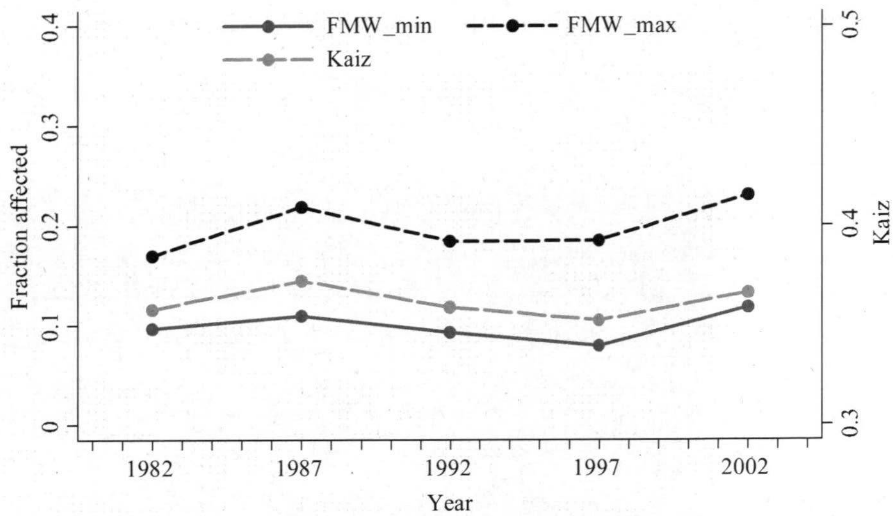

Figure 2. Time series of the fraction of minimum-wage workers and Kaiz index

analysis in the previous section reveals that a large percentage of young, elderly and married women receive wages near the minimum wage. To restrict our analysis to the demographic groups that are presumably heavily affected by the minimum wage, we focus on the following seven categories of workers: male teenagers (15–19 years old), male young adults (20–24 years old), male elderly (over 60 years old), female teenagers, female young adults, female elderly and middle-aged (25–59 years old) married women.

To establish the causal effect of a minimum-wage hike on employment, we examine the relation between the fraction of workers who are affected by the minimum-wage hike and the change in employment using prefecture-level variation. More specifically, we calculate the fraction of workers who are affected by a minimum-wage hike (fraction affected (*FA*)), in other words, workers whose wage is above the current minimum wage but below the revised minimum wage. This *FA* takes different values across prefectures, even if the amounts of the minimum-wage hike are homogeneous across prefectures, because the minimum-wage hike is more prevalent in the low wage prefectures, such as Okinawa, than in high wage prefectures like Tokyo. Card (1992) originally proposed using *FA* to examine the effect of the federal minimum-wage increase on employment, because the uniform increase of the minimum wage across states has different impacts across states depending on each state's wage distribution.

The following model is estimated to examine the effect of *FA* on the change in the employment rate of worker group k in prefecture i between years t and $t - 5$:

$$\Delta E^k_{i,t} = \beta^k_0 + \beta^k_1 FA_{i,t-5} + \beta^k_2 \Delta AW_{i,t} + \beta^k \Delta X^k_{i,t} + Y_t \gamma^k + e^k_{i,t}, \qquad (1)$$

where $\Delta E^k_{i,t}$ is the change in the employment rate of category k in prefecture i between years t and $t - 5$; $FA_{i,t-5}$ is the fraction of workers who are affected

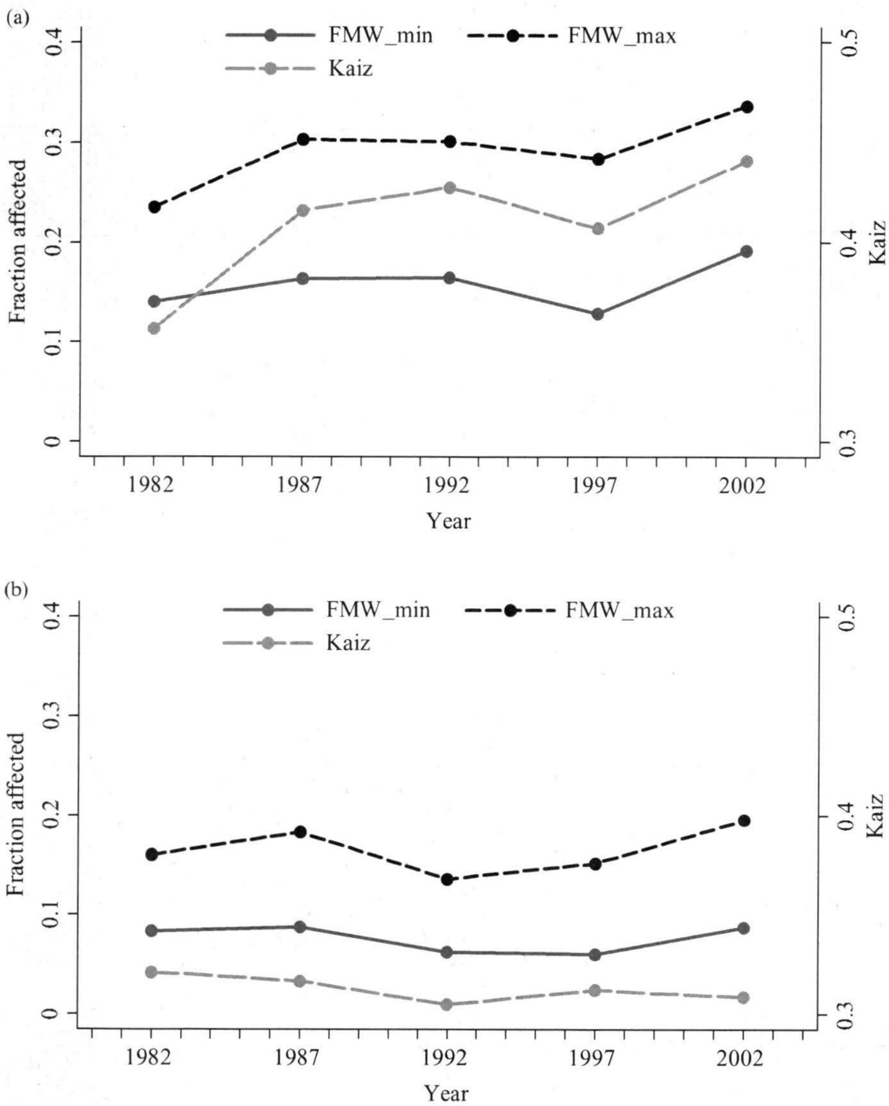

Figure 3. (a) Time series of fraction of minimum-wage workers and Kaiz index: Okinawa. (b) Time series of fraction of minimum-wage workers and Kaiz index: Tokyo

by the minimum-wage hike between t and $t-5$; $\Delta AW_{i,t}$ is the change in the average wage of middle-aged (25–59 years old) male workers; $X_{i,t}^{k}$ is a set of explanatory variables (the proportion of the population in the relevant categories and the unemployment rate of middle-aged males); and Y_t is a set of dummy variables for year t. We adopt two definitions of *FA*: *FA_min* and *FA_max*. *FA_min* is the fraction of workers who satisfy the condition,

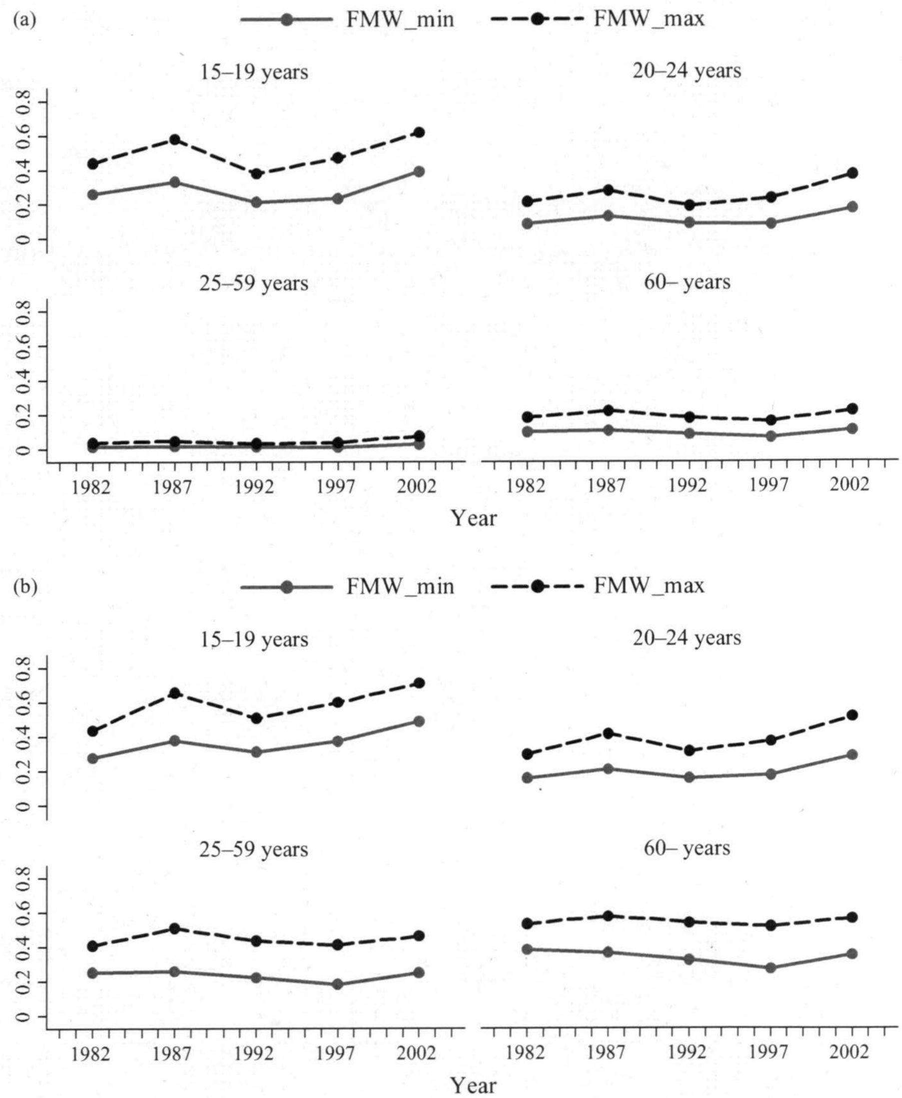

Figure 4. (a) Minimum-wage fraction maximum among junior-high school and high-school graduates: men. (b) Minimum-wage fraction maximum among junior-high school and high-school graduates: women

$MWIncome^{min}_{it-5} \leq Income_{it-5} < MWincome^{min}_{it}$. The minimum of the minimum-wage annual income, $MWIncome^{min}_{it}$, is calculated using the hourly minimum wage and the minimum value of annual work hours. *FA_max* is similarly defined based on the maximum of the minimum-wage annual income. Card and Krueger (1995) claim that an increase in the average wage relative to output price causes firms to cut employment through a reduction in output production.

Table 6. Summary statistics for the employment analysis sample

	Observations	Mean	Standard deviation	Minimum	Maximum
Fraction affected minimum (FA_min)	188	0.039	0.027	0.003	0.113
Fraction affected maximum (FA_max)	188	0.055	0.036	0.003	0.138
Employment rate					
Male, 15–19	235	0.162	0.028	0.097	0.240
Male, 20–24	235	0.751	0.072	0.552	0.895
Male, 60–	235	0.490	0.052	0.341	0.635
Female, 15–19	235	0.155	0.030	0.092	0.239
Female, 20–24	235	0.721	0.054	0.561	0.863
Female, 60–	235	0.239	0.043	0.144	0.351
Female, 25–29, married	235	0.659	0.104	0.415	0.858
Change in employment rate					
Male, 15–19	188	−0.005	0.028	−0.083	0.063
Male, 20–24	188	−0.021	0.044	−0.134	0.095
Male, 60–	188	−0.019	0.036	−0.113	0.077
Female, 15–19	188	−0.005	0.028	−0.079	0.065
Female, 20–24	188	−0.007	0.039	−0.108	0.131
Female, 60–	188	−0.007	0.021	−0.054	0.040
Female, 25–29, married	188	0.051	0.071	−0.032	0.298
Share of population					
Male, 15–19	235	0.040	0.006	0.028	0.061
Male, 20–24	235	0.034	0.008	0.020	0.059
Male, 60–	235	0.112	0.026	0.053	0.165
Female, 15–19	235	0.039	0.006	0.026	0.057
Female, 20–24	235	0.037	0.007	0.023	0.053
Female, 60–	235	0.150	0.035	0.071	0.231
Female, 25–29, married	235	0.216	0.079	0.047	0.308
Change in share of population					
Male, 15–19	188	−0.002	0.005	−0.017	0.012
Male, 20–24	188	−0.001	0.006	−0.018	0.011
Male, 60–	188	0.016	0.006	0.003	0.032
Female, 15–19	188	−0.002	0.005	−0.017	0.012
Female, 20–24	188	−0.002	0.005	−0.017	0.010
Female, 60–	188	0.019	0.007	0.004	0.037
Female, 25–29, married	188	−0.054	0.061	−0.180	−0.006
Change in average wage	188	0.318	0.443	−0.575	1.137
Unemployment rate					
Change in unemployment rate	188	0.006	0.011	−0.016	0.030

We control this scale effect by including the change in the average adult male's wage ($\Delta AW_{i,t}$) and the year dummy variables as explanatory variables. The estimation method is weighted least squares, with the weight being the inverse of the standard error of the dependent variable.

Table 6 reports the descriptive statistics for the regression analysis sample. The *FA* has a mean of approximately 4–6%, with sufficient variation across years and prefectures. Table 7 displays the fraction of workers affected by a minimum-wage hike for a 5-year period, and we can confirm that *FA* tends to be higher in rural, low-wage areas, such as Aomori and Okinawa, than in high-wage areas such as Tokyo. It also displays a variation over time within a prefecture.

Table 7. Fraction of workers affected by minimum-wage hike (%)

Year	1982–1987	1997–2002	Year	1982–1987	1997–2002
Hokkaido	3.18	0.81	Shiga	2.48	0.27
Aomori	6.22	4.15	Kyoto	6.19	1.34
Iwate	6.55	4.47	Osaka	5.32	0.98
Miyagi	4.63	0.52	Hyogo	5.76	1.11
Akita	6.27	3.46	Nara	2.20	0.26
Yamagata	6.72	3.07	Wakayama	2.93	0.53
Fukushima	5.54	1.37	Tottori	5.25	1.50
Ibaraki	2.27	0.36	Shimane	5.21	1.95
Tochigi	3.11	0.37	Okayama	2.60	0.53
Gunma	3.00	0.38	Hiroshima	2.72	0.52
Saitama	4.59	0.77	Yamaguchi	3.28	0.62
Chiba	4.74	0.75	Tokushima	5.19	0.65
Tokyo	2.98	0.71	Kagawa	4.36	0.80
Kanagawa	3.65	0.70	Ehime	5.15	0.60
Niigata	4.50	0.50	Kochi	5.56	0.51
Toyama	3.18	0.38	Fukuoka	2.95	0.68
Ishikawa	3.76	0.63	Saga	5.80	3.66
Fukui	3.98	0.50	Nagasaki	5.27	4.04
Yamanashi	2.92	0.60	Kumamoto	6.92	3.55
Nagano	2.74	0.51	Oita	5.02	3.72
Gifu	3.90	1.40	Miyazaki	6.87	4.34
Shizuoka	2.66	1.12	Kagoshima	5.88	4.03
Aichi	6.27	0.94	Okinawa	6.09	4.54
Mie	3.17	1.18			

The fraction of workers affected by the minimum-wage hike (*FA*) is defined as the number of workers affected by the minimum-wage hike (i.e. $MWIncome_{it}^{\min} \leq Income_{it} < MWincome_{it+5}^{\min}$) divided by the number of employed workers.

To illustrate the regression results for subsamples, Figure 5(a) and (b) visually reports the regression results for male and female teenagers between 1997 and 2002. The fraction of workers affected by the minimum-wage hike (*FA*) on the horizontal axis is the residual from the regression of *FA* on other explanatory variables. Both figures confirm a weakly negative relation between regression-adjusted *FA* and the change in the employment rate over the 5-year period.

Tables 8–10 report the comprehensive results of regressions by demographic groups. Columns (1) and (2) of Table 8 indicate that a higher *FA* results in a reduction of the employment rate among male teenagers. A 1% increase in *FA* reduces the employment rate by approximately 0.2 percentage points. Considering that the average employment rate of male teenagers is approximately 16% (Table 6) and the average *FA* is approximately 5%, the magnitude is small. Columns (3) and (4) indicate the negative impact of a higher *FA* on the employment rate of young adults, but the coefficients are not statistically significant. A higher *FA* does not affect the employment rate of elderly people.

Table 9 reports the results for women. The estimated coefficients are consistently negative and some of the coefficients are statistically significant. However, all results depend on the choice of *FA* variable, and it is rather

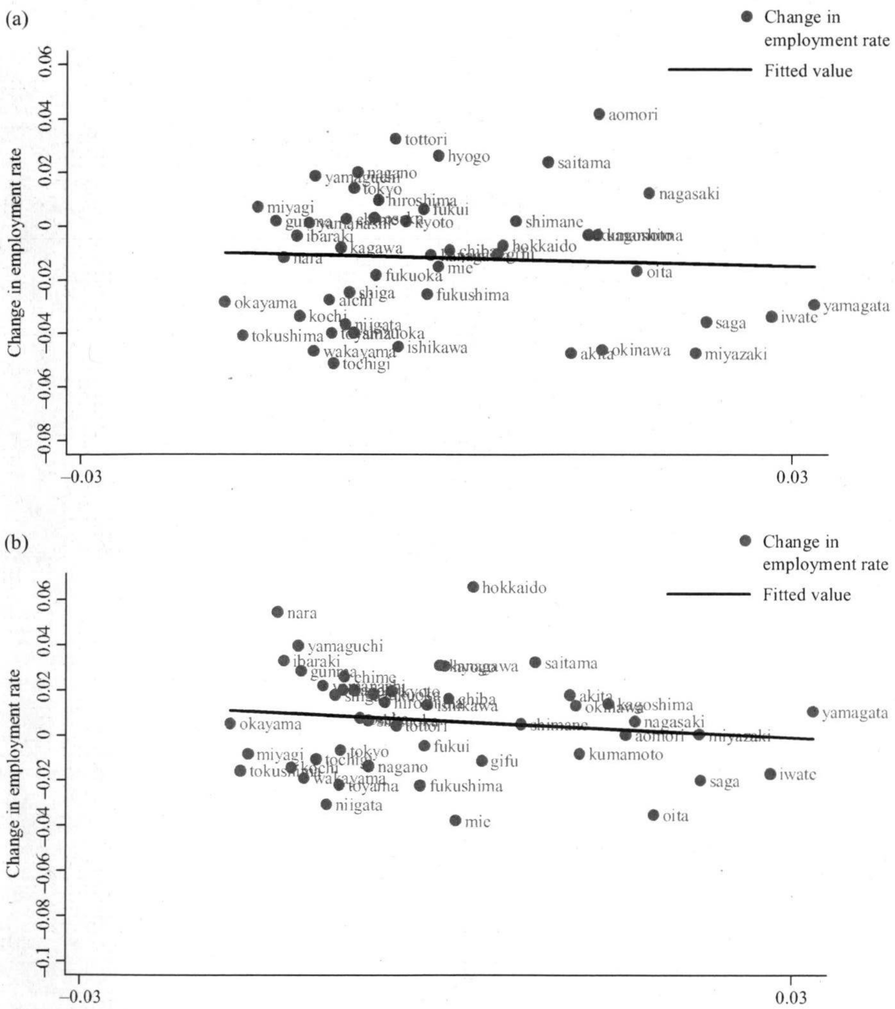

Figure 5. (a) Fraction of workers affected by the minimum-wage hike and change of employment rate, 1997–2002: male 15–19 year olds. (b) Fraction of workers affected by the minimum-wage hike and change of employment rate, 1997–2002: female 15–19 year olds

difficult to see a clear relation between *FA* and the employment rate from this table. Table 10 reports the results for married women in the 25–59 year age group who are likely to be employed as part-time workers. The estimated coefficients clearly indicate that the minimum-wage hike reduces employment among members of this group. A 1% increase in *FA* decreases the employment rate by 0.4–0.8 percentage points. Considering that the average of *FA* is approximately 5% and the employment rate of this group is roughly 66%, the magnitude is rather moderate.

Table 8. Fraction of workers affected by the minimum-wage hike and the change of employment rate, male dependent variable: change in employment rate

	(1)	(2)	(3)	(4)	(5)	(6)
Age group	15–19		20–25		60–	
Fraction affected (*FA*)	*FA_min*	*FA_max*	*FA_min*	*FA_max*	*FA_min*	*FA_max*
Fraction affected	−0.23	−0.20	−0.26	−0.26	−0.02	−0.00
	(0.11)	(0.12)	(0.15)	(0.14)	(0.11)	(0.11)
Change in population share	0.43	0.08	−1.31	−1.34	0.14	0.15
	(0.60)	(0.58)	(0.84)	(0.82)	(0.36)	(0.36)
Change in average wage	0.02	0.02	0.02	0.02	0.02	0.02
	(0.02)	(0.02)	(0.03)	(0.02)	(0.01)	(0.01)
Change in unemployment rate	−0.72	−0.70	−0.33	−0.27	−0.89	−0.88
	(0.45)	(0.47)	(0.71)	(0.70)	(0.39)	(0.38)
Constant	0.02	0.02	0.01	0.02	−0.04	−0.04
	(0.02)	(0.02)	(0.03)	(0.03)	(0.01)	(0.01)
Dummy for year	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*
Observations	188	188	188	188	188	188
R^2	0.35	0.35	0.34	0.35	0.69	0.69

Standard errors are reported in parentheses. The inverse of the estimated variance of the dependent variable is used as a weight for the weighted least squares estimation.

Table 9. Female dependent variable: change in employment rate

	(1)	(2)	(3)	(4)	(5)	(6)
Age group	15–19		20–25		60–	
Fraction affected (*FA*)	*FA_min*	*FA_max*	*FA_min*	*FA_max*	*FA_min*	*FA_max*
Fraction affected	−0.37	−0.13	−0.09	−0.23	−0.10	−0.11
	(0.10)	(0.11)	(0.12)	(0.12)	(0.10)	(0.08)
Change in population share	0.49	0.16	−1.43	−1.62	0.01	−0.01
	(0.64)	(0.64)	(0.79)	(0.80)	(0.22)	(0.22)
Change in average wage	−0.02	−0.01	−0.00	−0.01	−0.01	−0.01
	(0.01)	(0.02)	(0.02)	(0.02)	(0.01)	(0.01)
Change in unemployment rate	−0.25	−0.18	0.04	0.05	−0.10	−0.07
	(0.59)	(0.61)	(0.54)	(0.54)	(0.29)	(0.29)
Constant	0.01	0.01	0.03	0.05	−0.03	−0.03
	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.01)
Dummy for year	Y	Y	Y	Y	Y	Y
Observations	188	188	188	188	188	188
R^2	0.22	0.18	0.50	0.51	0.44	0.45

6. MINIMUM WAGE'S EFFECT ON THE CHOICES OF YOUTH AGED 16–17 YEARS

The effects of the minimum wage on the outcomes of youths are not limited to employment but also include schooling decisions. Cunningham (1981) and Ehrenberg and Marcus (1982) examine this subject by arguing that the minimum wage's effect on schooling decisions is complex because it affects both the opportunity cost of schooling and the return to schooling. The opportunity cost of attending high school is the wage that high-school-age youth expect to receive in the labour market. If the employment-reduction effect is limited, then

Table 10. Female, 25–59 years, married dependent variable: change in employment rate

	(1)	(2)
Fraction affected (*FA*)	*FA_min*	*FA_max*
Fraction affected	−0.43	−0.82
	(0.13)	(0.13)
Change in population share	0.49	0.81
	(0.60)	(0.58)
Change in average wage	−0.03	−0.07
	(0.02)	(0.02)
Change in unemployment rate	−1.32	−1.38
	(0.78)	(0.75)
Constant	0.11	0.18
	(0.02)	(0.03)
Dummy for year	Y	Y
Observations	188	188
R^2	0.76	0.78

a minimum-wage hike might increase the expected wage; in contrast, with a significant employment-reduction effect, the expected wage might decrease. The minimum wage also affects the return to schooling through a modification of the wage structure. If the minimum wage reduces the labour demand for low-skilled workers, the demand for skilled workers might be increased through a substitution effect, which may lead to a higher return to education. In our context, an increased minimum wage may reduce the demand for part-time jobs among high-school students or high-school dropouts, and increase the demand for high-school graduates. If a minimum-wage increase does not reduce employment, however, even high-school dropouts can earn a higher wage and are, therefore, less likely to return to education because of the minimum-wage increase. Overall, the minimum wage has a complex effect on high-school age youths' schooling decisions through the opportunity cost of attending high school and the relative value of returning to high school to receive a diploma.

In addition, Ehrenberg and Marcus (1982) emphasize the importance of household liquidity constraints. If a youth belongs to a household that is liquidity-constrained, he or she might finance their studies by working part time. In this scenario, increasing the minimum wage might increase employment as well as school attendance. These theoretical complexities suggest the importance of treating youths' decisions about schooling and employment jointly.

This section investigates the minimum wage's effect on young people's choice between employment and school enrolment. We focus on high-school age individuals, 16 and 17 years old, and classify them into the following exclusive categories to capture the possible joint decision of schooling and employment: (i) not in school and not employed; (ii) in school and not employed; (iii) in school and employed; and (iv) not in school and employed. The empirical model used in this section is the aggregate data version of Neumark and Wascher (1995), which is the same model that was used in the previous section except for the dependent variables. As in the employment analysis, we exploit the prefectural variation of the fraction of workers affected by the minimum-wage hike. The estimation model is described as follows:

$$\Delta p_{i,t}^{k} = \beta_0^k + \beta_1^k FA_{it-1} + \beta_2^k \Delta AW_{j,t} + \Delta X_{j,t}^{k}\beta^k + Y_t\gamma^k + \varepsilon_{j,t}^{k}, \tag{2}$$

where $p_{i,t}^{k}$ is the population of each category k divided by the relevant-aged population in prefecture i and year t; FA is the fraction of workers affected by the minimum-wage hike in the 5-year period; and AW, X and the other variables are the same as in the previous section. The estimation method is weighted least squares, in which the inverses of the standard errors of the dependent variables are used as weights. The above model is estimated based on a sample that includes only young people 16 and 17 years old. Some 15 and 18 year olds are also high-school age depending on the quarter of their birth, but the birth quarter is not recorded before the 1997 survey; therefore, we do not include 15 and 18 year olds in the sample.

Table 11 reports the summary statistics of the analysis sample for the choice of youths. Approximately 93% of 16–17 year-old youths are in school without employment, while 2.5% are in school and have part-time jobs. Of the remaining 5%, 2.5% are out of school and employed, whereas 1.7% are idle.

Table 12 presents the estimation results. For all estimations, the changes in proportions of youth who are in school but not employed are treated as the base category. Panel A reports the regression results using the fraction of workers who are affected by the minimum-wage changes as an independent variable. The change in the fraction of workers who are affected by a minimum-wage hike is defined by a range as in the previous analysis, because the definition of minimum-wage workers depends on whether we use the minimum hours of work or the maximum hours of work to define minimum-wage annual income. Columns (1) and (2) report the effect of the fraction affected on the change in the fraction of

Table 11. Summary statistics of analysis sample for high school age choice

	Observations	Mean	Standard deviation	Minimum	Maximum
In school, not employed	235	0.933	0.022	0.851	0.983
Δ In school, not employed	188	−0.919	0.024	−0.970	−0.847
Out school, not employed	235	0.017	0.008	0.002	0.058
Δ Out school, not employed	188	0.001	0.008	−0.024	0.028
In school, employed	235	0.025	0.018	0.002	0.100
Δ In school, employed	188	−0.004	0.014	−0.054	0.045
Out school, employed	235	0.025	0.011	0.006	0.071
Δ Out school, employed	188	−0.011	0.013	−0.062	0.023
Unemployment rate	235	0.023	0.012	0.006	0.070
Δ Unemployment rate	188	0.006	0.011	−0.016	0.030
Population share of age 16–17 years	235	0.047	0.010	0.028	0.075
Δ Population share of age 16–17 years	188	−0.002	0.005	−0.011	0.011
Average wage	235	2.950	0.718	1.637	4.613
Change in average wage	188	0.318	0.443	−0.575	1.137

Table 12. Effect of minimum wage on choices among high-school-age teenagers

	(1)	(2)	(3)	(4)	(5)	(6)
5 year change in state	Out of school not employed		In school employed		Out of school employed	
Fraction affected (*FA*) variable	*FA_min*	*FA_max*	*FA_min*	*FA_max*	*FA_min*	*FA_max*
Fraction affected	−0.00	0.04	0.27	0.30	0.11	0.13
	(0.03)	(0.03)	(0.07)	(0.06)	(0.05)	(0.04)
Change in population share	−0.17	−0.15	−0.08	0.38	0.04	0.24
	(0.26)	(0.25)	(0.35)	(0.32)	(0.34)	(0.33)
Change in average wage	0.00	0.01	0.01	0.01	−0.02	−0.01
	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)
Change in unemployment rate	0.07	0.07	0.20	0.17	0.01	0.00
	(0.14)	(0.14)	(0.21)	(0.22)	(0.20)	(0.20)
Constant	0.00	−0.00	−0.01	−0.03	−0.01	−0.02
	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.01)
Year dummy	*Y*	*Y*	*Y*	*Y*	*Y*	*Y*
Observations	188	188	188	188	188	188
R^2	0.11	0.11	0.18	0.22	0.17	0.18

Standard errors are in parentheses.

workers who are out of school and not employed. Regardless of the measurement of the fraction of workers affected by the minimum-wage hike, the minimum-wage hike did not significantly change the fraction of youths in this category. Columns (3) and (4) show that the fraction of youths who were employed while attending school increased when the minimum-wage hike affected more workers. To quantify its impact, let us think about the counterfactual situation of Okinawa not increasing its minimum wage between 1997 and 2002. The fraction affected was zero instead of the actual 4.54% reported in Table 7. The coefficient estimate is 0.30 (column (4)); therefore, the estimated effect on the increase in the probability of the choice to stay in school and be employed decreases by 1.359%. This is a large effect, given that only 2.5% of youths work while attending school (Table 11). Columns (5) and (6) report that the increase in the fraction of workers affected by a minimum-wage hike increases the probability of youths being out of school and employed. The estimated magnitude is approximately one-third of the effect on the choice of being in school and employed.

Overall, the higher the fraction of workers who are affected by the minimum-wage hike, the higher the fraction of youths in employment, either while attending high school or dropping out of high school. In particular, estimates indicate that a higher minimum wage relative to the regional wage distribution encourages high-school students to work as part-time workers. An examination of the long-term consequence of the minimum wage on skill formation and labour-market outcomes, as conducted by Neumark and Nizalova (2007) for the USA, is left for future research.

7. CONCLUSION

The present paper aimed to answer the following three questions based on Japanese data. First, do minimum-wage workers belong to low-income households? Second, what are the minimum wage's effects on employment among populations that are marginally attached to the labour market? Third, what are the minimum wage's effects on the employment and schooling choices of 16 and 17-year-olds?

Five waves of the ESS, from 1982, 1987, 1992, 1997 and 2002, were used to calculate the percentage of workers earning the minimum wage. The annual earnings and work-hours reported in range brackets enabled us to calculate the minimum and maximum number of minimum-wage workers. The minimum fraction of minimum-wage workers was approximately 3–4% between 1982 and 2002, while the maximum fraction was 6–10% during the same period for men. The corresponding figures for women were approximately 22 and 36–40%, respectively.

Female workers, workers with low educational backgrounds, workers in rural areas, and workers in retail or food industries were more likely than others to be employed at the minimum-wage level. Although workers with weaker labour-market characteristics were more likely to work at the minimum wage, those minimum-wage workers did not necessarily belong to disadvantaged households. In 2002, only around 20% of minimum-wage workers were heads-of-households and approximately half of them belonged to households with annual incomes of 5 million yen or more as a non-household head. This result confirms the finding of Tachibanaki and Urakawa (2007) that was based on a much smaller dataset.

The analysis results exploiting the cross-prefecture heterogeneity of the fraction of workers affected by a minimum-wage hike indicates that a minimum-wage hike reduces the employment of male teenagers and middle-aged married women. Its magnitude for male teenagers is estimated to be small, but that for middle-aged married women is moderate.

The analysis of the minimum wage's effect on the choices of high-school age youths shows us that an increase in the minimum wage encourages employment for young people whether or not they are attending high school. Although a higher minimum wage relative to regional wage decreases teenage male employment, it positively affects the employment of high-school age youths. The combination of these two pieces of evidence implies that a minimum-wage increase relative to the regional wage distribution will have a strong negative impact on male high-school graduates. The minimum wage's long-term effect on the skill formation and long-term labour-market outcomes of young people has yet to be examined. Overall, a policy supporting a rising minimum wage would not be helpful in alleviating poverty in Japan because such policy is not well targeted toward poor households and would reduce the employment of less-skilled workers. Adopting a more direct antipoverty policy, such as an earned-income tax credit, would be one viable policy alternative.[11]

[11] Recent studies by Leigh (2009) and Rothstein (2008), however, report that the expansion of earned-income tax credits suppresses the wages of low-skilled workers because it encourages their labor supply.

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Gender Salary Differences in Economics Departments in Japan

Ana Maria Takahashi[1]

Department of Economics

University of Utah

JOB MARKET PAPER

Abstract: By using unique data about academic economists in Japanese universities, we conduct the first detailed study of gender salary differences within Japanese academia. Despite the common belief among Japanese academics that there cannot be a gender salary gap within the Japanese academia, our empirical results show that female academic economists earn 7% less than comparable males, after controlling for rank and detailed personal, job, institutional and human capital characteristics. The coefficient for the female dummy has almost the same value, regardless of whether rank is included or excluded from the salary equation, suggesting that there is a significant gender salary gap *within* each rank, but there are no gender differences in rank attainment. Our results contrast with findings from previous studies in the US and in the UK where most of the gender salary differences stem from rank attainment differences. We provide possible explanations for why our results are different. Refereed articles, the most commonly accepted measures of productivity, have no statistically significant effect on salary. The fixed-term employment is associated with 24% lower annual salary and private university has a salary premium of 16%.

1 Introduction

There is now a large body of literature on labor market discrimination. Labor market discrimination is a situation in which an otherwise identical person is treated differently by virtue of that person's gender or race (Heckman 1998). A large part of the literature on labor market discrimination focuses on the problem of gender salary differences. This paper seeks to explore the problem of gender salary differences among Japanese academic economists by using an original data set collected via a mail survey.

A great number of studies have already examined gender salary differences within the academic profession in the US and the UK (Broder 1993; McNabb and Wass 1997; Ward

[1]Address: 949-7302, Niigata, Minamiuonuma, Urasa 1014-1, Flat 2007 A, Cell: 080-3397-3934; Tel/Fax: 025-777-5127, Japan; piciorea@economics.utah.edu

2001; Ginther and Hayes 2003; Blackaby et al. 2005). Most of these studies found that much of the gender salary differences in academia stems from the differences in rank attainment between males and females. By using data from Scottish universities, Ward (2001) finds that female academics receive 7% lower salary than males, after controlling for various productive and institutional characteristics, but without controlling for rank. When rank is controlled for, the estimated salary gap is reduced to 3.2% and becomes statistically insignificant. By using data from the Surveys of Doctorate Recipients in the US, Ginther and Hayes (2003) find negligible gender salary differences within rank, but substantial gender differences in rank attainment. Blackaby et al. (2005) use data from British universities, and find that the gender salary gap nearly halves when rank is controlled for.

In contrast to the abundance of literature in the US and the UK, there have been few studies about the gender salary gap within Japanese academia, despite a growing public interest in gender equality in Japan. In 1999, Japanese government enacted *the Basic Law for Gender Equal Society.* Consequently, in 2000, the Association of National Universities set out an *Action Plan* stipulating that each national university should increase the proportion of female academics to 20% by 2010. In 2008, the Ministry of Education, Sports, Science and Technology (MEXT) announced that it would provide 6 million yen in support, to selected universities, for each female academic hired[2]. Despite such an interest in achieving gender equality in academia, little is known about gender differences within Japanese academia. A study conducted by the EPMEWSE[3] (2008:18) shows that average salaries differ between male and female Japanese academics, after controlling for age. However, the study does not control for any other differences in characteristics between male and female academics. Therefore, a detailed study of gender salary differences is called for.

We conduct the first detailed study of gender salary differences within Japanese academia

[2]Asahi News, October 5, 2008.

[3]Japan Inter-Society Liaison Association Committee for Promoting Equal Participation of Men and Women in Science and Engineering (annex.jsap.or.jp/renrakukai/2007enquete/h19enquete_report_v2.pdf).

by using a data set that we collected via a mail survey administered in 2008. Our study focuses on academic economists. For each academic in our sample, our data contain detailed productivity characteristics, such as publication record and the amount of research grants, detailed information about career history, and detailed institutional characteristics.

There are at least two reasons why the study of gender salary gap within the Japanese academic labor market is of interest to researchers. First, there is a common belief among Japanese academics that there cannot be a gender salary gap within Japanese academia. In Japan, academics' salaries are determined by a rigid payment scheme, called the 'salary table', with distinct pay scales for each rank. It is commonly believed that age, experience, and education levels are the only criteria that determine the payment level. It is also believed that promotion is automatic, based on age and experience. If these beliefs were true, then salary would be a deterministic function of age, experience, and education, leaving little room for a gender salary gap. However, there have been no empirical investigations into whether such beliefs are in fact true. Therefore, whether or not there are gender salary differences within Japanese academia is still an open empirical question.

Second; there are important institutional differences between the US and Japanese academia; differences that may cause a very different pattern of gender salary gap in Japan compared to that in the US. In US academia, both salary determination and promotion decisions are departmental matters. Departments manage the budget for salaries and decide remuneration. Promotion decisions are also made at the department level, typically by a faculty committee. In contrast, in Japanese academia, only promotions decisions are made at the department level, typically by a faculty committee. The salary for each academic is typically determined by the university's personnel division (*jinji-bu*), and not by the department[4]. Thus, in Japanese academia, salary and promotion are determined by two different

[4]According to interviews conducted by the author. We interviewed several academics, including representatives of the Association of Private Universities of Japan (*Nihon Shiritsu Daigaku Kyoukai*), and of the Faculty and Staff Union of Japanese Universities (*Zenkouku Daigaku Kosen Kyoshokuin Kumiai*).

entities. This indicates that salary determination and promotion decisions are based on information with different levels of accuracy. Since salary is not decided at the department level, salary may be determined based on less accurate information about academic productivity. For example, the personnel division may not receive the full set of information about academic performance. Moreover, as the personnel divisions consist of non-academics, they may not be capable of assessing some aspects of productivity such as the quality of publications. On the contrary, promotion may be determined based on more accurate information, since it is determined at the department level by fellow academics. Information about academic productivity would be readily available to faculty, and research quality can be more accurately evaluated by academics than non-academics.

Since the accuracy of information is a central issue in theories of statistical discrimination, it is possible that statistical discrimination or prejudicial beliefs about female academics' productivity, if these exist, would manifest in terms of gender salary differences rather than in terms of promotion differences. Therefore, unlike in the US and the UK where promotion gaps are the main source of gender salary gap in academia, we may see a very different pattern of gender salary gap within Japanese academia.

We organize our paper as follows. Section 2 presents main theories of discrimination. Section 3 discusses relevant empirical literature. Section 4 briefly outlines the background information. Section 5 describes the empirical methodology. Section 6 presents the data and Section 7 presents the variables used for estimation. Section 8 contains the estimation results and Section 9 discusses the sample selection bias problem. Section 10 includes a discussion of the results and Section 11 concludes.

2 Theories of discrimination

2.1 Statistical discrimination theory

Theoretical economic literature advanced two major employer discrimination theories to explain different average salaries for members of two presumably equally productive groups: statistical discrimination theory and taste-based discrimination theory.

Phelps (1972) developed a model which assumes that the average productivity of females is lower than males'[5]. In addition to the observable productivity characteristics of an individual, employers use the individual's association with a particular group to predict the productivity of a worker. A female may receive lower salary than a male of the same productive characteristics because employers use the average characteristics of the female group to predict the female workers' productivity. This model has been criticized by Aigner and Cain (1977) who argued that this model is not a description of discrimination. In Phelps' model, the conclusion depends on the assumption that females and males have different average productivity. Once this assumption is removed, the average male and female salary will be the same. This is because, males with higher test scores (that is, higher ability) will earn higher salaries than females with the same scores, while males with lower test scores (that is, lower ability) will earn less than females with the same scores, thus completely offsetting the gender salary differences. This criticism led to a new statistical discrimination model.

Aigner and Cain (1977) developed a model in which the average productivity of males and females is the same, but an observed variable that indicates the ability of an individual (that is, test scores) is less informative for females. By introducing the assumption that the employer is risk averse, they showed that females would receive lower wages even when male' and female' average productivity is the same. In this model, females receive lower wages because the observed variable describing productivity is less informative for them and

[5]Throughout this section, we use 'females' synonymously with 'the disadvantaged group'.

because the risk averse employer needs to be compensated for taking the risk of employing workers whose ability is less certain.

In Arrow's model (1973), employers have prejudicial beliefs (preconceived ideas) that female workers are less productive than males. Arrow shows that, if it is costly to learn the workers' ability, prejudicial beliefs cause the average female wage to be lower than male'. Arrow further shows that such prejudicial beliefs affect workers behavior such that females invest less in human capital, thus perpetuating the beliefs.

Lundberg and Startz (1983) expanded Phelps' (1992) model by endogenizing human capital investment decision. In their model, employers only observe a test score which is a noisy indicator of a worker's marginal product. Furthermore, the test score is assumed to be less informative for females. Then, as in Phelps' (1972) model, the wage-test score profile is flatter for females, thus returns to human capital investments are lower for females. Workers invest in human capital to the point where marginal cost of investment just balances the incremental benefit from that investment. Thus, females invest less in human capital due to lower returns to human capital investments, causing the average female's wage to be less than male'.

2.2 Taste-based discrimination theory

Taste-based discrimination theory was first articulated by Becker (1957) and Arrow (1971). In this theory, discrimination arises from employer's distaste against working with a particular group of people such as females. If the employer hires those he distastes, he gains less utility. In a utility maximization framework, Becker shows that even when females and males are perfect substitutes in production, the short-run equilibrium wage for females will be lower than that of males'. This model also predicts that, in a competitive market, discriminatory firms would disappear. Non-discriminatory employers earn higher profits than discriminatory employers, since they are willing to hire workers who have been discriminated against and whose wages are lower. Thus, non-discriminatory firms are able to purchase discriminatory firms, indicating that there will be no discriminatory firms in the long-run.

However, due to persistent discrimination in the long-run, Becker's model became criticized. Several authors modified Becker's model to explain that discriminatory firms can in fact survive in the long run. Goldberg (1982) modified Becker's model to incorporate nepotism toward males, showing that nepotistic firms can survive in the long-run. In Goldberg's model, the nepotistic employer obtains extra non-pecuniary gains from hiring males. In a utility maximization framework, the short-run equilibrium wage is lower for females than for males. However, the price the neutral firm has to pay to purchase the nepotistic firm is not equal to nepotistic firm's profit alone, but profit plus the non-pecuniary gain the nepotistic firm enjoys. Although neutral firms have higher profits, their profits are not high enough to cover for the additional non-pecuniary gain. Therefore, nepotistic firms would survive in the long-run.

Black (1995) developed a model of employer discrimination by using a job search model in which the existence of discriminatory firms increases the search cost incurred by females. The presence of such a search cost gives firms monopsonistic power. Thus, firms by exploiting their monopsonistic power, offer lower wages for females. Rosen (2003) also uses a job search framework. Discriminatory firms pay lower wages for females as compared to neutral firms, but their hiring decision is suboptimal (non-profit maximizing). When the discrimination coefficient is not 'too high', discriminatory firms attain higher profit than neutral firms since the gains from lower wages outweigh the losses from suboptimal hiring. However, discriminatory firms have lower utility than neutral firms, indicating that discriminatory firms will disappear in the long run due to takeover. By introducing the separation between ownership and management, Rosen shows that discriminatory firms would have both higher profit and higher utility, thus they would survive in the long run.

3 Previous empirical literature

There is abundant literature that investigates the gender salary gap within academia in the US and the UK. Since the 1970s, studies on the gender salary gap in the academic labor market have been conducted by (Tuckman et al. 1977; Hasen et al. 1978; Hirsch and Leppel 1982). We summarize below more recent work. Broder (1993), by using a simultaneous equation model, estimates a salary equation using a sample of 362 male and 30 female academic economists in the US. The estimated coefficient for the female dummy indicates that females salaries are between 5% to 8% lower than males', after controlling for experience, productivity, institutional characteristics, and rank. Salary differences are more pronounced for the older cohorts. In addition, when she conducts a salary decomposition using the Oaxaca decomposition method (Oaxaca 1973), she shows that 25% of the raw salary differences could not be explained by the measured characteristics. She also finds that females are more likely to be in lower ranks; however, the difference in rank attainment is not statistically significant.

McNabb and Wass (1997) use data from all British universities for 1975, 1985 and 1992. The samples are very large; however, they lack measures for research output. In an Oaxaca decomposition, the authors obtain a gender salary gap of between 3.6 to 5.5% that could not be explained by differences in age, tenure, faculty affiliation, and rank. The unexplained salary gap would have been 7% larger if rank were not controlled for, suggesting that the under-representation of females at senior levels significantly contributed to the gender salary differences. Ward (2001) utilizes a sample of 482 male and 241 female from five Scottish universities to investigate the gender salary gap. Her results show that males have a 7.7% salary advantage over females, after controlling for numerous individual characteristics, but excluding rank. When rank is controlled for, the difference is reduced to 3.2% and is no longer statistically significant. When the Oaxaca decomposition is used, 3% of the salary difference