b^*_{Nt} \tag{3.15A}$$

Combining Inequalities (3.14A) and (3.15A) yields:

$$p_t > b^*_{Nt} \tag{3.16A}$$

Q.E.D.

Proof of Proposition 3.2.2

Substituting Equation (3.5) into Equation (3.13A) yields:

$$p_t - b_{N,t} + \delta \int_{p<2\sqrt{U(p)-C}} \left(\tfrac{1}{\bar{e}} \left(-a_{E,t+1}(p) + C\right)\right) dG(p) = 0 \tag{3.17A}$$

Then, the conditional mean of the regular pay in period t+1, conditional on being in the "layoff regime," i.e., $p_{t+1} < 2\sqrt{U(p_{t+1}) - C}$, is expressed as:

$$E[a^*_{E,t+1}(p)/p < 2\sqrt{U(p) - C}] = \tfrac{\bar{e}}{\delta} \left(p_t - b^*_{N,t}\right) \frac{1}{Pr\left(p<2\sqrt{U(p)-C}\right)} + C > C \tag{3.18A}$$

The first term in Equation (3.18A) shows that in order to maintain new employees' efforts, the lower the new employees' piece rate is, the higher the experienced employees' regular pay should be. There is a trade-off between a high regular pay for experienced employees and a high piece rate for new employees.[21] However, by

[21] Note that the optimal regular pay is presented in a form of conditional mean. This is because new employees do not care about the exact amount of the regular pay they will receive for each realization of p_{t+1}. However, they still care about their future regular pay. The only aspects they care about are how much pay they will receive on average during periods in which layoffs occur and during periods in which layoffs do not occur. Thus, as long as the conditional mean of a_E in each of the two layoff regimes satisfies the first-order conditions of the firm's maximization problem, any $a_E(p_{t+1})$ for $p_{t+1} \in [p^-, p^+]$ can be optimal for a firm. However, while hiring, the firm needs to specify not only the conditional mean of a_E in each of the two layoff regimes but also $a_E(p_{t+1})$ for all realizations of p_{t+1} because, otherwise, the firm would pay zero regular pay for any realization

Inequality (3.16A), we know that b^*_{Nt} is set below the output price; i.e., the first term in Equation (3.18A) is positive. Thus, the conditional mean of a_E, conditional on being in the "layoff regime," has a lower bound of C>0.

As confirmed in Equation (3.12), the no-quit constraint does not bind, which means that the following condition is satisfied even without the condition:

$$a_{E,t+1}(p_{t+1}) + b_{E,t+1}(p_{t+1})e_{E,t+1} - e^2_{E,t+1} \geq U(p_{t+1}) \tag{3.19A}$$

If we substitute Equation (3.13) into Equation (3.19A), Equation (3.19A) becomes:

$$a_{E,t+1}(p_{t+1}) \geq U(p_{t+1}) - \frac{1}{4}p^2_{t+1} \tag{3.20A}$$

We know from Equation (3.11A) that $U(p_{t+1}) - p^2_{t+1}/4 > C$ holds when $\gamma^*_{t+1}(p_{t+1}) = 1$, and thus, we can derive the following condition:

$$a_{E,t+1}(p_{t+1}) \geq U(p_{t+1}) - \frac{1}{4}p^2_{t+1} > C \quad \forall p_{t+1} \ \ s.t. \ p_{t+1} < 2\sqrt{U(p_{t+1}) - C} \tag{3.21A}$$

Thus, the lower bound of C exists not only for the conditional mean of the regular pay in period $t+1$, conditional on being in the "layoff regime," but also for every realization of $a_{E,t+1}(p_{t+1})$.

We have confirmed that by the first order conditions w.r.t. wages, Equation (3.8A) is satisfied. Since $\frac{\partial EU}{\partial e_{N,t}} = 0$ holds because of the effort supply function of new

of p_{t+1} in the second period, insisting that it originally set $a_E(p_{t+1})$=0 for that realization of p_{t+1}, and that the conditional means still satisfy the employment contract. In order for the firms to refrain from such moral hazards and gain the trust of their workers, $a_E(p_{t+1})$ for every realization of p_{t+1} should be precisely proposed when firms hire new workers and this should not be changed afterwards.

employees, Equation(3.8A) becomes:

$$\frac{\partial \mathcal{L}}{\partial e_{N,t}} = \frac{\partial \Pi}{\partial e_{N,t}} = 0 \tag{3.8A'}$$

This is nothing but Equation (3.17A). Thus, in the equilibrium, new employees' efforts are exerted until any further increase in $e_{N,t}$ will not change the total profits. As long as the probability of $p_{t+1} < 2\sqrt{U(p_{t+1}) - C}$ is positive, the optimal $b_{N,t}$ is set less than p_t (Proposition 3.2.1) because $a_{E,t+1}$ also plays the role of inducing new employee efforts. This means that an increase in $e_{N,t}$ will increase the profits from the first period because $(p_t - b_{N,t})e_{N,t} > 0$ holds. Then, to attain Equation (3.8A'), an increase in $e_{N,t}$ should decrease profits from the second period. Since an increase in $e_{N,t}$ will increase the probability of a worker being retained in the firm, $a_{E,t+1}$ should be high enough that the increase in the probability (of a worker being retained in the firm) will yield more costs ($a_{E,t+1}$) than benefits (C) from the view point of the firm. This makes $a_{E,t+1}$ in the "layoff regime" high enough that the no-quit constraint holds.

In contrast, regular pay under the "no-layoff regime" is free from constraint (3.12A) since it does not affect the new employees' efforts by virtue of the fact that workers can gain regular pay regardless of their efforts. (Under the "no-layoff regime," the second term in (3.12A) becomes zero, and hence $a_{E,t+1}$ is no longer included in the equation.) Consequently, regular pay under the "no-layoff regime" is free from the constraint of maintaining the workers' incentives, and it can be low provided that constraint (3.6A) is satisfied.[22]

Then, using the zero-profit constraint, the conditional mean of regular pay under

[22]Note that, under the "no-layoff regime", $a_{E,t+1}$ does not affect $e_{N,t}$, which means both the objective function and the Kuhn-Tucker constraint are linear with respect to $a_{E,t+1}$ under the "no-layoff" regime. Thus, we cannot conclude anything from Equation (3.12) for $a_{E,t+1}$ under the "no-layoff regime", and it is also possible for the firm to provide $a_{E,t+1}$ that induce workers to quit.

the "no-layoff regime" is expressed as:

$$E[a^*_{E,t+1}(p)/p > 2\sqrt{U(p) - C}] = -a^*_{N,t}\frac{1}{\delta Pr(p > 2\sqrt{U(p) - C})} - C\frac{Pr(p < 2\sqrt{U(p) - C})}{Pr(p > 2\sqrt{U(p) - C})} \tag{3.22A}$$

As is shown in Equation (3.22A), in the absence of performance-based layoffs, the firm can choose any combination of a^*_N and a^*_E as long as it satisfies the zero-profit constraint. This is because there is no difference between a^*_E in the "no-layoff regime" and a^*_N from the viewpoint of a firm in that neither of them affects the workers' incentives. Q.E.D.

Proof of Proposition 3.2.3

From Equation (3.13), we know that the piece rate paid to each unit of effort, b_E, is set equal to the output price. Thus, the amount of bonus is expressed as $b_{E,t+1}e_{E,t+1} = p^2_{t+1}/2$. Q.E.D.

Proof of Proposition 3.3

Since the first order condition with respect to $a_{N,t}$ does not change, even if promotions are allowed, Equation (3.11) still holds. The expression of (3.2A) before expanding the term, $\partial\mathcal{L}/\partial e_{N,t}$, does not change.

Given that $\partial e_{N,t}/\partial b_{N,t} = 1/2 > 0$ and given Equation (3.11), Equation (3.8A) still holds from Equation (3.2A). Then, conditions (3.3A) and (3.4A) do not change even if promotions are allowed. Given Equation (3.11) and (3.8A), the result in Equation (3.13) still holds: the piece rate is set equal to the output price.

When promotions are allowed, Equation (3.5A) becomes:

$$
\begin{aligned}
&\frac{\partial \mathcal{L}}{\partial \gamma_{t+1}(p_{t+1})} = \delta \left(\frac{e_{N,t}}{\bar{e}} - 1\right) \\
&\times \left(a_{E,t+1} + b_{E,t+1}e_{E,t+1} - e_{E,t+1}^2 + \frac{e_{N,t}}{\bar{e}_P}R(p) - U(p)\right) g(p_{t+1}) \\
&+\lambda n_{N,t}\delta \left(\frac{e_{N,t}}{\bar{e}} - 1\right) \left(p_{t+1}e_{E,t+1} - a_{E,t+1} - b_{E,t+1}e_{E,t+1} - \frac{e_{N,t}}{\bar{e}_P}R(p) + C\right) g(p_{t+1}) \\
&+\frac{\partial \mathcal{L}}{\partial e_{N,t}}\frac{\partial e_{N,t}}{\partial \gamma_{t+1}} + \mu_{1t} - \mu_{2t} = 0 \qquad \forall\, p_{t+1}
\end{aligned} \tag{3.23A}
$$

However, R in the first term and the second term in (3.23A) cancel out each other, and the two conditions, (3.11) and (3.8A) still hold. Thus, the optimal layoff decision (3.11A) is preserved, and Proposition 3.1 is true even when promotions are allowed.

Given Equation (3.11), $\partial\mathcal{L}/\partial e_{N,t}$ can be written as:

$$
p_t - b_{N,t} + \delta \int_p \left(\begin{array}{l} \gamma_{t+1}(p)\frac{1}{\bar{e}}\left(-a_{E,t+1}(p) + C - 2\frac{e_{N,t}}{\bar{e}_P}R(p)\right) \\ -\left(1-\gamma_{t+1}(p)\right)\frac{1}{\bar{e}_P}R(p) \end{array} \right) dG(p) = 0 \tag{3.24A}
$$

Considering that $\gamma_{t+1}^*(p_{t+1}) = 1$ holds for any p_{t+1} that satisfies $p_{t+1} < 2\sqrt{U(p_{t+1}) - C}$, and that $\gamma_{t+1}^*(p_{t+1}) = 0$ holds for any p_{t+1} that satisfies $p_{t+1} > 2\sqrt{U(p_{t+1}) - C}$, the conditional mean of the compensation in the second period other than the performance-based component conditional on $p_{t+1} < 2\sqrt{U(p_{t+1}) - C}$, becomes:

$$
\begin{aligned}
&E[a_{E,t+1}^*(p) + \tfrac{e_{N,t}}{\bar{e}_P}R(p)/p < 2\sqrt{U(p)-C}] \\
&\quad = \tfrac{\bar{e}}{\delta}\left(p_t - b_{N,t}^*\right)\frac{1}{Pr(p<2\sqrt{U(p)-C})} + C - \frac{\bar{e}}{e_{N,t}}\frac{1}{Pr(p<2\sqrt{U(p)-C})}\overline{R} \\
&where\ \overline{R} = Pr(p<2\sqrt{U(p)-C})\tfrac{e_{N,t}}{\bar{e}}\tfrac{e_{N,t}}{\bar{e}_P}E[R/p<2\sqrt{U(p)-C}] \\
&\qquad\qquad +Pr(p>2\sqrt{U(p)-C})\tfrac{e_{N,t}}{\bar{e}_P}E[R/p>2\sqrt{U(p)-C}]
\end{aligned} \tag{3.25A}
$$

The only difference between (3.18A) and (3.25A) is the last term in (3.25A); the last term in (3.25A) is added to the conditional mean of regular pay once promotions are allowed. $\overline{R}$ is the mean amount of the wage-increase (R) estimated at the beginning of the first period. As shown in (3.25A), if we allow for promotions, the regular pay in the "layoff regime" falls by an amount that is proportional to the wage increase when promoted. Thus, there is a trade-off between a high wage increase when an employee is promoted and the high regular pay of experienced employees in the "layoff regime" because both future factors can positively affect new employees' incentives. However, even if promotions are considered, the regular pay under the "layoff regime" is still set above layoff costs, C, i.e., Inequality (3.21A) still holds, because the no-quit condition does not bind, and the firm pays higher than necessary to keep workers in the second period.

In contrast, the regular pay in the "no-layoff regime" is still free from the constraint of maintaining the workers' incentives and can be low provided the zero-profit constraint is satisfied.

By substituting (3.25A) into the zero-profit constraint, the conditional mean of regular pay under the "no-layoff regime" is expressed as:

$$\begin{aligned}
&E[a^*_{E,t+1}(p) + \tfrac{e_{N,t}}{\bar{e}_P} R(p)/p > 2\sqrt{U(p) - C}] \\
&= -a^*_{N,t} \tfrac{1}{\delta Pr(p>2\sqrt{U(p)-C})} - C \tfrac{Pr(p<2\sqrt{U(p)-C})}{Pr(p>2\sqrt{U(p)-C})} + \tfrac{1}{Pr(p>2\sqrt{U(p)-C})} \overline{R} \\
&where\ \overline{R} = Pr(p < 2\sqrt{U(p) - C}) \tfrac{e_{N,t}}{\bar{e}} \tfrac{e_{N,t}}{\bar{e}_P} E[R/p < 2\sqrt{U(p) - C}] \\
&\qquad\qquad + Pr(p > 2\sqrt{U(p) - C}) \tfrac{e_{N,t}}{\bar{e}_P} E[R/p > 2\sqrt{U(p) - C}]
\end{aligned} \tag{3.26A}$$

The only difference between (3.22A) and (3.26A) is that (3.26A) has the third term. The mean pay in the "no-layoff regime", again, does not have a positive lower bound. Thus, even when promotion is allowed for, the main results of the previous section are preserved. Q.E.D.

Table 3.1.A: Changes in Consumer Price Index by Industry

Industry \ Year	2003	2004	2005	2006
Agriculture	99.6	103.2	100.0	102.0
Communications	108.6	107.3	100.0	96.4
Construction	100.6	100.3	100.0	100.3
Electricity, Gas, Heat Supply and Water	99.1	99.2	100.0	103.6
Fishery and Forestry	99.6	103.2	100.0	102.0
Manufacturing	100.2	100.3	100.0	100.6
Mining	99.5	99.6	100.0	101.0
Real Estate	100.3	100.1	100.0	100.0
Transport	100.0	100.0	100.0	99.6
Wholesale and Retail Trade	100.2	100.3	100.0	100.6

Note: The base year of the CPI is 2005, and the CPI for 2005 is fixed at 100 for each industry. The CPI data is obtained from the Consumer Price Index data by the Statistics Bureau at the Ministry of Internal Affairs and Communications (Statistics Bureau, Ministry of Internal Affairs and Communications, 2006). To avoid a potential mismatch between the CPI data and the industry categories used in the survey, the service industry and the "financing and insurance" industry are excluded from the sample. The price index for the mining industry is obtained from the price index data in the Corporate Goods Price Index (Bank of Japan, 2006).

Table 3.1.B: Descriptive Statistics for Estimation Sample of Layoff Regression

	All	Left Employer during Year t	No Change in Employer
CPI by industry	100.18	100.09	100.18
	(0.96)	(0.89)	(0.96)
Age	46.04	41.70	46.26
	(11.53)	(11.83)	(11.47)
Tenure	13.82	6.14	14.20
	(11.81)	(8.90)	(11.81)
Number of Children	1.49	1.41	1.49
	(1.09)	(1.14)	(1.09)
Years Needed To Be Experienced	2.12	1.98	2.13
	(1.19)	(1.00)	(1.20)
Firm Size	202.42	166.88	204.16
	(206.62)	(185.12)	(207.48)
Male	0.69	0.59	0.70
Married	0.85	0.75	0.86
Non-union Worker	0.63	0.79	0.62
Change/Left Employer during Year t	0.05	1.00	0.00
lnvoluntary Leave	0.01	0.17	0.00
Education Dummy Variables			
Junior High School	0.10	0.06	0.10
High School	0.53	0.58	0.53
Junior College	0.10	0.12	0.10
University	0.25	0.22	0.25
Graduate School	0.01	0.01	0.01
Payment-Type Dummy Variables			
Paid Monthly	0.73	0.57	0.74
Paid Weekly	0.00	0.00	0.00
Paid Daily	0.07	0.08	0.07
Paid Hourly	0.15	0.29	0.14
Paid Yearly	0.05	0.06	0.05
Observations	4638	217	4421

Note: Column 1 contains the descriptive statistics for individuals who were working as of January in year t. Those represented in Column 1 are divided into two groups: those who left/changed their employer during year t (Column 2) and those who continued with the same employer (Column 3). "Involuntary Leave" is a dummy variable that takes a value of 1 if the individual was laid off or left his/her employer due to the reason on the firm's side during year t. The CPI is the industry-level consumer price index during January of each year. The CPI for January 2005 is fixed at 100 for each industry. "Years Needed To Be Experienced" represents the average length of years it takes for workers to feel they are experienced in their field, and the average value is calculated for each of industry $\times$ occupation cells. To avoid a potential mismatch between the CPI data and the industry categories used in the survey, the service industry and the "financing and insurance" industry are excluded from the sample.

Table 3.1.C: Descriptive Statistics for the Estimation Sample of Wage Regressions

	All Workers	Experienced Workers	
		Tenure> X Years	Tenure>Three Years
Regular Pay (100 yen≈ 1 USD / month)	3273.01	3348.66	3442.42
	(1776.72)	(1788.03)	(1802.16)
Bonus (100 yen≈ 1 USD / year)	9604.75	10068.96	10514.09
	(9277.52)	(9413.45)	(9525.70)
CPI by industry	100.45	100.43	100.44
	(1.51)	(1.44)	(1.47)
Age	45.03	45.61	45.98
	(10.68)	(10.46)	(10.29)
Tenure	15.13	16.60	17.49
	(11.28)	(10.86)	(10.55)
Number of Children	1.49	1.52	1.55
	(1.06)	(1.05)	(1.05)
Years Needed To Be Experienced (X)	2.19	2.15	2.24
	(1.21)	(1.20)	(1.23)
Firm Size	273.13	277.86	281.22
	(204.70)	(204.06)	(204.33)
Male	0.77	0.78	0.80
Married	0.86	0.88	0.89
Non-union Worker	0.62	0.61	0.60
Education Dummy Variables			
Junior High School	0.07	0.07	0.07
High School	0.53	0.53	0.52
Junior College	0.10	0.10	0.10
University	0.28	0.28	0.29
Graduate School	0.02	0.02	0.02
Payment-Type Dummy Variables			
Paid Monthly	0.86	0.87	0.88
Paid Weekly	0.00	0.00	0.00
Paid Daily	0.04	0.04	0.04
Paid Hourly	0.08	0.07	0.06
Paid Yearly	0.02	0.02	0.02
Observations	3748	3389	3194

Note: Observations are restricted to those who earned positive values of regular pay and bonuses. For "experienced" worker groups, two criteria are used: those whose length of tenure is longer than X years and those whose tenure is longer than three years, where X represents the average length of years it takes for workers to feel they are experienced in their field (by industry). The CPI is the annual industry-level consumer price index in Table 3.1.A.

Table 3.2: Layoff Regression (Probit)

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	All Workers		Experienced Workers			
Involuntary Leave=1			Tenure> X Years		Tenure>Three Years	
Price	-0.0052	-0.0055	-0.0046	-0.0043	-0.0037	-0.0030
	(0.0014)	(0.0012)	(0.0020)	(0.0017)	(0.0018)	(0.0016)
Non-union Worker	0.0011	0.0011	0.0015	0.0015	0.0019	0.0019
	(0.0009)	(0.0009)	(0.0007)	(0.0007)	(0.0006)	(0.0006)
Male	0.0003	0.0003	0.00002	0.00002	0.0003	0.0003
	(0.0023)	(0.0023)	(0.0015)	(0.0016)	(0.0016)	(0.0016)
High School	-0.0045	-0.0045	-0.0035	-0.0034	-0.0058	-0.0057
	(0.0012)	(0.0011)	(0.0015)	(0.0015)	(0.0020)	(0.0020)
Junior College	-0.0017	-0.0017	-0.0013	-0.0013	-0.0022	-0.0021
	(0.0009)	(0.0009)	(0.0007)	(0.0007)	(0.0006)	(0.0006)
University	-0.0020	-0.0020	-0.0011	-0.0011	-0.0025	-0.0025
	(0.0009)	(0.0009)	(0.0008)	(0.0008)	(0.0009)	(0.0009)
Graduate School	0.0041	0.0040	0.0109	0.0110	0.0015	0.0016
	(0.0097)	(0.0094)	(0.0157)	(0.0157)	(0.0056)	(0.0057)
Experience	0.0004	0.0004	0.0004	0.0004	0.0006	0.0006
	(0.0002)	(0.0002)	(0.0001)	(0.0001)	(0.0003)	(0.0003)
Experience^2/100	-0.0006	-0.0006	-0.0007	-0.0007	-0.0010	-0.0010
	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0004)	(0.0004)
Tenure	-0.0007	-0.0007	-0.0003	-0.0003	-0.0003	-0.0003
	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Tenure^2/100	0.0012	0.0012	0.0006	0.0006	0.0006	0.0006
	(0.0003)	(0.0003)	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Married	-0.0062	-0.0061	-0.0056	-0.0056	-0.0033	-0.0033
	(0.0039)	(0.0039)	(0.0044)	(0.0045)	(0.0047)	(0.0046)
Number of Children	0.0007	0.0007	0.0007	0.0006	0.0001	0.0001
	(0.0005)	(0.0005)	(0.0006)	(0.0006)	(0.0006)	(0.0006)
ln(Firm Size)	-0.0004	-0.0004	0.0001	0.0001	0.0002	0.0001
	(0.0003)	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
Year Dummies	No	Yes	No	Yes	No	Yes
Payment-Type Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry Dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.1264	0.1267	0.1694	0.1697	0.1683	0.1707
Observations	4638	4638	3686	3686	3410	3410

Note: Marginal effects evaluated at the sample mean are reported. Standard errors, clustered at industry levels, are in parentheses under the regression coefficients. For "experienced" worker groups, two criteria are used: those whose length of tenure is longer than X years and those whose tenure is longer than three years, where X represents the average length of years it takes for workers to feel they are experienced in their field (by industry$\times$occupation). The dependent variable is an indicator function that takes a value of 1 if the individual was laid off or left his/her employer due to the reason on the firm's side during year t. Individuals who were working at the beginning of year t are used for the sample. All explanatory variables represent information reported at the beginning of year t. The reference group for the education dummy variables is "Junior High School." The CPI is the industry-level consumer price index during January of each year. The base year of the CPI is 2005, and the CPI for January 2005 is fixed at 100 for each industry.

Table 3.3: Regular Pay Regression

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	All Workers		Experienced Workers			
ln (Regular Pay)			Tenure> X Years		Tenure>Three Years	
	OLS	FE	OLS	FE	OLS	FE
Price	0.004	0.004	0.004	0.004	0.002	0.004
	(0.002)	(0.001)	(0.003)	(0.001)	(0.003)	(0.001)
Non-union Worker	0.448	0.676	0.430	0.496	0.162	0.585
	(0.408)	(0.208)	(0.436)	(0.140)	(0.432)	(0.118)
Price·Non-union Worker	-0.004	-0.007	-0.004	-0.005	-0.001	-0.006
	(0.004)	(0.002)	(0.004)	(0.001)	(0.004)	(0.001)
Male	0.578	-	0.581	-	0.581	-
	(0.029)		(0.025)		(0.025)	
High School	0.032	-	0.025	-	0.041	-
	(0.022)		(0.024)		(0.023)	
Junior College	0.093	-	0.097	-	0.122	-
	(0.031)		(0.033)		(0.037)	
University	0.201	-	0.184	-	0.204	-
	(0.033)		(0.039)		(0.039)	
Graduate School	0.387	-	0.384		0.405	-
	(0.082)		(0.078)		(0.072)	
Experience	0.030	0.066	0.030	0.062	0.029	0.060
	(0.003)	(0.013)	(0.002)	(0.018)	(0.003)	(0.016)
Experience^2/100	-0.058	-0.062	-0.061	-0.054	-0.057	-0.051
	(0.008)	(0.014)	(0.006)	(0.024)	(0.005)	(0.021)
Tenure	0.018	0.012	0.019	-0.073	0.020	-0.070
	(0.003)	(0.007)	(0.003)	(0.175)	(0.004)	(0.174)
Tenure^2/100	-0.013	-0.035	-0.014	-0.058	-0.017	-0.090
	(0.005)	(0.023)	(0.007)	(0.046)	(0.007)	(0.036)
Married	-0.016	-0.024	-0.007	-0.029	-0.000	-0.031
	(0.030)	(0.024)	(0.030)	(0.021)	(0.026)	(0.022)
Number of Children	0.011	0.014	0.010	0.009	0.014	0.009
	(0.024)	(0.025)	(0.025)	(0.024)	(0.024)	(0.024)
ln(Firm Size)	0.035	0.017	0.033	0.020	0.033	0.027
	(0.006)	(0.007)	(0.005)	(0.009)	(0.005)	(0.018)
Year Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Payment-Type Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry Dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.641	0.055	0.640	0.054	0.624	0.058
Observations	3751	3751	3389	3389	3197	3197

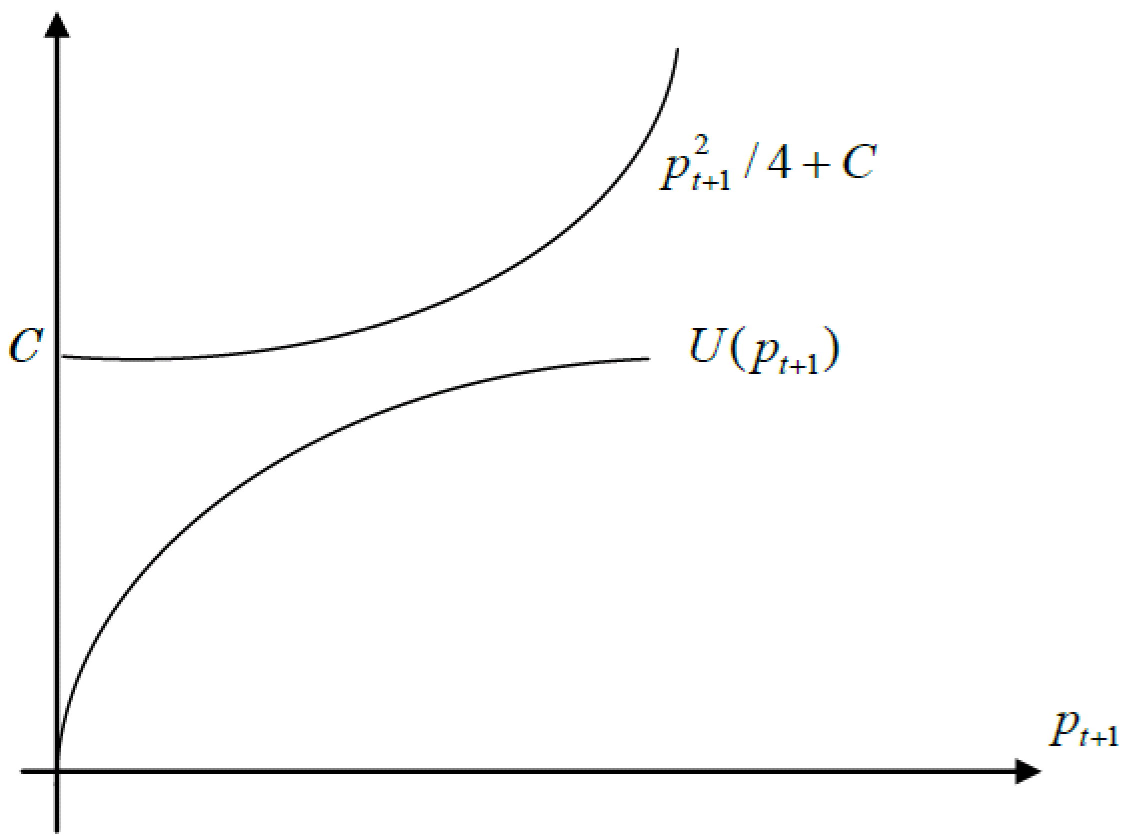
Note: Standard errors, clustered at industry levels, are in parentheses under the regression coefficients. For "experienced" worker groups, two criteria are used: those whose length of tenure is longer than X years and those whose tenure is longer than three years, where X represents the average length of years it takes for workers to feel they are experienced in their field (by industry×occupation). The reference group for the education dummy variables is "Junior High School." The CPI is the annual industry-level consumer price index in Table 3.1.A.

Table 3.4: Bonus Pay Regression

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable:	All Workers		Experienced Workers			
ln (Bonus Pay)			Tenure> X Years		Tenure>Three Years	
	OLS	FE	OLS	FE	OLS	FE
Price	-0.005	-0.002	0.001	0.007	0.0002	0.006
	(0.007)	(0.003)	(0.007)	(0.003)	(0.007)	(0.003)
Non-union Worker	-3.281	-0.806	-1.879	-0.084	-1.693	-0.083
	(0.772)	(0.221)	(0.885)	(0.146)	(0.946)	(0.182)
Price·Non-union Worker	0.032	0.008	0.018	0.0004	0.016	0.0004
	(0.008)	(0.002)	(0.009)	(0.001)	(0.009)	(0.002)
Male	0.712	-	0.700		0.720	
	(0.079)		(0.080)	-	(0.088)	-
High School	0.160	-	0.142		0.133	
	(0.038)		(0.041)	-	(0.049)	-
Junior College	0.260	-	0.268		0.265	
	(0.069)		(0.060)	-	(0.057)	-
University	0.458	-	0.426		0.416	
	(0.046)		(0.047)	-	(0.050)	-
Graduate School	0.773	-	0.726	-	0.751	-
	(0.172)		(0.176)		(0.167)	
Experience	0.017	0.086	0.017	0.104	0.025	0.118
	(0.010)	(0.011)	(0.010)	(0.011)	(0.014)	(0.012)
Experience^2/100	-0.057	-0.129	-0.062	-0.151	-0.073	-0.170
	(0.015)	(0.031)	(0.016)	(0.047)	(0.021)	(0.043)
Tenure	0.058	0.075	0.048	0.146	0.039	0.125
	(0.008)	(0.023)	(0.007)	(0.089)	(0.015)	(0.090)
Tenure^2/100	-0.065	-0.174	-0.037	0.107	-0.022	0.187
	(0.009)	(0.072)	(0.009)	(0.060)	(0.022)	(0.086)
Married	0.109	0.140	0.117	0.096	0.126	0.100
	(0.044)	(0.120)	(0.050)	(0.131)	(0.035)	(0.134)
Number of Children	-0.011	-0.025	-0.012	-0.009	-0.010	-0.013
	(0.038)	(0.043)	(0.041)	(0.045)	(0.042)	(0.046)
ln(Firm Size)	0.118	0.024	0.123	0.041	0.127	0.041
	(0.036)	(0.018)	(0.037)	(0.013)	(0.037)	(0.012)
Year Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Payment-Type Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Industry Dummies	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.562	0.049	0.562	0.029	0.549	0.026
Observations	3751	3751	3389	3389	3197	3197

Note: Standard errors, clustered at industry levels, are in parentheses under the regression coefficients. For "experienced" worker groups, two criteria are used: those whose length of tenure is longer than X years and those whose tenure is longer than three years, where X represents the average length of years it takes for workers to feel they are experienced in their field (by industry$\times$occupation). The reference group for the education dummy variables is "Junior High School." The CPI is the annual industry-level consumer price index in Table 3.1.A.

Figure A.3.1: Graphical Explanation of $p_{t+1}^2/4 + C$ and $U(p_{t+1})$



Note: Figure A.3.1 depicts $p_{t+1}^2/4+C$ and $U(p_{t+1})$ as functions of p_{t+1}. When the curve, $p_{t+1}^2/4+C$, is above $U(p_{t+1})$, $\gamma_{t+1}^*(p_{t+1}) = 0$ is satisfied, and performance-based layoffs do not occur. Even if $U(p_{t+1})$ is allowed to vary with p_{t+1}, as long as the utility function of unemployed workers is a concave function of p_{t+1}, i.e., as long as $U'(p_{t+1}) \geq 0$ and $U''(p_{t+1}) \leq 0$ are satisfied, performance-based layoffs do not occur for sufficiently large p_{t+1}.

CHAPTER IV

Labor-Market Responses to Standard Hours Reduction: Evidence from Japan (with Daiji Kawaguchi and Hisahiro Naito)

4.1 Introduction

Dismal labor-market outcomes after the financial shock of 2008 revived an argument that favored work-sharing policy. Proponents of work-sharing policy argue that the exogenous reduction of hours worked creates employment through the substitution of hours for employment. Previous studies in labor economics, however, cast doubt on work-sharing policy, pointing out two major challenges. The first challenge is the increase of hourly wage as a consequence of the reduction of standard hours. The second challenge is the effect of a production scale-down induced by increased labor cost, which may overwhelm the substitution effect from hours to employment.

Regarding the first challenge, in many historical cases, the reduction of standard hours reduced hours worked but did not reduce weekly or monthly wages of existing workers enough to keep the hourly wage constant. For example, when labor unions and employer federations agreed to reduce standard hours from 40 to 35 per week in Germany during period from the mid 1980s to the mid 1990s, they typically agreed not to reduce weekly or monthly pay (Hunt, 1999).[1] When the French government reduced the standard hours from 40 to 39, the law prohibited employers from reducing weekly or monthly pay (Crepon and Kramarz (2002)). Other studies also report an increase in the hourly rate associated with the reduction of standard hours in Europe, where pay is largely determined through union-employer bargaining and the reduction of weekly or monthly wages faces strong resistance.[2] Skuterud (2007) offers unique evidence by analyzing the case of Quebec, Canada, which reduced the standard hours of non-unionized workers from 44 to 40 between 1997 and 2000 without prohibiting employers from cutting weekly or monthly pay. Even in this setting, he did not find an increase of employment in Quebec compared with neighboring states. Since his evidence is not completely free from interregional equilibrium effects, an evaluation of a national policy that covers a wide range of workers in a flexible wage-setting economy is still needed to fully assess the validity of work-sharing policy.

This paper contributes to the literature by assessing the impacts of Japan's exogenous reduction of standard hours from 48 in 1988 to 40 in 1997 on labor-market outcomes. Studying Japan's case is valuable because wages are determined flexibly in a decentralized employer-employee negotiation. Moreover, the Japanese wage is known to respond flexibly to exogenous shocks, because bonus payments comprise around 25 to 33 percent of total annual compensation (Steinberg and Nakane (2011)). Firm-specific human capital resulting from labor-market friction is often pointed out as an important feature of the Japanese labor market (Hashimoto and Raisian (1985)).

[1] This agreement was what a labor union finally won after a protracted strike, and it is true that there was a great compromise on the side of employers in reaching the agreement. Instead of accepting the demands of the labor union, employers obtained greater flexibility in the use of standard hours. For example, after the agreement, standard hours no longer had to be spread evenly over each day of the week, and could vary from week to week as long as they averaged to the agreed-upon number over a certain number of months. Additionally, standard hours were allowed to vary across employees as well as long as they averaged to the agreed-upon number (Hunt, 1999).

[2] Nymoen (1989) for Norway, Pencavel and Holmlund (1988) and Skans (2004) for Sweden, Estevao and Sa (2008) and Chemin and Wasmer (2009) for France, Raposo and van Ours (2008) for Portugal. Apart from the work-sharing argument, Hamermesh and Trejo (2000) analyze the effect of introducing an 8-hours-per-day limit for men in 1980 on top of a pre-existing 40 hours-per-week limit in California.

Consequently, both employers and employees have strong incentives to protect their relation, and hence future returns on their investments in specific human capital, by adjusting wage compensation, including bonus payments (Hashimoto (1979)). Although Japanese labor law prohibits employers from cutting compensation without workers' consent unless employers experience hardship (and courts apply strict criteria to recognize the hardship), many unions/workers' representatives and employers agree to contracts so that the bonus payment depends on the firm's performance. In this institutional setting, the bonus payment may well be adjusted to the reduction of standard hours so that the hourly wage is kept constant. Several studies examine the effect of scheduled hours, which are voluntarily determined by individual labor contracts, on hours worked in the Japanese context using macro data (Brunello (1989), Niimi (1998) and Saito and Tachibanaki (2002)), but none of these studies explicitly examine the impacts of the reduction of standard hours. Moreover, isolating the policy impact from macroeconomic business cycle effects was difficult in these studies because they relied on macroeconomic data.

Figure 4.1 presents the time series of the average of standard hours weighted by workers' composition by industries and establishment sizes for mining, construction, manufacturing, and public-utility industries. The figure also includes the average hours worked of workers in these four industries. Both of the two series decline over the 1989-2002 period in the same way, but this does not necessarily imply the policy effects on average hours worked because the decrease in hours worked also coincided with a recession of the Japanese economy from 1991.[3] With this concurrent time trend, it is difficult to draw a conclusion about the policy impact on the observed reduction in hours worked. To overcome this difficulty, we exploit the heterogeneous timings of the reduction in standard hours across industries and establishment sizes to examine how the reduction in standard hours affected labor-market outcomes of

[3]See Motonishi and Yoshikawa (1999), Hayashi and Prescott (2002) and Kobayashi and Inaba (2006) for descriptions and explanations of the prolonged stagnation.

interest.

To preview the results in this paper, an analysis based on the Basic Survey of Wage Structure shows that standard hours had a moderate effect on hours worked. The most-preferred specification indicates that about 31 percent of the reduction in hours worked between 1989 and 2002 is explained by the reduction in standard hours. In addition, the reduction in hours worked was not followed by a decrease in either monthly wages or annual bonuses of existing workers. As a result, overall labor cost per hour increased, and we do not see any evidence of job creation. These results are similar to those in Europe, where pay is largely determined through union-employer bargaining and the reduction of weekly or monthly wages faces strong resistance. Thus, we can conclude that the idea of work-sharing did not work even in Japan, where the wage setting is flexible because of decentralized wage negotiation and bonus payments.

The rest of this paper unfolds as follows. Section 4.2 introduces the legal setting of Japanese standard hours and its historical change. Section 4.3 describes the data set. Section 4.4 explains the estimation strategy. Section 4.5 evaluates the effects of standard hours on hours worked, wage, and employment. Section 4.6 adds an estimation of hourly wage rate and interprets the results based on existing theory. The last section concludes and derives implications.

4.2 Legal Institution on Standard Hours in Japan

The Labor Standard Act (LSA) in Japan prohibits employers from employing workers exceeding daily and weekly standard hours. The current standard hours are 40 hours per week and 8 hours per day (LSA Section 32). Employers can set the work hours beyond these legal limits only under an agreement with a workers' representative that represents the majority of employees (LSA Section 36). Overtime under this agreement must be compensated by a wage premium of at least 25 percent

(LSA Section 37).[4][5]

By the early 1980s, Japan had accumulated a current-account surplus, and the Japanese government faced diplomatic pressure to reduce it. In response, the Japanese government attempted to reduce hours worked by reducing standard hours.[6] Standard hours were historically set at 48 per week. In December 1985, a study group organized by the Ministry of Labor published a report that suggested 45 hours per week and 8 hours per day as new standard hours. Following the issuance of this report, the Central Labor Standard Commission, which consisted of representatives of the public interest, employers, and employees, started to discuss revising the standard hours. After numerous meetings among the three parties, the commission finally recommended that standard hours be set at 46 per week for the time being, followed by 44, and eventually reduced to 40. In addition, the recommendation requested a lenient moratorium period for small- and medium-sized firms, reflecting the opinion of employers' representatives. The gradualism reflected in this recommendation represents a compromise between representatives of employers and employees. In accordance with this recommendation, the LSA was revised in 1987 and implemented from April 1, 1988.

This revision set the standard hours at 46 per week immediately after April 1, 1988, but this change did not affect many workers, because establishments with 100 or fewer employees in many industries were exempted from this change. An addi-

[4] Regardless of whether it is regularly scheduled overtime or not, any work hour that exceeds the standard hours must be compensated by the overtime wage premium. In the case that the scheduled work hours set by each firm is below the legal standard hours, hours that exceed the firm-specific scheduled hours but are still within the legal standard hours do not need to be compensated by the overtime premium.

[5] See (Sugeno, 2002, Chapter 3 Section 5) for an overview of the Japanese legal system on standard hours. (Hamaguchi, 2004, Chapter 12 Section 2) describes the legal process of reducing the standard hours between 1987 and 1997. Umezaki (2008) also describes the process of the LSA revision by interviewing two government officials who played a central role in the 1987 LSA revision.

[6] Unlike other counties such as France and Germany, work-sharing was not the direct goal of reducing standard hours in Japan. Instead, diplomatic pressures, mainly from the US, which suffers a significant current account deficit against Japan, drove the Japanese government to start this attempt.

tional revision of the Cabinet order in December 1990 further reduced it to 44 from April 1, 1991. The LSA was further revised in 1993 to implement 40 hours per week from April 1994. In this reduction process, particular moratorium periods were given to industries with long work hours and smaller establishment sizes, mainly because Japan's Chamber of Commerce and Industry, which represents the interests of owners of small- and medium-scale businesses, actively lobbied for lenient moratorium periods for small businesses. The resulting transition is summarized in Table 4.1. The moratorium periods ended by March 1997, by which time the standard hours had become 40 hours per week uniformly across industries and establishment sizes with only a few exceptions.[7]

Another important legislation that may have affected hours worked was the Temporary Act for the Promotion of Work Hour Reduction (*Jitan Sokushin Ho*) enacted on September 1, 1992. The Act was effective for five years and offered three legal provisions. First, the Act promoted the formation of an establishment-level committee, consisting of employer and employee representatives, for work-hour reduction. The agreement in the committee that is submitted to a local labor standard office becomes a legally binding contract as a usual labor agreement (*Roushi Kyoyaku*). Second, the Act provided an exemption from the application of the Antitrust Act for employers to collude in an effort to reduce the hours worked. This exemption was provided because of the perception that establishments had *exceedingly* competed for quick service, and this excessive competition had resulted in long hours worked. Third, the law established a subsidy up to 3 million yen (about 30 thousand US dollars) to promote labor-saving capital investment for establishments with up to 300 regular employees. The subsidy was paid only to establishments that implemented projects that aimed to achieve the goal of reducing hours worked, in order to compensate for two-thirds of their expenses. Hence, the subsidy was not allowed to be used for

[7]Exceptions apply to small establishments that usually hire less than 10 employees in commerce and service industries.

compensating workers directly (Ida (2005)).

Since the subsidy policy covers only establishments that hire 300 or fewer employees, the policy impact estimated in this study partly includes that of the Temporary Act for the Promotion of Work Hour Reduction. Standard hours are longer for establishments that hire 300 or fewer employees, and therefore, our estimates for the impact of standard hours on hours worked is attenuated by the Temporary Act for the Promotion of Work Hour Reduction.

4.3 Data

The data set used in this study is micro data from the Basic Survey on Wage Structure (BSWS), compiled annually by the Japanese government between 1989 and 2002. This survey is conducted in June of every year and includes workers randomly sampled from almost all regions and industries in Japan except for agriculture. The data set includes about 1.5 million workers randomly sampled from 60-70 thousand establishments for each year. The survey covers all establishments with 10 or more employees in both private and public sectors and establishments in private sectors with 5 to 9 employees.

Establishments are randomly sampled in proportion to the size of prefectures, industries, and the number of employees from the Establishment and Enterprise Census (EEC hereafter), which lists all establishments in Japan.[8] Human-resource managers of randomly selected establishments are asked to extract workers' information from their payroll records according to the given instructions for random sampling, includ-

[8]This list is revised every 2-5 years. Of the years relevant to our analysis, the lists were revised in 1986, 1991, 1994, 1996, 1999, and 2001. The establishments included in the BSWS 1989-1992 are randomly sampled from the 1986 EEC list, the 1993-1995 establishments are from the 1991 list, the 1996-1997 establishments are from the 1994 list, the 1998-2001 establishments are from the 1996 list, and the 2002-2003 establishments are from the 1999 list. Although about one half of the establishments are sampled in two consecutive years when establishments are sampled from the same list, only about 1/10 of establishments are sampled at the time of the list revision. The establishments included in the sample are significantly replaced at the times of the list revision: 1993, 1996, 1998, and 2002.

ing the sampling probability, which depends on the establishment's size and industry.

The main variables used in this analysis are scheduled hours worked and overtime hours worked, scheduled monthly payment, overtime payment in June, bonus payment of the previous calendar year[9], full-/part-time status, type of worker (production worker or non-production worker), as well as the firm's attributes, including the number of regular workers (*Joyo Rodo Sha*)[10], and industry.

Managerial workers are exempt from standard hours (LSA Section 41),[11] but the legal definition of managerial workers is rather vague and the exact identification of exempted workers from the statistical occupation code is impossible. Thus, we keep those workers who are potentially exempted workers in the sample. To examine the effect of this sample selection, we check the robustness of our results by restricting the sample to production workers in a subsequent analysis.

The timing for the reduction of standard hours differs by industry and establishment size. Thus we identify the number of standard hours that applies to a specific worker in a specific year by the information of industry and establishment size that a specific worker works for. Although this is a key step in the empirical analysis, the legal industry classification in the Labor Standard Act and the statistical industry classification in the BSWS do not overlap for many industries. Thus, to avoid the biases that might arise from mismatching a worker and a number of standard hours, we restrict our analysis to workers who are in manufacturing, mining, and construction industries, because legal and statistical industrial codes perfectly overlap for these three industries.

[9]The definition of each term is in the Appendix.

[10]Those workers who satisfy one of the following three criteria are classified as regular workers: 1. on contracts that do not clearly specify a contractual time period, 2. on contracts that last more than a month, or 3. on contracts that last less than a month, but on which the workers worked 18 or more days in the last two months. This classification includes part-time workers if one of the criteria above is satisfied.

[11]Agricultural workers and workers in charge of continuous watch services (i.e., security guard, school or apartment janitor) are also exempted. These workers are not included in the analysis sample in this study.

Discontinuity in standard hours serves as a "key" in our estimation, and we estimate the local average policy effect across discontinuity points at establishment sizes of 100 and 300.[12] In this estimation, we eliminate establishments with less than 30 workers from the sample because this contributes to reducing a computational burden but does not affect our estimation; there is no discontinuity in standard hours at the establishment size of 30 among the industries in our sample.

As was confirmed in Section 4.2, the fraction of establishments that were affected by the policy change had been very restricted prior to 1991, and hence the revision in 1991 was actually the first "effective" implementation of the standard-hour reduction. For this reason, we treat the years 1989 and 1990 as the control group among the available sample years. The subsequent reduction of standard hours took place in April 1991, April 1993, April 1994, and April 1995, before a complete transition to 40 hours per week in April 1997. Therefore, the time between 1991 and 1996 is treated as transition period. In addition, the "after" policy-change period, 1997-2002, is included to allow for lagged policy effects.

Table 4.2 reports descriptive statistics of the analysis sample. The average of standard hours, weighted by the number of employees, decreased from 46.86 in 1989

[12]It is true that there is also a discontinuity point at establishment size of 9, but in this paper, we do not consider the policy effect across this discontinuity point. The reasons for this are as follows: first, the discontinuity in the standard hours at establishment size of 9 occurs only from April 1994 to March 1995, and hence it can be thought that omitting the policy effect among these small establishments would not make a great difference in the results. Second, in order to estimate the policy effect around the discontinuity point at establishment size of 9, it is necessary to include establishments with 9 or fewer employees in the sample. However, there is a technical concern about including establishments with 9 or fewer employees in the sample due to the fact that a different sampling rule is applied to this small establishment size. The sampling rule for an establishment with 10 employees or more and for an establishment with 5-9 employees is as follows: *1) Establishments with 10 employees or more: Private establishments and establishments of public corporations under Specified Independent Administrative Institutions or the Local Public Corporation Labour Relations Law. 2) Establishments with 5-9 employees: Private establishments that belong to firms with 5-9 employees.* According to these sampling rules, firm size is also included as one of the sampling conditions for establishments with 5-9 employees, while it is not included for establishments with 10 or more employees. As a result, establishments with 5-9 employees are more likely to be characterized by information unique to small-scale firms than those of similar sizes but with the number of employees just above 10, because these establishments with 10 or more employees in the sample could belong to large firms as well, while it is not possible for those with 5-9 employees, according to the sampling rules.

to 40.00 in 2002. Weekly hours worked also decreased from 44.58 in 1989 to 40.54 in 2002. Although the hours worked decreased during the period, monthly scheduled cash earnings and annual bonuses increased. (For the earnings data in Table 4.2, nominal values are reported.) Most of the increase in monthly earnings and annual bonuses occurred during the first three years of the sample period in response to the rapid increase in price of goods during Japan's bubble economy that lasted from the late 1980s to the beginning of 1991. (The increase in hourly wage rate reflects both effects of the reduced work hours during the sample period as well as the increase in monthly earnings stated above.)

Figure 4.2 compares the distribution of hours worked (= scheduled hours worked + overtime hours worked) in 1989 and 2002. The 1989 distribution of hours worked was bimodal, peaking at 42 and 48 hours, but the 2002 distribution became almost unimodal, peaking at 40 hours. The 2002 distribution shifted to the left and was less dispersed than the 1989 distribution. This evidence reflects the fact that the variation in the initial standard hours among industries and firm sizes vanished after 1997 because of the uniform standard hours of 40 hours per week.

4.4 Estimating Local Average Policy Effects

4.4.1 Identification

Challenges in estimating the average policy effect are two-fold. First, the reduction of standard hours coincided with the macroeconomic downturn starting from 1991. Second, the policy change was partly endogenous, because the moratorium period was given to establishments with 100 or fewer employees, or establishments with 300 or fewer employees in industries characterized by long hours worked.[13] As confirmed in

[13]Since our sample does not include employees in establishments with 30 employees or less, only the establishment sizes of 101 and 301 serve as thresholds in the analyses. The threshold applied differs across industries and years (e.g., 101 for the manufacturing sector and 301 for the construction sector in 1989). If establishments with 30 or fewer employees were included, establishments with 10

Table 4.1, the standard hours are heterogeneous across industries and establishment sizes. LSA categorizes establishments into five groups, based on the number of people they employ, as follows: 1-9 employees, 10-30 employees, 31-100 employees, 101-300 employees, and 301+ employees. During the moratorium period, establishments that fall into a larger-size category were likely to have shorter standard hours. To isolate the policy effect from these potentially contaminating factors, we estimate the following econometric model:

$$y_{jts} = \alpha \bar{h}_{jts} + f(s) + \beta T_t + D_j \delta + D_j T_t \zeta + T_t g(s) + D_j h(s) + u_{jts} \tag{4.1}$$

where $f(s) = \phi_1 s + \phi_2 s^2 + ... + \phi_p s^p$, $g(s) = \gamma_1 s + \gamma_2 s^2 + ... + \gamma_p s^p$, and $h(s) = \eta_1 s + \eta_2 s^2 + ... + \eta_p s^p$. y_{jts} is the outcome of interest (hours worked, wages, and employment) of industry j (33 categories), year t (1989,...,2002), and establishment size s (31,32,...). Standard hours are denoted as $\bar{h}_{jts}$ defined by industry, establishment size, and year, as tabulated in Table 4.1. T_t is a time trend term corresponding to year t. For the industry dummy variables, 33 industry categories from mining, construction, manufacturing, as well as electricity, gas, heat supply, and water industries are used. D_j is a vector that contains industry dummy variables, and δ, ζ, η_1, η_2,..., and η_p are coefficient vectors that contain coefficients estimated for industry j (2,..,33).

The continuous function $f(s)$ captures the effect of establishment size on hours worked. It is expected that the dummy variables D_j capture the heterogeneity across industries, and T_t captures macroeconomic shocks. The effect of industry-specific and size-specific macroeconomic shocks are expected to be captured by $D_j T_t$ and $T_t g(s)$, respectively, and size-industry effects are also controlled for by $D_j h(s)$. The coefficients of $f(s)$, $g(s)$, and $h(s)$ are allowed to differ below and above the discontinuity point.

Including these variables circumvents the problem arising from the policy endo-

employees would also serve as a threshold.

geneity resulting from the political process of policy implementation because they control for the heterogeneity in hours worked that differ by industry, year, and size. The policy effect, α, could differ across discontinuity points, industry, and year, but the specification assumes a single parameter. We intentionally employ this single-parameter specification because we are interested in the local, average policy effect across discontinuity points, industries over years to learn the policy's aggregate impact on the Japanese economy.

The dependent variable, y_{jst}, is the cell mean of individual observations y_{ijst} within an industry j, establishment size s in year t. We collapsed individual data in this way to avoid overstating the precision for $\hat{\alpha}$ that is caused by a potential misspecification of the regression function. This aggregation captures the effect of standard hours on labor-market outcomes, because standard hours are defined over industry and size of establishment. Since the precisions of the cell means are different, reflecting the number of individual observations for each cell mean, all of the regression analyses use the number of observations in cell j divided by the average number of observations in a cell as the weight (Lee and Card (2008)).

The standard hours $\bar{h}_{jst}$ discontinuously change with respect to establishment size at 100 and 300 employees within a specific industry in a specific year. Thus the above regression model estimates the local, average treatment effect of standard hours across industries over years via the regression discontinuity design.

4.4.2 Choice of Regression Function

The continuous functions $f(s)$, $g(s)$, and $h(s)$ with polynomial order p are estimated in Equation (4.1) using observations around discontinuity such that $s \in [100 - b, 100 + b]$ or $s \in [300 - b, 300 + b]$, where b is the bandwidth of the local linear regression. We choose the combination of the polynomial order p and the bandwidth b that minimizes the mean squared error of the prediction by cross-validation (Lee and Lemieux (2010)). The cross-validation value is calculated as follows, using all observations in the neighborhood of the discontinuity points.[14]

$$CV(p, b) = (1/N) \sum_{sjt} (y_{sjt} - \hat{y}_{sjt}(p, b))^2, \tag{4.2}$$

where p is the polynomial order and b is the bandwidth of the local linear regression model to predict $\hat{y}_{sjt}(p, b)$. The combination of p and b that minimizes this cross-validation function is chosen as the most-preferred functional form. The result of the cross-validation using hours worked as y_{sjt} is reported in Table 4.3. In this case, the combination of $p = 1$ and $b = 90$ is chosen as the most-preferred function. Although we do not present all the results here, we implemented the same cross-validation for the other dependent variables.

4.5 Results

4.5.1 Possibility of Establishment-Size Manipulation

Before estimating the effect of standard hours on labor-market outcomes, we examine whether there are any signs of manipulation of establishment size to avoid the strict standard hours. As shown in Table 4.1, establishment sizes of 100 and 300 serve as the thresholds at which standard hours change.[15] Thus, these two thresholds could potentially give establishments an incentive to reduce the number of employees to below the thresholds so that they could enjoy longer standard hours.

First, we explore the possibility of endogenous establishment-size decision by looking at simple statistics. Table 4.4 shows that the fraction of each size category is stable over time. In particular, the fraction of establishments with 100 or fewer employees

[14] For the discontinuity around 100, the lower bound is set at the 95th percentile of observations between 31 and 100, and the upper bound is set at the 5th percentile of observations between 101 and 200. The upper and lower bounds around 300 are similarly set.

[15] Note that employees in establishments with 30 or fewer employees are not included in our sample.

did not increase during policy changes, i.e., 1992 and 1996, and thus does not support the view that firms manipulated employment to escape stricter regulations.

Next, we examine changes in the distribution of establishment size around the two thresholds, 100 and 300. Figure 4.3 reports the histogram of establishment size. If establishments manipulated their size to avoid stricter regulations for larger establishments, we should observe a disproportionately larger number of establishments with size 100 than 101 and more with 300 than 301.

It is not surprising that we observe a spike at 100 because it could serve as a round number; thus it is possible that some firms reported their establishment size as 100 even though their real size is a little larger or smaller than that. If the spike at 100 became larger when establishments with 31-100 employees faced longer standard hours than larger establishments with 101 or more employees, however, it would offer evidence that firms intentionally decreased their number of employees a little to enjoy longer standard hours. We will conduct an additional analysis to test this possibility, because it is difficult to conclude based only on Figure 4.3 whether the spike at the thresholds became larger during the policy changes. The estimated model is:

$$\begin{aligned} lny_{jts} &= \sigma_1 I(s \in [101, 100+b]) + \sigma_2 I(s \in [301, 300+b]) \\ &+ k(s) + \tau T_t + D_j \xi + D_j T_t \nu + T_t l(s) + D_j m(s) + u_{jts} \end{aligned} \tag{4.3}$$

where $k(s) = \kappa_1 s + \kappa_2 s^2 + ... + \kappa_p s^p$, $l(s) = \lambda_1 s + \lambda_2 s^2 + ... + \lambda_p s^p$, and $m(s) = \mu_1 s + \mu_2 s^2 + ... + \mu_p s^p$. The original observation units are establishments, and they are collapsed to year$\times$ industry$\times$size cells. The dependent variable is the logarithm of the number of establishments within each cell. $I(s \in [101, 100+b])$ is a dummy variable that takes a value of one if the establishment is on the right side of the threshold of 100; similarly, $I(s \in [301, 300+b])$ is a dummy variable that takes a value of one if the establishment is on the right side of the threshold of 300.

The estimation sample includes observations with establishment size, $s \in [100 -$

$b, 100 + b]$ and those with $s \in [300 - b, 300 + b]$. Results with b=3, 5, and 10 are reported in Table 4.5. A linear function is chosen as the optimal function for $k(s)$, $l(s)$, and $m(s)$ by the cross-validation method.[16]

If establishments manipulated the number of employees to avoid stricter standard hours, the estimates of σ_1 and σ_2 should be negative. In Table 4.5, however, neither of these estimates is significantly negative with any bandwidth, which means that we do not see any evidence of firms' manipulation of employment in response to the change in standard hours.

4.5.2 Effect on Hours Worked

Now, we come back to the estimation of Equation (4.1) using hours worked as a dependent variable. The estimation result is reported in Table 4.6. Although results using a linear trend term, *year*–1989, are reported in Table 4.6, we also tried using a trend-term of functional forms, such as a logarithm, logistic, and polynomials and found that the results illustrated in Table 4.6 are robust against the functional forms of trend terms. Columns 1 through 4 suggest that a 1-hour reduction of standard work hours reduces weekly hours worked by 0.16-0.18 hours. By examining the results in Table 4.6, we can see that this result is robust against the combinations of interaction terms. The estimates in Column 4 in Table 4.6 are used to draw Figure 4.4. Establishments that hire 100 or fewer employees were subject to standard hours of 43.25, on average; 101 to 300 employees, 42.64 hours; and 301 and more, 41.75 hours. This discontinuity of standard hours creates a discontinuity of hours worked

[16] In this regression, the conventional heteroskedasticity-consistent standard error is used. It is known that the conventional heteroskedasticity-consistent standard error in the cell -weighted regression proposed in Lee and Card (2008) is equivalent to the clustered standard error formula in the micro-level regression. Thus, the robust standard errors used in the cell-weighted regression can be interpreted as standard errors clustered at year×industry×size category (by one). For the sake of consistency, in the remaining regressions in which the observation units are establishments, the standard errors are clustered at year×industry×size category (by one). In these micro-level regressions, changing the clustering level in standard errors (e.g., to year×industry) did not change the significance/insignificance of the result.

at establishment sizes of 100 and 300.

The size of the estimated coefficient, 0.158, implies that an 8-hour reduction of standard hours between 1988 and 1997 contributed to reducing the hours worked by 1.264 hours. Mean hours worked in 1989 was 44.58 hours (Table 4.2), implying that the law contributed to reducing the hours worked by 2.84 percent. The average hours worked was reduced by 4.04 hours between 1989 and 2002 (Table 4.2); thus the legislative change explains about 31 percent of the reduction of hours worked. Section 4.6 discusses the reason for this modest change in hours worked.

To address the concern that the sample includes the exempted workers discussed in Section 4.3, we estimate the hour equation using only production workers that do not include the exempted workers. The results are robust among the four columns in Table 4.6, and in Figure 4.5 we visually illustrate the result of Column 4 in Table 4.6. The results are similar to those in Figure 4.4 obtained from all workers, while the estimated policy effect for production workers is larger than that for all workers.

4.5.3 Effects on Wage

Previous studies of Continental Europe report that the reduction of standard hours reduced hours worked, while it did not reduce weekly or monthly wage. We now examine if this was the case for Japan.

There are several reasons to have a prior expectation that wages in Japan adjusted flexibly to the reduction of hours worked. First, wage determination is decentralized in Japan. The unionization rate was 25.9 percent in 1989 and declined monotonically to 21.5 percent in 2002. Japanese unions are organized at each company and union-employer wage negotiation takes place at the company level. Consequently, there is no standardized industry or occupation wage that applies to nonunion workers. The statutory minimum wage was defined over hourly rate of pay or daily pay for 8 hours, but not over weekly or monthly wage, and thus it does not affect our discussion.

Second, the majority of Japanese employees receive bonus payments, typically twice a year in summer and winter. According to Table 4.2, the average annual bonus amount reported in 1989 was 902,294 yen (about USD 9,022.94). The total annual bonus payment amounted to 3.5 times the employee's monthly salary in 1989, on average. As is mentioned in Section 4.1, the Japanese bonus amount is known to be relatively flexible because Japanese workers typically recognize that the amount can fluctuate depending on firms' performance (Hashimoto, 1979; Kawaguchi and Ohtake, 2007).

To examine the effect of standard hours on monthly wages, the log of cash earnings in June are regressed on standard hours.[17] Figure 4.6 draws the result of the local linear regression of the most-preferred specification at the discontinuity points. Workers in larger establishments earn higher monthly wages, on average, but the regression line is almost continuous at establishment sizes of 100 and 300. The estimated effect of standard hours on monthly wages is virtually zero, and it is not statistically significant. Table 4.6 also suggests that this result is robust against the combinations of interaction terms. Although the reduction of standard hours reduced hours worked, it did not decrease monthly wages.

Next, to assess the impact on the annual bonus amount, the log of bonus payments is regressed on standard hours. The annual bonus payment of the previous calendar year is reported by respondents in the survey. Since Japanese employees typically receive bonuses twice a year (summer and winter), the reported bonus amounts in the survey represent, in most cases, the sum of the summer bonus and the winter bonus. Note that any revision of standard hours occurred in April and that the bonuses are paid after April. Thus, in the regression, the log of annual bonus payments of the previous calendar year is regressed on standard hours after April in the previous year. According to Table 4.6, after we control for size-time effects, the coefficient of

[17] Bonus is not included in this variable.

the standard hours becomes insignificant. Figure 4.7 draws the result of Column 4 in Table 4.6. The mean log annual bonus amount increases as the establishment size increases, and the discontinuous changes in the bonus amount at the establishment sizes of 100 and 300 are virtually zero.

These results from Table 4.6 and Figures 4.6 and 4.7 suggest that even in Japan, where wages are determined relatively flexibly, neither monthly wage nor bonus payments responded to the reduction in standard hours in a significant way. This evidence suggests that the reduction of standard hours increased overall labor cost per hour. These results are exactly the same as the findings in many European countries, where pay is largely determined through union-employer bargaining, and the reduction of weekly or monthly wages faces strong resistance (Hunt (1999) for Germany, Crepon and Kramarz (2002), Estevao and Sa (2008) and Chemin and Wasmer (2009) for France, Nymoen (1989) for Norway, Pencavel and Holmlund (1988) and Skans (2004) for Sweden, and Raposo and van Ours (2008) for Portugal).

4.5.4 Effects on Employment

Whether or not the reduction in standard hours contributed to work-sharing has been a topic of research interest in many countries, and in this section we also examine it.

The BSWS is not designed as a panel study and basically cannot capture the change in the number of employees over time. It, however, allows us to recover the number of employees in each industry, establishment size, year cell using the frequency weight that tells how many individuals an observation in the sample represent. In addition to this measure for the overall employment, this survey also includes information on the number of workers who are directly recruited from schools at each establishment. Utilizing these measures of employment, we examine how employment changed in response to the reduction in standard hours.

First, we regress two different measures of the total employment: ln (Number of employees)[18] and the share of the total number of employees among cells (%) on standard hours andother covariates. The original observation units are individuals, and they are collapsed to 33 industries×3 size categories (31-100, 101-300, 301+)×14 years cells.[19] If the work-sharing argument applies, shorter standard hours should increase employment.

Table 4.7 reports the regression results, and it returns insignificant estimates for the coefficient of the standard hours in all the columns. This analysis in which employment is regressed on standard hours using repeated cross-sectional data, however, is likely to suffer from the bias caused by reverse causality because standard hours vary depending on the number of employees in an establishment. Accordingly, a positive employment shock could increase the likelihood of an establishment to be in a size category with lower standard hours, which would bias the estimates downwardly. Thus, there is a possibility that insignificant estimates illustrated in Table 4.7 would have been significantly positive without the endogeneity bias.

To circumvent this obvious reverse causality, we next utilize information on the number of workers who are directly recruited from schools at each establishment. The number of workers directly recruited from schools captures the change of employment most sensitively, because the adjustment cost is presumably minimal.[20] Table 4.6 shows the results of regressions in which the ratio of the number of newly hired workers to the number of existing workers is regressed on standard hours.[21] According to this

[18]The number of employees in each cell is calculated by summing up the frequency weight within a cell.

[19]The reason we did not take a regression discontinuity for this regression is that in a regression discontinuity model, the X axis is establishment size with a bin width of 1 and the Y axis is the total number of employees at each size, i.e., (the number of establishments at each size)×(the establishment size). If we use the establishment size with a bin width of 1 as X, X almost overlaps with Y because the number of establishments at each size is very small. Thus, in Table 4.7, we collapsed the data into larger cells and used the total number of employees within each cell for Y.

[20]Reflecting this fact, how new hiring from schools is adjusted to exogenous changes, such as recession or an aging society, has been the subject of many studies. Here, we examine the adjustment of hiring new graduates in terms of the change in standard work hours.

[21]The reason for using the fraction, and not the number of newly hired workers from schools,

table, after we control for size-time effects, the coefficient of standard hours becomes insignificant.

Figure 4.8 visually confirms the result of Column 4 in Table 4.6. Note that establishments that hire no school graduates are also included in the sample to capture behavioral changes, such as stopping or starting hiring from schools, among establishments. The fraction of establishments with no hiring from schools is large, particularly for establishments with less than 100 employees (61% for establishments with 31-100 employees, 29% for 101-300, and 12% for 301+). Reflecting this fact, Figure 4.8 shows the fraction of newly hired workers from schools increases as the establishment size becomes larger. If we exclude firms with no hiring from schools, the figure indicates the opposite trend (i.e., the fraction of new employees are smaller for large establishments because the number of existing employees are larger for these establishments.) In either case (with or without hiring of new school graduates), however, the estimated coefficient implies that the reduction of standard hours did not affect recruitment from schools.

All information combined, although the reduction of standard hours reduced the total hours worked, neither monthly wage nor bonus payments of existing workers decreased in response to the policy change. As a result, the overall labor cost per hour increased, and we do not see any evidence of job creation. This implies that the idea of work-sharing did not work even in an economy where wages are flexibly determined.

4.5.5 Effects on Full- and Part-time Compositions

An unintended consequence of the reduction of standard hours may be to encourage firms to substitute full-time workers with part-time workers. Firms incur fixed

is that there is a possibility that the same reverse causality problem (the regression with total employment) might occur when using the number of new hires. (Since new hires increase depending on the number of existing employees, we used the fraction of newly hired workers from schools within each establishment, instead of its number.)

costs to employ full-time workers, such as social security co-payments, recruitment, and firing costs, including retirement allowances. The reduction of hours worked by a full-time worker increases the fraction of fixed costs per hour. Since the fixed cost for a part-time worker is lower for various reasons, such as the exemption from social security payments or less commitment to long-term employment, the reduction of standard hours may create an incentive for firms to substitute full-time workers with part-time workers.

To assess whether the reduction of standard hours led to this unintended consequence, the fraction of part-time worker is regressed on standard hours. Table 4.6 indicates that the effect of standard hours on the fraction of part-time workers is insignificant, and this result is robust against the combinations of interaction terms. Figure 4.9 reports the result of the local linear regression of the most-preferred specification estimated around the discontinuity points (Column 4 in Table 4.6). The larger an establishment, the lower is its fraction of part-time workers. The estimated coefficient, however, does not suggest the discontinuity of the fraction created by the discontinuous change of standard hours. Therefore, the claim that the reduction of standard hours increased the fraction of part-time workers is not empirically supported.[22]

4.6 Why hours did not respond much

The analysis in Section 4.5.2 reveals that the legislative change explains only about 31 percent of the reduction in hours worked. This change in hours worked appears

[22]We recognize that this regression might suffer from an endogeneity problem that arises from the definition of part-time workers. Since part-time workers in Japan are defined as those whose scheduled workweek is less than that of regular employees, the cut-off hours are equal to standard work hours in many cases. Thus, the cut-off hours declined over the period with the reduction in standard hours, which could cause the endogeneity in this estimation, and it is possible that the coefficient of standard work hours is upwardly biased. Since even with this potential upward bias the estimates are not significantly positive, we cannot deny the possibility that the reduction in standard hours increased the fraction of part-time workers only from this study. More consideration should be given to this topic in future studies.

to be somewhat modest. We recognize this small response of weekly hours worked as evidence that reflects the fact that the fraction of firms affected by the policy change was originally restricted. The policy affected only firms with an initial workweek longer than the revised standard hours. It is also expected that firms in which the average workweek was originally longer than the initial standard hours would not decrease the workweek significantly in response to the policy change, since they were already willing to pay overtime prior to the policy change.[23] In our sample, 16% of all the establishments in 1989 initially set the workweek shorter than 40 hours/week. Thus, for these firms, the policy change would have no impact. In addition, 27% of establishments in the sample originally set the workweek longer than 48 hours; these establishments can be considered as those that do not mind paying overtime, and it is likely that the policy change did not matter for them.

Considering that nearly one half of the sample comes from the group for which the policy change is unlikely to be effective, it is easier to understand the small estimates in Section 4.5.2. In other words, there is a possibility that the workweek dropped largely for firms that actually decreased it, but this effect was attenuated by the large number of firms that were not affected by the policy change. Therefore, the small estimate in Section 4.5.2 does not necessarily imply that the change of the workweek within the same firm was small.

While acknowledging this fact as the principal reason for the moderate estimated impact, in this section we further explore the possibility that the fixed job model by Trejo (1991) can explain the reason for it.

Trejo (1991) presents the following two types of theoretical models: the fixed wage and the fixed job models. In the fixed wage model, the straight-time hourly wage rate is exogenously given. In response to the change in standard hours, firms choose the combination of the number of employees and the weekly hours worked by

[23] These differences in choosing work hours above or below standard hours depend on the difficulty of substituting hours for employment and the quasi-fixed cost of employment.

each employee, given fixed capital and an exogenous output level. In the most relevant case in which the initial number of standard hours does not bind but the number of revised standard hours does, this model predicts that cost-minimizing firms move to a new kink at the revised standard hours, i.e., 40 hours in the firm's cost function. In this case, firms reduce work hours through the substitution effect (between hours of work and employment) and the scale effect. In the fixed job model, firms and workers agree on a set of weekly work hours and earnings, and both are indifferent to the possible combinations of straight-time and overtime wage rates, as long as the same weekly work hours and earnings are attained. Thus, regardless of the initial levels of the agreed-upon work-hour and earning sets, the straight-time hourly wage rate can be adjusted in accordance with legislative changes in overtime premiums and standard hours. Such adjustments maintain the initial levels of weekly work hours and earnings. Thus, if this fixed-job model is valid in our data, the modest change in hours worked could be attributed to the flexible adjustment of the straight-time hourly wage rate. Mathematically, weekly earnings Y can be written as:

$$Y = w[\bar{h} + p(h - \bar{h})] \tag{4.4}$$

where w represents the straight-time hourly wage rate, and p denotes the overtime premium rate.[24] In the fixed job model, w is adjusted to achieve the same combinations of weekly earnings, Y, and weekly work hours, h. During the policy changes, standard hours $\bar{h}$ decreased. Hence, if the Japanese labor market is explained by the fixed job model, w should decrease in response to the reduction in standard hours, $\bar{h}$, for the same Y and h to be realized. To test if this model is applicable here, we estimate the following model:

$$lnw_i = \alpha\bar{h}_{jts} + X_i\beta + D_s\phi + \rho\, T_t + D_j\,\delta + D_jT_t\,\zeta + T_tD_s\gamma + D_{js}\eta + u_i \tag{4.5}$$

[24]The Labor Standard Act in Japan requires $p \geq 1.25$.

The dependent variable is the log of the straight-time hourly rate of individual i in industry j, year t, and establishment size s. In this paper, we adopt the setting used in Trejo (1991). Equation (4.5) is estimated at the individual level, with explanatory variables, such as female variables, weekly hours worked and its square, educational dummy variables, and potential years of experience and its square. The reason weekly hours are included as X in the equation above is that validating the effectiveness of the fixed job model necessitates comparing the straight-time hourly rates of individuals who work the same number of hours per week. Trejo (1991) states that the above-mentioned wage equation is neither a labor-demand nor a labor-supply function, but rather the type of equilibrium wage-hour locus described by Lewis (1969) and Rosen (1974). Since wage rates vary with hours of work in the hedonic equilibrium context, weekly hours and its square are included, as X in Trejo (1991).

Table 4.8 shows the regression results. In all specifications, the coefficient of standard hours is significantly negative, indicating that the reduction in standard hours is accompanied by an increase in the straight-time hourly wage rate. This result rejects the fixed job model that predicts a reduction in the hourly wage rate in response to a decrease in standard hours.[25]

4.7 Conclusion

This paper examined the effect of reducing standard hours on hours worked, wages, and employment based on Japanese data. The number of weekly standard hours

[25] Note that the estimated coefficients of standard hours in Table 4.8 do not directly suggest that the fixed wage model is also rejected in our data. The model above regards straight-time hourly wage rates as a control variable and also compares them among workers who worked the same hours. Therefore, the model setting employed here itself does not fit the idea of the fixed wage model in which hourly wage rates are assumed to be exogenous. Unfortunately, our data do not contain the necessary information to test the validity of the fixed wage model, because testing the model requires information about the employer's quasi-fixed employment costs (Barkume (2010)), and it has been shown that having this information is essential to test the model (Ehrenberg (1971) and Ehrenberg and Schumann (1982)). If we interpret the modest change in hours worked within the framework of the fixed wage model instead, factors such as large fixed costs, a small-scale effect, and a small substitution between capital and labor could explain the change.

declined from 48 in 1988 to 40 in 1997.

We exploit the discontinuity of standard hours by establishment sizes and the gradual phase-in of stricter standard hours by establishment size and industry to estimate the policy impacts on hours worked, wages, and employment. The analysis applying a regression discontinuity design revealed that the reduction of standard hours modestly reduced hours worked. About 31 percent of the reduction of hours worked between 1989 and 2002 is attributable to the policy change. This reduction of hours worked was not associated with a decrease of either monthly wage or bonus payment. As a result, overall labor cost per hour increased, and we do not see any evidence of job creation.

The analysis of Japan's nationwide policy change adds new evidence to the literature that the work-sharing policy did not work even in an economy where wages are flexibly determined.

Appendix: Definitions of Terms Used

Standard Hours : Legal limits posed on weekly work hours by the Labor Standard Act. Employers can set the hours worked to exceed these legal limits only under an agreement with a workers' representative that represents the majority of employees. Overtime under this agreement must be compensated by at least a 25 percent wage premium.

Hours Worked : Calculated as "Actual number of scheduled hours worked + Actual number of overtime hours worked." In order for this to be comparable with the standard hours regulated at a weekly level, "hours worked" in this paper refers to hours worked by employees per week.

Scheduled Cash Earnings : Amount of contractual cash earnings, not including overtime allowance. This amount is composed of earnings in cash only, not of earnings in kind.

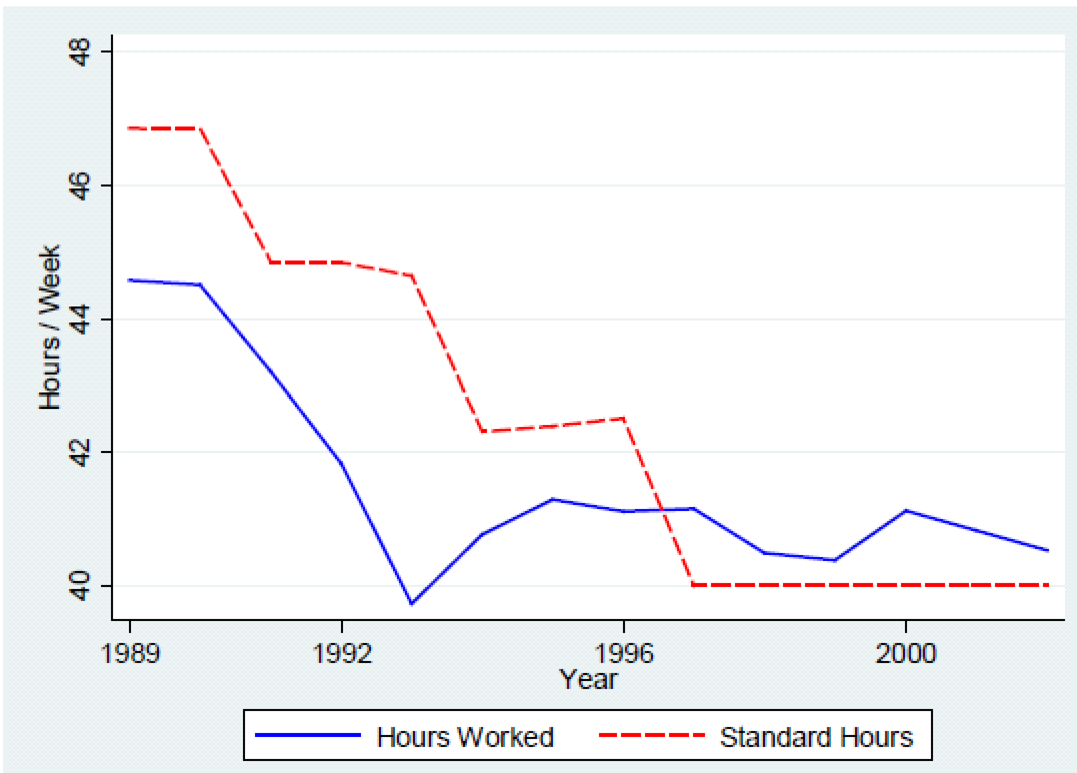
Overtime Allowance : Allowances paid for overtime worked.

Monthly Wages : Scheduled Cash Earnings+Overtime Allowance

Straight-time Hourly Wage Rate : Hourly wage rate for scheduled hours worked. In the data, this is calculated as (Scheduled Cash Earnings - Commuter Allowance - Regular Attendance Allowance - Family Allowance) / Scheduled Hours.

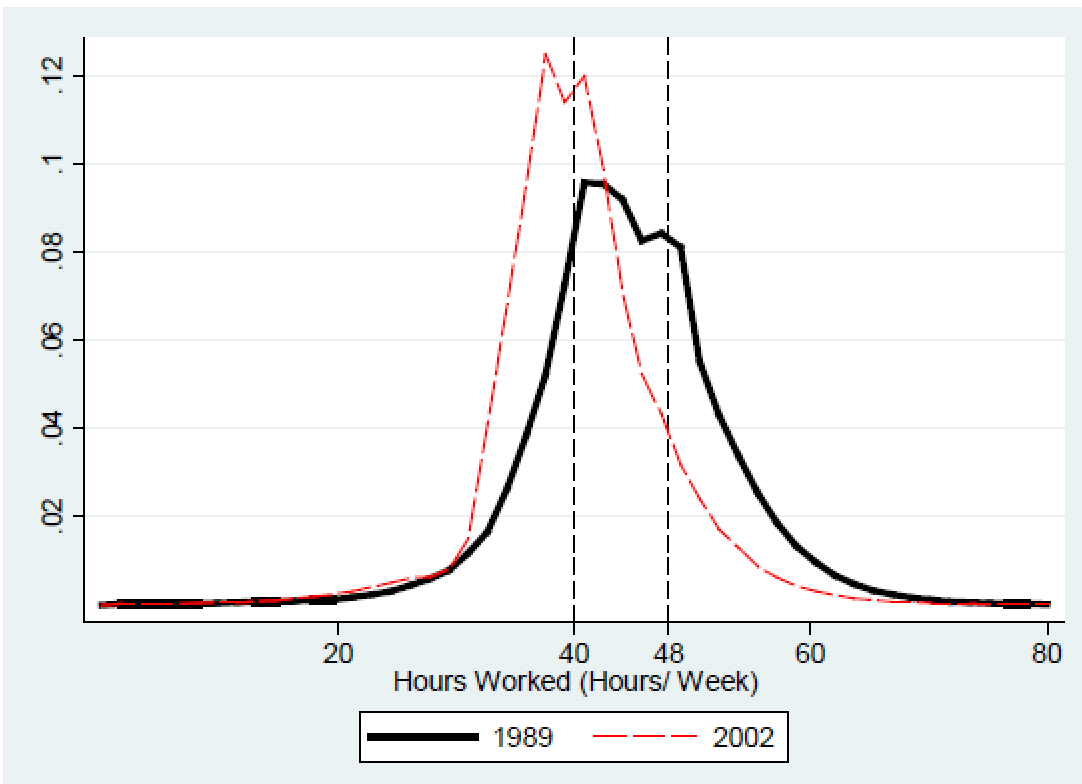
Bonus Payment : The majority of Japanese employees receive bonus payments in addition to their regular salary, typically twice a year in summer and winter. The bonus payment refers to the total bonus amount they receive during the period from January 1 to December 31. The survey is carried out annually in June, and respondents report the annual bonus payment for the previous year.

Figure 4.1: Weighted Average of the Standard Hours and Hours Worked in 4 Industries (Mining, Construction, Manufacturing, and Electricity, Gas, Heat Supply, and Water), 1989-2002



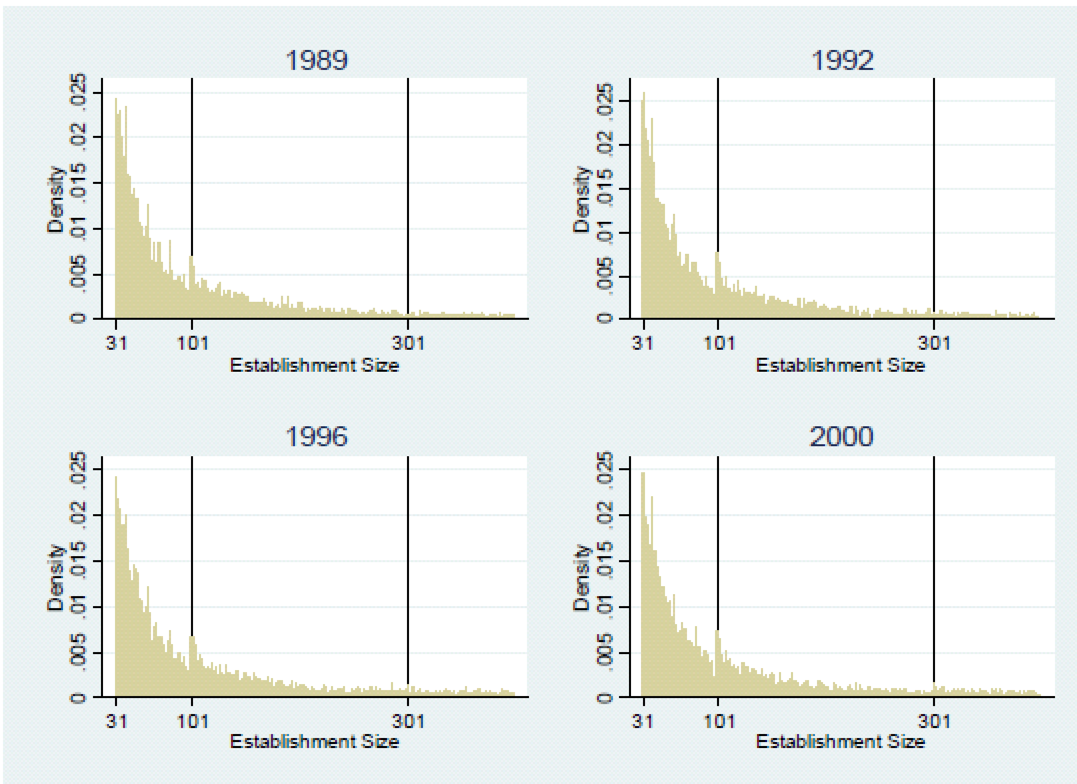
Note: Observation units are individuals. The average standard hours and hours worked in each year are calculated from microdata of Basic Survey of Wage Structure. Sample is restricted to workers in establishments with 30 or more employees in four industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water.

Figure 4.2: Distribution of Hours Worked, 1989 and 2002



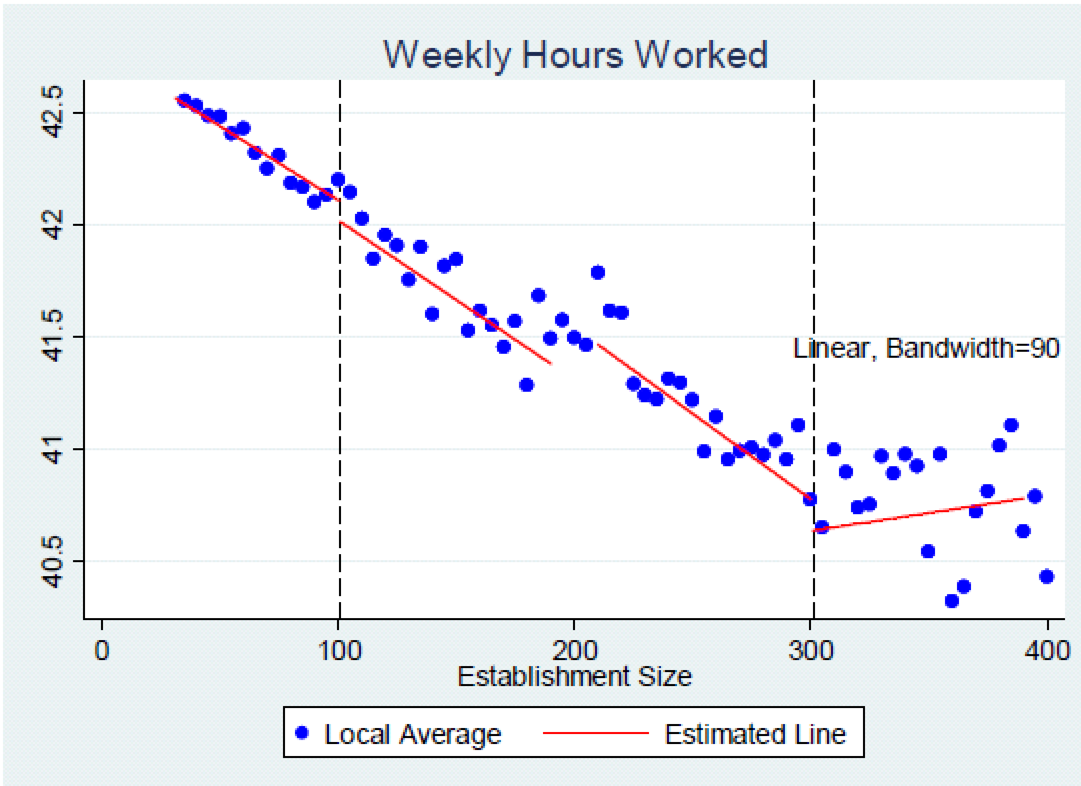
Note: Observation units are individuals. The density of each work hour is calculated from microdata of the Basic Survey of Wage Structure. Sample is restricted to workers in establishments with 30 or more employees in four industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water.

Figure 4.3: Distribution of Establishment Size, 1989, 1992, 1996, and 2000



Note: Observation units are establishments. The density of each establishment size is calculated from microdata of the Basic Survey of Wage Structure. Sample is restricted to workers in establishments with 30 or more employees in four industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water. The bin width used in the figures is one.

Figure 4.4: Impact of Standard Hours on Weekly Hours Worked, 1989-2002

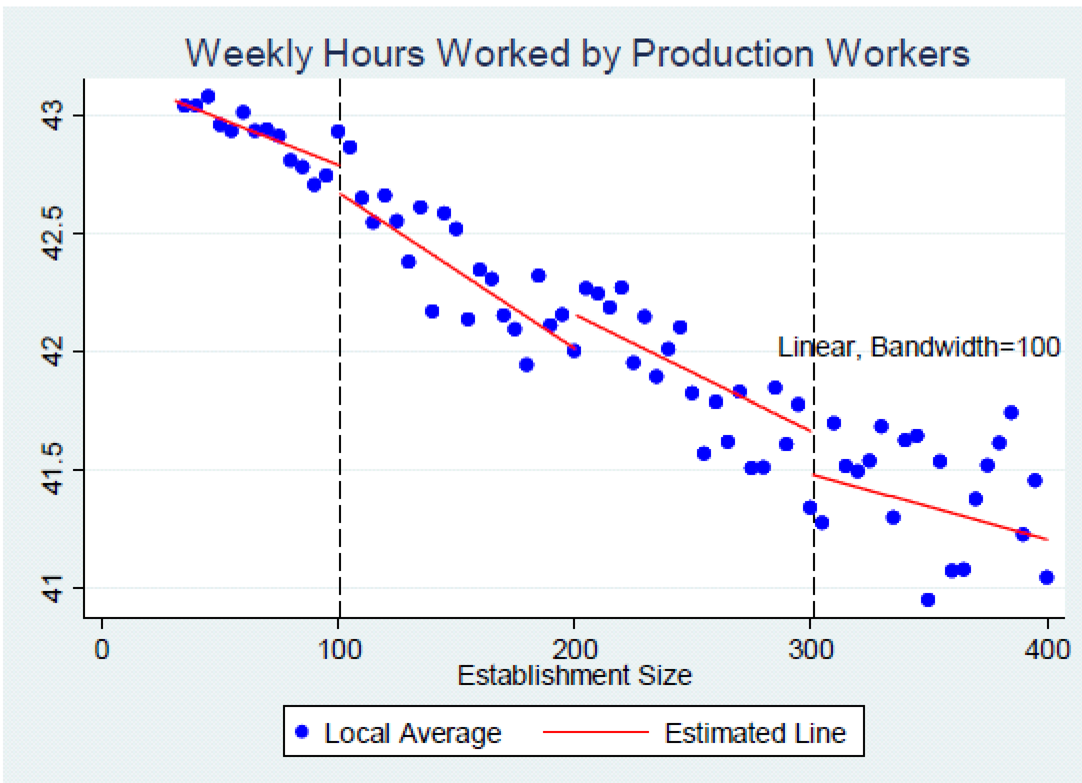


Note: Observations units are year×industry×size. The estimates are obtained from Column 4 in Table 4.6. The dots in the figure are the predicted values obtained after controlling for confounding factors. (They are the predicted values from the regression of y on establishment-size dummies with bin width of 5, a time-trend term (T), industry, and industry×T. The sample means of T and industry dummy variables are substituted throughout all the size levels to plot the dots.) The red line refers to the estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.012)}{0.158}\,\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 58616, R^2 = 0.340$$

The conventional heteroskedasticity-consistent standard errors are reported in parentheses. The sample means of standard hours, which are 43.25 for 31-100, 42.64 for 101-300, and 41.75 for 301+, and sample means of T and industry dummy variables are substituted to draw the curve in Figure 4.4. Each variable except for standard hours is fixed at its sample mean throughout all the firm-size levels, and only the standard hours were allowed to change at the thresholds of firm size. Thus, the magnitude of discontinuities in the figure corresponds to the coefficients estimated in Column 4 in Table 4.6.

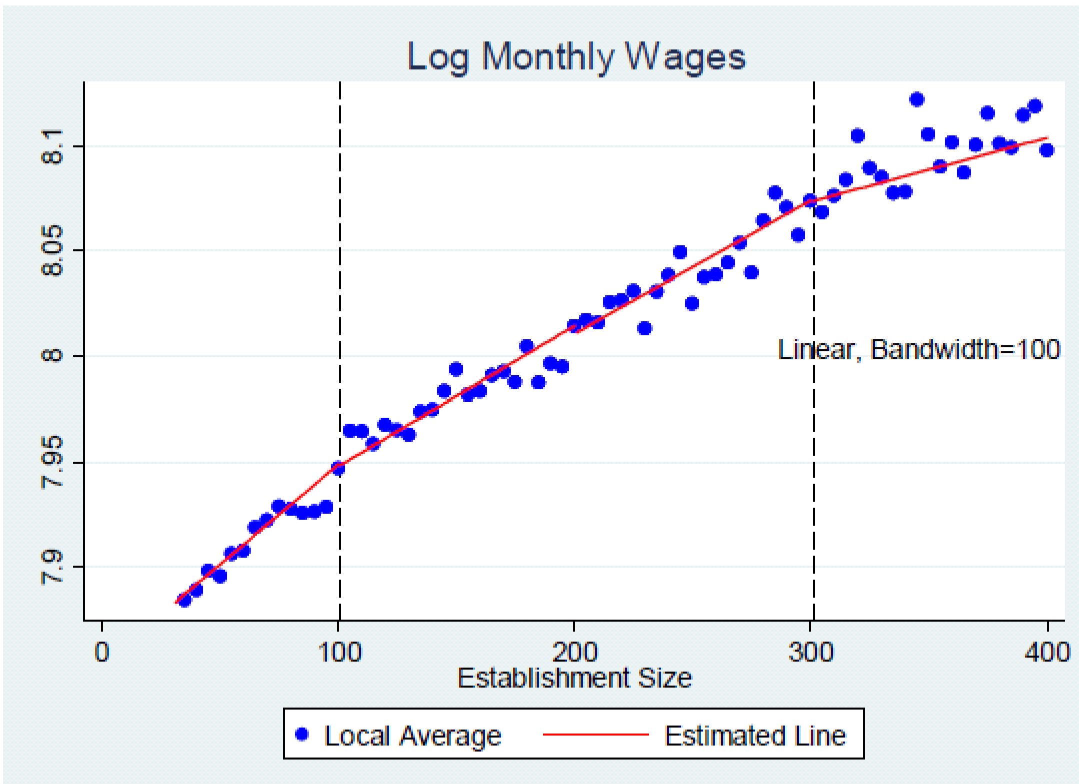
Figure 4.5: Impact of Standard Hours on Weekly Hours Worked by Production Workers, 1989-2002



Note: The same note applies as in Figure 4.4 except for
Estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.016)}{0.207}\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 52533, R^2 = 0.262$$

Figure 4.6: Impact of Standard Hours on Log Monthly Wages, 1989-2002

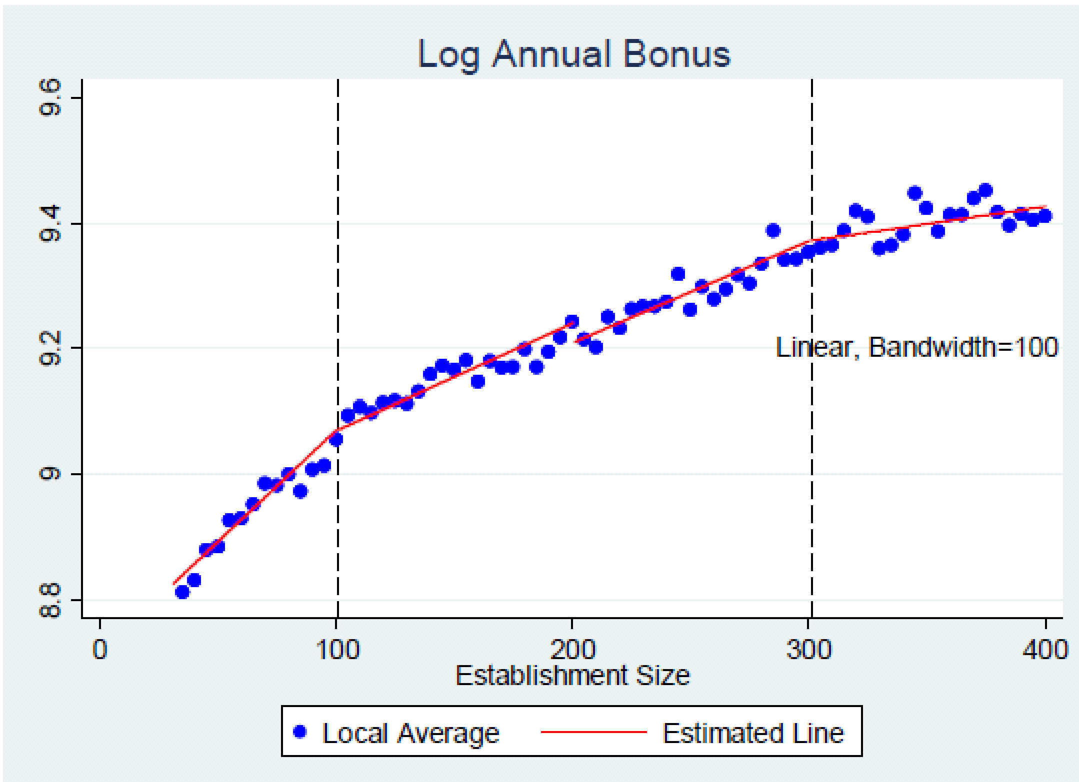


Note: Monthly wages are calculated as scheduled cash earnings plus overtime allowance. The real values, calculated as nominal wages divided by the consumer price index (CPI), are used. The same note applies as in Figure 4.4 except for

Estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.001)}{0.0003}\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 61749, R^2 = 0.634$$

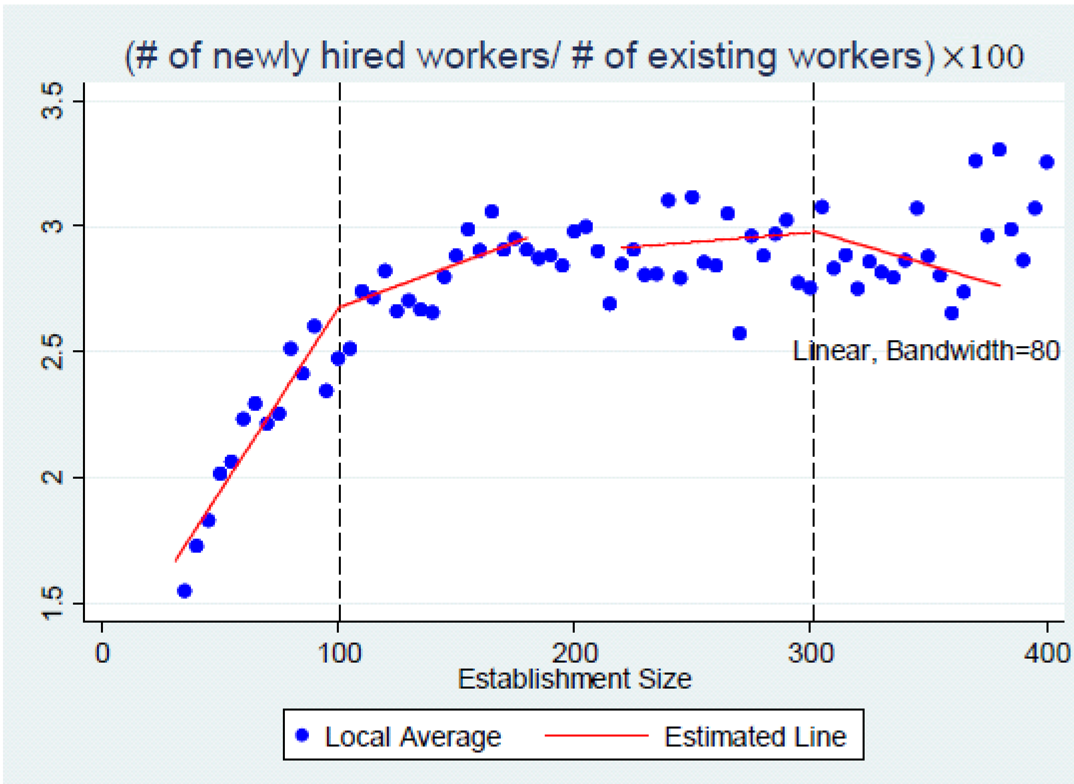
Figure 4.7: Impact of Standard Hours on Log Annual Bonus, 1989-2002



Note: The real values, calculated as nominal bonus amounts divided by the consumer price index (CPI), are used. The same note applies as in Figure 4.4 except for
Estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.002)}{-0.001}\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 61527, R^2 = 0.544$$

Figure 4.8: Impact of Standard Hours on the Ratio of the Number of Newly Hired Workers to the Number of Existing Workers, 1989-2002

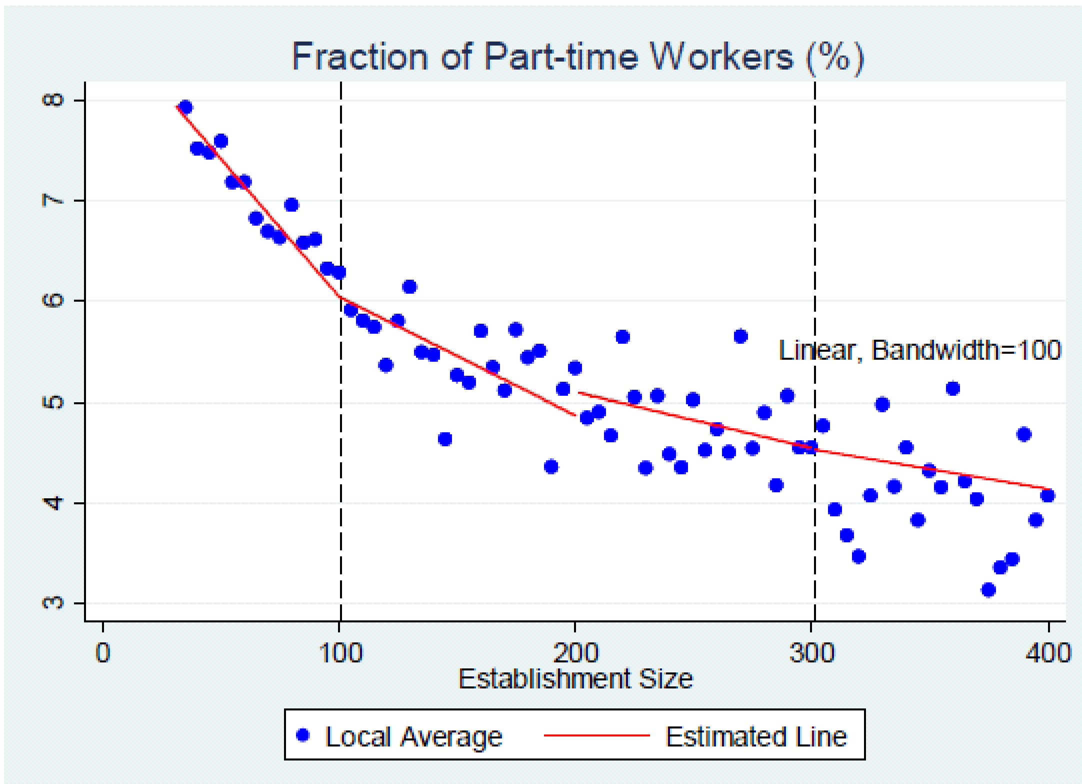


Note: The dependent variable is (the number of newly hired workers)×100/ (the number of existing workers) for each establishment. The minimum value of the dependent variable is zero. Observation units are establishments. The estimates are obtained from Column 4 in Table 4.6. The dots in the figure are the predicted values obtained after controlling for confounding factors. (They are the predicted values from the regression of y on establishment size dummies with a bin width of 5, a time-trend term (T), industry, and industry×T. The sample means of T and industry dummy variables are substituted throughout all the size levels to plot the dots.) The red line refers to the estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.009)}{-0.008}\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 135174, R^2 = 0.068$$

Standard errors, clustered at year×industry×size category (by one), are reported in parentheses. The sample means of standard hours, which are 43.25 for 31-100, 42.66 for 101-300, and 41.73 for 301+, and sample means of T and industry dummy variables are substituted to draw the curve in Figure 4.8. Each variable except for standard hours is fixed at its sample mean throughout all the firm-size levels, and only the standard hours were allowed to change at the thresholds of firm size. Thus, the magnitude of the discontinuities in the figure corresponds to the coefficients estimated in Column 4 in Table 4.6.

Figure 4.9: Impact of Standard Hours on Fraction of Part-time Workers (%), 1989-2002



Note: The real values, calculated as nominal bonus amounts divided by the consumer price index (CPI), are used. The same note applies as in Figure 4.4 except for
Estimated equation at the discontinuity:

$$\hat{y} = \underset{(0.034)}{0.018}\overline{h_{jst}} + \hat{f}(s) + \hat{\beta}T_t + D_j\hat{\delta} + D_jT_t\hat{\zeta} + T_t\hat{g}(s) + D_j\hat{h}(s), N = 61749, R^2 = 0.346$$

Table 4.1: History of Standard Hours by Industry and Establishment Sizes Included in the Analysis Sample

	-1988/3	1988/4-1991/3	1991/4-1993/3	1993/4-1994/3	1994/4-1995/3	1995/4-1997/3	1997/4-
Manufacturing							
1-9	48	48	46	46	46	44	40
10-30	48	48	46	46	44	44	40
31-100	48	48	46	46	44	44	40
101-300	48	46	44	44	44	44	40
301+	48	46	44	44	40	40	40
Mining							
1-9	48	48	46	46	46	44	40
10-30	48	48	46	46	44	44	40
31-100	48	48	46	46	44	44	40
101-300	48	48	46	44	44	44	40
301+	48	48	46	44	44	44	40
Construction							
1-9	48	48	46	46	46	44	40
10-30	48	48	46	46	44	44	40
31-100	48	48	46	46	44	44	40
101-300	48	48	46	44	44	44	40
301+	48	46	44	44	40	40	40

Note: The numbers in the first column (1-9, 10-30, 31-100, 101-300, and 301+) represent establishment sizes, i.e., the number of employees in each establishment. The legal industry classification in the Labor Standard Act (LSA) and the statistical industry classification in the Basic Survey of Wage Structure (BSWS) do not overlap for many industries. Thus, to avoid the biases that arise from mismatching between a worker and a standard hour, we restrict our analysis to workers who are in manufacturing, mining, and construction industries because legal and statistical industrial codes perfectly overlap for these three industries. The industry category in Table 4.1 is that used in the LSA. The "manufacturing" category in the LSA includes both the manufacturing industry and the "electricity, gas, heat supply, and water" industry in BSWS. Thus, although the industry category names used are different between the LSA and the BSWS, it is still possible to perfectly identify the standard hours that apply to a specific worker for the three industries above.

Table 4.2: Descriptive Statistics of the Analysis Data, 1989-2002, 4 Industries: Mining, Construction, Manufacturing, and Electricity, Gas, Heat Supply and Water, All Employees in Establishments with 30 or More Employees, 1989-2002

Sample	Total	1989	1992	1996	2000	2002
Standard Hours	42.55	46.86	44.85	42.52	40.00	40.00
	(2.79)	(0.99)	(0.99)	(1.93)	0.00	0.00
Hours Worked	41.57	44.58	41.84	41.12	41.13	40.54
	(7.35)	(7.87)	(7.28)	(7.11)	(7.14)	(7.04)
Scheduled Hours Worked	38.18	40.17	38.58	37.90	37.76	37.33
	(5.37)	(5.67)	(5.59)	(5.24)	(5.02)	(5.03)
Overtime Hours Worked	3.39	4.41	3.27	3.22	3.37	3.21
	(4.52)	(5.15)	(4.34)	(4.33)	(4.57)	(4.47)
Monthly Scheduled Cash Earnings	2787.63	2287.28	2611.57	2883.95	3004.04	3003.08
(100 yen $\approx$ 1 USD)	(1369.87)	(1142.81)	(1288.43)	(1389.54)	(1409.23)	(1405.64)
Monthly Overtime Allowance	317.58	327.31	293.12	316.61	336.41	324.70
(100 yen $\approx$ 1 USD)	(437.06)	(415.58)	(413.12)	(438.81)	(462.78)	(457.02)
Annual Bonus in Previous Calendar Year	11296.43	9022.94	11288.99	11783.71	11158.88	11104.16
(100 yen $\approx$ 1 USD)	(9360.74)	(7815.31)	(9597.12)	(9445.77)	(9199.96)	(9307.42)
Straight-time Hourly Wage Rate	17.21	13.40	15.92	17.93	18.63	18.88
(100 yen $\approx$ 1 USD)	(10.20)	(7.76)	(9.15)	(11.32)	(10.17)	(10.51)
Establishment Size	525.23	547.92	579.14	505.14	462.61	450.61
	(1192.51)	(1304.22)	(1387.01)	(1099.40)	(1016.09)	(974.16)
Fraction of Newly Hired Workers	2.45	3.10	3.04	2.39	1.52	1.46
from Schools (%)	(3.59)	(4.30)	(4.08)	(3.47)	(2.60)	(2.51)
Fraction of Part-time Workers (%)	5.26	5.24	5.22	4.93	5.87	6.27
Establishment Size: (%)						
301+	37.04	35.63	36.48	37.18	35.47	34.94
101-300	24.40	23.95	23.67	23.57	24.43	24.28
31-100	38.57	40.42	39.85	39.25	40.10	40.77
Industry Distribution: (4 categories) (%)						
Mining	1.01	1.42	1.21	0.95	0.93	0.84
Construction (3 categories)	9.45	6.99	6.88	8.28	13.24	12.35
Manufacturing (22 categories)	82.43	86.94	87.31	83.28	78.34	78.80
Electricity, Gas, Heat supply,	7.10	4.66	4.61	7.49	7.49	8.01
and Water (4 categories)						
Observations	6221155	459642	454879	463382	436067	399720

Note: Standard deviations are in parentheses. All monetary compensation is denominated in 100 yen (approximately one US$). The industry category in Table 4.2 is that used in the Basic Survey of Wage Structure. Sample is restricted to workers in establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water.

Table 4.3: Cross-validation Functions and Optimal Bandwidth for Local Linear Regressions of Hours Worked on Size of Establishment

	Values for Cross Validation			
Bandwidth	Linear	Quadratic	Cubic	Quartic
10	16.406	24.488	226.465	12786.800
20	15.422	157.439	170.463	195.530
30	14.876	15.836	18.101	3538.984
40	14.632	15.621	19.072	25.056
50	14.565	14.971	16.406	21.587
60	14.484	14.855	16.706	20.742
70	14.405	14.949	24.053	4645.501
80	14.386	14.651	18.252	132.341
90	14.365	14.572	17.876	143.322
100	14.371	14.538	15.567	35.797
The optimal combination of a bandwidth and a polynomial function	Linear, Bandwidth=90			

Note: A result that has the minimum CV is chosen as the optimal combination of a bandwidth and a polynomial function. The value of cross validation is given as $CV(p,b) = \{\Sigma_{s,j,t}(y_{sjt} - \hat{y}_{sjt}(p,b))^2\}/N$, where $\hat{y}_{sjt}(p,b)$ is predicted from the local linear regression of a polynomial order, p, and a bandwidth, b, and N is the number of observations within 5 percentage points of each cutoff point on either side of the discontinuity (bandwidth b).

Table 4.4: Distribution of Establishments among Establishment-Size Regions (1989, 1992, 1996, and 2000)

Establishment Size Year	1989	1992	1996	2000
301+	22.4%	22.7%	23.9%	22.7%
101-300	25.2%	25.1%	24.9%	25.5%
31-100	52.4%	52.2%	51.2%	51.8%
Total	100%	100%	100%	100%

Note: Observation units are establishments. The distribution of establishments among 3 establishment-size regions (31-100, 101-300, and 301+) is calculated from microdata of Basic Survey of Wage Structure. Sample is restricted to establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply and water.

Table 4.5: Impact of the Reduction in Standard Hours on Firm's Employment Decision, 1989-2002

Dependent Variable: ln(Number of Establishments)$_{jts}$	(1)	(2)	(3)
Bandwidth (b)	3	5	10
$I(s \in [101, 100 + b])$	0.003	0.080	0.161
	(0.043)	(0.039)	(0.034)
$I(s \in [301, 300 + b])$	0.085	0.095	0.082
	(0.059)	(0.053)	(0.042)
Industry	Yes	Yes	Yes
Trend	Yes	Yes	Yes
$k(s)$	Yes	Yes	Yes
Industry×Trend	Yes	Yes	Yes
Trend×$l(s)$	Yes	Yes	Yes
$m(s)$×Industry	Yes	Yes	Yes
R^2	0.452	0.411	0.353
N	2554	3990	7591

Note: The original observation units are establishments, and they are collapsed to year×industry×size cells. The dependent variable is the logarithm of the number of establishments within each cell. $I(s \in [101, 100 + b])$ is a dummy variable that takes a value of one if the establishment is on the right side of the threshold of 100; similarly, $I(s \in [301, 300 + b])$ is a dummy variable that takes a value of one if the establishment is on the right side of the threshold of 300. Cell size-weighted regressions, using (the number of observations in cell j)/(the number of total observations/the number of cells) as a weight (Lee and Card (2008)), are performed. The conventional heteroskedasticity-consistent standard errors are reported in parentheses. The estimation sample includes observations with establishment size, $s \in [100 - b, 100 + b]$ and those with $s \in [300 - b, 300 + b]$. Results with b=3, 5, and 10 are reported in Table 4.5. A linear trend term, calculated as year-1989, is used in the regressions. A linear function is chosen for $k(s)$, $l(s)$, and $m(s)$ by the cross-validation method. The coefficients of these continuous functions are allowed differ among 4 establishment-size regions: 31-100, 101-200, 201-300, and 301+. In the regression, $I(s \geq 201)$ is also included, and hence a gap in the number of establishments between the regions around s=101 and 301 is allowed. A constant term is also included. Samples are restricted to establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply and water.

Table 4.6: Impact of Standard Hours on Weekly Hours Worked, Earnings, and Employment, 1989-2002

Dependent Variables	(1)	(2)	(3)	(4)
Weekly Hours Worked	0.179	0.176	0.153	0.158
	(0.012)	(0.012)	(0.013)	(0.012)
R^2	0.317	0.323	0.323	0.340
N	58616	58616	58616	58616
Weekly Hours Worked by Production Workers	0.232	0.227	0.208	0.207
	(0.016)	(0.016)	(0.017)	(0.016)
R^2	0.242	0.245	0.246	0.262
N	52533	52533	52533	52533
Log Monthly Wages	0.001	0.001	0.0004	0.0003
	(0.001)	(0.001)	(0.001)	(0.001)
R^2	0.616	0.622	0.622	0.634
N	61749	61749	61749	61749
Log Annual Bonus	0.003	0.004	-0.0001	-0.001
	(0.002)	(0.002)	(0.002)	(0.002)
R^2	0.522	0.528	0.529	0.544
N	61527	61527	61527	61527
Number of Newly Hired Workers from School×100/	-0.084	-0.081	-0.007	-0.008
Number of Existing Workers	(0.009)	(0.009)	(0.009)	(0.009)
R^2	0.053	0.057	0.062	0.068
N	135174	135174	135174	135174
Fraction of Part-time Workers (%)	0.040	0.033	0.032	0.018
	(0.034)	(0.034)	(0.035)	(0.034)
R^2	0.314	0.324	0.324	0.346
N	61749	61749	61749	61749
Industry	Yes	Yes	Yes	Yes
Trend	Yes	Yes	Yes	Yes
$f(s)$	Yes	Yes	Yes	Yes
Industry×Trend	No	Yes	Yes	Yes
Trend×$g(s)$	No	No	Yes	Yes
$h(s)$×Industry	No	No	No	Yes

Note: For all the estimations except for that of the fraction of newly hired workers from school, observation units are year×industry×size. The conventional heteroskedasticity-consistent standard errors are reported in parentheses. In these estimations, cell size-weighted regressions, using (the number of observations in cell j)/(the number of total observations/the number of cells) as a weight (Lee and Card (2008)), are performed. A linear trend term, calculated as year−1989, is used in the regressions. For the estimation of the fraction of newly hired workers from school, observation units are establishments, and standard errors are clustered at year×industry×size category (by one). For all the estimations, the sample includes observations with establishment size, $s \in [100-b, 100+b]$ and those with $s \in [300-b, 300+b]$. The combination of the polynomial order (p) and the bandwidth of the local linear regression (b) is the one that attains the minimum value in the cross-validation from the specification of Column 4. The coefficients of the continuous functions $f(s)$, $g(s)$, and $h(s)$ are allowed to differ among 4 establishment-size regions: 31-100, 101-200, 201-300, and 301+. In the regressions, $I(s \geq 201)$ is also included, and hence a gap in the number of establishments between the regions around s=101 and 301 is allowed. A constant term is also included. Samples are restricted to establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply and water. Monthly wages and annual bonus are corrected for inflation.

Table 4.7: Impact of Standard Hours on Total Employment, 1989-2002

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent Variable	ln(Number of Employees)$_{jts}$			Share of Total Employees among Cells		
Standard Hours	-0.021	-0.054	0.042	-0.005	-0.009	0.005
	(0.023)	(0.074)	(0.022)	(0.007)	(0.017)	(0.003)
Constant	7.815	9.400	3.886	0.200	0.408	-0.225
	(1.163)	(3.662)	(1.082)	(0.319)	(0.821)	(0.142)
Industry	Yes	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes	Yes
Size	Yes	Yes	Yes	Yes	Yes	Yes
Industry×Year	Yes	Yes	Yes	Yes	Yes	Yes
Year×Size	No	Yes	Yes	No	Yes	Yes
Size×Industry	No	No	Yes	No	No	Yes
R^2	0.881	0.884	0.995	0.703	0.705	0.998
N	1345	1345	1345	1345	1345	1345

Note: The original observation units are individuals, and they are collapsed to 33 industries×3 size categories (31-100, 101-300, 301+)×14 years cells. The dependent variables are the log of the total number of employees in each cell for Column 1-3 and the share of the total number of employees among cells (%), i.e., (the number of employees in each cell)$_{jts}$×100/(the total number of employees in the sample) for Columns 4-6, respectively. The number of employees in each industry, establishment size, year cell is recovered using the frequency weight that tells how many individuals an observation in the sample represent. The conventional heteroskedasticity-consistent standard errors are reported in parentheses. For these estimations, cell size-weighted regressions, using (the number of observations in cell j)/(the number of total observations/the number of cells) as a weight (Lee and Card (2008)), are performed. Out of the total number of the cells, i.e., 1386 cells, 41 cells are dropped from the sample because of no observation. Samples are restricted to establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water.

Table 4.8: Impact of Standard Hours on ln(Straight-time Hourly Wage Rate), 1989-2002

	(1)	(2)	(3)	(4)
Standard Hours	-0.0038	-0.0036	-0.0022	-0.0018
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Female	-0.4950	-0.4943	-0.4941	-0.4890
	(0.0008)	(0.0008)	(0.0008)	(0.0008)
Hours	-0.0006	-0.0006	-0.0007	-0.0009
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Hours^2/100	-0.0136	-0.0136	-0.0135	-0.0130
	(0.0004)	(0.0004)	(0.0004)	(0.0004)
Experience	0.0402	0.0401	0.0401	0.0403
	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Experience^2/100	-0.0605	-0.0604	-0.0604	-0.0606
	(0.0002)	(0.0002)	(0.0002)	(0.0002)
High School	0.0548	0.0546	0.0544	0.0555
	(0.0008)	(0.0008)	(0.0008)	(0.0008)
Junior College	0.2414	0.2409	0.2405	0.2424
	(0.0012)	(0.0012)	(0.0012)	(0.0012)
University	0.3056	0.3052	0.3052	0.3059
	(0.0011)	(0.0011)	(0.0011)	(0.0011)
Constant	2.6545	2.6532	2.5783	2.6013
	(0.0265)	(0.0342)	(0.0355)	(0.0371)
Industry	Yes	Yes	Yes	Yes
Trend	Yes	Yes	Yes	Yes
Size	Yes	Yes	Yes	Yes
Industry×Trend	No	Yes	Yes	Yes
Trend×Size	No	No	Yes	Yes
Size×Industry	No	No	No	Yes
R^2	0.658	0.659	0.659	0.664
N	6221155	6221155	6221155	6221155

Note: Observation units are individuals. The dependent variable is the log of straight-time hourly wage rate. Straight-time hourly wage rate is calculated as: (Scheduled Cash Earnings−Commuter allowance−Regular attendance allowance−Family allowance)/Scheduled Hours. Earnings data are corrected for inflation. Observation units are individuals. Standard errors, clustered at year×industry×size category (by one), are reported in parentheses. Samples are restricted to workers in establishments with more than 30 employees in 4 industries: mining, construction, manufacturing, and electricity, gas, heat supply, and water. The reference group for the education dummy variables is junior high school.

CHAPTER V

Conclusion

Using the theoretical and empirical tools from labor economics and public finance, this dissertation explores new evidence related to several topics that affect the Japanese labor market.

Chapter II examines how the 2004 tax reform in Japan affected labor supply of married women. This chapter uses quantile difference-in-difference estimations and the relatively new decomposition techniques, the Firpo, Fortin, and Lemieux decomposition and the DiNardo, Fortin, and Lemieux decomposition, to explore the effect this tax reform had on the distribution of work hours and incomes. These analyses found that the 2004 tax reform in Japan contributed to increasing work hours (and hence the incomes) of low-income married women. However, the tax reform resulted in the conventional kink at 1.03 million yen in the household budget line becoming more conspicuous, making it a more attractive income choice than before. Due to this change in the household budget line, the annual income of some of the medium- to high-income married women decreased to 1.03 million yen when they experienced both the 2004 tax reform and an increase in their husbands' incomes. As a result, the decrease in incomes caused by a large negative husband's income effect offset the increase in income among the low-income women, which resulted in a slight decrease in the average income among married women.

Chapter III provides a theoretical and empirical analysis of the effect of performance-based layoffs on wage rigidity in the context of performance pay. If the presence of the threat of performance-based layoffs is considered, high wages can be a device to induce workers to deliver high levels of work effort in both the preceding periods and in the current period. This gives firms an incentive to raise the future regular pay so as to maintain the workers' current efforts, which results in the downwardly rigid regular pay of experienced employees under the threat of performance-based layoffs. Together with the finding that layoffs are more likely to occur during recessions, the theoretical implication regarding the regular pay provides an explanation for the downward rigidity of regular pay during recessions. Although this is the result of the fixed component of wages, when we consider the performance-based component of wages, the outcomes are different. The firms are discouraged from maintaining their workers' current incentives at a higher level due to the lower value of productivity during recessions, which results in the bonus that increases or decreases wages proportionally to the output price. Thus, firms base wages less on workers' performance during recessions by paying lower bonuses. As a result, wages during recessions become both "downwardly rigid" and "rigid" (inflexible) with respect to performance. The empirical results obtained from the Japanese panel data set supported these theoretical implications. In particular, this type of explanation for wage rigidity can be applied to unionized workers whose layoff decisions are likely to be performance-based in countries such as Japan where performance-based pay has been widely employed.

Chapter IV utilizes the regression discontinuity approach to explore the impact of Japan's reduction of weekly standard working hours (from 48 hours/week in 1988 to 40 hours/week by 1997) on labor-market outcomes. This analysis exploits the discontinuity of standard working hours, based on establishment sizes, and the gradual phase-in of stricter standard working hours, based on establishment size and industry, to estimate the policy impacts on hours worked, wages earned, and employment. The analysis applying a regression discontinuity design revealed that the reduction of standard working hours modestly reduced the number of hours worked. About 31 percent of the reduction in the number of hours worked between 1989 and 2002 is attributable to the policy change. This reduction was not associated with a decrease of either monthly wage or bonus payment. As a result, overall labor costs per hour increased, and we do not see any evidence of job creation. Thus, even in Japan, where wages are determined flexibly, we observed evidence that is similar to what has been found in European countries, such as Germany and France, where pay is largely determined through union-employer bargaining.

These three chapters provide new information about the Japanese labor market that is useful for evaluating policies concerning spousal exemptions, wages, and standard working hours. Although the focus of this dissertation is on the Japanese labor market, the discussion and the approach adopted in these chapters can also be extended to cases in other countries. It is hoped that, in future studies, the approaches applied here will also add to an understanding of the labor market in other countries.

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