A Market Test for Sex Discrimination: Evidence from Japanese Firm-Level Panel Data

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Abstract

This paper examines the empirical implications of Becker's classical theory of employer discrimination. If the male-female wage differential is due to employer discrimination, then non-discriminatory employers will hire more women and enjoy a higher profit than discriminatory employers. This proposition is tested using Japanese firm-level panel data from the 1990s. The empirical results, based on pooled OLS, indicate that an increase in the proportion of women employed by a firm enhances its operating profit. The female proportion could be endogenous in the profit equation, however, because firms may adjust their female labor in response to positive productivity shocks. To deal with this possible endogeneity, investment or intermediate input amounts are used as proxy variables for productivity shocks, as suggested by Olley and Pakes (1996) and Levinsohn and Petrin (2003). The findings based on these proxy variable estimations confirm the robustness of the finding that hiring more women results in higher profit. This result fails to reject the hypothesis that employer discrimination is a source of the male-female wage gap. However, the size of estimated effect of female proportion on profit is 1/20 of the predicted coefficient calculated based on the assumption that all the observed gender wage gap is due to gender discrimination. This result suggests that the large portion of gender wage gap is due to gender productivity gap. In addition, those firms that hire more women do not necessarily grow faster than firms that hire fewer women. The data do not reject the static implication of Becker's hypothesis, but do reject its dynamic implication.

1 Introduction

Women's average wage is lower than that of men in many countries, and the difference in average wages is very persistent. In March 1996 in the US, White males' average hourly wage was 18.96 US dollars per hour, while that of White females was 12.25 dollars.[1] In 1995 in Japan, males' average hourly wage was 2,960 Yen, while the females' average was 1,760 Yen.[2]

Researchers have discussed whether this difference reflects either a male-female productivity differential or discrimination against women, which is defined as receiving a lower wage than would be indicated by their productivity. Many labor economists have regressed wage on proxies of productivity, such as educational background or job experience, and examined the unexplained male-female wage gap. The unexplained differential has been attributed to sex discrimination. However, there are productivity differences across persons that cannot be observed by econometricians, and these differences may differ systematically across sexes. For example, women with household responsibilities may be less productive in the workplace and thus earn less than men, reflecting their lower productivity (Becker (1985)). Therefore, obtaining a good measure of workers' productivity is very important for making the

[1] Based on Current Population Survey, March 1996. See Altonji and Blank (1999).

[2] According to statistics from the International Labor Office, the average monthly earnings (including bonus payments) of wage earners and salaried employees in the non-agricultural sector was 496.0 thousand yen for men and 252.8 thousand yen for women, whereas hours of work per week was 38.7 for men and 33.2 for women in 1995. Assuming there are 4.33 weeks in a month, the hourly rate of pay was 2.96 thousand yen for men and 1.76 thousand yen for women. US$1 was 114.30 Yen in June 1997.

regression-adjusted wage gap compelling evidence for sex discrimination. Although many studies have found evidence of sex discrimination, this evidence is not necessarily definitive because the perfect measure of each worker's productivity is anywhere from difficult to impossible to obtain. Consequently, it is reasonable to claim that the discussion based on wage regression faces a big challenge. (See Altonji and Blank (1999) for a survey of the literature on sex discrimination.)

To circumvent the limitation of the wage regression method, several recent studies have implemented a "market test" to identify labor market discrimination against minority workers. This method tests the empirical implications of Becker (1971)'s employer discrimination hypothesis: If majority and minority workers are equally productive but minority workers are paid less, firms can earn more profit by hiring more minority workers. To test the hypothesis, some measure of a firm's profitability is regressed on the minority employee proportion of the total number of employees. If discrimination against women does not exist, then the female employee proportion should not affect the firm's profitability. However, if discrimination exists, then having a higher minority proportion should increase profitability. Hellerstein, Neumark, and Troske (2002), Szymanski (2000), Sano (2005) and Kodama, Odaki, and Takahashi (2005) have tested this theoretical prediction. The latter two studies utilized Japanese data and did not reject Becker's taste discrimination theory in their cross-sectional estimates. Ashenfelter and Hannan (1986), Black and Brainerd (2004), and Black and Strahan (2001) have examined whether firms in more competitive environments hire more women.[3]

Of the previous work, Hellerstein, Neumark, and Troske (2002) is the closest to that presented in this paper. These researchers implemented a market test of sex discrimination using US employer-employee matched data. They found that, holding other variables constant, an increase in the female proportion increased the operation income/sales ratio. Their test did not reject Becker (1971)'s taste-based discrimination theory. Although their study made a path-breaking contribution to the literature, considering the source of variation in the female proportion is important. If the variation is created by the difference in managers' discriminatory preferences, then their test appropriately tests the null of the non-existence of taste discrimination. However, if the variation in female proportion is created by the temporary productivity or demand shock to the firms, then the female proportion becomes endogenous and the correlation between profit and female proportion does not reject the null of the employer taste discrimination model.

In addition, the female proportion may be endogenous in the US due to the Equal Employment Opportunity Act and the different degrees of law enforcement across firm sizes. Holzer (1998) reported that larger firms hire

[3]The other recent method for overcoming the limitation of the residual method is estimating the production function that regards male and female labor as separate inputs. Then the implied ratio of the female and male marginal products is compared with the ratio of the female and male wages. Hellerstein, Neumark, and Troske (1999) found evidence consistent with female discrimination using US establishment data, while Hellerstein and Neumark (1999) did not find such evidence using Israeli data. Both papers were based on cross-sectional estimates.

disproportionately more African-American workers than smaller firms. He speculated that both stronger law enforcement for larger firms and the exemption of small firms from law enforcement explained his finding. If the same discussion applies to women, then larger firms that arguably enjoy high profitability may hire disproportionately more women to comply with equal opportunity laws.

This paper implements the market test for sex discrimination using Japanese firm-level panel data. This paper's contributions to the literature are threefold. First, closer attention is paid to the endogeneity of female proportion induced by productivity shocks using a proxy variable method proposed by Olley and Pakes (1996) and Levinsohn and Petrin (2003). In addition, the panel data structure allows a fixed-effects estimation, which arguably mitigates the estimation bias under certain conditions. Although these estimators are not free from biases, the OLS estimator presumably estimates the lower bound of the true coefficient, and the fixed- effects estimator presumably estimates its upper bound. Second, the survey covers not only manufacturing, but a wide range of industries, assuring the generality of the results. Third, Japan is an ideal country for implementing the market test because a) the male-female wage gap is much larger than in Western counties, making the market test powerful if the gap is due to discrimination;[4] and b) the Equal Employment Opportunity legislation that prohibits discrimination against

[4]See Antecol (2001) and Blau and Kahn (2003) for international comparisons of the male-female wage gap.

women in recruitment, hiring, job assignment, and promotion became effective in April 1999.[5] This weak law enforcement mitigates the endogeneity bias due to feedback from high profit to higher female employment; in the US, firms that enjoy higher profits may comply with anti-discrimination laws due to the higher cost of being sued.

The results of this paper indicate that those firms with more female employees earn higher profits. This result is robustly found in several specifications. Thus, Becker's static prediction is not falsified. It should be noted, however, that the size of estimated coefficient is 1/20 of the predicted coefficient calculated based on the assumption that all the observed gender wage gap is due to gender discrimination. This result suggests that the large portion of gender wage gap comes from gender productivity gap. In addition, the test rejects the hypothesis that non-discriminatory employers grow faster than discriminatory employers. Thus, Becker's dynamic prediction is rejected.

The rest of this paper is organized as follows. Section 2 lays out theoretical model that motivates the empirical strategy, Section 3 discusses the empirical strategy, Section 4 discusses the data, Section 5 shows the estimation results, and Section 6 concludes.

[5]The original equal opportunity legislation, effective in April 1986, only required that employers make an effort not to discriminate against women.

2 Theory

The theoretical model of employer taste discrimination is introduced as the basis for the market test. Suppose firm managers obtain utility by earning higher profits and hiring less women and have following utility function:

$$U = pf(M, F) - w_M M - w_F F - dF, \tag{1}$$

where p is the product price, M is the number of men, F is the number of women, w_M is the wage for men, w_F is the wage for women, and d is the manager's psychic cost of hiring women denominated by a monetary term. The managers dislike having women in the workplace and feel as if they pay $(w_F + d)$ by hiring an additional female worker. The discrimination coefficient d is assumed to be heterogeneous across managers. The product and labor market are perfectly competitive and the managers are price takers.[6] The managers solve their utility maximization problem by choosing M and F. The solution to the problem is denoted as $M^*(p, w_M, w_F, d)$ and $F^*(p, w_M, w_F, d)$. The profit function becomes

$$\pi(p, w_M, w_F, d) = pf(M^*, F^*) - (w_M M^* + w_F F^*). \tag{2}$$

[6]Some readers might think it is more natural to model that discriminatory managers are willing to pay less to women than to men. However, this type of modeling under the perfectly competitive labor market assumption results in two extreme types of firms, those hiring either only women or only men. This is because those managers who are willing to pay more for women than the market female wage hire women only and those managers who want to pay less for women than the male market wage hire only men. To avoid these corner-solution cases, I must assume some kind of labor market friction and adopt the labor market search model. This might be an interesting modeling strategy, but is likely to be inconsistent with the goal of the market test.

If the product price and wages are identical across firms, conditioned on time, region, and industry, the profit is only the function of the discrimination coefficient d, conditioned on time, region, and industry. It is obvious that $\partial\pi/\partial d < 0$ because when $d = 0$ the managers maximize profits. It is worth mentioning that non-discriminatory employers are predicted to earn higher profits through two mechanisms. First, non-discriminatory employers produce profit-maximizing output by hiring more women until their value of the marginal product of labor is equivalent to their wage. Second, non-discriminatory employers produce a given output at less cost because they choose a cost- minimizing input mix.

Because d is not directly observable, female proportion $F/(M + F)$ is used as a proxy variable for d. The comparative statics show that $\partial(F/(M+F))/\partial(d) < 0$ if $f_{MF} < -\frac{M}{F} f_{FF}$, which I assume to hold in the following analysis.[7] Using this result, the discrimination coefficient is written as the function of female proportion $d = d(\frac{F}{M+F}; p, w_M, w_F)$ and $\partial d(\cdot)/\partial \frac{F}{M+F} < 0$. The profit function results in

$$\pi(p, w_M, w_F, d(\frac{F}{M+F}; p, w_M, w_F)). \tag{3}$$

The employer taste discrimination model predicts $\frac{\partial\pi}{\partial(F/(M+F))} = \frac{\partial\pi}{\partial d} \frac{\partial d}{\partial\{\frac{F}{M+F}\}} > 0$ and the model is rejected if $\frac{\partial\pi}{\partial(F/(M+F))} = 0$.

[7]If males and females are too complementary in production, then this condition is violated. If males and female are highly complementary, then an increase in the discrimination coefficient reduces female employment and that results in a reduction of males' marginal productivity. Thus males' optimal input is reduced at a more rapid pace than the reduction of the number of female workers. This extreme case is less likely to occur in reality.

In its empirical implementation, the product price and wages are assumed to be held constant by including time, regional, and industry dummy variables. If a firm experiences an idiosyncratic price (demand) shock that is not captured by these dummy variables, then the price shock can create a variation in female proportion because $\frac{\partial(\frac{F}{F+M})}{\partial p}$ could be either positive or negative, depending on the shape of the production function and the current female proportion.[8] The price shock naturally affects the profit rate. Thus, the market test could falsely reject the null of the employer taste discrimination model. This endogeneity of the female proportion in the profit equation is taken care of by using the proxy variable method, which presumably captures price shock, as suggested by Olley and Pakes (1996) and Levinsohn and Petrin (2003).

3 Empirical Strategy

3.1 Static Test

3.1.1 Cross-sectional Estimation

I estimate the following equation to examine whether a higher proportion of female employees increases a firm's profit.

$$profit_{it} = \beta_0+\beta_1 x_{1it}+\beta_2 x_{2it}+\beta_3 x_{3it}+ind_{it}\beta_4+reg_{it}\beta_5+year_t\beta_6+v_{it}+u_{it}, \quad (4)$$

where subscripts i and t are indexes for firm and year, respectively. The dependent variable $profit_{it}$ is a proxy for profitability, defined as operating

[8]The result of the comparative statics is $\frac{d(\frac{F}{F+M})}{dp} = \frac{1}{M+F}\frac{1}{f_{FF}f_{MM}-f_{FM}^2}\{\frac{M}{M+F}(f_{FM}f_M - f_{MM}f_F) + \frac{F}{M+F}(f_{FM}f_F - f_{FF}f_M)\}$. We cannot determine the sign of this expression.

income/total sales, which is essentially the price-cost margin. Here, operating income is calculated by subtracting the sales cost and overhead cost from total sales. Operating income does not correspond to economic profit without subtracting the opportunity cost of capital. The discrepancy between operating income and economic profit depends on each firm's amount of capital. To deal with this issue, I include a fixed assets / total sales ratio, denoted as x_{2it} in the regression. The variable x_{1it} is the proportion of female employees among the total employees, including part-time workers. To examine whether the heterogeneity of full- and part-time workers affects the results, the workers are decomposed into full-time male, part-time male, full-time female, and part-time female workers. If discrimination against women exists, then a high female proportion will result in high profit; accordingly, a positive β_1 rejects the null hypothesis of no sex discrimination. The variable x_{3it} is the firm's age. Since older firms tend to hold obsolete capital and the assets/total sales ratio may not capture the capital in the efficiency unit and older firms may hold higher-level intangible capital, such as research and development knowledge or an established brand name, it is important to control for this variable. To capture industrial and regional product, wage heterogeneity, and time-specific macro shock, industry dummies, prefecture dummies, and year dummies are included, respectively.[9] The variable v_{it} is the idiosyncratic demand or productivity shock to a specific firm, and u_{it} is

[9]Industry dummies and regional dummies are allowed to be time variant because a non-negligible number of observations experienced a change in industry code and location.

the idiosyncratic error term that contains all information not captured by v_{it}. The error term v_{it} captures the effect of the demand (or productivity) shock p_{it} on profit.

It is worth noting that the profit function is only the function of output/input prices, which is presumably captured by the industry, prefecture, and year dummies in the above specification, as far as firms attempt to maximize their profits in a perfectly competitive market. If firms maximize their profits, then the female proportion of workers should not appear in the profit function because it is already chosen optimally in terms of profit maximization, and the envelope theorem should apply. Thus, the female proportion of workers appears in the profit function as a proxy variable for the taste discrimination that is heterogeneous across firms. This reasoning explains why the profit function contains few explanatory variables except for the female proportion of workers.

The assumption

$$E(v_{it}|x_{1it}, x_{2it}, x_{3it}, ind_{it}, region_{it}, year_t) = 0$$

and

$$E(u_{it}|x_{1it}, x_{2it}, x_{3it}, ind_{it}, region_{it}, year_t) = 0$$

assures that the OLS estimator is unbiased and consistent. The first assumption excludes the situation in which an idiosyncratic demand or productivity shock to the firm's profit disproportionately affects the number of female and male employees. If female employment adjusts to idiosyncratic shock more quickly, as documented by Houseman and Abraham (1993), then v_{it} and x_{1it} are positively correlated and the OLS estimator of β_1 is upward biased. This bias is serious because the test may spuriously reject the null hypothesis of no discrimination against women. Thus, I suggest a remedy for this bias in the following discussion of the proxy variables estimation and the fixed-effects estimation. I assume that $E(u_{it}|x_{1it}, x_{2it}, x_{3it}, ind_{it}, region_{it}, year_t) = 0$ always holds.

3.1.2 Controlling the productivity shock by proxy variables

As discussed before, the variation in the female employee proportion across firms may result not only from a variation in employers' intent to discriminate based on sex, but also from a variation in firms' demand or productivity shock. If the firms' unobserved heterogeneity also affects their profits, then the female proportion is endogenous; this makes the OLS estimator biased.

To control for demand or productivity shocks, the proxy variables for these shocks are introduced into the profit equation. In the context of the production function estimation, Olley and Pakes (1996) examined the conditions under which investment could be used as a proxy for unobserved productivity shocks. Their idea was that those firms that experience positive productivity shocks invest in their capital expecting that productivity shocks are serially correlated, and current positive productivity shocks make firms believe that there will be successive positive productivity shocks in the future. Under general conditions, they showed that $i_{it} = i(k_{it}, v_{it})$ and

$\partial i_{it}/\partial v_{it} > 0$ where i is investment amount, k_{it} is capital stock amount and v_{it} is the demand or productivity shock. As far as firms invest a strictly positive amount, productivity shock is expressed as an inverse function of investment and capital $v_{it} = g(i_{it}, k_{it})$. This function is parameterized as

$$v_{it} = \gamma_1 \frac{i_{it}}{k_{it}} + \gamma_2 (\frac{i_{it}}{k_{it}} - \overline{\frac{i_{it}}{k_{it}}})^2, \tag{5}$$

where $\bar{x}$ is the sample mean of x. This function is substituted into (4), and the resulting model is estimated using the observations with positive investment as an analysis sample.[10]

The drawback of Olley and Pakes (1996) is that not a small number of firms report zero investment. Thus, a large number of firms should be dropped from the analysis sample. To overcome this drawback, Levinsohn and Petrin (2003) suggested using the amount of intermediate inputs as a proxy variable for productivity shocks. The intermediate good's demand function is given as $m_{it} = m(v_{it}, k_{it})$ given k_{it} is a state variable that cannot be instantaneously adjusted. For the intermediate input to be a valid proxy for productivity shock, m_{it} should be monotonic in v_{it} for given any k_{it}. When the monotonicity condition is satisfied, the productivity shock is expressed as a function of the intermediate input and capital stock $v_{it} = h(m_{it}, k_{it})$.

[10]The coefficients for the asset/sales ratio (β_2), γ_1, and γ_2 are identified due to the functional form assumption. Olley and Pakes (1996) considered the two-step estimation method to identify the production function parameters under a very flexible functional form for the investment function. This consideration is not necessary here because the goal is to estimate the profit function controlling for productivity shocks. The aim of the analysis is to identify the coefficient for the female proportion, which is identified without relying on the functional form assumption of the investment function.

This function is approximated as

$$v_{it} = \delta_1 \frac{cost_{info}}{cost} + \delta_2 \left(\frac{cost_{info}}{cost} - \overline{\frac{cost_{info}}{cost}}\right)^2 + \delta_3 \left(\frac{cost_{info}}{cost} - \overline{\frac{cost_{info}}{cost}}\right) \times \frac{asset}{sales}, \quad (6)$$

where $cost_{info}$ is the cost of information processing and communication. This intermediate input is chosen because it is commonly used for both manufacturing and service sectors. This cost is normalized by the total production cost, *cost*. The presumption here is that information-related costs respond more rapidly to product demand or productivity shocks than total cost, which seems reasonable. To allow for the interaction with asset amount, the cost share of info. & comm. cost is interacted with the asset/sales ratio.

3.1.3 Capturing heterogeneity by fixed effects

If factors that affect female employment proportion other than employers' intent to discriminate are time constant, then the endogeneity problem can be resolved by allowing an arbitrary correlation between $v_{it} = c_i$ and female proportion by applying a fixed-effects estimation to (4). The fixed effects estimator identifies the causation from female employment to profit if x_{it} varies over time by a change in employers' taste of sex discrimination or some other factors not correlated with the firms' profits. More formally, the sufficient condition for the fixed- effects estimator to be unbiased is the strict exogeneity of the independent variables from idiosyncratic error (i.e., $E(u_{it}|x_{1i}, x_{2i}, x_{3i}, ind_i, region_i, year_t, c_i) = 0$, where x_i is the vector containing the whole history of x_{it} of individual i.). This assumption rules out the inter-temporal correlation of u_{it} and explanatory variables, as well as any contemporaneous correlation. For example, if the idiosyncratic shock to profit and female proportion are positively correlated, then the market test falsely rejects the null hypothesis of no discrimination. To deal with this endogeneity problem, I include industry-year dummies based on an arguable assumption that the demand shock each firm faces is mostly captured by industry- and year-specific shocks. If most of the variation in demand shock across firms is explained by industry- and year-specific shocks, then including this industry-year interaction arguably removes the presumable upward bias in the fixed-effects estimator.

It is, however, worth noting that the fixed-effects transformation leaves a small variation in female proportion because the large variation in the female proportion comes from between-firm variation rather than from within-firm variation. Also, if discriminatory taste does not change over time, the within-firm variation in the female proportion only picks up the product demand or productivity fluctuation. Thus, the fixed-effects estimator could suffer from the upward bias more seriously.

3.2 Dynamic Test

The empirical strategy thus far tests the static implication of Becker (1971)'s hypothesis. Next, I examine whether firms with a high female proportion grow faster because non-discriminatory employers earn higher profits than discriminatory ones, as claimed in Becker (1971). Using the sales growth and the number of employees as proxies for a firm's growth, I estimate the following equation:

$$\frac{y_{i1999} - y_{i1992}}{(y_{i1999} + y_{i1992})/2} = \gamma_0 + \gamma_1 x_{1i1992} + \gamma_2 age_{i1992} + ind\gamma_3 + reg\gamma_4 + u_i, \quad (7)$$

where y_{it} is sales or the number of employees in firm i and time t, x_{1i1992} is the female proportion of firm i in year 1992, and age_{i1992} is the firm's age in year 1992. The industry and prefecture dummies also are included. This dependent variable, which was first suggested by Davis and Haltiwanger (1992), takes a value between -2 and 2. When a firm disappears during the period, 0 is assigned for y_{i1999} and the value takes -2; when a firm grows from zero sales to positive sales, the value takes 2. Using this measure enables me to include those firms that disappeared during the analysis period. This is important because firms may be extinct due to their discriminatory behavior. The long-term implication of Becker (1971)'s hypothesis predicts positive γ_1, thus testing the null hypothesis of $\gamma_1 = 0$ is our interest. The sample consists of only those firms that were observed in 1992.

4 Data

I used the basic survey of firms' activity collected by the Ministry of Economy, Trade, and Industry (METI) of the Japanese government to implement the test. This is a firm-level census survey that covers all firms hiring more than 50 employees and holding more than 30 million yen in capital. The available data cover 6 years, 1992 and every year between 1995 and 1999; and the sample size is about 25,000 firms for each year. The survey was not conducted in 1993 and 1994. Each firm is assigned a permanent firm ID, and each firm is matched over time based on this ID. From the data sets, I extracted each firm's performance measures, such as total sales, sales cost, and overhead cost, data on the firm's employees, such as the number of full-time employees with sex breakdown, the number of part-time employees with sex breakdown, the book value of its fixed assets, the year the firm was founded, the prefectural location of the firm, and the three-digit code indicating the industry in which the firm operates. There were originally 152,857 firm-year observations in the 6 years of data, but after excluding observations with missing sales information or inconsistent employee records, there remained 152,774 firm-year observations. Among these observations, 108 observations were dropped due to missing values for the asset/sales ratio, and thus there remained 152,666 observations. The proxy for profitability is the operating income ratio, defined as (total sales - sales costs - overhead costs)/total sales. Since this variable takes an extreme value due to the very small amount of total sales, the observations whose operating income ratio was below -100 percent were dropped from the sample, and there were 152,606 firm-year observations remaining.[11] The sample size is 152,606 observations with 36,445 companies. This unbalanced panel is the analysis sample for implementing the static test. The dynamic test requires observations for at least the year 1992. The remaining 24,319 companies that satisfied this

[11]It turned out that this sample restriction was critical in order to obtain the following results.

restriction are used in the dynamic analysis.

The survey record unfortunately does not distinguish missing values and zeros, except when a firm did not answer the entire survey. Since replying to the survey is compulsory due to the Statistics Law and because the METI exerts its best effort to fill in missing values with a follow-up phone survey, missing values are presumably rare. Thus, all values of zero in the record are treated as actual zeros.

This study uses all the industries as an analysis sample. The aim is to increase the estimation's efficiency by maximizing the number of observations. Industry-level regression often is preferred for the estimation of production functions. The limitation of cross-industry regression is not as severe as in the estimation of the profit function because the parameters in the production function are not required to be identical across industries. However, the effect of female proportion on profit might differ across industries due to differences in production technologies. Thus, the estimated coefficient should be interpreted as the average effect across heterogeneous industries.

The permanent difference in the profit level across industries for arbitrary reasons is captured by the industry dummies. Thus, the coefficient for female proportion is identified by the within- industry, across-firm variation of female proportion. Also, the macro and regional economic shocks on profit level are captured by the year and prefecture dummies.

The descriptive statistics of the analysis sample used for the static test are reported in Table 1. The operating income to sales ratio is 2.42 percent, which is a standard figure for Japanese data. This number is about one-tenth of the US figure reported in Hellerstein, Neumark, and Troske (2002), but it is known that the profit rate of Japanese firms is lower than that of US firms.[12] The unweighted mean of firms' sales growth index, defined as $(sales_{1999} - sales_{1992})/((sales_{1999} + sales_{1992})/2)$, was -0.63 between 1992 and 1999, reflecting the long-term recession during that period, and the unweighted mean of the employment growth index similarly defined was -0.59. The female proportion of each firm's total employees was 32 percent at the mean, which was figured by adding 23 percent for full-time workers and 9 percent for part-time workers.

Table 2 closely examines the yearly labor adjustment. The unweighted mean of the yearly log employment difference is -0.01; this number implies that firms downsized their employees by 1 percent annually on average. An examination of the 10th percentile (-0.14) and the 90th percentile (0.13) of the distribution, however, shows a large amount of heterogeneity in the labor adjustment. The number of female employees declined by 2 percent annually at the mean, and this number indicates that female workers suffered from recession more severely than male workers during the 1990s. An examination of the 10th percentile and the 90th percentile of the distribution reveals

[12]See, for example, Gugler, Mueller, and Yurtoglu (2004) for an international comparison of the return to investment. The average annual growth rate of firm values in the US was 0.124 between 1985 and 2000 for 8,591 firms, while the rate in Japan was 0.064 during the same period for 2,219 firms. Note that Hellerstein, Neumark, and Troske (2002) restricted their analysis sample to manufacturing establishments, but the current sample contains all industries.

that the labor adjustment of female workers was more heterogeneous than males' labor adjustment. Female workers are more likely to be subject to employment adjustment, probably because they hold less firm-specific human capital and, accordingly, less relation-specific rent. As evidence of this, the average tenure in 1997 was 13.3 years among male workers and 8.2 years among female workers (see Table 16 in Japan Institute of Labour (2002)).

The relationship between the change in log total employment and the change in log female employment is plotted in Figure 1. If the observation is on the 45 degree line, then the female proportion of total employment did not change over time. Thus, basically, the deviation from the 45 degree line is used to identify the coefficient of the female proportion in the fixed effects estimation. This scatter diagram implies that the data contain sufficient variation in the female worker proportion over time within a firm. The regression coefficient of the change in log female employment on the change in log total employment is 1.22, with a standard error of 0.003. This coefficient implies that the female proportion increases by 2.2 percent on average when total employment grows by 10 percent.[13]

Regarding the across-firm distribution of female proportion, Becker's taste discrimination theory predicts that firms in competitive industries have no

[13]This finding may raise some doubt that the female proportion is endogenous in the profit equation, since female employment may be more sensitive to demand shock than male employment. However, when I included total employment as an additional explanatory variable in the profit equation as a proxy variable for demand shock, the coefficient for female proportion reported in the next section did not change essentially. Those results are not reported here.

room to indulge their preferences for sex discrimination. Accordingly, firms in competitive industries are likely to hire more women. This presumption, on which the market test of taste discrimination is based, is refuted if there is no correlation between the degree of competition for each industry and the industry average of female proportion.

The null hypothesis of no-correlation between the concentration and female proportion is tested by estimating the following model.

$$x_{1ijt} = \theta_1 + \theta_2 hi_j + reg_{it}\theta_3 + year_t\theta_4 + e_{it}, \qquad (8)$$

where x_{1ijt} is the female proportion of firm i in industry j at year t and hi_j is the Herfindahl index for industry j in 1992. Becker's framework is refuted if $H_0 : \theta_2 = 0$. The Herfindahl index in the initial year is used because the year-to-year Hefindahl index variation can pick up the effect of an industry-specific business cycle effect, and the business cycle can affect the female proportion.

The regression results appear in Table 3. The results indicate that firms in more concentrated industries are less likely to hire women. This implies that Becker's framework is not rejected by the data. The variation of female proportion across industries could be a result of heterogeneous technologies; thus this test is not a definitive test for the employers' taste discrimination. However, at the very least, the presumption of the market test is not rejected.

5 Estimation Results

5.1 Static Results

5.1.1 Basic Results

Table 4, Column (1) reports the results of the OLS estimation of the equation (4). The coefficient for the female proportion is 0.36 (s.e.=0.18). The magnitude of the coefficient implies that a 10 percent increase in the female proportion increases the profit ratio by 0.036 of a percentage point, while the average profit rate is 2.42 percent.

Table 4, Column (2) reports the result of the OLS estimation with Olley and Pakes (1996)'s proxy variables for demand or productivity shocks. Some observations are dropped from the column (1) due to missing or zero investment. The investment/fixed asset ratio does not enter the model significantly. This is a sign that the investment does not effectively capture productivity or demand shocks. The coefficient for the female proportion becomes larger in this specification.

Table 4, Column (3) is for the OLS result with Levinsohn and Petrin (2003)'s proxy variables for productivity or demand shocks. Now, the coefficient for the linear term of the cost share of information processing and communication costs is positive and significant. This suggests that this variable works well as a proxy variable for positive productivity or demand shocks. The coefficient for the female proportion becomes larger than the OLS estimate, and it is statistically significant. It is rather difficult to explain this positive change because we expected an upward bias of the OLS estimator *a priori*, but the confidence intervals of the two estimates are overlapping due to the large standard errors for each coefficient. This shows the robustness of the OLS results even after controlling for productivity shocks using the cost share of information processing and communication.

The fixed-effects estimates reported in Table 4 Column (4) are as large as 1.74 (s.e. = 0.27). These fixed-effects coefficients, which are larger than the OLS coefficients, may result from a strong, positive correlation between idiosyncratic error and female proportion. Suppose the female proportion across firms varies due to heterogeneity in the employers' preference for discrimination and temporary demand shocks. As implied by Table 2 and explicitly discussed in Houseman and Abraham (1993), female employment is more sensitive to product demand shocks. If the employer's preference for discrimination does not vary over time, then the source of within- variation of the female proportion is all due to temporary demand shocks. Also, the endogeneity of the female proportion causes a downward bias for the capital coefficient because the female proportion and the fixed asset / total sales ratio are positively correlated (Table 3 Column (2)).

The estimation result of the model that allows for the AR(1) serial correlation of the idiosyncratic error term appears in Table 4, Column (5). The coefficient becomes larger, but this could be due to the violation of strict exogeneity of the idiosyncratic error term. In this estimation, the Cochrane-Orcutt estimation is applied to the within-transformed variable. The explanatory variable becomes $(x_{it} - \bar{x}_i) - \rho(x_{it-1} - \bar{x}_i)$ and the idiosyncratic error term becomes $(u_{it} - \bar{u}_i) - \rho(u_{it-1} - \bar{u}_i)$, where ρ is the AR(1) coefficient of the idiosyncratic error process: $u_{it} = \rho u_{it-1} + white\ noise$. Thus, if firms reduce their female proportion one period after a positive shock, the error term and the female proportion are positively correlated, given $\rho > 0$. Accordingly, the coefficient for the female proportion becomes positively biased. Strict exogeneity should hold for the consistency of the fixed-effects estimator in the first place, but the violation of this assumption biases the FEAR(1) estimator more seriously than the FE estimator.

The analysis thus far has ignored the possibility that current, positive demand shock increases the current female proportion via more rapid female labor adjustment. To deal with this potential endogeneity of x_{1it}, I include industry-year dummies in the regression; the dummies capture the demand shock as far as demand shock is common across firms operating in the same three-digit industry in a specific year. Table 4, Column (6) shows the results of the fixed-effects estimation with industry-year dummies. The estimated coefficient becomes smaller (1.59, s.e.=0.27), and this change from the previous fixed-effects estimate suggests a positive correlation between the current industry-year positive demand shock and the female proportion.

Olley and Pakes (1996) and Levinsohn and Petrin (2003) articulated that the fixed-effects coefficients for the input that can be easily adjusted are more upward biased than OLS coefficients in the estimation of the production function. This is because the within- variation of inputs and idiosyncratic shock are strongly correlated. This more severe upward bias for the shock-sensitive inputs makes the downward bias for shock-insensitive inputs more serious because shock-sensitive inputs and shock-insensitive inputs generally have a positive correlation. Their discussion is relevant here because positive productivity shock positively affects profit level and is likely to have a positive correlation with the female proportion. In fact, we observe negative coefficients for the fixed asset / sales ratio as fixed-effects estimates. Thus the fixed-effects estimates on female proportion could be upward biased, and they can be understood as the upper bounds of the true causal effect of female proportion on profit.

Although the possible upward bias of the fixed-effects estimator has been discussed, the OLS estimator instead could be downward biased. Suppose some firms have persistently high profits due to their market power and their corporate governance structures allow their managers to enjoy the rent. Then, their managers have more room to indulge their preferences for discrimination without going bankrupt due to the firms' persistent market power. Thus, time-constant high profitability and female proportion could be negatively correlated and the OLS estimator for female proportion in the profit equation could be downward biased.

To summarize the results, the OLS estimate does not seem to be seriously upward biased by productivity shock. Considering the possible downward bias of the OLS estimator, the OLS estimate should reflect the lower bound of the true causal effect of female proportion on profit. Table 4 Column

(1) can be understood as the most preferable estimate for the lower bound. According to this estimate, when the female proportion increases by 10 percentage points, and 30 percent of the workers are women in an average firm, the profit rate increases by 0.036 of a percentage point, while the average profit rate is about 2.4 percent. Using semi-elasticity, a 10-percentage-point increase in female employment increases profit by 1.7 percent. This semi-elasticity is close to the US semi-elasticity of 2.1 percent found in Hellerstein, Neumark, and Troske (2002). The results in Table 5 imply that a large part of this profit-enhancing effect of hiring women comes from hiring part-time women.

To make sense of the size of the estimated, most-preferred coefficient of 0.36, a thought experiment to calculate the upper bound of the effect of female proportion on the profit rate is useful. Suppose that the unadjusted male-female wage differential of 0.59 in 1995 (See footnote 2) is all due to sex discrimination, and male and female workers are perfect substitutes in production. The ratio of the wage bill to total sales is 0.185 at the median in my data. If the proportion of females is exogenously increased by 10 percentage points, while the proportion of male is decreased by 10 percentage points, keeping the output and total employment level constant, then the profit rate should increase by 0.76 (=0.185*(100-59)*0.1) of a percentage point. This number is about 20 times as large as the preferred estimated effect, which is 0.036. This result implies that the large part of the male-female wage differential is not due to discrimination, because if it were, the estimated coefficient should have been 20 times larger.

5.1.2 Full-time and Part-time Decomposition

The results thus far have treated full- and part-time workers equivalently. Critics may claim that the previous results merely reflect the causation from a high part-timer proportion to a higher profit because most of the part-timers are women (about 85 percent in the analysis sample). To address this possibility, I re-classified workers into four categories: male full-time workers, female full-time workers, male part-time workers, and female-part time workers, and created a proportion of each type of workers in relation to total workers. Column (1) in Table 5 shows the result of the OLS regression of profit ratio on the full-time female proportion, part-time male proportion, and part-time female proportion, along with the other explanatory variables. The coefficient for the full-time female proportion is 0.30 (s.e.=0.22), and this result is consistent with the previous finding, although it becomes statistically significant. The positive effect of female part-time proportion on profit is also statistically significant.

Table 5, Column (2) reports the estimation results of the specification that includes investment variables as proxy variables for demand or productivity shocks. The coefficients for the proxy variables are not significant. This insignificance suggests that the investment variables do not capture the demand or productivity shocks effectively. Table 5 Column (3) reports the estimation results that include the cost share of information-related expenses. The coefficient of this cost share is positive, and the coefficient for the quadratic term is negative. Thus, the cost share of information- related expenses has a concave relationship with profit level, and this presumably captures the demand or productivity shocks to the firm. The coefficients are almost identical to the coefficients reported in Table 5 Column (1). Thus we can confirm the robustness of the OLS estimates.

Once the correlation between time-constant firm heterogeneity and the explanatory variables is taken into account by the fixed-effects estimation, the coefficient for the full-time female proportion is 1.53 (s.e.=0.30), as appears in Column (4) of Table 5. The coefficient for the part-time female proportion becomes as large as 1.98 (s.e. = 0.30). The coefficients for each variable get even larger in the fixed-effects AR(1) model estimation. The results do not change essentially when the industry-year dummies are included, as reported in Column (6) of Table 5. These large coefficients for the fixed-effects models might partly reflect the fact that firms encountering positive demand shock adjust their labor force by hiring more female, part-time workers. The negative coefficients for the fixed assets / total sales ratio are also a sign of the bias. Overall, then, we should regard the fixed effects estimates with some doubt.

5.1.3 Higher Profit from More Sales or Less Cost?

Being less discriminatory against women increases the profit level through two mechanisms. First, it increases the output to the profit-maximizing level (scale effect). Second, it decreases labor costs by substituting men with women, given the output level (substitution effect). To shed light on the mechanism behind the relationship between higher women's proportion and higher profit level, the following analysis examines the effect of women's proportion on the total wage bill conditioned on the output level.

It is preferable to divide the effect of female proportion on the profit rate found in the previous section into the scale and the substitution effects. However, it is not easy to do that in this application because the profit rate ($\frac{Sales-Cost}{Sales}$) is not linearly separable into a sales part and a cost part. Instead, the analysis in this section tests the null hypothesis that the substitution effect does not exist because the substitution effect is an appealing and intuitive explanation of why firms with more women earn higher profits.

If there is no substitution effect, then the female proportion should not affect the total wage bill once the output level is conditioned. This null hypothesis is tested by estimating the following equation:

$$\log(wagebill)_{it} = x_{1it}\alpha_1 + \alpha_2 \log(output)_{it} + ind_{it}\alpha_3 + reg_{it}\alpha_4 + year_t\alpha_5 + u_{it}. \tag{9}$$

The variable x_{1it} is the female proportion of the employees. If the substitution effect is not present, then $\alpha_1 = 0$ is expected.

This model is estimated by an OLS pooled regression because the between-firm variation of female proportion, which is likely to capture the variation of the employers' taste, is exploited in the estimation. To capture the industry- and regional-level differences in the wage bill, the three-digit industry and

47 prefecture dummy variables are included. The macroeconomic shocks are captured by the year dummy variables.

Table 6 reports the regression results. The results in Column (1) imply that the increase of female proportion does not affect the wage bill. However, once the employees are decomposed into part-time and full-time, those firms with a higher proportion of full-time female workers have lower wage bills, as reported in Column (2). Thus the null hypothesis of no-substitution effect is rejected. The positive coefficients of part-time proportions for both male and female are found. Those firms that employ more part-time workers may use labor-intensive technologies.

It is rather difficult to come up with a story explaining why firms using labor-intensive production technologies hire fewer women. Because the elasticity of labor demand should generally be higher when the cost share of labor is higher, those firms with labor-intensive technologies should hire more women because wages are lower for women than for men. Thus, the negative coefficient for the full-time female proportion offers consistent evidence for the substitution effect.

5.2 Dynamic Results

In this section, I examine whether a firm that hires more women will grow at a faster rate due to higher profits by estimating the equation (7) from the previous section. If non-discriminatory firms purge discriminatory firms from markets, then a perfectly competitive market eliminates discrimination against women, as predicted by Becker (1971). All firms that reported valid numbers of sales, employees, workers' composition, and firm age in 1992 are used to estimate the equation (7); in other words, the sample selection is based on the criteria that the firms employed more than 50 employees, held at least 30 million yen in capital in 1992, and reported valid numbers for the dependent and independent variables. Notice that the growth index defined by $(sales_{1999} - sales_{1992})/((sales_{1999} + sales_{1992})/2)$ takes -2 in the case of disappearance from the sample in 1999. The firms that disappeared between 1992 and 1999 remain in the sample, and there is no issue related to sample selection.

Table 7, Column (1) tabulates the results of the OLS regression of the growth index of sales, conditioned on the firms' characteristics in 1992. The female proportion negatively affects firms' survival; a 1 percentage point increase in the female proportion decreases the growth index by 0.20, while the average of this index is -0.63. Similar results are obtained when the detailed re-classification of workers is used as a set of explanatory variables. Column (2) shows that a percentage point increase in the female proportion reduces the firms' sales growth index by 0.26. Similar results are obtained for the employment growth regression, and the estimation results are reported in Columns (3) and (4).[14]

In the cross-sectional analysis, those firms with higher proportions of

[14]Considering the mean reversion of the sales and employment growth, the specifications with initial value of sales and employment were estimated. These specifications resulted in almost identical coefficients for initial female ratio variables.

female employees are confirmed to earn higher profits. However, those firms with more women do not necessarily grow faster. The evidence is consistent with a hypothesis that the market structure is not competitive enough or that discriminatory firms have a certain market power that prevents them from being pushed away from the market. If the corporate governance structure is such that the shareholders' benefit is maximized, then the rent due to market structure or market power goes to the shareholders, and managers have no room to indulge their preference for discrimination. The empirical results in this section suggest that markets are not competitive enough to eliminate sex discrimination, and corporate governance structures seem to leave room for managers to enjoy the rent.

6 Conclusion

Becker (1971) proposed a model of employer discrimination, which indicated that discriminatory employers hire fewer women than would be necessary to attain the profit-maximizing level, in order to indulge their preference for discrimination. This paper examined whether firms with a higher female proportion earn higher profits to test the empirical implication of this theory. Firm-level, panel data from Japan covering a six-year period were used to implement the test. The empirical results robustly showed that a rise in female employment increases the profit ratio, defined by the operating income/sales ratio. This result fails to reject the hypothesis that employer discrimination is a source of the male-female wage gap. However, the size of estimated coefficient was 1/20 of the predicted coefficient calculated based on the assumption that all the observed gender wage gap was due to gender discrimination. This result suggested that the large portion of gender wage gap came from gender productivity gap.

Becker (1971)'s employer discrimination hypothesis has a long- term implication that employers who do not discriminate purge discriminatory employers because the former firms earn more profits. This long-term implication was tested by examining whether the firms hiring more women grow faster than the firms hiring fewer women. The result of this test indicated that the initial level of female employment does not induce firm growth, which is measured by sales growth or growth in the number of employees. This implies the rejection of the long-term prediction of Becker (1971)'s original employer taste-based discrimination theory.

The rejection of the long-term implication of Becker (1971)'s theory of employer discrimination could be due to two factors. First, the market structure might not be competitive and the corporate governance structure may be such that the rent is imputed to discriminatory employers. Second, another discrimination mechanism that allows the long-term persistence of minority discrimination could be in motion.

Considering the rejection of the long-term prediction of employer discrimination theory in the US and Japanese data, it is promising to empirically test alternative models that predict the persistence of discrimination. A model based on coworker discrimination, such as Sasaki (1999)'s, would be a strong alternative.

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Table 1: Descriptive Statistics, All Firms, 1992, 1995-1999

	Mean	S.D.
Operating Income / Sales (%)	2.42	5.94
Sales Growth Index 1992-1999 (N=24319)	-0.63	0.92
Total Employment Growth Index 1992-1999 (N=24319)	-0.59	0.93
Proportion Full-time Female	0.23	0.16
Proportion Part-time Male	0.02	0.05
Proportion Part-time Female	0.09	0.15
Fixed Assets / Total Sales	0.35	0.89
Firm Age	35.24	14.44
Investment / Fixed Assets (given >0, N=125162)	0.14	2.33
Information processing & communication cost / cost (given >0, N=144617)	0.005	0.02
Herfindahl Index in 1992	0.029	0.028
Number of Firms	36445	-
Number of Observations	152606	-
Decomposition based on the number of repeated observations		
1	5469	-
2	8568	-
3	9789	-
4	13700	-
5	28430	-
6	83658	-

Note: The sample includes observations with operating income / sales ratios above -100 percent. Only for the sales and employment growth rate, the observations available for 1992 are selected. Sales growth index is defined as

Sales Growth Index = (Sales in 1999 – Sales in 1992) / ((Sales in 1999 + Sales in 1992) / 2).

This number takes values between -2 and 2. Total Employment Growth Index is similarly defined.

Table 2: Labor Adjustment during the 1990s: 1992, 1995-1999, Pooled

	Mean	Standard Deviation	Median	10th Percentile	90th Percentile	N
Δ Log Total	-0.01	0.18	-0.01	-0.14	0.13	116146
Δ Log Total (Weighted)	-0.02	0.21	-0.02	-0.14	0.10	116146
Δ Log Female	-0.02	0.30	-0.02	-0.27	0.22	116146
Δ Log Female (Weighted)	-0.04	0.32	-0.03	-0.25	0.16	116146

Note: Firm size before labor adjustment is used as the weight for weighted statistics. The sample is restricted to those firms observed in two consecutive years. The years 1992 and 1995 are treated as consecutive years.

Table 3: The Determination of Female Proportion across Industries
Dependent Variable: Proportion Female

	(1)	(2)
Model	OLS	OLS
Herfindahl Index	-0.81	-0.76
	(0.05)	(0.05)
Firm Age	-	-0.001
		(0.00009)
Asset / Sales	-	0.004
		(0.002)
Observations	118319	118319
Number of Firms	24319	24319
R^2	0.03	0.04

Note: Prefecture and year dummies are included.

Table 4: The Determination of Profit Rate (Operating Income / Total Sales, Percentage)

	(1)	(2)	(3)	(4)	(5)	(6)
Model	OLS	OLS	OLS	FE	FEAR(1)	FE
Proportion Female	0.36	0.57	0.44	1.74	2.79	1.59
	(0.18)	(0.19)	(0.18)	(0.27)	(0.31)	(0.27)
Firm Age	-0.01	-0.01	-0.009	-	-	-
	(0.00)	(0.0001)	(0.002)			
Fixed Assets / Total Sales	0.06	0.20	0.07	-0.21	-1.61	-0.19
	(0.05)	(0.09)	(0.05)	(0.02)	(0.05)	(0.02)
Investment / Fixed Assets	-	0.002	-	-	-	-
		(0.02)				
(Inv. / FA –mean(Inv./FA))^2	-	0.007	-	-	-	-
/1000		(0.03)				
Information Cost / Cost	-	-	0.03	-	-	-
			(0.007)			
(I.C. / C – mean(I.C./C))^2	-	-	-0.001	-	-	-
			(0.0002)			
(F.A./Tot. Sales)	-	-	-0.02	-	-	-
* (I.C. / C – mean(I.C./C))			(0.005)			
Ind. * Year Dummies	No	No	No	No	No	Yes
R^2	0.07	0.08	0.08	-	-	-
Number of Firms	36445	33426	35589	36445	30970	36445
Number of Observations	152606	125162	144617	152606	116161	152606

Note: Standard errors are in parentheses under the regression coefficients. OLS standard errors are robust against panel clustering. All specifications include three-digit industry, prefecture, and year dummies.

Table 5: The Determination of Profit Rate (Operating Income / Total Sales, Percentage)

	(1)	(2)	(3)	(4)	(5)	(5)
Model	OLS	OLS	OLS	FE	FE AR(1)	FE
Proportion Full-time Female	0.30	0.60	0.37	1.53	2.68	1.43
	(0.22)	(0.23)	(0.23)	(0.30)	(0.34)	(0.30)
Proportion Part-time Male	-0.80	-0.62	-0.72	0.50	1.36	0.29
	(0.40)	(0.43)	(0.40)	(0.41)	(0.44)	(0.41)
Proportion Part-time Female	0.43	0.55	0.54	1.98	3.04	1.78
	(0.20)	(0.21)	(0.20)	(0.30)	(0.34)	(0.30)
Firm Age	-0.01	-0.01	-0.009	-	-	-
	(0.00)	(0.002)	(0.002)			
Fixed Assets / Total Sales	0.06	0.20	0.07	-0.21	-1.61	-0.19
	(0.05)	(0.09)	(0.05)	(0.02)	(0.05)	(0.02)
Investment / Fixed Assets	-	0.002	-	-	-	-
		(0.02)				
$(\text{Inv.} / \text{FA} - \text{mean}(\text{Inv./FA}))^2$ /1000	-	-0.007	-	-	-	-
		(0.03)				
Information Cost / Cost	-	-	0.03	-	-	-
			(0.007)			
$(\text{I.C.} / \text{C} - \text{mean}(\text{I.C./C}))^2$	-	-	-0.0001	-	-	-
			(0.00002)			
(F.A./Tot. Sales) * (I.C. / C – mean(I.C./C))	-	-	-0.02	-	-	-
			(0.005)			
Ind. * Year Dummies	No	No	No	No	No	Yes
R^2	0.07	0.08	0.08	-	-	-
Number of Firms	36445	33426	35589	36445	30970	36445
Number of Observations	152606	125162	144617	152606	116161	152606

Note: Standard errors are in parentheses under the regression coefficients. OLS standard errors are robust against panel clustering. All specifications include three-digit industry, prefecture, and year dummies.

Table 6: The Determination of Wage Bill
Dependent Variable: log (wage bill)

	(1)	(2)
Method	OLS	OLS
Female Proportion	-0.01	-
	(0.02)	
Full-time Female	-	-0.08
		(0.02)
Part-time Male	-	0.05
		(0.04)
Part-time Female	-	0.06
		(0.02)
Log (Sales)	0.75	0.75
	(0.00)	(0.00)
Observations	152604	152604
R-squared	0.79	0.79
Number of Firms	36445	36445

Note: Standard errors are in parentheses under the regression coefficients. OLS standard errors are robust against panel clustering. All specifications include three-digit industry, prefecture, and year dummies.

Table 7: The Determination of Sales Growth

Dependent Variable: $(Sales_{1999}-Sales_{1992})/([Sales_{1992}+Sales_{1999}]/2)$

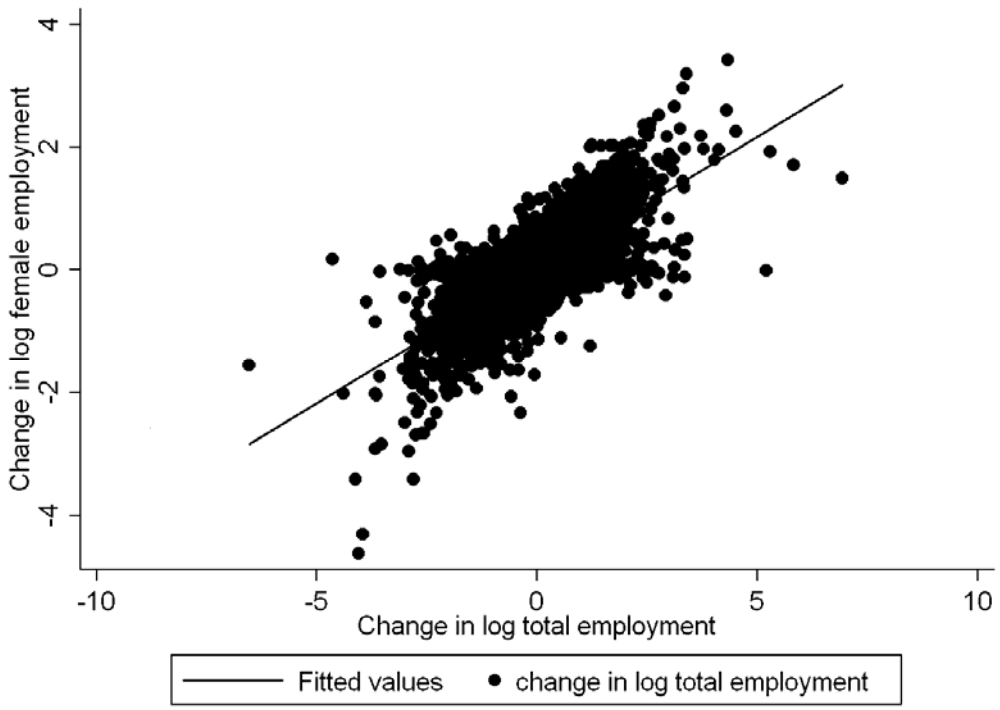
or $(Employment_{1999}-Employment_{1992})/([Employment_{1992}+Employment_{1999}]/2)$

The Range of Dependent Variable: -2 to 2

	(1)	(2)	(3)	(4)
Dependent Variable	Sales		Employment	
Independent Variables in 1992				
Proportion Female	-0.20	-	-0.26	-
	(0.04)		(0.04)	
Proportion Full-time Female	-	-0.26	-	-0.31
		(0.05)		(0.05)
Proportion Part-time Male	-	-0.05	-	-0.38
		(0.15)		(0.15)
Proportion Part-time Female	-	-0.09	-	-0.18
		(0.05)		(0.05)
Firm Age	0.003	0.003	0.003	0.004
	(<0.000)	(<0.000)	(<0.000)	(<0.000)
Number of Observation	24319	24319	24319	24319

Note: All specifications include three-digit industry dummies and prefecture dummies. The standard errors are in parentheses. Among the 24,319 firms for which all the explanatory variables are observed in 1992, 17,431 firms were observed in 1999. All the firms that are not observed in 1999 are assigned zero sales and employment.

Figure 1: The Change in Log Total Employment and Log Female Employment, 1992-1999



Note: Regression: y=-0.01 (0.001) + 1.22 (0.003) x, R^2=0.53, N=116146