

Wage Distribution in Japan: 1989-2003

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October 6, 2007

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Abstract

This study examines the recent trend of the Japanese wage distribution based on a micro-level data set from the *Basic Survey on Wage Structure* (1989-2003). We perform several decomposition analyses of changes in the distribution of the hourly wage. We observe that lower returns to education and years of tenure contribute to a diminishing income disparity between groups for both sexes. A larger variance within a group contributes to the wage disparity for males, while an increased heterogeneity of workers' attributes contributes to the wage disparity for females. The Dinardo, Fortin, and Lemieux decomposition confirms the basic findings with a parametric variance decomposition.

Key Words: Wage Distribution, Wage Equation, Variance Decomposition, DiNardo-Fortin-Lemieux Decomposition, Japan.

JEL Classification Code: J31

1 Introduction

Economic inequality recently has become a major focus of political debate in many developed countries. In such debates, Japan's experience over the last 15 years has started attracting international attention (Economist, the (2006) and OECD (2006)). The long-lasting stagnation of the Japanese economy during the 1990s, rapid globalization and technological change, and the recent economic reform toward deregulation and a more market-oriented economy are claimed to be responsible for widening income inequality (see Bungei Shunju (2006) for the summary of the debates). While the debate about the reason for broadened inequality is heating up, the premise of the debates—that Japan has experienced widening inequality—has not yet been decisively confirmed.

The trend of income inequality in Japan, which has the world's second-largest economy, also has attracted much attention from labor economists because it can offer a testing ground for determining whether the recent trends in income inequality during the 1990s in the US (Autor et al. (2007)), the UK (Goos and Manning (2007)), and Canada (Boudarbat et al. (2003)) can be explained by such global factors as skill-biased technological change (SBTC) or globalization of the economy.

From the trend of wage inequality during the 1980s in the US, there once seemed to be a consensus that SBTC or the globalization of the economy is the main source for increasing wage inequality (Katz and Murphy (1992))

and Juhn et al. (1993)). However, “revisionists” started casting doubts about the traditional hypotheses because, during the 1990s, the trend of increasing inequality had slowed down (Card and DiNardo (2002)), while the speed of SBTC or globalization is believed to have accelerated. DiNardo et al. (1996) and Lee (1999) proposed that the increased inequality through the 1980s can be largely explained by a falling real minimum wage or union density. Lemieux (2006) suggested that the compositional change of the labor force can well explain the increase in residual wage inequality during the 1980s and 1990s. In addition, several studies had reported that Canada, which presumably went through technological and international environmental changes similar to those in the US, had not experienced increased wage inequality during the 1980s (Morissette et al. (1994), Doiron and Barrett (1996), DiNardo and Lemieux (1997), Card et al. (1999) and Card et al. (2003)). This fact created skepticism toward the SBTC or globalization hypothesis.¹

In response to this “revisionist” view, traditional SBTC proponents started emphasizing the polarization of the labor market (Goos and Manning (2007), Autor et al. (2005), and Autor et al. (2007)). They claim that information technology (IT) has replaced routine, white-collar jobs and increased the demand for abstract information-processing skills. The theory predicts a diversion of wage inequality in the upper tail of the wage distribution, such as the 90-50th percentile wage differentials, and Autor et al. (2007) in fact confirm the continued widening of the 90-50th percentile wage gap during the 1990s. In addition, a recent Canadian study by Boudarbat et al. (2003) that

examines the change in the wage distribution in the 1990s finds an increasing trend of income inequality, in particular, an increasing return to education. When evaluating the controversy between “revisionists” and “SBTC proponents,” examining what happened to the income distribution in the world’s second-largest economy during the 1990s is important.

Similar to other developed countries, Japan has experienced the penetration of IT (Shimizu (1999), Kohara and Ohtake (2001) and Kawaguchi (2006)) and an increased amount of competition with newly emerging economies, in particular, with neighboring Asian countries (Higuchi and Genda (1999), Head and Ries (2002), Fukao (2002) for the effect of foreign direct investment and Sakurai (2004) for the effect of trade) over the last 15 years. Thus, if SBTC and globalization explain the trends in wage inequality in the US, we should observe a similar wage dispersion in Japan during the 1990s. In fact, Moriguchi and Saez (2006) report that an increase in the concentration of income in the top 0.1 percent of the income distribution was not observed during 80s and 90s in Japan, contrary to the finding from the US.

In contrast to the high level of interest in the recent trend of income inequality in Japan, there has not been a consensus regarding whether income inequality widened during the 1990s and onward. On the one hand, Tachibanaki (2005) claims that income inequality widened during the 1980s and 1990s. In contrast, Ohtake (2005) claims that the increase in income inequality is partly due to the aging population, based on earlier works by Ohtake and Saito (1998) and Iwamoto (2000)²; the degree of income in-

equality is intrinsically high among elderly people, and the aging population mechanically widens income inequality.

Regarding wage inequality, Genda (1998) reports a non-increase in wage inequality during the 1980s and early 1990s, mainly due to a decrease in between-group wage differentials. Shinozaki (2002) analyzes the change in the wage distribution during the 1990s and finds that wage inequality remained the same during the 1990s. Ohtake (2005) points out the decline of the between-group wage differentials and the increase of the within-group wage differentials in the 1990s.³

Given the research results for Japan introduced above, this study focuses on the wage distribution, exploiting the micro data of 1989-2003 from the *Basic Survey on Wage Structure (BSWS)*. Focusing on the wage rate is important because most hypotheses on income dispersions, such as the SBTC hypothesis and the globalization hypothesis, predict a dispersion of the wage rate rather than total income. This study decomposes the log wage variance among finely defined groups of workers and attributes the change in the log wage variance to between- and within- group changes. Then, the entire distribution of the log wage is estimated by a kernel-density estimation and the change in the wage distribution is decomposed into the change in the workers' skill distribution and the change in the factor prices, using DiNardo et al. (1996)'s technique (DFL decomposition, hereafter).

We show that the modest decline in the variance of the log wage in the first half of the 1990s is attributable to the smaller variance between groups

caused by lower returns to education and years of tenure for both sexes. The expansion of the variance among males after 1997 is explained by a larger variance within the group: a larger residual variance in the wage equation, while the shape of the wage distribution had been stable between 1989 and 2003. The DFL decomposition reveals that this stability is the product of two opposing effects; the between-group variance has decreased, while the within-group variance has increased. For females, the log-wage variance increase after the mid-1990s was mainly caused by greater heterogeneity in terms of workers' skill attributes, in particular, years of job tenure. The DFL decomposition further confirms that the change in the skill distribution contributed much to the change in the wage distribution from 1989 to 2003.

Our empirical examination proceeds as follows. Section 2 provides a brief data description, and Section 3 reports the estimation results of the wage equations for full-time workers and examines the change in the estimated parameters on workers' attributes, which is the basis for the remaining explorations. Section 4 performs a variance decomposition of wage rates for full-time workers into the change in the estimated coefficients in the wage equation, the variance of workers' attributes, and the residual variance. Section 5 turns to an application of the DFL decomposition. The last section summarizes our empirical findings and discusses their policy implications.

2 Data

The data set used in this study is micro-level data from the *Basic Survey on Wage Structure (BSWS)*, compiled annually by the Japanese government between 1989 and 2003. This survey includes observations randomly chosen from almost all regions and industries except for agriculture in Japan. The annual number of observations is approximately 1.5 million workers from 60-70 thousand establishments. The sample includes all establishments with 10 or more employees in both private and public sectors and all establishments that belong to private firms with 5 to 9 employees.

The establishments in the sample are randomly chosen in proportion to the size of prefectures, industries, and number of employees from the Establishment and Enterprise Census (*EEC* hereafter) that lists all establishments in Japan. This list is revised every 2-5 years. Of the years relevant to our analysis, the lists were revised in 1986, 1991, 1994, 1996, 1999, and 2001. The *BSWS* 1989-1992 sample is randomly picked from the 1986 *EEC* list, the 1993-1995 sample is from the 1991 list, the 1996-1997 sample is from the 1994 list, the 1998-2001 sample is from the 1996 list, and the 2002-2003 sample is from the 1999 list. While the sampling is based on the same list, about half of the establishments are chosen in two consecutive years, but only about 1/10 of the establishments in the sample are picked at the time of the list revision. We should recognize the large discontinuity of the analysis sample at the times of the list revision: 1993, 1996, 1998, and 2002. The

randomly selected establishments were asked to extract their workers' information from their payroll records.⁴ The establishment and individual files were merged using an establishment identification number.

The unit of analysis is an individual worker with relevant information from the establishment to which he/she belongs. In addition to variables related to wages, the available information includes each worker's age, sex, educational attainment, full-time/part-time status, type of work or job, employment status (with or without permanent status), working days/hours, as well as the firm's attributes, including the number of permanent workers (*Joyo Rodo Sha*)⁵, firm size, industry, and location.

Regarding wages, the individual data include the contracted hours of work and overtime hours between June 1 and June 30, and the total amount of payment for the corresponding period, including overtime pay and allowances, such as those for dependents and transportation. The data also include the total bonus payment between January 1 and December 31 of the previous year. The wage rate in this study is defined as the hourly rate of pay, which is calculated by dividing the total payment in June plus one twelfth of the total bonus payment in the previous year by the sum of the contracted hours of work and the overtime work hours in June.

Table 1 reports the summary statistics of the main variables used in this study. The sample presented in the table is confined to full-time workers (*Joyo Ippan Rodo Sha*) in the private sector.⁶ We proceed with our analysis without placing any limitations on workers' age.⁷

The average nominal wage rate increased until 1994 and then leveled off after that year, which is common to both male and female workers. Although it decreased slightly after 1999, the average wage rate in real terms deflated by Consumer Price Index at the 2000 price was almost constant due to deflation.

The average years of tenure, defined as the years an employee had worked for the current firm, *extended* between 1989 and 2003 by 1.6 years for male and 2.2 years for female workers. Their years of tenure expanded mainly because of the aging of the population and the extension of the mandatory retirement age in the 1990s.⁸

We observe an increase in educational attainment between 1989 and 2003. The share of junior-high-school graduates ($Education \leq 9$) decreased, while that of university graduates ($Education \geq 16$) increased by 10 percentage points, which is common to both male and female workers. The share of senior-high-school graduates ($Education=12$) decreased and that of two-year-college graduates ($Education=14$) rose substantially for female workers, though those shares were mostly unchanged for male workers.

Regarding wage inequality, Figure 1 draws four measures of log wage inequality: the difference between the 90th and 50th percentiles, the difference between the 50th and 10th percentiles, the variance of the log wage, and the Gini coefficient. We pay attention to 90th-50th percentile and 50th-10th percentile differences, instead of the 90th-10th percentile difference, because the non-parallel movement of these trends is one of the main grounds for the SBTC proponents as reported for the US in the 1990s by Autor et al. (2007).

Panels A and B of Figure 1 show that the inequality in the log wage declined until the mid-1990s for all measures, kept the same level in the late-1990s, and started increasing from 2000 for both sexes.⁹

It is rather surprising that the log wage inequality declined throughout the 1990s, while Japanese academics and the media were debating about the issue of increased wages and income inequality (Bungei Shunju (2006) and Ohtake (2005)). In the following decomposition analysis, we offer a solution for this puzzle. We also should note that the 90-50 percentile difference and 50-10 percentile difference move almost in parallel between 1989 and 2003, and this finding contrasts with the finding for the US by Autor et al. (2007).

3 Estimation of a wage equation for full-time workers

In this section, we employ the following wage equation to explore changes in the returns to each attribute of workers' human capital.

$$y_{it} = x_{it}\beta_t + u_{it}, \quad E(u|x) = 0 \quad (1)$$

The dependent variable (y_{it}) is the logarithm of hourly wage rates in real terms. Subscripts i and t refer to the i th individual and year t . The vector x_{it} is a vector of the explanatory variables that are reported in Table 1: dummies for educational attainment, squared potential years of experience, squared years of tenure, the interaction term between potential years of experience and years of tenure, the logarithm of sizes of firms and estab-

lishments, as well as eight industry dummy variables. The error term has a zero conditional expectation. We apply the OLS method to male and female workers separately. All standard errors are calculated to be robust against the presence of heteroscedasticity.

Regression results of the wage equations for males are reported in the first four columns of Table 2. First, we should pay attention to the declining return to education during 1989-2003, as evidenced by the smaller coefficients on the education dummies in the later sample year. In particular, we observe a large decline of the return to two- and four-year college education between 1989 and 2003. This decline may be explained by the increase in the supply of educated workers, as clearly evident in Table 1.¹⁰ The last four columns of Table 2 report the wage regression results for females. We find a sharp contrast for female workers. Although the dummy for senior-high-school graduates ($Education=12$) decreased, as observed in male workers, the coefficient on the dummy for two-year-college graduates ($Education=14$) remained unchanged, and that on the university-graduate dummy ($Education=16$) had a much smaller decline than that of male workers. Second, Table 2 shows smaller coefficients on the years of tenure for both male and female workers. This finding is consistent with the fact that the average years of job tenure increased between 1989 and 2003.

These findings are consistent with what we report in Table 1: A larger proportion of workers with higher educational levels and longer years of tenure made the returns to education and tenure smaller. The average years of

education becomes longer partly due to supply factors, such as an increase in parents' income (Arai (1998)) and an increase in the number of college graduates because the Ministry of Education deregulated new college openings and expanded their capacity (1991 revision of University Establishment Standard (*Daigaku Setti Kijun*)). Workers' job tenure has extended, partly because of the aging of the population and the extension of the mandatory retirement age. The extended years of job tenure among females can be explained by the change in the social norm that now encourages women to stay in the labor force after marriage and child bearing (Kawaguchi and Miyazaki (2007)). These supply shocks may have decreased the equilibrium return to education and job tenure.

4 Decomposition of the variance in the wage rates for full time workers

In this section, we decompose the variance in the wage rates into that within a group and that between groups (i.e., the composition of groups).

The variance in the logarithm of wage rates is decomposed as follows without covariance between x and u because of the assumption $E(u|x) = 0$ in (1).

$$Var(y_t) = \beta_t' Var(x_t) \beta_t + Var(u_t). \quad (2)$$

The change in $Var(y)$ in period τ from the base period 1989 is decomposed

as follows.

$$\begin{aligned}
Var(y_\tau) - Var(y_{89}) &= \beta'_\tau Var(x_\tau)\beta_\tau - \beta'_{89} Var(x_{89})\beta_{89} + Var(u_\tau) - Var(u_{89}) \\
&= [\beta'_\tau Var(x_\tau)\beta_\tau - \beta'_{89} Var(x_\tau)\beta_{89}] \\
&\quad + [\beta'_{89} Var(x_\tau)\beta_{89} - \beta'_{89} Var(x_{89})\beta_{89}] \\
&\quad + [Var(u_\tau) - Var(u_{89})]. \tag{3}
\end{aligned}$$

The first term corresponds to changes in the wage structure that are captured by the changes in the estimated coefficients in the wage equation β . The second term corresponds to the changes in the variance of workers' attributes, which are captured by the change in the variance of the explanatory variables in the wage equation $Var(x_t)$. The last term corresponds to the changes in the variance of the error term $Var(u_t)$.¹¹

Figure 2 reports the results of the decomposition in the change of the variance of log real wage rates for male and female workers. The temporal change of the variance is decomposed into the following three components: 1: the change due to the change in β , 2: the change due to the change in $Var(x_t)$, and 3: the change due to the change in $Var(u_t)$. We performed the decomposition for male and female workers separately.

Figure 2 Panel A reports the decomposition for male workers. The actual coefficient of the variance in the wage rates declined, which is largely explained by the smaller wage variance among groups. This is caused by the smaller disparities among workers with different degrees of educational attainment or years of tenure, as reported in Table 2. At the same time, we

notice that the variance in residuals, which stands for wage inequality within a group, expands after 1997. This can be interpreted as an increase of the return to unobserved skill (Juhn et al. (1993)) or the change in the population weight (Lemieux (2006)). The latter possibility is pursued in depth later in this section.

In sum, we find that the components move in two opposite directions: a smaller variance between groups and a larger variance within groups. The effect of a decrease in the between-group variance dominates the effect of the increase in the within-group variance, and this results in a decrease in the total variance. One might argue that no expansion in wage inequality is observed by emphasizing the changes in the return to skills, but we should notice that this is only one part of the whole picture.

Similar to male workers, the variance of the log wage declined after the 1990s for female workers, as Figure 2 Panel B indicates, and the structural change in β is the main cause of this change. Contrary to male workers, the variance in residuals is on a slightly declining trend, but the variance in x contributes to a larger variance in the wage rates. This suggests that while the disparity between groups was smaller due to the smaller returns of workers' skill attributes to wages, a larger variance in skills expanded the wage disparity for female workers.

Now, we attempt to examine how much of the change in the residual variance is due to the change in workers' skill composition in the labor market in the spirit of Lemieux (2006). To attain this goal, we further decompose the

residual variance into the change in the distribution of x and the change in the mapping from x to the residual variance. To implement the decomposition, we assume the following functional form of heteroscedasticity:

$$\text{Var}(u_\tau|x) = \exp(x\gamma_\tau) \quad (4)$$

Under this assumption, $\text{Var}(u)$ is rewritten as

$$\text{Var}(u_\tau) = E_{x|t=\tau}[\exp(x\gamma_\tau)] \quad (5)$$

This decomposition is identical to the idea of Lemieux (2006). This application of the law of iterated expectation articulates that the change in the residual variance is decomposed into the change in the distribution of x and the change in γ_τ , which stands for structural change. In other words, the residual variance can change mechanically due to a change in the population structure without structural change.

Exploiting this feature, we can calculate the artificial variance that has the variance structure of year τ , but the distribution of the attribute is that of year 1989 as follows, using an argument similar to that of DiNardo et al. (1996):

$$\begin{aligned} \text{Var}(u)_{x=89}^\tau &= \int \exp(x\gamma_\tau) f(x|t=89) dx \\ &= \int \exp(x\gamma_\tau) \frac{P(t=89|x)f(x)}{P(t=89)} \frac{P(t=\tau)}{P(t=\tau|x)f(x)} f(x|t=\tau) dx \\ &= E_{x|t=\tau}[\theta \exp(x\gamma_\tau)] \end{aligned} \quad (6)$$

where $\theta = \frac{P(t=89|x)}{[1-P(t=89|x)]} \frac{[1-P(t=89)]}{P(t=89)}$. The numerator in the first term is the propensity score to be in the 1989 sample, given x . For example, consider a

dummy variable for senior-high-school graduates as a part of x . The propensity score for 1989 is higher than that for 2003 among high-school graduates because the share of senior-high-school graduates was larger in 1989 in the labor market. The proportion $P(t = 89)$ is the share of the 1989 sample among all observations.

Using the counterfactual variance above, the change in the residual variance can be decomposed as

$$Var(u)_{x=\tau}^{\tau} - Var(u)_{x=89}^{89} = [Var(u)_{x=\tau}^{\tau} - Var(u)_{x=89}^{\tau}] + [Var(u)_{x=89}^{\tau} - Var(u)_{x=89}^{89}]. \quad (7)$$

The first term corresponds to the change in the residual variance due to the distributional change of x , and the second term corresponds to the change in the residual variance structure. In other words, the first term results from the change in the population weight, while the second term is interpreted as the within-group change because the change occurs within a group indexed by x .

To implement the above decomposition of the variance of wage rate residuals, we pool the observations in 1989 and year τ and apply a probit estimation to regress the dummy variable that takes 1 if the observations are in 1989 sample on x_i . The propensity score from this probit regression is used to calculate the estimated value of $\theta_{\tau i} (= \hat{\theta}_{\tau i})$. The residual of the wage regression \hat{u}_i is taken from the wage regression whose results are reported in Table 2 and discussed in the previous section. Then, the logarithm of \hat{u}_i^2 in year τ is regressed on x_i to obtain coefficients γ_{τ} . The results of this regres-

sion are reported in Table 3. We calculate the exponential of the estimated value, which corresponds to $\exp(x_i\gamma_\tau)$ in (4) (we call this \hat{v}_i^τ). Based on this predicted residual \hat{v}_i^τ , we calculate \hat{V}_{89}^τ as the weighted average of \hat{v}_i^τ in year τ using $\hat{\theta}$ as the weight.

Figure 3 presents the results of the decomposition expressed in the equation (7) for male and female workers, respectively. Figure 3 Panel A shows that the residual variance declined in the first half of the 1990s for male workers and then began to increase after 1997. The “V-shaped” trend is also observed even after making the distribution of x constant at its 1989 level. This suggests that the variance in residuals expanded, even after removing the effect of the skill distribution change. This finding contrasts with the finding for the US by Lemieux (2006). The increase in the residual variance may suggest an increase in the return to unobserved skills after 1997, as pointed out by Juhn et al. (1993) in the US context.

Figure 3 Panel B shows that the residual variance for females has a stable trend, but if the skill distribution were that of 1989, the residual variance would have declined monotonically. This indicates that the residual variance increased in the late 1990s, mainly due to a shift in the population weight toward groups with intrinsically larger residual variance. This is natural because more-educated and longer-experienced workers tend to have a higher within-group variance, as evidenced by Table 4. After removing the effect of full-time workers’ compositional change, we can conclude that the within-group residual variance was stable among female workers throughout

the 1990s. This result contrasts with the result for males, and this may be because of the increase in the real value of the minimum wage, as pointed out by Abe and Tamada (2007). The real value of the minimum wage steadily increased in the late 1990s, and this may well have contributed to the compression of the lower end of the wage distribution.

Overall, our findings demonstrate that there were two opposing trends for the log wage variance. One is declining wage inequality across groups, mainly due to the declining returns to education and tenure. The other, opposite trend is increased inequality within a group of workers indexed by the workers' attributes vector x among males and the increased heterogeneity of the labor force (increased $Var(x)$) among females. The first trend dominates the second trend and, as a result, the overall log wage variance declined during the first half of the 1990s for both sexes. Behind the declining log wage variance, within-group variance increased for males, and this increase in within-group variance cannot be explained by a shift in the population weight due to such causes as aging.

This phenomenon for male workers could well explain why Japanese people have a nagging sense of increased inequality (Genda (2006)), although we cannot confirm it from the trend of the aggregate statistics. As Clark and Oswald (1996) show, people tend to care more about their relative wage position within a reference group rather than their position in an aggregate distribution. As for females, behind the declining log wage variance, the variance has increased due to the increased heterogeneity of female workers.

This can be understood as a transitional phenomenon. In 1989, females uniformly had relatively low skills, but some females started to have high skills, such as more years of education or job tenure. As a result, the female labor force became more heterogeneous in 2003.

5 Changes in the wage distribution: The DiNardo, Fortin, and Lemieux decomposition

This section examines changes in the wage distribution using a kernel-density estimation. The merit of the kernel-density estimation is that we can confirm the change in the shape of the wage distribution without sacrificing any information. We further decompose the change in the distribution into the part due to the distributional change of the attributes x and the part due to the relational change between the attributes x and wage rates, employing the DFL decomposition (DiNardo et al. (1996) and Lemieux (2002)).

We implement the kernel-density estimation using the Epanechnikov kernel and an optimal bandwidth.¹² The kernel density estimation procedure is fairly well known, and we do not illustrate the procedure here.

The wage distribution in 1989 can be understood as the product of the relation between wage and attributes and the distribution of x as follows:

$$f^{1989}(y) = \int f^{1989}(y|x)h(x|t = 1989)dx, \quad (8)$$

where $f^{1989}(y|x)$ is the wage determination mechanism in 1989 that maps workers' and firms' attributes x to the distribution of the log wage, which is

denoted as y . The probability density $h(x|t = 1989)$ indicates the distribution of attributes in year 1989. Similarly, the distribution of wage in 2003 is

$$f^{2003}(y) = \int f^{2003}(y|x)h(x|t = 2003)dx. \quad (9)$$

The counterfactual wage distribution that is determined by the product of the wage determination mechanism in 2003 and the attributes distribution in 1989 is expressed as

$$f_{1989}^{2003}(y) = \int f^{2003}(y|x)h(x|t = 1989)dx. \quad (10)$$

This counterfactual wage distribution is calculated by DiNardo et al. (1996)'s method.¹³ This method is fairly well known and thus we do not illustrate this procedure either.

We note that the gap between the counterfactual distribution and the actual distribution in 1989 is captured by the change in the wage determination mechanism (change in β in equation (1)) and the residual distribution. In contrast, the gap between the counterfactual distribution and the actual distribution in 2003 is caused by the change in the distribution of workers' attributes (the change in the distribution of x in equation (1)), holding β and the residual distribution constant.

Figure 4 reports the actual wage distributions in 1989 and 2003, and the counterfactual wage distribution, assuming that workers' and firms' attributes had remained at their 1989 levels. Several interesting findings emerge as follows.¹⁴

First, for all years, the actual 2003 distribution is located to the right of the 1989 distribution, and this indicates that the real wage increased during the sample period. In addition, the counterfactual wage distribution, if the skill distribution is fixed at 1989 level, is located to the right of 1989 actual distribution. This shift comes from the change in the skill price between 1989 and 2003. This tendency is particularly strong at the lower end of the wage distribution. This indicates that there had been some skill price changes that compressed the wage distribution at its lower end. The increase in the real value of the minimum wage due to deflation in the late 1990s and early 2000s may partly explain this, particularly for females.¹⁵

Second, the deviation between the actual distribution and the counterfactual distribution is particularly large in the top half of the wage distribution. This gap implies that the emergence of this chunk of high-wage workers for both sexes cannot be explained by the change in skill prices. Thus the increase of high-wage workers is likely to be caused by an increase in the number of workers who have high skills.

Third, the degree of the shift in the wage distribution to the right is larger among females than among males during the period. The faster wage growth of female workers than male workers is consistent with the gender wage convergence among full-time workers during the 1990s (Kawaguchi (2005)). As the relative positions of the actual wage distributions and the counterfactual distributions show, the gender wage convergence occurred partly because of the convergence in attributes x and partly due to the change in skill prices.

We do not necessarily find that the gender wage convergence was caused by the rapid wage growth of highly paid workers; rather, the convergence happened because of an arguably uniform shift of the wage distribution among female workers.

6 Conclusion

This paper examined the change in the wage distribution among full-time workers during the 1990s in Japan. We take advantage of a rich, micro-level data set from the *Basic Survey on Wage Structure* (1989-2003) to perform an in-depth analysis of the changes in the variance and distribution of the wage rate.

Although simple aggregate statistics may give the impression that wage inequality did not change during the period, the decomposition analysis reveals that the seemingly steady trend is a product of two opposing trends: 1. declining between-group (defined by education, experience, tenure and firm/establishment size) wage inequality for both sexes, and 2. increasing within-group inequality among male workers and increasing observed heterogeneity among female workers.

The declining between-group wage differentials are largely due to the decline of the return to education and job tenure, perhaps because of the increase in college-educated and long-tenured workers. Our results show that the aging of the population in itself cannot explain the recent increase in within-group wage inequality among males. As for females, gradual increases

in college-educated and long-tenured workers make the female full-time workers more heterogeneous. This contributes to the increase of wage inequality among them.

The increase of within-group inequality among male workers may well explain why the general public in Japan feels that inequality has increased in the last 15 to 20 years, though we can hardly confirm it in the aggregate statistics. An individual presumably perceives inequality by comparing his/her wage within a group to which he/she belongs rather than comparing it to the total distribution. The increase in the heterogeneity of female workers is consistent with the emergence of career-oriented women in the last 20 years, and this trend will continue mechanically as long as these career-oriented women stay on the current track.

Several research topics are left for future research. First, the explicit demand-supply analysis should be implemented, and the results should be reconciled with results from other developed countries. Second, how the increased within-group inequality among males and the increased heterogeneity among females translate into household-level income inequality through marriage should be investigated. Third, if an increase in temporal income inequality is associated with an increase in temporal income mobility, lifetime income may not become unequal. The change in temporal mobility should be further investigated. These additional considerations are indispensable for designing social welfare programs because current low-wage workers do not necessarily belong to low-income households in a dynamic sense.

Acknowledgment

An earlier version of this paper was prepared as background work for the Ministry of Economy, Trade and Industry (METI) led by Kyoji Fukao and was published as a part of Kawaguchi et al. (2006). We thank the METI and the Ministry of Health, Labor and Welfare for providing us with the micro-level data. We appreciate comments from two anonymous referees, Dwayne Benjamin (editor), Yukiko Abe, Joe Chen, Kang Shik Choi, Hiroyuki Chuma, Yoshio Higuchi, Hidehiko Ichimura, Jin Hwa Jung, Miki Kohara, Ayako Kondo, Hiroshi Ohashi, Isao Ohashi, Yoshiro Tsutsui, Fumio Ohtake, Yoshiyasu Ono, Takehisa Shinozaki, Satoshi Shimizutani, Tsuyoshi Tsuru, Yaohui Zhao, and seminar participants at Hitotsubashi University, Osaka University, the 2007 WEAI Pacific Rim Conference at Peking University, the University of Tokyo, the Spring 2007 Japan Economic Association Meeting at Osaka Gakuin University, the National Institute of Social Security and Population, Tohoku University, and the 2007 EALE annual conference at the University of Oslo. We also appreciate editorial assistance from Donna Maurer.

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Notes

¹Gottschalk and Smeeding (1997) also reports heterogeneous trends in income inequality across developed countries in the 1980s.

²Ohtake and Saito (1998) used *National Survey on Family Income and Expenditure: 1979, 84, 89* and Iwamoto (2000) used *Comprehensive Survey of Living Conditions of the People on Health and Welfare: 1989-95*. See Oshio (2006) for another discussion.

³The stable wage distribution found by Shinozaki (2002) and the increased income inequality found by Ohtake (2005), Tachibanaki (2005) and Oshio (2006) can be reconciled by the decline of work hours among the low-income group. According to the Comprehensive Survey of Living Conditions of the People on Health and Welfare, the average weekly work hours among males in households that earned more than 10 million yen annually had been stable, around 7 hours, between 1981 and 2001, while male workers from households with annual incomes of 3 million yen or below declined from 6.5 to 4 hours during the same period. We thank Miki Kohara for pointing this out.

⁴A person in charge of personnel matters in each establishment was asked to randomly choose a number of workers from its pool of employees based on the given instructions for random sampling, including the sampling probability, which depended on the establishments size and industry.

⁵Those workers who satisfy one of the following three criteria are classified as permanent workers: 1. On contracts that do not clearly specify a contractual time period, 2: On contracts that last more than a month, or 3: On contracts that last less than a month, but on which the workers worked 18 or more days in the last two months. This classification includes part-time workers if one of the criteria above is satisfied.

⁶The proportion of full-time workers without a explicit contract period consists of 81% of total employees among males in 1992. The same figures are 75% for males in 2002, 67% for females in 1989, and 47% for females in 2003. The calculation is based on *Basic Survey of Employment Structure*.

⁷Our results are unchanged even after limiting the sample to workers

whose ages are 15-65.

⁸Setting the mandatory retirement age below 60 became illegal in 1994 (Araki (2002)). According to the Survey of Employment Management, in 1989, 93.0 percent of firms set a certain age as the mandatory retirement age. The number gradually increased to 96.9, 97.1, and 97.5 percent in 1994, 1999, and 2003, respectively. Among the firms that set a mandatory retirement age, those that set the mandatory retirement age at or above 60 was 61.9, 84.1, 99.2, and 98.9 percent in 1989, 1994, 1999, 2003, respectively.

⁹Panels A and B of Figure 1 show a jump in all measures for males and the 50-10 percentiles difference for females in 1993. This sudden jump in the inequality measure was probably due to the revision of the list of establishments. We attempted without success to correct for this gap by using a re-weighting procedure. This point should be noted as a caveat.

¹⁰Noro and Ohtake (2006) found a stable trend in the four-year- college / high-school graduatesf wage differentials in their most-preferred specification after controlling for the supply increase in educated workers. Thus, the declining return to education can be explained by the rapid increase of the four-year- college graduate supply.

¹¹An alternative decomposition is possible as:

$$\begin{aligned} Var(y_\tau) - Var(y_{89}) &= [\beta'_\tau Var(x_{89})\beta_\tau - \beta'_{89} Var(x_{89})\beta_{89}] \\ &+ [\beta'_\tau Var(x_\tau)\beta_\tau - \beta'_\tau Var(x_{89})\beta_\tau] \\ &+ [Var(u_\tau) - Var(u_{89})]. \end{aligned}$$

This decomposition renders qualitatively and quantitatively similar results.

¹²The choice of kernel function and bandwidth does not essentially affect the results because of the extremely large sample.

¹³The propensity score for calculating the sample weight was estimated by the Logit model.

¹⁴We implemented the Kolmogorov-Smirnov test to formally test the null hypothesis that the difference between the distributions is due to the sampling error. For all years and sexes, the null hypotheses were strongly rejected.

¹⁵The Japanese economy experienced deflation in the late 1990s and early 2000s. Overall, the price level remained almost unchanged between 1993 and 2003. In contrast, the minimum wage has increased almost every year and it cumulatively increased by about 13 percent between 1993 and 2003.

Table 1: Basic statistics

Sample: Full-time workers in establishments that hire 10 or more employees (full- and part-time workers combined) or in single establishments of firms that hire 5-9 employees.

Sample	Male		Female	
	1989	2003	1989	2003
Year	1989	2003	1989	2003
Wage Rate	20.73	26.69	12.05	17.42
(100 Yen)	(14.24)	(16.09)	(12.40)	(11.45)
Real Wage Rate	23.08	27.18	13.42	17.74
(100 Yen, 2000 CPI)	(15.85)	(16.39)	(13.81)	(11.66)
Potential	21.29	21.86	18.51	19.07
Experience	(12.45)	(12.35)	(14.05)	(13.35)
Tenure	12.48	14.06	7.30	9.51
	(10.01)	(11.02)	(7.39)	(8.71)
Firm Size	1431	1391	1105	1140
	(1920)	(1836)	(1758)	(1704)
Establishment	348	302	223	228
Size	(1012)	(854)	(647)	(642)
Education≤9	0.24	0.08	0.23	0.07
Education=12	0.52	0.51	0.57	0.51
Education=14	0.04	0.09	0.15	0.28
Education≥16	0.20	0.31	0.04	0.14
Mining	0.02	0.01	0.01	0.00
Construction	0.06	0.09	0.02	0.03
Manufacturing	0.46	0.39	0.45	0.32
Utilities	0.02	0.03	0.01	0.01
Transportation	0.15	0.11	0.04	0.05
Wholesale	0.09	0.08	0.11	0.10
Finance	0.05	0.05	0.11	0.10
Real Estate	0.01	0.01	0.01	0.01
Services	0.14	0.23	0.24	0.38
Observations	771824	720047	345645	267365

Note: Standard deviations are in parentheses. Observations are available every year between 1989 and 2003, but the descriptive statistics are reported for only two years.

Table 2: Estimates of wage equations for full-time workers

Sample	Male		Female	
Year	1989	2003	1989	2003
Education=12	0.144 (0.001)	0.114 (0.002)	0.155 (0.002)	0.123 (0.003)
Education=14	0.278 (0.002)	0.217 (0.002)	0.297 (0.002)	0.309 (0.003)
Education \geq 16	0.459 (0.001)	0.404 (0.002)	0.496 (0.004)	0.471 (0.004)
Experience	0.042 (0.000)	0.039 (0.000)	0.008 (0.000)	0.012 (0.000)
Experience ² /100	-0.073 (0.000)	-0.066 (0.000)	-0.012 (0.001)	-0.024 (0.001)
Tenure	0.028 (0.000)	0.023 (0.000)	0.056 (0.000)	0.042 (0.000)
Tenure ² /100	-0.018 (0.001)	-0.005 (0.001)	-0.018 (0.001)	-0.008 (0.001)
Exper·Tenure/100	-0.004 (0.001)	-0.012 (0.001)	-0.069 (0.001)	-0.048 (0.001)
ln (Establishment Size)	0.021 (0.000)	0.029 (0.000)	0.032 (0.001)	0.047 (0.001)
ln (Firm Size)	0.059 (0.000)	0.058 (0.000)	0.064 (0.001)	0.037 (0.001)
Observations	771824	720047	345645	267365
R-squared	0.65	0.60	0.54	0.51

Note: Robust standard errors in parentheses. The reference category for educational attainment is junior-high-school graduates. Industry dummy variables and a constant are included but the results are not reported. The wage equations are estimated based on the sample every year between 1989 and 2003, but the regression results are reported for only two years.

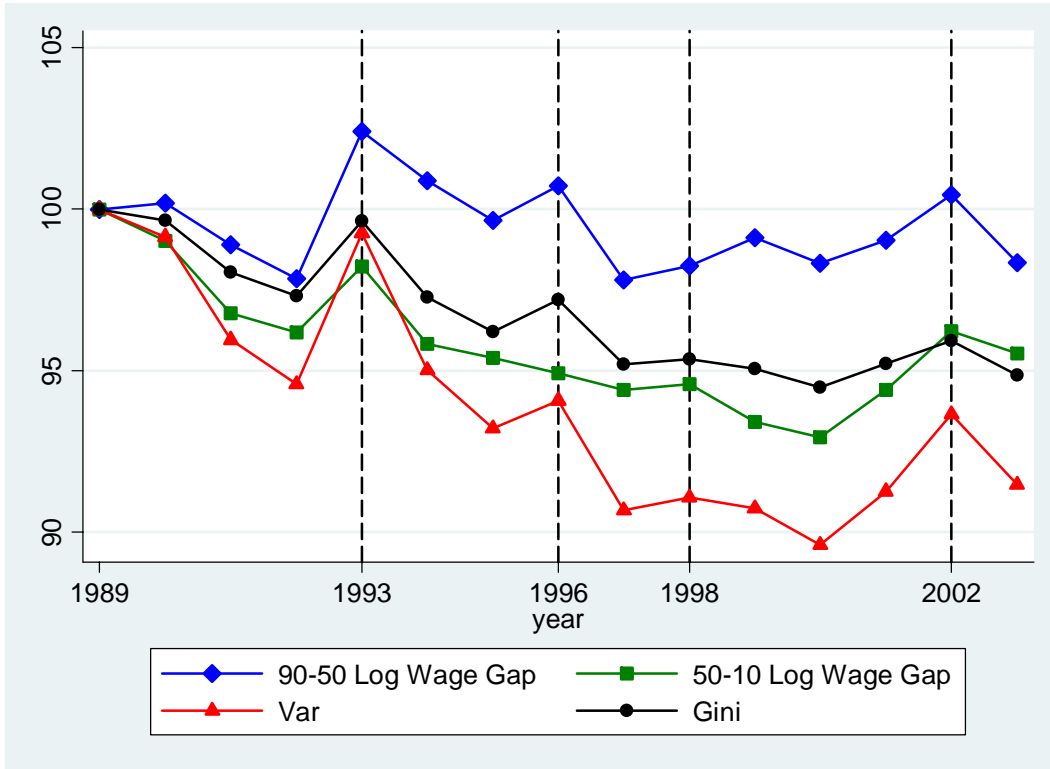
Table 3: Regression of the residuals squared in the log wage equation on explanatory variables

Sample	Male		Female	
	1989	2003	1989	2003
Year	1989	2003	1989	2003
Education=12	0.233 (0.007)	0.180 (0.011)	0.340 (0.011)	0.232 (0.020)
Education=14	0.256 (0.015)	0.268 (0.014)	0.394 (0.016)	0.222 (0.022)
Education≥16	0.538 (0.010)	0.537 (0.012)	0.785 (0.022)	0.550 (0.024)
Experience	0.051 (0.001)	0.063 (0.001)	0.078 (0.001)	0.079 (0.002)
Experience ² /100	-0.059 (0.002)	-0.061 (0.002)	-0.156 (0.003)	-0.153 (0.004)
Tenure	-0.059 (0.001)	-0.056 (0.001)	-0.087 (0.002)	-0.070 (0.002)
Tenure ² /100	0.029 (0.004)	0.022 (0.004)	-0.117 (0.006)	-0.087 (0.007)
Exper·Tenure/100	0.109 (0.004)	0.085 (0.005)	0.383 (0.008)	0.305 (0.008)
ln (Establishment Size)	0.014 (0.003)	0.022 (0.003)	-0.039 (0.004)	0.010 (0.005)
ln (Firm Size)	-0.100 (0.002)	-0.054 (0.002)	0.004 (0.003)	-0.006 (0.003)
Observations	771824	720047	345645	267365
R-squared	0.04	0.04	0.04	0.04

Note: Robust standard errors in parentheses. The reference category for educational attainment is junior-high-school graduates. Industry dummy variables and a constant are included but the results are not reported. The residual squared equations are estimated based on the sample every year between 1989 and 2003, but the regression results are reported for only two years.

Figure 1: Change in the log wage inequality

Panel A: Male



Panel B: Female

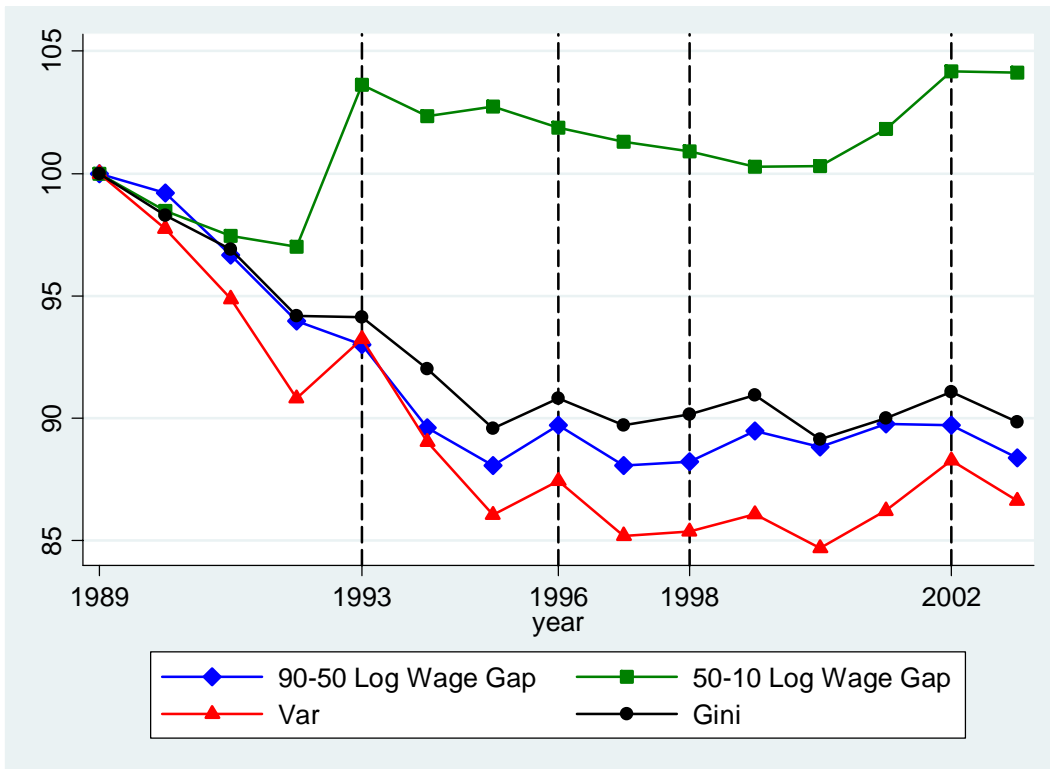
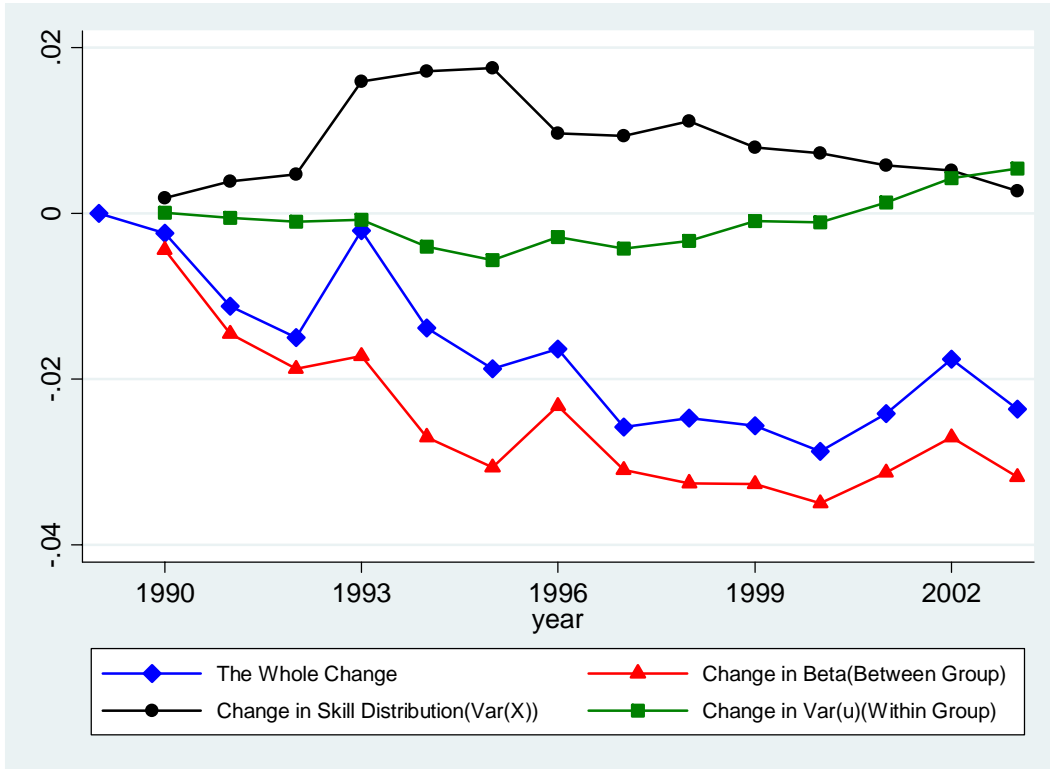


Figure 2: Decomposition of the variance of log wage
 Panel A: Male



Panel B: Female

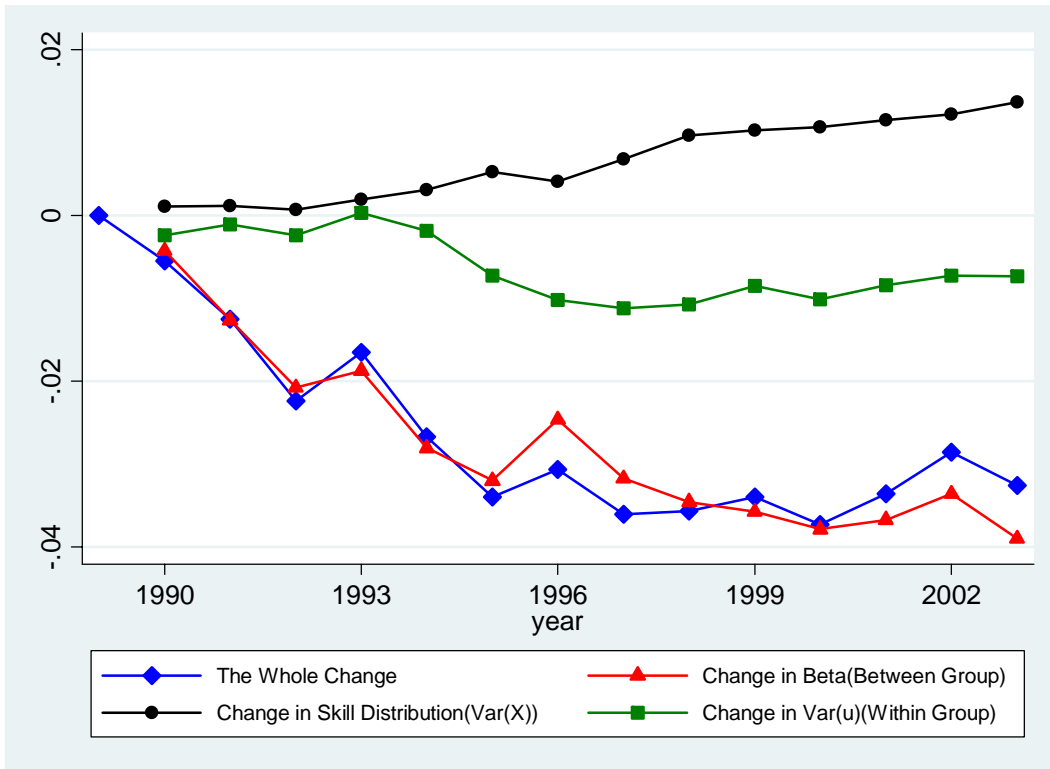
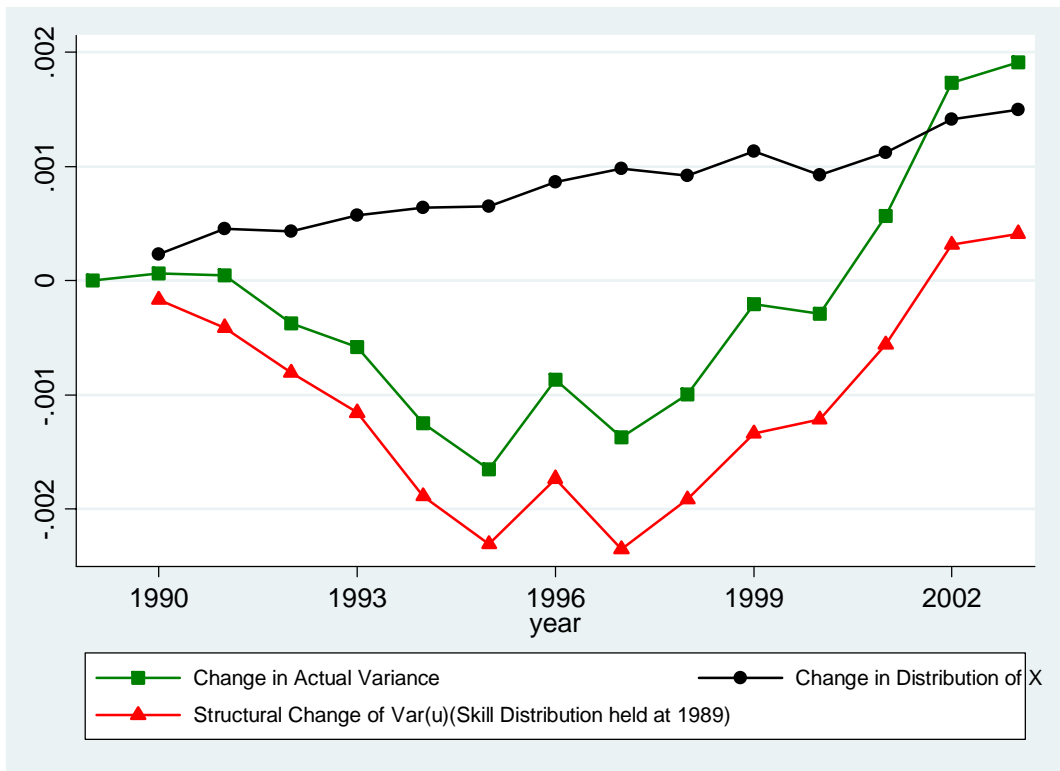


Figure 3: Decomposition of the residuals in the variance of the wage rate

Panel A: Male



Panel B: Female

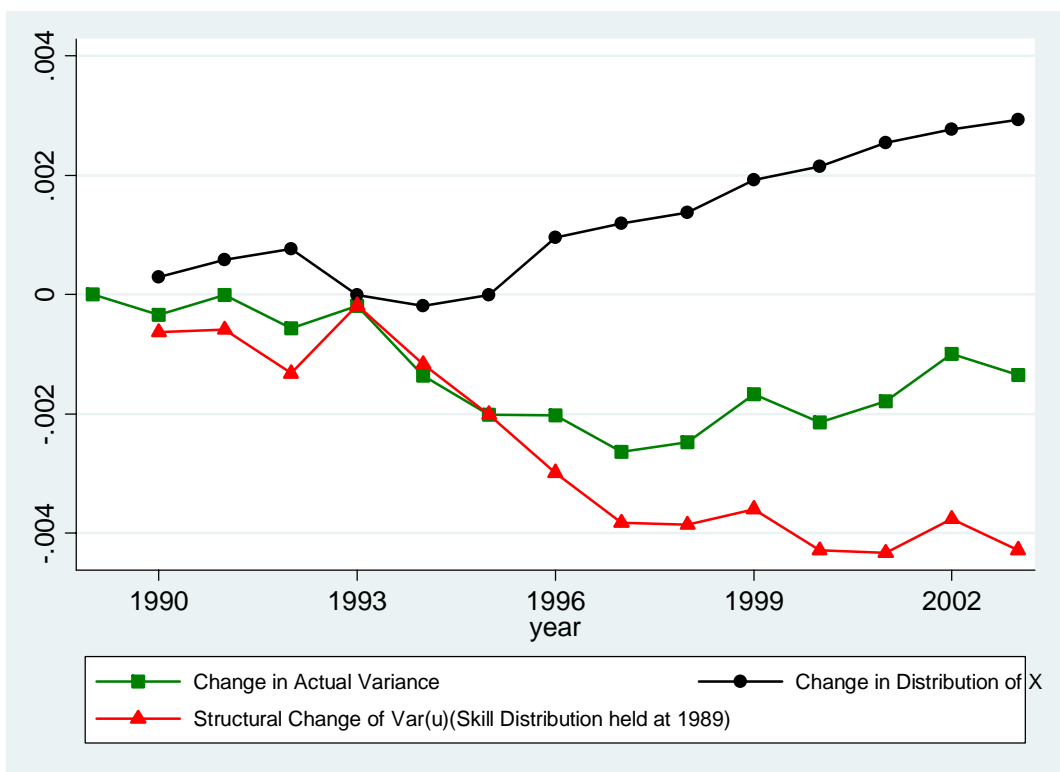
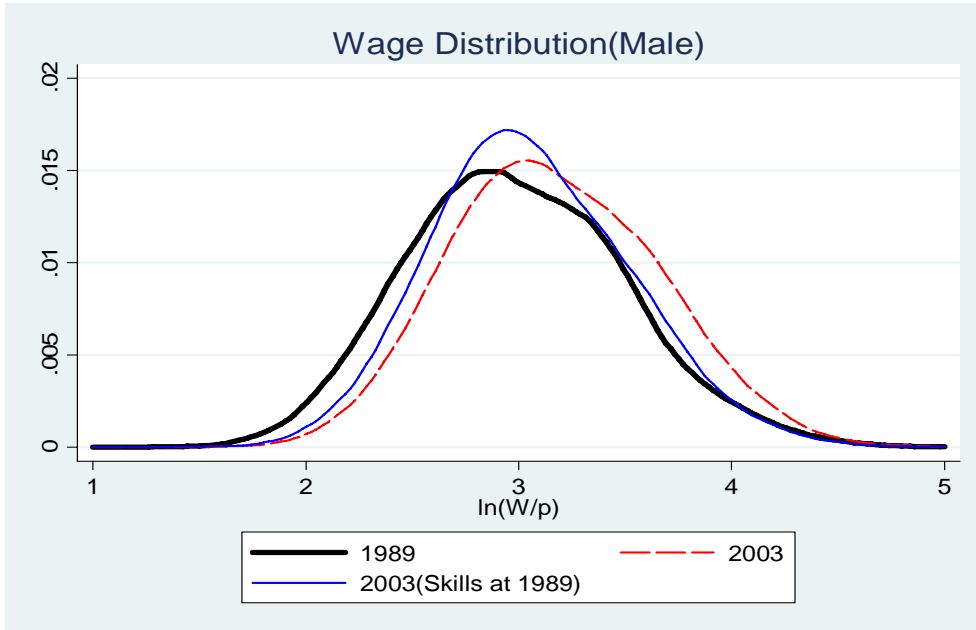


Figure 4: Real hourly wage distributions for full-time workers

Panel A: Male workers



Panel B: Female workers

