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ABSTRACT

Since monetary union with western Germany on 1 July 1990, eastern female monthly wages have risen by 10 percentage points relative to male wages, but female employment has fallen 5 percentage points more than male employment. Using the German Socio-Economic Panel to study the years 1990-1994, I show that along with age, the wage of a worker in 1990 is the most important determinant of the hazard rate from employment. Differences in mean 1990 wages explain more than half of the gender gap in this hazard rate, since low earners were more likely to leave employment, and were disproportionately female. The withdrawal from employment of low earners can explain 40% of the rise in relative female wages. Competing risks analysis reveals that the wage has its effect through layoffs, and hence through labor demand, which is consistent with the hypothesis that union wage rises have caused the least productive to be laid off. There is no evidence that reduction in child care availability is a major factor in reducing female employment rates.

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Since monetary union with western Germany on 1 July 1990, the monthly wages of full time female workers in eastern Germany have risen from 81% of male wages to 90% of male wages in 1994. Wages of all female eastern workers have risen from 74% to 84% of male wages. These wage ratios suggest that the transition has been extremely beneficial to women, particularly in view of the fact that average real monthly wages rose 85% in this period. The employment rate for women 18-60 has fallen in the same period from 84% to 63%, however, compared to a drop from 94% to 78% for men, a 5 percentage point larger drop. But if this simply reflects a switch to leisure or household production that women were prevented from exercising under communism, and the faster drop in employment is merely a convergence to western norms, this may not be a cause for alarm: the transition could still be characterized as favorable to women. Conversely, the larger employment drop for women could reflect discriminatory layoff practices or involuntary withdrawal from the labor market due to the reduction in child care availability.

In this paper I examine the changes in wages and employment for 1990-1994 using the German Socio-Economic Panel. I argue that as far as employment declines are concerned, gender is to some extent a red herring. The most important determinant of the hazard rate from employment (other than age) is the wage an individual was receiving in 1990. Thus low wage individuals have disproportionately left employment, and low wage earners are disproportionately female. Differences in mean 1990 wages account for more than half of the male/female hazard rate gap. This in turn means that the average female wage has risen relative to the male wage due to the exit of low earners from employment. I find that this selection can account for 40% of the fall in the gender wage gap.

I analyze the competing risks of losing a job due to layoff, and losing a job due to other reasons (retirements, quits etc), and find that the effect of the wage is exclusively through the risk of layoff, suggesting that a labor demand story is at work. This evidence is consistent with the idea that union wage raises have caused the less productive to be laid off. The presence of children age one or younger increases the risk for women of ending employment for "non-layoff" reasons, but due to the small numbers of such children this is not an important factor overall. The presence of older children does not increase women's hazard rate. The results suggest that decreased child care availability is unlikely to have been important in reducing female employment.

1 Background and Hypotheses

Monetary union between east and west Germany, and hence the beginning of economic transition, took place on 1 July 1990. The immediate change in the wage system at this time was the conversion of East German wage contracts into West German marks at a rate of one for one (Krueger and Pischke 1995). The western trade unions gradually took over the eastern wage bargaining system, and the conversion to the western system was accomplished in most industries over the course of 1991. As this happened, it became common for the unions to negotiate step-wise wage increases designed to achieve convergence to the wage level of the equivalent western industry by 1994. The contracts varied greatly according to the location and health of the industry, however, and in December 1992 the monthly wages in effect varied from 57.5% of western levels in the clothing industry to 83% for construction in east Berlin. When factors such as the longer working week in the east are taken into account, the ratios somewhat slower. (See Bispinck and Meissner 1993.) Subsequently, however, as employment continued to fall, employers exercised their right to reopen negotiations, and unions agreed to delay convergence in some cases. Some newly formed firms declined to join the relevant employer federation and were hence not bound by the bargaining, while some other firms simply paid less than the bargained wage. Thus a survey of industrial firms by the Deutsches Institut für Wirtschaftsforschung (German Institute for Economic Research) in the winter of 1993/4 found that 30% of firms were paying less than the originally bargained wage in their industry, compared with only 10% for western industry in a separate 1993 survey. Only 40% of eastern firms reported belonging to their employer federation. However, large firms were more likely to belong to the employer federation and to respect the bargained wage, so the proportion of workers affected by payment of wages below bargained wages is lower. (Scheremet 1995.)

Output dropped precipitously (about 50%) and employment fell from about 9 million in 1990 to 6 million in 1992 (Licht and Steiner 1994). The government adopted two measures designed specifically to deal with this: early retirement and public works jobs. A special program in place from October 1990 (unification) until December 1992 allowed retirement onto western pensions by men at age 57 (initially, then 55) and women at age 55. This replaced a scheme introduced by the communist government which had allowed retirement by men at 60 and women at 55. By 1993 866,600 people had retired early. In 1993 about 250,000 people were in public works jobs. The western system of short-time work, where workers are put on reduced hours, and have their lost hours compensated at the replacement ratio of unemployment benefits, was widely used, particularly in the first year of the transition (215,000 workers were affected in 1993). Government training programs have also been used (382,000 participants in 1993). (Kühl 1994.)

The evolution of female wages relative to male has been little analyzed. The relatively greater decline of female employment has received some attention by other analysts. The hypothesis mentioned above, that this represents a natural convergence of female participation to western norms is not given serious attention. The women leaving employment appear to remain attached to the labor force: Holst and Schupp (1992), for example, report for 1991 that of women under 57 who were not working, 39% wished to resume working as soon as possible. One hypothesis is that managers and works councils, who select which workers should be laid off, believe that layoffs of women are a lesser evil, while another is that that women quit due to lack of child care availability and social pressure to leave jobs for men (Brinkmann et. al. 1993). The negative effect of wages on the hazard rate from employment has been noted by Licht and Steiner (1994), but without making a link to male/female differences and without distinguishing between layoffs and other separations. More general analysis of eastern employment declines frequently discusses the role of the union wage raises, with some observers believing that wage subsidies to counter the raises would have an important beneficial effect (Bedau 1996, Bellmann 1994, Begg and Portes 1992).

Under the German Democratic Republic child care was readily and cheaply available, but the number of available child care places dropped after monetary union. When considering child care for the youngest children, however, it is critical to recall that fertility dropped dramatically in the first two years of transition, so that there was a very important fall in the number of children (aged 3 or younger) requiring creches. Since fertility may have declined in a way not orthogonal to child care availability for the prospective parent, investigation of the effect of child care decline for the very young is difficult. Potentially more promising is to investigate the effect of the rapid reductions in availability of Schulhort places (after-school care for children in classes 1-4), since the existence of children of this age is exogenous to the transition.

In 1989 Schulhort places were available for 81% of children of that age, 94% children of kindergarten age had access to kindergartens, and 82% of children age 1-3 had access to creches (assuming, as was usual, that mothers stayed home for the first year of the baby's life). Time-series on the subsequent decline are more difficult to obtain. Still assuming that children 0-1 do not need creches, the availability per child 1-3 had fallen to 70% by December 1991, while the availability of kindergarten places had scarcely fallen by this time. Based on four of the six regions, availability of Schulhort places fell from 94% in 1989 to 76% in September 1990, to 53% in December 1991, and to 32% by mid 1992. (See Bundesministerium für Familie, Senioren, Frauen und Jugend 1994.)

2 Data and Sample

The data used are from respondents to the German Socio-Economic Panel (GSOEP) who resided in eastern Germany when that region was first surveyed in June 1990, immediately before monetary union.[2] Individuals who moved to western Germany or who subsequently worked in western Germany are not removed from the sample, and hence the analysis refers to eastern Germans rather than to eastern Germany. By 1994 4.8% of the sample resided in the west, and a further

[2] The public use data set is employed, which means, for example, that no regional indicators are available.

6% commuted to work in the west (more than half to west Berlin). The data refer to the period 1990-1994.

The wage used is gross earnings in the month prior to the interview, not adjusted for end of year bonuses (principally because the bonuses have many more missing values). Wages for respondents living in the west are adjusted using the western consumer price index (1985=100). The eastern consumer price index is adjusted so that a meaningful comparison of the price levels between east and west may be made, using the results of a 1991 study of the relative price levels in the east and west (Krause 1994). The wages of those living in the east are thus deflated with this adjusted index, so that the real wages of those in the east and west should have comparable purchasing power parity.

In the analysis of the evolution of the wage distribution, the sample in each year is restricted to those for whom the wage is meaningful. Thus those working in agriculture and fishing (15% of 1990 employment) and the self-employed are excluded. Also excluded are apprentices and others under 18, those over 60, anyone reporting zero weekly working hours, and those for whom the wage or any of the covariates used were missing. Individuals who reported either full or part-time work (including public works jobs, but excluding training programs) or short-time work (with non-zero hours) are included. 3% of the 1994 sample had a public works job, down from 5% in 1992. In the analysis of the evolution of employment, the sample used consists of those satisfying these conditions at the time of the first interview, in June 1990. These individuals are then followed whether or not they continue to satisfy the restrictions.

For analysis of employment, I use the "calendar" section of the GSOEP: respondents 16 or older indicate for the calendar year preceding the interview what their labor force status was month by month, allowing the construction of spells. I take individuals who indicated that they were employed in June 1990 (and satisfied the restrictions described above) and consider the spell to continue until they begin a spell of non-employment of more than four months in length. I call the resulting spells "attachment" spells. I do this because I consider job changes a desirable response to transition, and, given the magnitude of the upheaval experienced by eastern Germany, I am not unduly concerned by an individual who experienced a short spell of non-employment between jobs. Notice that this will also allow women taking only short maternity leaves to maintain an uninterrupted spell. The calendar file from the 1995 interview year was used to construct spells through 1994, but all other information comes from interview years 1990-1994. Notice that the setup of the spells means that time and duration cannot separately be identified.

The variable means for the 1990 and 1994 samples are shown in Tables 1a and 1b. Particularly marked changes occurred in the educational distribution for women, and in the industry and firm size mix. Education (highest attained) is broken into four categories: tertiary education ("university"), "apprenticeship" (implying participation in the dual-system classroom and work in a firm), "vocational training" (meaning classroom vocational training), and "general schooling", meaning only general classroom education (this last category is extremely small). The proportion of women with an apprenticeship declined sharply, in contrast to the proportion of men. For the industry mix note especially the ten percentage point rise in women working in government and men working in construction. The fertility decline is also apparent in the child variables (which indicate the presence in the household of one or more children in the age group).

3 Results

3.1 Evolution of the Wage Distribution: Cross-Section Analysis

Figure 1 shows (Epanechnikov) kernel density estimates of the distribution of real wages in 1990 and 1994, by sex, and for both eastern and western Germany. Notice that these graphs do not exploit the panel aspect of the data. The top left panel plots the distributions for eastern Germany. The increase in mean of the distribution is dramatic for both sexes, while there is also an important increase in dispersion. The different experiences of men and women are also striking, however. Contrary to what one might expect given the communist equality rhetoric, the mean of the female distribution in 1990 was substantially below that of the male distribution. The shapes of the two distributions were very similar, however. By 1994, the shapes of the male and female distributions had diverged, while the means had converged. Unlike in the first two years of transition, by 1994 the top of the distribution was not dominated by those working in the west: only 19% of the top male decile and 10% of the top female decile were living or working in the west.

The bottom left graph shows that the eastern developments in male/female differences are not the result of convergence to western standards. The graph for the west, which is plotted for German citizens only to avoid the use of weights (western foreigners are oversampled), shows that, due to the low hours of women, the western female mean is far below the male mean in both 1990 and 1994. Mean eastern male wages were 44% of western levels in 1990 and 72% in 1994. For females, the percentages were 56% and 105%. (These weighted means include foreigners.) (See Burda and Schmidt 1997 for more details on male east/west convergence.)

In order to separate the effects of hours and hourly wages, a simple adjustment for hours is made to the monthly wage. The wage level is regressed on hours worked per week and its square, as well as a spline in hours. Predicted wages for the average hours per week in the east and the west are calculated, and the distributions plotted in the right two graphs. In the case of the east, the 1994 male and female hours-adjusted distributions are very close to each other. In the west they remain very different, with the female distribution much less dispersed and with a much lower mean than the male distribution.

The evolution of the eastern distributions can be examined further with the aid of quantile regression. Quantile regression predicts the value of the dependent variable at a particular quantile, rather than at the mean. Therefore a log wage quantile regression at the .75 quantile with only a sex dummy as a covariate shows the percent difference between the male and female 75th percentile wage. As with mean regression, other covariates can then be added to assess whether differences in the covariates can explain the differences in the wage at the 75th percentile.

I run a series of log wage quantile regressions for the .1,.25,.5,.75 and .9 quantiles, including first only a sex dummy, and then a more complete set of controls: age, age squared, education dummies, tenure, firm size dummies and 22 industry dummies (more detailed categories than used in the employment analysis below and reported in the means). I then repeat all the regressions controlling also for hours by including weekly hours and its square, and dummies for 30 hours or below, 31-40 hours and 41-45 hours per week. (Monthly hours are not available.) I do not include marital status or its interaction with sex. When the regressions were run including these variables, the coefficient on marital status was never significant, while the coefficient on its interaction with sex was significant only for low quantiles when hours were not controlled for (when it was negative and large), suggesting the interaction was proxying for hours. The omission of marital status is consistent with Blau and Kahn (1997). Appendix table A1 reports the results of the median regressions (.5 quantile).

Figure 2 displays in four panels the coefficients on the sex dummies from the many regressions. The top left panel plots for 1990 and 1994 the coefficient on the sex dummy, and the 95% confidence intervals, from the regression where sex is the only covariate. The quantile regressions in this case essentially pick out points from the plots of Figure 1 (this would be exactly true if Figure 1 plotted log wages). This plot shows large unconditional gains for women, which are significant at all except the .1 quantile. The top right panel plots the coefficients from the regressions where controls for hours are added. The 1990 coefficient is now very similar across quantiles, indicating that the upward slope of the 1990 line in the top left panel represented differences in hours worked across quantiles. The unconditional gains between 1990 and 1994 in what may be viewed as hourly wages are larger in the middle of the distribution, and indeed the gains at the .25 and .9 quantiles are no longer significant. Relative increases in female hours at the bottom of the wage distribution and reductions in the middle explain most of this. The narrowing of the 1990-94 gain at the .9 quantile appears to be due to changes between 1990 and 1994 in the coefficients on hours, however.

The two lower panels add the controls other than hours to the regressions of the two upper panels. These two panels also suggest clear gains in the middle of the distribution. The point estimates indicate smaller conditional than unconditional gains for women between 1990 and 1994 (the gap between the lines is smaller in the lower panels). This implies either that returns to characteristics have evolved in such a way as to favor women, or that women's characteristics have improved relative to men's, either due to choices such as industry choice or due to selection in who leaves employment. [3]

The change in the wage gap can be decomposed using the technique of Blau and Kahn (1997), inspired by Juhn, Murphy and Pierce (1991). The decomposition relies on properties of the mean which do not carry through to median etc, so quantile analysis is not pursued here. The procedure is based on the coefficients obtained from male wage regressions for the years of interest, implicitly assuming that female returns would be the same in the absence of discrimination. Thus β_{90} and β_{94} are obtained from the regressions for males:

$$ln\ w_{it} = X_{it}\beta_t + \sigma_t\theta_{it} \tag{1}$$

where θ_{it} is a standardized residual and σ_t is the standard deviation of wage residuals. The male-female wage gap for year t is:

$$D_t = \overline{ln\ w_{mt}} - \overline{ln\ w_{ft}} = \Delta X_t\beta_t + \sigma_t\Delta\theta_t, \tag{2}$$

where Δ represents the difference in male-female averages in X_t and θ_t. The change in the gap between 1994 and 1990 can therefore be decomposed:

$$D_{94} - D_{90} = (\Delta X_{94} - \Delta X_{90})\beta_{94} + \Delta X_{90}(\beta_{94} - \beta_{90}) + (\Delta\theta_{94} - \Delta\theta_{90})\sigma_{94} + \Delta\theta_{90}(\sigma_{94} - \sigma_{90}) \tag{3}$$

The first term reflects the contribution of changes in characteristics of females relative to males at given returns, the second the effect of changes in returns to characteristics given that the

[3] Note that by "selection" I do not necessarily mean self-selection, but simply non-random exits from employment.

distribution of characteristics differs between men and women. These are known as the observed X's and observed prices effects. The third term represents relative movements in position in a distribution of residuals with a given standard deviation (known as the gap effect, in essence the mystery term). The fourth term represents how an increase in the standard deviation of residuals affects wages given the different relative positions of men and women in the residual distribution. The second and fourth terms thus represent observed and unobserved components of the effects of changes in wage structure, while the first and third terms represent the effect of changes in observed and unobserved relative characteristics of men and women ("gender-specific" effects).

As explained in detail by Blau and Kahn (1997), the $\Delta\theta_t\sigma_t$ terms are calculated based on the mean of the residuals for women computed from the male wage equation. The mean of the male residuals from the male wage equation is zero, hence $\Delta\theta_t\sigma_t$ is equal to the negative of the mean female residuals. The term $\Delta\theta_{90}\sigma_{94}$ is trickier, as it involves calculating what the mean 1990 female residuals would be if the standard deviation of residuals were that of 1994 (for men the mean is again zero). Each woman is thus assigned first a percentile in the 1990 male residual distribution, based on her 1990 residual, then she is assigned the residual that corresponds to that percentile in the 1994 distribution. The negative of the mean of these is $\Delta\theta_{90}\sigma_{94}$.

The results of this decomposition are shown in Table 2. In the first column the male regression underlying the decomposition has only the age, education and tenure variables as controls. The second column adds firm size dummies and industry dummies. The third column adds controls for hours (as in the wage regressions above). The effect of changes in average hours and return to hours should be interpreted cautiously, since the variation in male hours is in a different region from the variation in female hours, so the βs estimated in the male regression may not be meaningful. The results of the underlying regressions are not reported.

Overall, focusing on the columns without hours controls, changes in observable X's account for between -0.035 and -0.044 of the -0.115 change in the gap. The "gap" component, the unexplained improvement in women's position, is more important, contributing about -0.10. Changes in observed prices are a significant brake on reductions in the gap if industry and firm size are taken into account, tending to increase the gap by 0.05. Despite increased inequality in male residuals, the change in the distribution of the residuals ("unobserved prices") does not have a large or necessarily even detrimental effect on the wage gap. The division between wage structure effects and gender-specific effects is similar across specifications, with gender-specific factors being the clear driving force. Any effects of selective withdrawal from employment of women could appear in either the components of observed X's or could be part of the "gap" component, if the selection occurs based on unobservables.

Common to the three columns is the contribution of changes in education of a 0.03 reduction in the male-female gap. The changes in the distribution of education were noted in the table of means, Table 1a. The proportion of women with an apprenticeship declined substantially, while the proportion remained constant for men. The return to an apprenticeship is lower than the return to the other common type of education, vocational training (this may be confirmed in the median regressions of Table A1). Hence the reduction in their number for women has increased the relative female wage. Since few individuals change their educational status and new entrants are small in number, this change in educational mix is likely to be the effect of selective exits from employment (this is confirmed in the employment analysis below). Changes in the return to education have had a smaller impact on the wage gap. Age and tenure contribute little to the wage gap change either through changes in mean or return.

The fact that men have shifted to smaller firms faster than women has tended to reduce the wage gap, since workers in smaller firms earn less. The increasing wage premium for workers in large firms has favored men, however, who were initially more likely to work in large firms. The net effect is to reduce the wage gap by less than 0.01. The change in the industry mix has worked against women, as have the changes in returns to working in different industries. Changes in the industry (and firm size) mix could be a mixture of selection effects due to workers in declining industries losing or quitting jobs and perhaps leaving employment, and choices made by individuals able to make job to job transitions. Changes in hours and returns to hours appear to have worked in women's favor.

Thus, from this decomposition only the -.03 contribution of education can clearly be identified as a selection effect. Although this is a substantial magnitude, it seems likely that some part of the gap effect also represents selection, and the analysis of employment below will shed light on this.

3.2 Evolution of Wages and Employment: Longitudinal Statistics

In this section I exploit the panel dimension of the data. Figure 3 provides some insight into longitudinal movements of wages and employment by plotting wage growth (percent change in average non-zero wages) from 1990 to 1994 and the 1994 employment rate for each of twenty 1990 wage quantiles. The sample is the group satisfying the 1990 sample restrictions.

In the two left hand panels of the figure the quantiles are computed separately for the male and female distributions, and hence each point is based on about 60 individuals. We see that wage growth is extremely high (roughly between 50% and 200%) and falling in quantile (i.e. those with initially low wages experienced larger rises). [4] The female curve lies above the male curve. The panel below this which plots employment rates indicates the remarkable fall in employment, and except in the highest quantiles this fall was larger for women.

The impression thus obtained is altered significantly, however, if instead of plotting the variables against the quantiles calculated separately for men and women, the variables are plotted against the wage quantiles computed for the pooled group of men and women, as in the two right hand graphs. The implication of this method is that the points represent different numbers of individuals, and the low quantile points for men, in particular, are based on very small numbers. [5] Although the wage growth plots are not dramatically altered (the male and female curves now cross), the employment plots are: the female line is neither clearly above nor below the male line. The message is clear: men and women with similar wages had similar rates of employment exit – women have had a much larger decline in employment largely due to having had lower wages in 1990. This point will now be developed more formally.

[4] Notice that if the longitudinal aspect of the data is not exploited the results are very different. If the comparison is made between wages in the 1990 bottom quantile and the 1994 bottom quantile, and similarly for the other pairs of quantiles, wage growth is found to be *rising* in quantile, and considerably smaller. This is the difference between my results and those of Steiner and Puhani (1996).

[5] In fact, the points for the two lowest quantiles have been suppressed for men, since only one man from the bottom decile in 1990 was still working in 1994.

3.3 Analysis of Attachment Spells

I turn here from analyzing employment to analyzing "attachment" spells (as defined in the data section). The hazard model used to analyze the durations of attachment spells is the Cox partial likelihood proportional hazards model. The hazard is assumed to be of the form

$$\lambda(t; z) = \lambda_0(t)e^{z(t)\beta}, \tag{4}$$

where z(t) contains potentially time-varying covariates describing the workers or their job characteristics, and $\lambda_0(t)$ is the baseline hazard, which is allowed to be non-parametric (if the covariates are measured as deviations from means, the baseline hazard may be interpreted as the hazard for the mean worker). The likelihood function is

$$L = \Pi L_i, \tag{5}$$

where

$$L_i = \frac{e^{\sum_j z_{ij}(t)\beta}}{[\sum_{h\epsilon R^i} e^{z_h(t)\beta}]^{m_i}} \tag{6}$$

and where m_i is the number of individuals leaving a spell at time i, the set R^i contains all observations that could have left at time i, and z_{ij} are covariates for the jth observation leaving at time i (see Kalbfleisch and Prentice 1980, p.74). The Peto-Breslow approximation is used to deal with ties (more than one spell ending in a given period). While covariates may be time-varying, to some extent I prefer to use 1990 values, since these are essentially exogenous to the transition. The exceptions are presence of children, marital status, and eligibility for early retirement, which must be time-varying to make sense even if some endogeneity is present.

The top left panel of Figure 4 plots the Kaplan-Meier survival curves by sex, and reveals as expected that the female survival rate is much lower than the male rate. The top right panel plots survival curves by highest attained education and by sex. The graph reveals that the big difference between the male and female survival curves is caused by the much lower survival rate for females with an apprenticeship. This confirms that the reduction in the proportion of women with an apprenticeship, with its beneficial effect on the wage gap, is indeed due to selective exit from employment.

The bottom two panels of Figure 4 begin the analysis of the cause of the survival rate gap. In the bottom left panel the survival curves are adjusted for wage differences: that is, a Cox partial likelihood proportional hazards regression is run for each sex separately with the log wage in 1990 as sole covariate. The survival curves are then evaluated and plotted for the mean wage in the pooled male/female sample. This adjustment removes a large proportion of the male/female survival rate gap. (We shall see below that this comes entirely through the different wage means, not through a different response of men and women to the wage.) The bottom right panel performs a similar adjustment but this time for the presence of a child one year old or less in the household. This does little to reduce the male/female gap: this is due to the fact that few households have children aged 0-1. (Notice that the adjustment here allows for a time-varying covariate.)

In Table 3 I report the results of hazard rate regressions for a pooled sample of men and women. In column (1) I include only a sex dummy as a covariate. Notice that baseline hazard is constrained to be the same for men and women (unlike in Figure 4). I report exponentiated coefficients (e^{β}), so the number reported implies that the female hazard rate is 68% higher than the male rate. In column (2) I add the log of the 1990 wage, which reduces the gap in the hazard rate to 28%, a reduction of 60%. In column (3) I add hours of work per week, to see if the effect of the wage partly reflects a tendency of low hour workers to leave employment. This does not appear to be the case, and since hours per week are insignificant even when further controls are added, I exclude it from the further specifications reported.

In column (4) I add more complete covariates (children, marital status, age, tenure and education), without allowing the coefficients to vary by sex, to see if the coefficient on sex becomes insignificant, but it does not. The significant coefficients are the age dummies, indicating as expected an extremely high hazard rate for those aged 56 or older in 1990, the dummy for general schooling, and the married dummy (which indicates a lower hazard rate for married people). In column (5) I add two time-varying covariates indicating whether a person was eligible either for the communist early retirement scheme (plan 1) or the post-communist scheme (plan 2). These dummies have large and significant coefficients, and naturally reduce the explanatory power of the older age dummies, which nevertheless remain generally significant. In column (6) I add dummies for 1990 firm size and industry (twelve categories), which does not affect the other coefficients in any interesting way.

We would expect some of the effects to vary by sex, however, such as the effect of children, and from Figure 4 we expect the effect of the education dummies to differ. Results not reported show that in fact interacting the apprenticeship dummy with sex suffices to make the sex coefficient insignificant. In columns (1) and (4) of Table 4, rather than introducing interactions, I estimate separate hazard rates by sex, which allows the baseline hazard to vary by sex (I omit the firm size and industry dummies). Here we see that the point estimates on the wage coefficient are very close for men and women. Also, the presence of a child age 0-1 in the household significantly increases the hazard rate (by a factor of two) for women. Women are not affected by the presence of older children, and men are not affected by the presence of children of any age. This suggests quite strongly that changes in child care availability cannot be a large component of the gap in the male/female survival rate, since I showed in Figure 4 that the implications of the presence of a child age 0-1 are small, despite the large coefficient. The coefficient on the married dummy is insignificant for women, but implies a significant reduction in the hazard for men.

In view of the importance of the decline in the proportion of women with apprenticeships on the mean wage, it is worth examining the education coefficients. Although the coefficient on the apprenticeship dummy is larger and more significant for women, it is not significant at conventional levels. If, however, the regression is rerun with the wage omitted, the coefficient becomes significant and larger (the coefficient is 1.5, the t-statistic 3.1 – the complete results are not reported). The apprenticeship dummy is thus proxying for something captured more completely by the wage, which suggests that the effect of selectivity on the wage gap is likely to be larger than the "observed X's" education component suggested in the analysis above.

A calculation may be performed to gauge the impact on the wage gap of the exit of low wage earners. The baseline hazard is computed for the model of column 4 in Table 3, and hence, using the estimated coefficients, the the predicted hazard and survival probability in each period for each person. The average predicted survival probability by the time of the 1994 interview is .69 for men and .54 for women. If the 31% of men and 46% of women with the lowest survival probabilities are dropped, the gender gap in mean 1990 wages for the remaining workers is .209 log points, only slightly larger than the actual 1994 gender gap of .191 among those surviving to the 1994 interview. [6] The importance of the 1990 wage may be assessed by assigning women the wage from their corresponding percentile in the male distribution, and recomputing the survival probabilities based on the old coefficients. The average predicted survival probability rises to .60. If the 31% of men and 40% of women with the lowest survival probabilities are dropped, the mean 1990 wage for those who remain is .259 log points, implying that selection based on wage accounts for 42% of the reduction in the gender gap. This calculation suggests that although some of the "gap" effect computed above is due to selective exits from employment, a large part is not.

The GSOEP gathers information each year on job changes since the previous year, and inquires as to the reason for any job ending. This information can be matched to the end of attachment spells. The number of missing values for the reason for the job end is quite high, however. Furthermore, as the 1995 wave has not been used, the reason for most jobs ending in 1994 is missing (only 36 uncensored spells (3%) end in 1994, however). Hence, a job end reason is available for 77% of the uncensored spells. The reasons for these spells ending are tabulated in Table 5. More than half the spells clearly ended involuntarily (due to firm closing, layoff or firing), while most of the others could be interpreted as ending voluntarily or involuntarily. Retirement is an important ending reason (23% of spells), although the retirements are clustered in the first two years after monetary union, when the early retirement program was available. Although the reasons given differ somewhat by sex, they neither suggest that women have been

[6] This would appear to imply that selection explains all of the change in the wage gap: however, there is a sizeable fall in the gender gap among those surviving to the 1994 interview.

disproportionately and possibly discriminatorily laid off, as some observers have suggested, nor do they suggest that women are demonstrating similar preferences to western women by withdrawing voluntarily from employment.

This information may be exploited to carry out competing risks analysis. I divide reasons for a spell end into two groups: clearly involuntary ("layoffs") and the rest ("other" or "non-layoffs"). I then assess the effects of the variables on the risk of being laid off (in which case the non-layoff spells are considered censored) and the risk of leaving a job for other reasons (in which case the layoff spells are considered censored). In both cases spells where the ending reason is unknown are recorded as censored. These results are reported for women and men separately in columns (2),(3), (5) and (6) of Table 4.

The two most important results are that the wage has its effect entirely through layoffs, while children have their effect entirely through "other" (and only for women). The fact that the wage affects only the risk of layoffs and not the "other" risk appears to rule out a labor supply story whereby those with low wages have chosen to leave employment. On the contrary, it provides strong circumstantial evidence that a labor demand story is at work, and that low wage individuals may have been laid off due to the union's raising their wage.

All the regressions in Table 4 have been rerun with firm size and industry dummies as additional regressors (these results are not shown). Somewhat surprisingly, this changes the other coefficients little (the apprenticeship dummy coefficient becomes less significant, the tenure coefficient more significant), and the industry dummies are jointly insignificant for men and women in the risk of a spell ending due to non-layoff reasons.

Although the results are not reported, the hazard rate regressions have been estimated for employment spells rather than attachment spells. The results are similar, although the coefficients on age dummies are smaller and all child variable coefficients are insignificant. The unadjusted gender difference in the hazard rate falls by 65% when the 1990 wage is controlled for, to only 17%.

4 Conclusions

Women's wages in eastern Germany rose by ten percentage points relative to men's wages in the four years following monetary union. I demonstrate that 40% of this gain may be accounted for by selective withdrawal of women from employment. More than half of the large disparity between men and women in terms of the hazard rate to non-employment is explained by differences in their initial wages. Workers with low pay in 1990 were much more likely to be laid off subsequently, and women had considerably lower pay than men in 1990. Initial wages do not affect the risk of losing a job for reason other than layoff, which implies that a labor demand rather than labor supply story is at work. For women, the presence of a child age one or younger increases greatly the risk of leaving a job for non-layoff reasons, yet few enough women had children of this age that it is not an important factor. The presence of older children does not increase the female hazard rate. The results suggest that a decline in child care availability is unlikely to have been important in causing more women to leave employment.

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Table 1a: Means of Individual Characteristics 1990 (June) and 1994
(Standard Deviations in Parentheses)

Variable	1990		1994	
	Women	Men	Women	Men
Log(wage)	7.10 (.38)	7.42 (.29)	7.75 (.46)	7.96 (.34)
Child 0-1 ?	0.03 (.16)	0.08 (.27)	0.00 (.04)	0.02 (.13)
Child 2-6 ?	0.22 (.42)	0.21 (.41)	0.15 (.35)	0.19 (.39)
Child 7-11 ?	0.28 (.45)	0.26 (.44)	0.26 (.44)	0.26 (.44)
Age <=35 ?	0.43 (.49)	0.42 (.49)	0.40 (.49)	0.39 (.49)
Age 36-45 ?	0.27 (.45)	0.26 (.44)	0.34 (.47)	0.32 (.47)
Age 46-50 ?	0.13 (.34)	0.14 (.35)	0.11 (.31)	0.11 (.32)
Age 51-55 ?	0.12 (.33)	0.12 (.32)	0.12 (.32)	0.13 (.34)
Age 56-60 ?	0.05 (.22)	0.06 (.24)	0.04 (.19)	0.05 (.21)
Tenure (months)	142 (114)	161 (129)	85 (100)	87 (112)
General schooling?	0.06 (.24)	0.03 (.16)	0.04 (.19)	0.03 (.16)
University?	0.09 (.28)	0.13 (.34)	0.12 (.33)	0.14 (.35)
Apprenticeship?	0.61 (.49)	0.62 (.48)	0.52 (.50)	0.63 (.48)
Vocational training?	0.25 (.43)	0.22 (.41)	0.32 (.47)	0.20 (.40)
Married ?	0.76 (.43)	0.77 (.42)	0.75 (.44)	0.76 (.43)
Hours work per week	40.3 (8.0)	45.9 (6.9)	39.9 (7.9)	45.4 (7.8)
Observations (escapes)	1047 (462)	1014 (297)	671 (--)	757 (--)

Table 1b: Means of Individuals' Job Characteristics 1990 (June) and 1994 (Standard Deviations in Parentheses)

Variable	1990		1994	
	Women	Men	Women	Men
Firm size < 20 ?	0.12 (.33)	0.08 (.27)	0.24 (.43)	0.25 (.43)
Firm size 20-199 ?	0.29 (.45)	0.21 (.41)	0.32 (.47)	0.34 (.47)
Firm size 200-1999 ?	0.35 (.48)	0.36 (.48)	0.23 (.42)	0.23 (.42)
Firm size 2000+ ?	0.24 (.42)	0.35 (.48)	0.20 (.40)	0.19 (.39)
Mining/quarrying/energy	0.04 (.20)	0.10 (.29)	0.02 (.15)	0.06 (.23)
Chemicals/synthetics	0.04 (.20)	0.05 (.22)	0.02 (.15)	0.03 (.18)
Iron/steel	0.04 (.20)	0.07 (.26)	0.01 (.12)	0.06 (.24)
Mechanical engineering	0.03 (.17)	0.10 (.30)	0.02 (.13)	0.07 (.26)
Electrical engineering	0.08 (.27)	0.09 (.29)	0.02 (.14)	0.04 (.20)
Wood/paper/leather/textiles/food	0.10 (.30)	0.09 (.28)	0.04 (.19)	0.05 (.22)
Construction	0.03 (.16)	0.13 (.33)	0.03 (.17)	0.23 (.42)
Retail/wholesale trade	0.15 (.35)	0.05 (.22)	0.14 (.35)	0.11 (.31)
Transportation	0.07 (.26)	0.13 (.34)	0.06 (.24)	0.11 (.31)
Private services	0.05 (.21)	0.02 (.15)	0.14 (.34)	0.04 (.20)
Education/science/sport	0.15 (.36)	0.07 (.26)	0.16 (.37)	0.05 (.23)
Health	0.14 (.35)	0.03 (.17)	0.15 (.35)	0.02 (.14)
Government	0.08 (.28)	0.07 (.25)	0.18 (.39)	0.12 (.32)
Observations (escapes)	1047 (462)	1014 (297)	671 (--)	757 (--)

Table 2: Decomposition of Change in Gender Gap 1990-1994

	Controls for age, education, tenure		
	only	and firm size, industry	and hours
A. Descriptive statistics			
Male residual standard deviation 1990	0.25	0.25	0.23
1994	0.30	0.28	0.27
Mean female residual percentile 1990	21	21	25
1994	33	35	39
B. Decomposition of change			
Change in differential (1994-1990):	-0.115		
Observed X's:	-0.035	-0.044	-0.053
Education	-0.032	-0.032	-0.031
Age and tenure	-0.003	0.001	0.001
Firm size	--	-0.025	-0.025
Industry	--	0.011	0.008
Hours	--	--	-0.006
Observed prices:	0.009	0.046	0.028
Education	0.009	0.008	0.005
Age and tenure	0.001	-0.003	-0.003
Firm size	--	0.017	0.018
Industry	--	0.024	0.020
Hours	--	--	-0.013
Gap:	-0.103	-0.096	-0.083
Unobserved prices:	0.013	-0.021	-0.007
Sum gender-specific:	-0.138	-0.140	-0.136
Sum wage structure:	0.023	0.025	0.020

Note: Decomposition is based on mean regression. See text for details.

Table 3: Determinants of Single Risk Hazard Rate Out of Employment Attachment
(Exponentiated coefficients; t-statistics in parentheses)

	(1)	(2)	(3)	(4)	(5)	(6)
Sex (female=1)	1.68 (7.0)	1.28 (3.0)	1.28 (2.9)	1.43 (4.2)	1.35 (3.5)	1.42 (3.8)
Log($wage_{90}$)	--	0.44 (-8.4)	0.44 (-7.0)	0.54 (-5.3)	0.54 (-5.3)	0.55 (-4.9)
Hours per $week_{90}$	--	--	1.00 (-0.1)	--	--	--
$Child_t$ 0-1 ?	--	--	--	1.19 (0.7)	1.19 (0.7)	1.21 (0.8)
$Child_t$ 2-6 ?	--	--	--	0.86 (-1.3)	0.86 (-1.3)	0.87 (-1.1)
$Child_t$ 7-11 ?	--	--	--	0.89 (-1.1)	0.89 (-1.1)	0.91 (-0.9)
Age_{90} <=35 ?	--	--	--	1.27 (2.1)	1.26 (2.0)	1.24 (1.9)
Age_{90} 46-50 ?	--	--	--	0.96 (-0.3)	0.97 (-0.3)	1.00 (0.0)
Age_{90} 51-55 ?	--	--	--	2.33 (6.8)	1.33 (1.9)	1.38 (2.1)
Age_{90} 56-60 ?	--	--	--	8.24 (14.7)	1.75 (2.6)	1.87 (2.8)
$Tenure_{90}$ (months)	--	--	--	1.00 (-0.3)	1.00 (-0.6)	1.00 (-1.7)
$University_{90}$?	--	--	--	0.88 (-0.8)	0.87 (-0.9)	0.82 (-1.2)
$Apprenticeship_{90}$?	--	--	--	1.10 (1.0)	1.09 (0.9)	1.03 (0.2)
General $schooling_{90}$?	--	--	--	1.67 (3.0)	1.65 (3.0)	1.62 (2.7)
$Married_t$?	--	--	--	0.80 (-2.4)	0.81 (-2.2)	0.84 (-1.9)
Eligible for early pension 1_t?	--	--	--	--	5.63 (6.8)	6.22 (7.1)
Eligible for early pension 2_t?	--	--	--	--	5.20 (9.6)	5.52 (9.9)
Firm size, industry $dummies_{90}$?	no	no	no	no	no	yes
Chi-square (d.f.)	49.8 (1)	115 (2)	115 (3)	400 (14)	499 (25)	557 (31)

Notes:
a. Estimation is by Cox partial likelihood proportional hazard.
b. There are 2061 spells, of which 759 are uncensored.

Table 4: Competing Risks and Analysis by Sex
(Exponentiated coefficients; t-statistics in parentheses)

	Women			Men		
	All Exits	Layoffs	Other Exits	All Exits	Layoffs	Other Exits
Log(wage$_{90}$)	0.56 (-4.1)	0.49 (-3.5)	0.98 (-0.1)	0.58 (-2.5)	0.51 (-2.0)	1.19 (0.4)
Child$_t$ 0-1 ?	2.28 (2.4)	0.54 (-0.6)	3.55 (2.6)	1.03 (0.1)	1.31 (0.6)	0.82 (-0.3)
Child$_t$ 2-6 ?	0.98 (-0.2)	0.99 (-0.0)	1.05 (0.2)	0.79 (-1.1)	0.82 (-0.7)	0.53 (-1.3)
Child$_t$ 7-11 ?	0.92 (-0.6)	0.92 (-0.4)	0.58 (-2.0)	0.91 (-0.5)	1.02 (0.1)	0.63 (-1.1)
Age$_{90}$ <=35 ?	1.23 (1.5)	0.90 (-0.5)	3.56 (3.8)	1.18 (0.8)	0.70 (-1.2)	2.32 (1.9)
Age$_{90}$ 46-50 ?	0.96 (-0.2)	0.77 (1.1)	1.05 (0.1)	1.03 (0.1)	1.31 (0.9)	0.53 (-0.9)
Age$_{90}$ 51-55 ?	1.13 (0.6)	1.15 (0.6)	0.88 (-0.2)	1.71 (2.2)	1.47 (1.1)	2.30 (1.6)
Age$_{90}$ 56-60 ?	1.23 (0.7)	1.38 (0.7)	1.00 (0.0)	2.21 (2.5)	4.09 (3.8)	3.14 (2.0)
Tenure$_{90}$	1.00 (-1.1)	1.00 (-0.6)	1.00 (-1.6)	1.00 (0.4)	1.00 (-2.4)	1.00 (1.9)
University$_{90}$?	0.84 (-0.7)	0.99 (-0.0)	0.82 (-0.5)	0.90 (-0.5)	1.13 (0.4)	0.64 (-1.1)
Apprenticeship$_{90}$?	1.25 (1.7)	1.47 (1.9)	1.07 (0.3)	1.07 (0.4)	0.94 (-0.2)	1.70 (2.0)
General schooling$_{90}$?	1.55 (2.1)	1.85 (1.8)	0.98 (-0.0)	4.12 (4.8)	2.25 (1.6)	7.69 (3.7)
Married$_t$?	0.87 (-1.2)	1.02 (0.1)	0.84 (-0.9)	0.62 (-2.6)	0.74 (-1.1)	0.63 (-1.3)
Eligible for early pension 1$_t$?	9.22 (6.8)	--	54.60 (6.0)	4.15 (2.2)	--	9.25 (2.7)
Eligible for early pension 2$_t$?	5.19 (6.9)	--	62.91 (6.7)	5.57 (6.9)	--	21.8 (7.2)
Spells (exits)	1047 (462)	1047 (203)	1047 (153)	1014 (297)	1014 (122)	1014 (105)
Chi-square (15)	243	39.1	318	241	34.3	287

Notes:
a. Estimation is by Cox partial likelihood proportional hazard.

Table 5: Reasons for End of Attachment Spell (%)
(Number of spells in parentheses)

Reason	Women	Men	All
Firm closed	23.3 (83)	22.0 (50)	22.8 (133)
Laid off/fired	33.7 (120)	31.7 (72)	32.9 (192)
Retired	4.8 (17)	1.3 (3)	3.4 (20)
Took early retirement	16.0 (57)	26.4 (60)	20.1 (117)
Fixed term contract ended	3.9 (14)	6.2 (14)	4.8 (28)
Quit	5.9 (21)	5.3 (12)	5.7 (33)
Parental leave	9.8 (35)	1.8 (4)	6.7 (39)
Other	2.5 (9)	5.3 (12)	3.6 (21)
Total	100% (356)	100% (227)	100% (583)

Table A1: Median Wage Regressions 1990, 1994
(t-statistics in parentheses)

	1990		1994	
Sex	-0.256 (-17.9)	-0.182 (-15.8)	-0.145 (-8.9)	-0.076 (-4.0)
Age	0.015 (3.2)	0.016 (4.6)	0.027 (4.7)	0.027 (4.1)
Age^2 (*100)	-0.021 (-3.6)	-0.021 (-4.7)	-0.031 (-4.3)	-0.031 (-3.8)
Tenure (*100) (months)	0.045 (6.3)	0.033 (6.0)	0.026 (3.3)	0.025 (2.9)
University?	0.148 (6.0)	0.144 (7.7)	0.199 (8.0)	0.180 (6.4)
Apprenticeship?	-0.250 (-15.0)	-0.236 (-18.7)	-0.191 (-10.8)	-0.158 (-7.9)
General schooling?	-0.437 (-12.7)	-0.411 (-15.8)	-0.372 (-8.4)	-0.316 (-6.2)
Firm size 20-199 ?	0.081 (3.4)	0.077 (4.2)	0.131 (6.7)	0.138 (6.3)
Firm size 200-1999?	0.093 (4.0)	0.076 (4.2)	0.214 (9.9)	0.208 (8.4)
Firm size 2000+ ?	0.123 (4.9)	0.115 (6.1)	0.248 (10.1)	0.228 (8.2)
Industry dummies?	yes	yes	yes	yes
Hours controls?	no	yes	no	yes
R^2	0.29	0.39	0.26	0.33
Observations	2061		1428	

Notes:
a. Dependent variable is log of real monthly wage.
b. Omitted educational category is vocational training.
c. Controls for hours are: hours per week and its square, and dummies for 30 or fewer hours, 31-40 hours, and 41-45 hours.

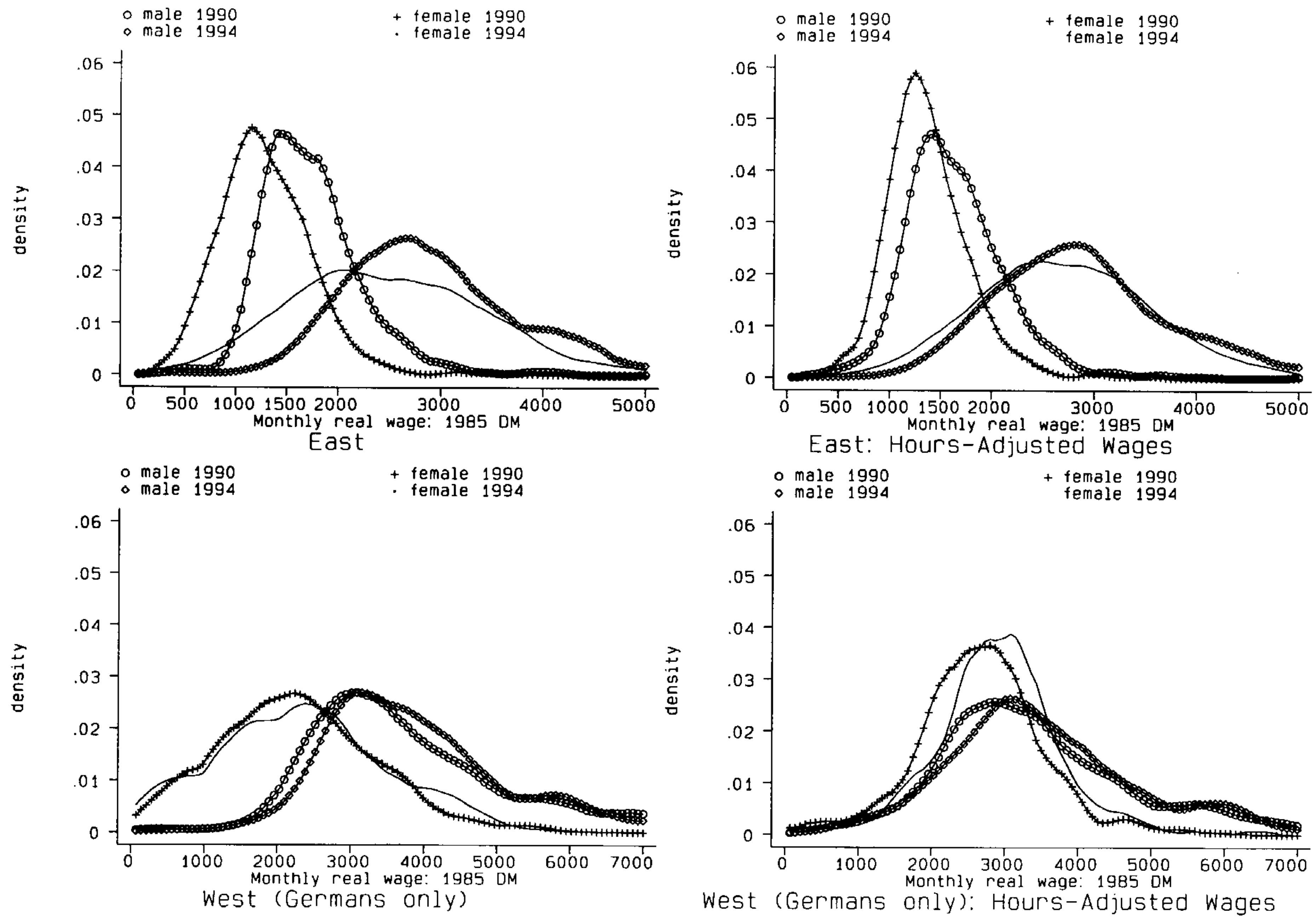

Figure 1: Distribution of Monthly Wages in East and West

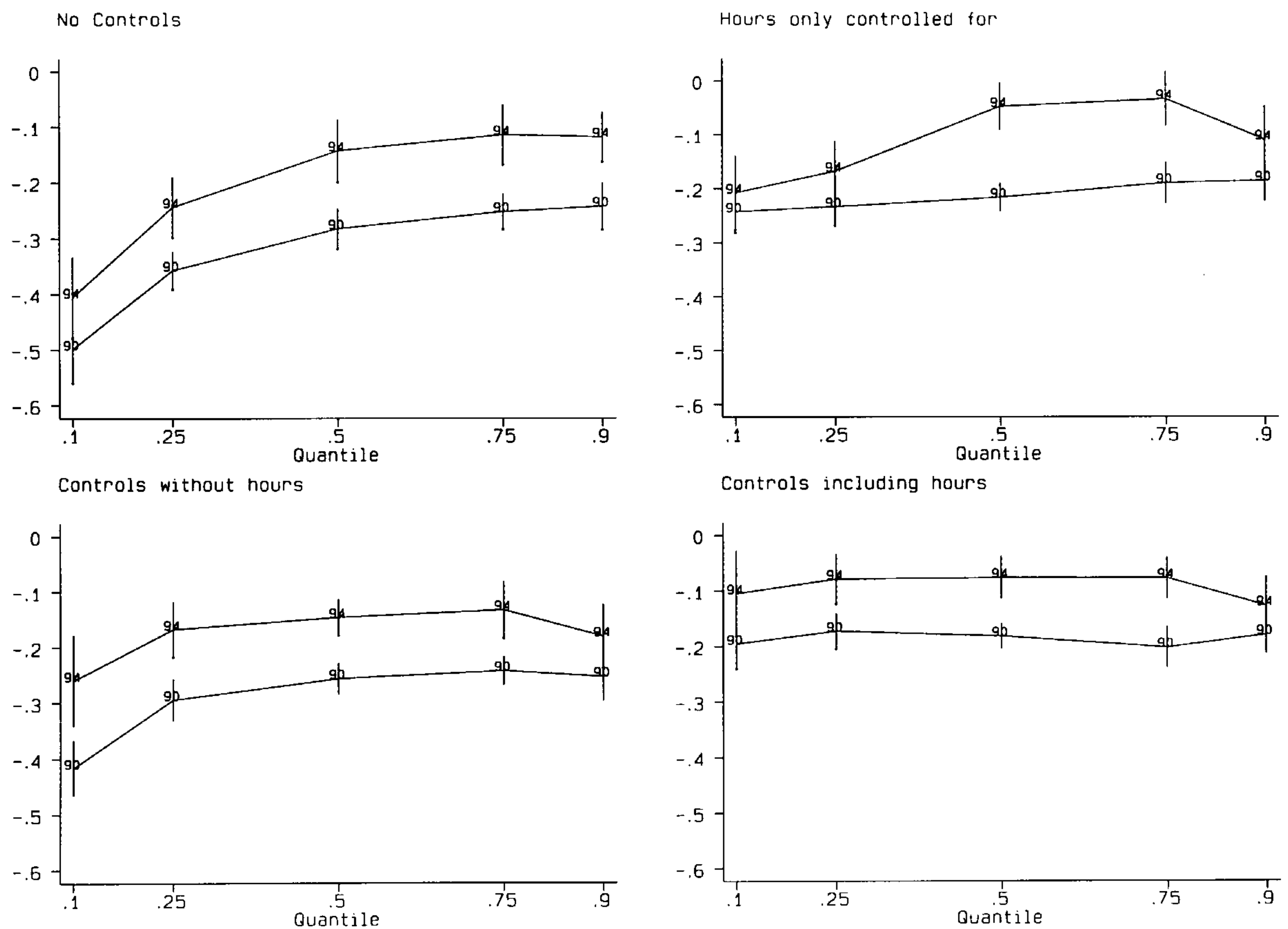

Figure 2: Coefficient on Sex Dummy in Log Wage Regression 1990,94

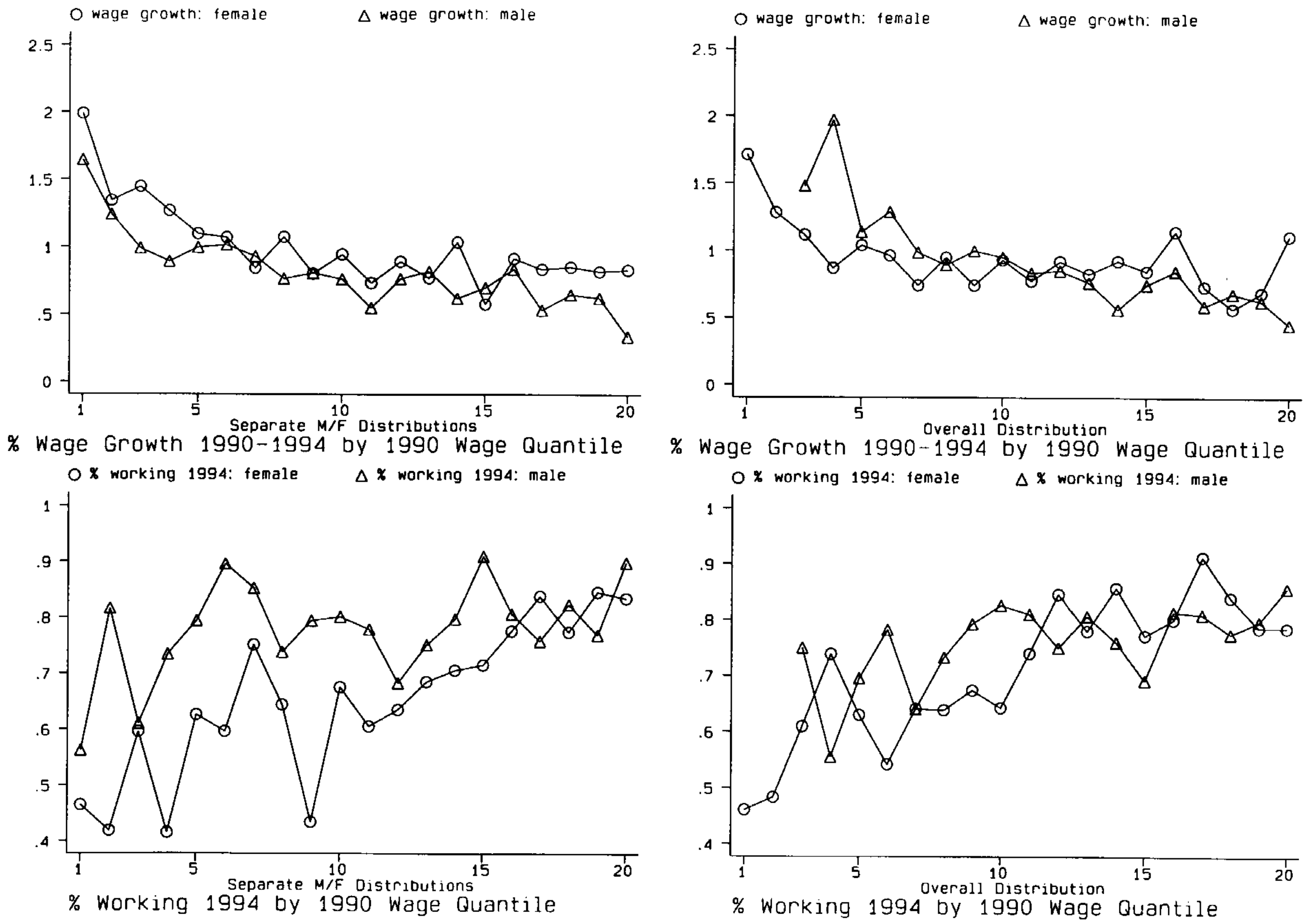

Figure 3: 1994 Wage and Employment by 1990 Wage Quantile

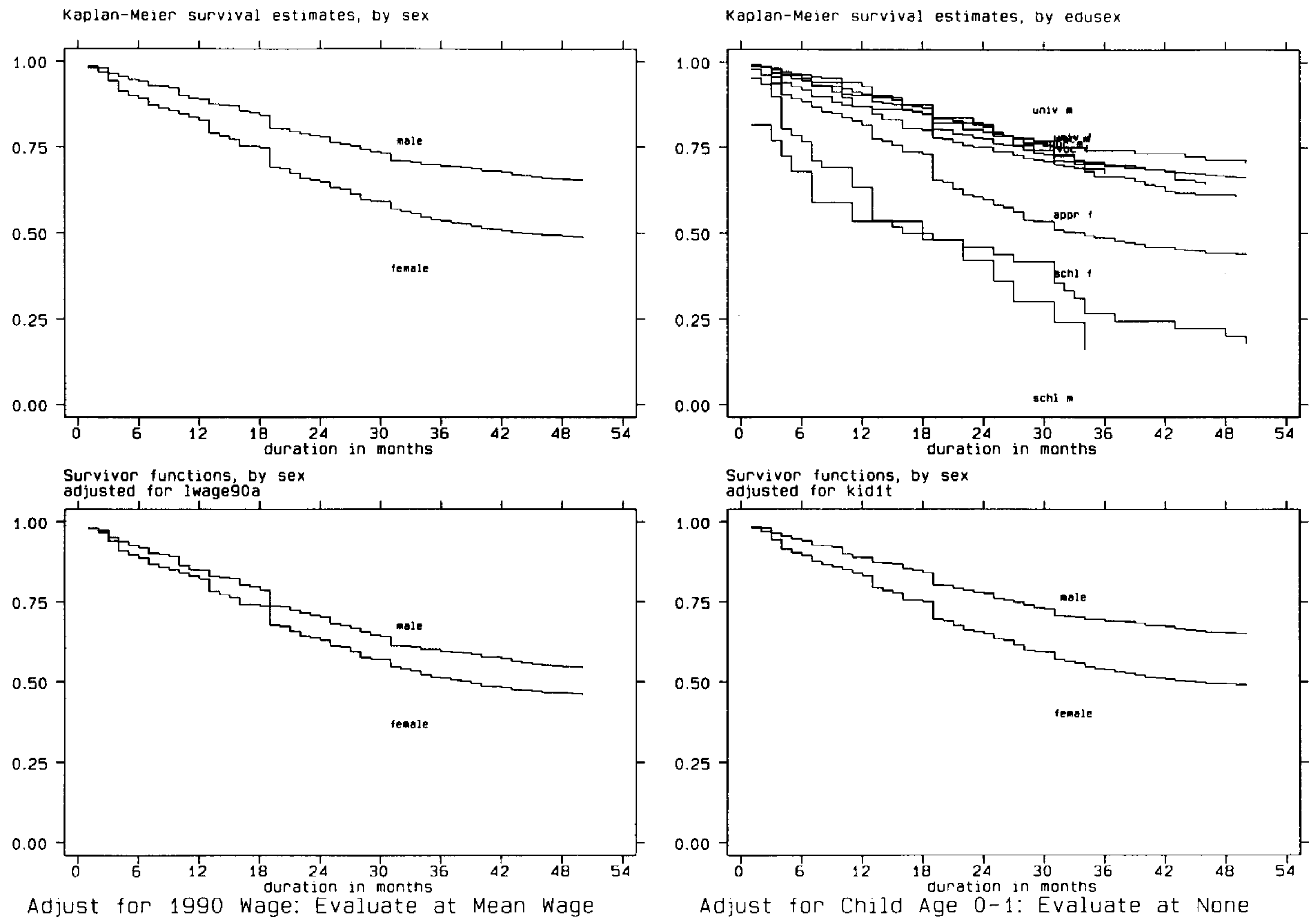

Figure 4: Attachment Spells in Progress June 1990