remains unexplained, after controlling for numerous characteristics and rank. Rank is found to be the largest contributor to the gender salary gap, accounting for 40% of the raw gender salary gap.

Ginther and Hayes (2003) uses data from the Surveys of Doctorate Recipients in the US. They find a raw gender salary gap of 11.3% in 1993. Most of this gap is explained away when they control for rank, and various personal, institutional and productivity characteristics, leaving only a negligible amount unexplained. However, they find that being female decreases the probability of being promoted to tenure by as much as 6.8%. More recently, Blackaby et al. (2005) estimate the gender pay gap in UK economics departments by using a sample of 291 male and 60 female. They find that females receive 9.4% less than males, after controlling for ethnicity, publications weighted by the quality of the journal, age, education, and various institutional characteristics, but without controlling for rank. When rank is controlled for, the gender salary gap sunstantially reduces to 5.5%, but it remains statistically significant. Thus, most of the previous studies from the US and the UK found that much of the gender salary differences in academia stems from differences in rank attainment between males and females.

4 Compensation scheme in Japanese universities

In this section we provide background information regarding the compensation scheme in Japanese universities. There are three types of universities in Japan: national, public and private. National universities are established and funded by the central government. Public universities are established by local governments, and funded by both the local and the central governments. Private universities are established by private entities and are financially self-supporting. The total annual salary of academics includes: monthly salary for 12 months, and, depending on university, bonuses and other allowances[6]. Bonuses are typically

[6]Salary does not include research funds, travel funds for research or allowances for housing and dependants. However, it does include transportation and adjustments for the cost of living in certain cities.

provided twice a year, and the total amount can be the equivalent of 4-6 month of salary.

Salary in academia is determined based on a relatively rigid payment scheme, called the salary table. The salary table is set by each university based on negotiations between the university and its own union[7]. There could be a few different salary tables depending on the types of employment contract within a university. For example, there could be one type of salary table for those hired on the typical life-time employment basis, and a different type of table for those hired on a fixed-term basis. Although all types of universities have the freedom to set their own salary table, national universities typically follow the guidelines provided by the National Personnel Authority (NPA) (*Jinjiin*)[8]. Most private universities set their own salary tables, however, there are some private universities that follow the NPA guidelines.

Table 5 presents an excerpt from a salary table guideline provided by the NPA. The salary table shows how salaries progress depending on rank. A full-time academic usually starts as a lecturer or assistant professor, and then moves up the ladder to the rank of full-professor. The salary table contains classes and divisions, where the former refers to rank. Class 5 is a full-professor, class 4 is an associate professor, class 3 is an assistant professor, and class 2 is a lecturer. Each class (rank) contains a number of divisions, or payment scales, with each academic assigned to a division. The assignment is typically decided by the university's personnel division run by non-academic staff. The precise criteria that are used to determine the initial division and how each academic progresses through divisions are not specified in the salary table. There is a common belief among academics that the initial division (that is, at hiring) is determined by age, experience, and education level only, and that, each academic progresses one division annually. If this belief is true, then salary would be a deterministic function of age, experience, and education, leaving little room for

[7]Each university has its own union.

[8]The National Personnel Authority is a specialized, neutral, third-party organization dealing with public employee management.

gender salary differences. However, to date there has been no research investigating whether the belief is indeed true.

In April 2004 national and public universities were 'corporatized' based on the *National University Corporation Law* enacted in 2003. Prior to this, the salary tables at national and public universities were determined according to public servants' payment schemes, set by the government. Thus, salary tables were the same at every national university in the country, notwithstanding some allowances that varied by university. After the 'corporatization', national and public universities gained the freedom to set their own salary tables. However, as mentioned above, national and public universities typically follow the guidelines provided by the NPA. Their transformation into corporations meant a change in the legal status of academics, by removing their public employee status, and by allowing university management more freedom in setting compensation schemes. However, 'corporatization' did not mean a change in ownership of universities, and, national and public universities are still owned and sponsored by the government.

5 Empirical methodology

Theories of discrimination suggest that female academics may earn lower salary than male academics with comparable productive characteristics. In order to estimate gender salary differences, we control for various productivity characteristics of each academic. Two methods are typically used to empirically investigate the gender salary gap. On the one hand, there is the Oaxaca-Blinder decomposition (Oaxaca 1973; Blinder 1973), which requires running separate earning regressions for males and females. On the other hand, one can estimate a 'Mincer type' human capital earnings regression (Mincer 1974) for the pooled sample with a female dummy included.

Although the Oaxaca decomposition method has the advantage that it allows all the coefficients to be different for males and females, its application becomes difficult when the number of females is small, as has occurred in previous studies. For example, Blackaby et al. (2005) used the pooled regression method since their sample only contains 60 female. As will be noted in section 6, our data contains 337 academics out of which 58 are female. We, therefore, consider that the number of females is not large enough to apply the Oaxaca decomposition. Thus, we have chosen the pooled regression method for our empirical analysis.

Below is our earning equation:

$$Log(Annual\ Salary) = \alpha' Z_i + \beta_1 (AssocProf)_i + \beta_2 (FullProf)_i + \gamma (Female)_i + \epsilon_i \quad (1)$$

Z_i is the vector of variables that directly affect annual salary (i.e., human capital characteristics and other objective salary determinants). $(AssocProf)_i$ is a dummy variable that takes the value 1 if the academic is an associate professor, 0 otherwise. $(FullProf)_i$ is a dummy variable that takes the value 1 if the academic is a full-professor, 0 otherwise. $(Female)_i$ takes the value 1 if the academic is a female and 0 otherwise. Thus, the coefficient γ captures the gender salary gap, after controlling for human capital characteristics and other objective salary determinants.

This method of estimation faces at least three econometric challenges. First, human capital variables may be noisy and biased proxies for actual human capital accumulation. For example, education is a widely used human capital proxy. If the actual human capital is lower for females than males at a given level of education, we tend to overestimate the gender salary gap. In a study of the white-black wage gap, Neal and Johnson (1996) show that education is a biased proxy for human capital, and find that, when unbiased human capital measures such as AFQT[9] scores are used, the estimated salary gap decreases. In our paper, we tackle this problem by incorporating detailed productivity characteristics for each academic, such as publications and the amount of external grants obtained, in

[9] Armed Forces Qualification Tests (US)

addition to traditional measures of human capital such as education and experience. In fact, one of the benefits of analyzing the academic labor market is that, besides the traditional measures of human capital, various other productivity characteristics can be controlled for, thus minimizing potential biases.

Second, rank attainment may be affected by discrimination. In such a situation, the female dummy variable does not capture the combined effect of discrimination stemming from both salary discrimination and rank attainment discrimination. One way to mitigate this problem is to estimate the earning equation without the rank variables (McNabb and Wass 1997; Ward 2001; Moore at al. 2007). When the rank variables are not included, the earning equation can be thought of as a reduced-form equation, in which case, the female dummy captures the total effect of discrimination stemming from salary and promotion discrimination combined. Another method is to estimate a salary equation and an ordered logit rank equation, separately, as below:

$$Salary\ equation: Log(AnnualSalary) = \alpha' Z_i + \beta(Rank)_i + \gamma(Female)_i + \epsilon_i \tag{2}$$

$$Rank\ equation: y_i^* = \beta' Z_i + \theta(Female) + \mu_i \tag{3}$$

where (Rank)=1 if associate professor, 2 if full-professor and 0 if assistant professor and below. y^* is a latent variable such that, $(Rank)_i$=0 if $y_i^* < c_1$, $(Rank)_i$=1 if $c_1 \leq y_i^* < c_2$, and $(Rank)_i$=2 if $y_i^* \geq c_2$. Therefore, the total salary gap stemming from discrimination in salary and promotion *combined* can be written as:

$$Total\ gender\ salary\ gap = \gamma + \beta[P(Rank = 2|Female) - P(Rank = 2|Male)] \tag{4}$$

where $P(Rank = 2|Male)$ is the probability that a respondent is a full-professor given that the respondent is a male. $P(Rank = 2|Female)$ is defined in a similar manner. These probabilities can be computed at the sample average of all the explanatory variables. β is the coefficient for the rank variable in equation (2) above. Thus, if females have lower probability of being promoted to full-professor, the second term of (4) will capture the drop in salary due to the lower rank attainment for females. This method has the advantage of being able to directly estimate the difference in promotion probability. In this study, we use both methods described above.

Third, self-selection into the academic labor market might be a potential source of bias in the female coefficient. Heckman (1998) shows that the estimated wage gap between blacks and whites reduced in the 1990s, not because discrimination disappeared, but because a considerable number of blacks, who potentially earn lower wages, dropped out of the labor force. Similarly, in our case, if female graduates whose potential salary is lower in academia decide not to enter the academic labor market, we would underestimate the potential gender salary gap. The available techniques to correct for selection bias, such as the Heckit model, require information about the graduates who did not join academia. Unfortunately, we do not have such data. Thus, direct elimination of selection bias is not possible in our case[10]. However, we investigate whether selection bias is a problem by utilizing statistics for male and female PhD graduates in Japan over the period 1969 to 2007[11].

6 Data

Data utilized in this project have been obtained from a survey we administered via a postal questionnaire[12]. Past studies used publicly available data (Broder 1993; McNab and Wass 1997; Ginther 2004; Blackby et al. 2005) or undertook independent surveys (Ward 2001). In Japan, salary information is confidential and there are no national or private statistics on academics collected regularly. Therefore, we undertook a mail survey in order to collect data. Our survey method is presented below.

[10]Similarly, the majority of papers on gender salary differentials in academia in the US and the UK were not able to correct for sample selection bias.

[11]Unfortunately we only have statistics for PhD graduates from Japan.

[12]The questionnaire was distributed only in Japanese. A copy in Japanese (or its English translation) is available from the author upon request.

First, from the website of the Ministry of Education, Culture, Sports, Science and Technology (MEXT) we obtained an official list of all four-year universities in Japan. The list contained 747 universities: 87 national, 76 public and 584 private universities. We accessed each university website provided in the list[13] in order to collect the names of academics in economics and economics-related departments[14]. Due to the facts that some universities do not list faculty names and some universities do not have economics departments, we were able to collect only 4353 names from only 132 universities. Many Japanese economics departments also employ faculty specializing in language education. We eliminated such faculty where possible. In addition, we excluded universities that accept only female students.

Next, from the 4353 collected names, we selected 1863 academics and mailed them questionnaires directly. Ideally, the selection method should be random. However, this could have led to a very small female sample. In order to increase the number of female observations, we selected all the female-sounding names (287 names). Nonetheless, the rest of the selected academics (1576 names) are randomly chosen. Questionnaires were sent from April to June 2008 and participants could reply either by mail or online. Two remainders were sent by mail in July and August, and an additional reminder was sent to approximately 600 academics by email. At the end of our survey period, we received 363 responses (252 by mail and 111 online). Thus, we achieved a rate of response of 19.5%. This response rate is not too high but not too low either, as compared to other previous studies that used similar mail surveys of academics; for example, Moore et al. (2007) achieved a response rate of 13%, while Ward (2001) obtained a response rate of 30%. Our sample contains 299 male and 64 female; however, due to some incomplete responses, the usable sample is 337 (of which 58 are female). Our female sample is comparable in size with female samples previously used in the literature. Broder (1993) used a sample with 30 female and 362 male; Ginther (2004)

[13]Out of 747 listed universities only 449 had a link to a website. Missing links are mostly for private universities.

[14]Often, economics departments are combined with business departments to form a larger department. In this case, names from the business departments are also included.

had 90 female in her sample, while Blackaby et al. (2005) had a sample with 60 female and 291 male.

The percentage of females in our sample is 17.2%. Based on the statistics provided by MEXT[15] Statistics of School Education (*Gakkou Kihon Chousa*), the percentage of females in economics departments in Japan was 12.6% in 2007[16]. Thus, we over-sampled females. Over-sampling of females was purposely done in order to increase the precision of estimates. In the labor market discrimination literature, over-sampling of minority groups is not uncommon. For example, Neal and Johnson (1996) use data from the NSLY[17] which over-samples blacks. Moreover, Monk and Robinson (2000) in a study of gender and racial earning differences in academic markets over-sampled females. We thus believe that over-sampling of females does not affect the relevance of our results. Nevertheless, we advise caution when generalizing results.

One may be concerned with respondent biases. For example, if only those who identified with the purpose of analyzing gender inequalities replied, our estimates would be biased. However, we believe that such a bias does not exist in our sample, since in our cover letter, we did not emphasize that the data will be used only to analyze gender salary differences. Moreover, 96.44% of the respondents (94.83% female respondents) replied that they did not feel discriminated against in their salaries. Thus, there is no reason to believe that our sample is affected by respondent biases.

One may be also concerned with the over-representation of full-professors due to non-random responses, since in typical mail surveys of academics, such a problem is not uncommon (Ward 2001; Blackaby et al. 2005). In our sample, 63% of respondents are full-professors. However, according to MEXT Statistics of School Education (*Gakkou Kihon Chousa*) data, in 2007, 60% of academics in economics departments in Japan were full-

[15]Ministry of Education, Sports, Science and Technology (MEXT)
[16]Based on the author's calculations.
[17]National Longitudinal Surveys of Youth (US)

professors. Thus, compared to past studies, the difference is relatively minor for our sample[18]. Finally, we under-sampled private universities because the MEXT does not provide website links for a significant number of private universities. According to the MEXT Statistics of School Education (*Gakkou Kihon Chousa*), 73% of academic economists work in private universities while only 59% of our sample is in private universities.

7 Variables and descriptive statistics

Table 1 shows definitions of our variables. The dependent variable is the logarithm of the total annual salary which includes the 12-month salary plus bonuses and allowances from the current institution. Other sources of income, such as money earned from other institutions or consulting fees are not included. We have a great number of control variables and classified them according to four characteristics: personal, job, institutional and human capital.

Personal characteristics Becker (1991) suggests that married adults posses more human capital than unmarried adults. We include a dummy variable for married respondents in order to control for the effect of marital status on salary. Ward (2001) finds that the presence of young children increases salary by 8%. We, therefore, include the number of children under age 6 in order to control for the effect of the presence of young children on salary[19].

Job characteristics (AssocProf) is a variable that takes the value 1 if the respondent is an associate professor, 0 otherwise; (FullProf) is a variable that takes the value 1 if the respondent is a full-professor, 0 otherwise. The excluded category includes assistant professors and lecturers. Japanese universities typically hire academics on a lifetime employment basis (i.e., no term-specified contracts). However, since 1997, with the enactment of the *Legislation of the Fixed-Term System for Faculty Members*, the fixed-term contract has been

[18] In Blackaby et al. (2005) 28.5% of their sample is full-professor, while the representation of full-professors in the population is only 18.8% (UK). Similarly, Moore et al. (2007) have 37.3% full-professors in their sample from the UK, while the the representation of full-professors in the population is only 18.8%.

[19] We only included the number of children under age 6, since our preliminary estimates did not show a significant effect of older children.

introduced. In order to control for the effect of fixed-term employment on salary, we include a dummy variable for the fixed-term contract.

Few studies control for the effect of teaching load on salary. Taylor et al. (2006) reports that time spent on teaching has a negative effect on research productivity. In order to control for the possibility that teaching load has a direct impact on salary, we control for the total number of courses taught during the previous year. This variable can also be thought of as capturing institutional differences as well, since top universities are expected to be more research-oriented, thus requiring less teaching from their academics. We also control for the number of courses taught for the first and second time, since these courses may take longer preparation time. According to McDowell et al. (2001) and Koplin and Singell Jr. (1996), females economists usually prefer fields like labor economics as opposed to fields like theory or quantitative methods. In fact, our sample also suggests that female academics are most represented in labor economics. In order to separate the effect of field choice from the effect of being female, we include a dummy variable for labor specialization[20].

We also control for being in an administrative position (e.g., the Dean of department or the Chair of the university), since holding such positions may increase the salary. Survey participants were asked to report the percentage of their time allocated for administrative duties. The variable (Admin) takes the value 1 if the respondent spends more than 50% of his/her working time on administrative duties. This variable is a proxy for holding administrative positions. Finally, we control for cohort effects for those who entered the academic labor market in the 1980s, the 1990s, between 2000-2003 and 2004 onward. The cohort that entered academia from 2004 onward is expected to capture the possible effects of 2004 national and public university 'corporatization'. The 2000-2003 cohort dummy captures the possible effect of the 2000 *Action Plan* (see Introduction) that stipulates that national uni-

[20]In our preliminary estimations, we included dummies for various other fields. However, this did not affect the results; only the labor field was found to be highly statistically significant.

versities should increase the number of female academics. In the 1990s there is a significant convergence in the percentage of male and female PhD graduates joining the academia (see Figure 2B). The 1990s cohort dummy captures the effects of possible changes in the labour market conditions that caused such a convergence. The cohort dummy for the 1980s capture the effects of specific labor market conditions at that time, for example, the enactment of the *1985 Law of Equal Employment Opportunity of Men and Women.*

Institutional characteristics Estimated gender salary differences could arise if females are over-represented in universities with lower remuneration. Thus, we control for various institutional characteristics. We include dummy variables for private and public universities, with national universities being the reference group. We also include a dummy variable for business department. To proxy for the quality of the department, as in other studies (Taylor et al., 2006), we use a dummy variable that takes the value 1 if the department is offering a PhD degree.

The variable (IntGrant) shows the amount of internal research grant received from the current institution in the previous academic year. According to our data, 80% of the respondents said that each academic in their department receives the same amount of internal grant. Thus, the amount of internal grant can be viewed as either showing the financial status of the university, or its research orientation. We also consider a special competitive grant called Center of Excellence grant (COE). (COE) is the variable showing the amount of research grant each respondent received from the COE grant. COE grants are provided by MEXT[21] and offer funds at the department level. According to Arimoto (2007), in 2006, 51% of selected projects belonged to the top ten national universities. Thus, we expect that this type of grant shows the quality of the department and of the university.

[21]Ministry of Education, Science, Sport and Technology

Human capital characteristics We control for standard human capital measures such as experience and education. The variable (Seniority) is the total number of years of experience at the current university; (Experience) is the total number of years of experience as an academic. We also include non-academic experience measured as the total number of years worked full-time outside academia. We also included the squared terms of (Seniority), (Experience) and non-academic experience, in order to capture that the rate of return on human capital continues at a diminishing rate. In order to control for the effect of career breaks on salary, we include a dummy variable that takes the value 1 if the academic took a leave during his/her academic career. This variable would capture the effect of human capital depreciation due to inactivity. The variable (PhD) is the dummy variable indicating that the respondent has a PhD degree. This variable captures the effect of education on salary. In order to capture the differences in the PhD programs from which the respondents graduated, we also include an additional dummy variable for a PhD degree obtained overseas.

To control for additional differences in productivity, we include the amount of competitive external grant that an academic received in the previous year[22], and the number of publications. The number of refereed articles is the most accepted measure of scholarly research output (Taylor et al., 2006). Previous literature found that the returns on salary from non-refereed publications are quite low (Oster and Hamermesh 1998; Moore et al. 2007). However, since Japanese academics in our sample produce a considerable amount of work other than refereed articles, we control for various types of publications. Publications are classified according to their types: referred single authored articles, refereed co-authored articles, working papers, single authored books, co-authored books, books edited, book chapters and textbooks. Publications are reported for the whole career. In the prior literature, the quality of research output is controlled for by distinguishing articles published in top

[22]The amount is per individual; however the grant might have been obtained jointly. When grant was obtained jointly with other researchers, respondents were asked to report the amount based on their own contribution to the project.

journals. In our survey, however, in order to preserve the anonymity of the respondents, we did not ask the name of the journal of publication. Therefore, we cannot directly adjust for the quality of the publication. However, we asked the survey participants to report the number of publications according to the location of the publisher. Thus, each type of publication is further divided into subtypes depending on whether it was published in Japan or in the US/Europe. We expect publications in the US or Europe to be more cited than those published in Japan, since these are published mostly in Japanese. Thus, we can capture potential differences in the impact of the research output.

Descriptive statistics are presented in Table 2. The average annual salary is 10.5 million yen for males and 8.8 million yen for females. The average age is 50.7 for males and 43.3 for females. The average seniority is 13.5 years for males and 8.6 for females. The average experience is 18.3 years for males and 10.9 for females. The average non-academic experience for both males and females is about 3 years. 67% of males and 43% of females are full-professor, while 27% of males and 43% of females are associate professors. In sum, males are older on average, have more experience, and obtain higher average salary.

Only 2% of males took career breaks, while as much as 12% of the females did so. Females are more likely than males to be found in private universities (66% of females and 57% of males). It appears that there are no significant differences between males and females in terms of the highest educational achievements; 65% of males and 64% of females have PhD degrees. However, the number of males with a PhD degree from overseas is slightly higher than that of females (11% of males and 7% of females). The greatest number of females was hired after the year 2000. Figure (1) plots the salary profiles according to seniority and experience. Both females' seniority-salary profile and experience-salary profile lie below that of males'.

8 Estimation results

8.1 Salary equation

Table 3 reports the OLS results for four model specifications, each using different publication record measurements. For each model we present separate results, with rank variables included and excluded. OLS 1 includes the most detailed publication record. When rank variables are included, the coefficient for the female dummy is -0.071 and it is statistically significant at the 1% significance level, indicating that females earn 7% less than males, after controlling for detailed personal, job, institutional and human capital characteristics.

The effects of other control variables are also of interest. Age appears to be a significant determinant of salary. An increase in age by one would increase salary by 1.6% at the age of 40. Married academics earn, on average, 5% more than those not married. The number of children under 6 years old has a positive effect on salary. Salary would increase by 3% for each additional child in this age break. The past literature tends to find positive effects of children for the male sample (Bellas and Toutkoushian 1999), while insignificant effects for the female sample (Ward 2001; Barbezat 1987), suggesting that males with children might be acting as breadwinners. Our positive coefficient for children might suggest that the effect of children for the male sample is dominating the result. Indeed, when we summarize the salary by the presence of young children (under 6 years old) for respondents younger than 40 years old, males with young children earn more than males without young children (8.1 million yen and 7 million yen respectively); while females with young children earn slightly less than females without young children (7.1 million yen and 7.4 million yen respectively).

Full-professor rank has a positive and statistically significant influence on salary. The coefficient of 0.18 suggests that there is an 18% salary gap between assistant professors and full-professors. The coefficient for (AssocProf) is positive, indicating a 6% salary gap between associate professors and assistant professors; however, the coefficient is not statistically significant. The coefficient for (FixTerm) is statistically significant at the 1% significance level. The result indicates that the salary would be 24% lower if the contract is of limited duration. The effect of the total number of courses is negative and marginally significant. Past literature found a negative impact of teaching on research productivity (Taylor et al. 2006). Our results suggest that teaching load may have a direct and negative impact on salary. Alternatively, we can interpret the result as showing that teaching-oriented universities are paying less. Academics who are specialized in labor economics have a salary premium of about 7%. As mentioned before, in our sample, the highest concentration of females is in the labor field. One may expect that females might be concentrated in a lower paying specialization, however, in our sample, the labor field enjoys a higher salary. We observe no significant cohort effects.

The coefficient for (PrivUniv) is positive and highly significant, showing a salary premium of 16% for those working in private universities. There is also a salary premium of 7% for those who work in PhD-granting departments. This premium could stem from greater outside funding for PhD-granting departments as suggested by Koplin and Singell Jr. (1996), or due to unobserved differences in the quality of academics who work in such departments. The amount of internal grant and the COE grant do not have statistically significant effects on salary.

Now let us turn our attention to the effects of human capital characteristics on salary. The coefficients for (Seniority) and its square are small and statistically insignificant. The estimated coefficients suggest that an increase in seniority of one year would increase salary by only 0.3%, evaluated at seniority equals to 10. The small effect of seniority is similar to Ransom's (1993) findings[23] that assert that the small effect of seniority on salary is the result of monopsonistic power universities have in the academic labor market. The effects

[23]In Ransom's study the effect of seniority is negative. A more detailed analysis of this issue is beyond the purpose of this study.

of both academic and non-academic experience are insignificant. Career break variable has a negative coefficient, but it is statistically insignificant. Nonetheless, this variable has a significant impact on the estimated coefficient for females. Although not reported here, exclusion of this variable would increase the gender salary gap from 7.1% to 7.8%.

The coefficient for (PhD) is positive (0.023), but statistically insignificant. In Japan, until recently, a PhD was granted as a life-time work achievement rather than at the completion of a doctoral dissertation. Therefore, not having a PhD does not indicate lower human capital. However, a PhD obtained abroad seems to have a sizable effect on salary. There is a salary premium of about 5.6% for a doctorate obtained outside Japan. The amount of external grant has a significant coefficient[24]. The estimated coefficient indicates that earning a one million yen grant is associated with an increase in salary of 1% or about 100 thousand yen evaluated at the mean salary[25].

All the publication variables have insignificant effects on salary, except for book chapters published in the US and Europe. In virtually all of the previous studies, the number of publications, especially refereed articles, is found to be a significant determinant of salary. In our paper, the results are different and appear to confirm the common belief among Japanese academic economists that publications are not determinants of salary, since salary is determined by the rigid salary table[26]. However, there could be other factors that influenced our results for the publication variables. For example, there could be error-in-variables endogeneity in publication measures[27].

[24] The amount of external grant may be an endogenous variable due to unobserved ability. However, the majority of papers on salary differences in academia do not control for such biases.

[25] We would like to note here that, in Japan, external research grants do not add to salary and do not provide summer salary, as in the US.

[26] Although not reported here, we experimented with other measures of publications. We use the publication rates defined as the number of publications divided by total experience as an academic. We found that the effects of these measures of publications are insignificant and the coefficient for female was unaffected by the choice of such a specification.

[27] In a preliminary estimation we controlled for error-in-variables endogeneity in the refereed articles by applying a 2SLS stage procedure to OLS 4. Given concerns related to the choice of instruments, we do not fully report the results here. We used as instruments parents' education and their interaction terms and spouse's education. Hansen's J-statistics did not reject the overidentifying restrictions (p-value=0.60). The C-statistics test rejected the exogeneity of refereed articles (p-value=0.01). We obtained a small statistically

When OLS 1 is estimated without controlling for rank variables, the female coefficient decreases only slightly in absolute value, from -0.071 to -0.069, and remains statistically significant at the 1% significance level. Thus, the gender salary gap is almost unaffected by exclusion of rank variables, indicating that there is a large salary gap *within* each rank, but there are no significant differences in rank attainment between genders. Most of the prior studies from the US and the UK found that much of the gender salary differences stems from the fact that female academics are over-represented in lower ranks (Ward 2001; Ginther 2004), and that there is a little salary gap within each rank. Thus, our results show the entirely opposite pattern. Most of the coefficients for other variables appear to be unaffected by the exclusion of rank variables as well.

In order to study the robustness of the results to our measures of research output, OLS 2 uses a less detailed publication record. In this specification, the publication record is not classified according to the location of the publisher. The coefficient for the female dummy decreases slightly in absolute value, from -0.071 to -0.068, but remains statistically significant at the 1% significance level, for the model with rank variables included. When rank variables are excluded, the coefficient is -0.065 and significant. None of the coefficients for the measures of publications are significant, and, virtually all the other coefficients are unaffected by the change in the definitions of measures of the publication record.

OLS 3 employs more aggregated measures of publications. In OLS 3 we do not distinguish between single authored and co-authored publications as we did in OLS 2. However, before adding the single and co-authored publications in order to obtain the total number of publications, we divide the number of co-authored publications by 2, assuming that the number of co-authors is usually two. When rank variables are included, the coefficient for female is -0.069 and is highly statistically significant. When rank variables are excluded, the

significant effect of refereed article on salary; one extra refereed article would increase annual salary by about 0.8% or 8 thousand yen, evaluated at the sample mean salaries. However, the estimated coefficient for the female dummy was almost unaffected (-0.070) and remains statistically significant at the 1% level. Results are available from the author upon request.

coefficient is -0.067 and is also highly significant. The publications variables do not appear to have a significant effect on salary, on either model, with or without rank variables. The other coefficients are qualitatively and quantitatively similar to OLS 1 and OLS 2.

Finally, OLS 4 uses only the total number of refereed articles (TotRefArticles) as measure of publication record. This model is relevant, since refereed articles have been considered the most accepted measure of research output in the prior literature. The female coefficient is -0.07 when rank variables are included, and it is statistically significant at the 1% significance level. The coefficient for refereed articles is small (0.0003) and statistically insignificant. When rank variables are excluded, the female coefficient is -0.069, and still highly significant. There are no noticeable differences in the coefficients between this model and the previous models.

In sum, all models indicate that there is a significant gender salary gap within each rank. The estimated coefficient for female ranges between -0.068 to -0.071, after controlling for detailed personal, job, institutional, human capital characteristics and rank. The coefficient for female is statistically significant in all models. When rank variables are excluded, the female coefficient decreases only slightly in absolute value, ranging between -0.065 to -0.070, but it remains statistically significant in all models. Thus, our results indicate that there is a significant gender salary gap *within* each rank, but there is no significant gender difference in rank attainment.

8.1.1 Additional results

We would like to discuss below results not reported in Table 3. Figure 1 shows that there is a greater gender gap later in the career, after about 25 years of experience. Such situation could have been caused either by (i) the presence of cohort effects or (ii) because the gender salary gap widens later in the career. In order to check the latter possibility, we included in OLS 1 an interaction term ($Female * Dummy(Experience \geq 25)$). The coefficient for the interaction term is negative, but not statistically significant. The coefficient for the female dummy drops only slightly in magnitude to -0.067(p-value=0.015). Therefore, we do not find evidence that the gender salary gap widens with experience.

In order to see if the gender salary gap decreases or increases with new cohorts, we included interaction terms between female and cohort dummies. Since the majority of females in our sample entered the academia after 2000 (49% of females), we include in OLS 1 interactions between female, and (Cohort00-03) and (Cohort04). The coefficients for both interaction terms are positive, but insignificant; 0.005(p-value=0.93) and 0.04(p-value=0.57), respectively. Thus, we did not find evidence that the gender gap is decreasing with new cohorts.

We also ran OLS 1 separately for private university and national university samples. The female coefficient is -0.05(p-value=0.23) for the private university sample; and -0.071(p-value=0.14) for the national university sample. Although the gender salary gap appears to be smaller for private universities, the large standard errors make the comparison difficult.

Blackaby et al. (2005) show that the number of outside job offers explains the gender salary gap for the UK academic economists. We do not have information on outside job offers; however, we do have information regarding the number of universities each academic worked at. We thus include a variable to control for the number of previous academic jobs. The average number of universities academics in our sample previously worked at (excluding the current university) is 0.68 for males and 0.53 for females. The coefficient for this variable is insignificant, 0.02(p-value=0.86), and the coefficient for the female dummy does not change appreciably in value and remains significant, 0.071(p-value=0.01).

8.2 Rank attainment equation

We estimate an ordered logit rank attainment model using the same specification as in OLS 1 of the salary equation. Table 4 shows the results. Contrary to our expectations, the coefficient for the female dummy is positive (0.092), indicating that females are 0.2% *more* likely than males to be a full-professor (see the marginal effect in Table 4). However, the coefficient is insignificant and the effect is small. Thus, there is almost no difference in rank attainment between males and females. Despite the common belief that promotion is a deterministic function of age and experience, the coefficients for age and experience are not statistically significant. Having a PhD would increase the probability of being a professor by 6%, holding all other characteristics constant. The coefficient for (FixTerm) is negative and significant. Most of the coefficients for publications are insignificant, however, working papers published in Japan and co-authored books published in Japan have positive and statistically significant effects on rank attainment.

The logit estimation results showing that age and experience are not significant determinants of promotion are puzzling. As we are concerned that the results might have been affected by our choice of model, we also estimate the same rank equation by using OLS, the second column in Table 4 showing those results. The female coefficient is small and statistically insignificant (0.02), indicating that there is little difference in rank attainment between genders. However, age and academic experience appear to be significant determinants of the rank attainment.

Although the results show that there is not much of a gender difference in rank attainment, it is still useful to compute the total gender salary gap defined in equation (4). Based on the results of the ordered rank equation, the marginal effect is $P(Rank = 2|Female)$-$P(Rank = 2|Male)$=0.002. The coefficient for rank, β, is 0.010 and it is reported in the OLS 5 in Table 3. Thus, the total salary gap is $-0.071 + 0.10 \times 0.002 = -0.0708$. Since, according to our results, females are more likely than males to be full-professors, the gender salary gap reduces when we combine gender salary differences with promotion differences. However, because the differences in rank attainment are small, virtually the entire salary gap can be attributed to the salary gap *within* each rank.