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German Wage Underpayment: An Investigation into Labor Market Inefficiency and Discrimination

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Summary

Using stochastic panel wage frontiers, this paper estimates the relative underpayment of females and males in the reunified Germany. The estimates are initially applied to discrimination analysis. It finds that females have higher underpayment and that the male-female differential increased over the period 1991–1993. However, the paper suggests that the estimates of underpayment reflect other concerns, such as dynamic monopsony. Labor market inefficiency must be taken into account before discrimination analysis is possible.

1. Introduction

It is becoming increasingly common, in the study of wage rates, to adopt a technique known as the stochastic frontier approach, which has more commonly been used to measure cost and production efficiencies. For instance, Robinson and Wunnava (1989) use it to measure gender discrimination in the United States, thereby offering a replacement to the standard decomposition technique developed by Blinder (1973) and Oaxaca (1973). Its advantage over ordinary least squares is that it allows for the presence of labor market inefficiencies. This attribute is attractive because theory predicts that workers will be underpaid in labor markets characterized by imperfect information (Hofler and Murphy 1992). To optimize lifetime income in the presence of asymmetric information, job searchers will adopt a reservation wage strategy and will accept wages below the maximum wage that could be earned, given their human capital attributes. Labor market inefficiency is then the consequence of these job market frictions and firms derive monopsony power. This paper estimates stochastic panel frontier models in order to evaluate the validity of the stochastic frontier approach.

This paper uses data from the 1991-1993 waves of the German Socio-Economic Panel (GSOEP) to estimate German underpayment between 1991 and 1993. This period is chosen because German labor markets changed substantially after reunification in 1990. One of the principal concerns that emerged in the public debate during this period was how the less skilled East German workers would be absorbed into the highly efficient West German labor force. Under the planned socialist economy, all economically active workers under 50 years old were employed. By the end of 1991, the percentage of workers employed had stabilized at 80 percent after dipping as low as 70 percent during the year (Lechner 1997, 5). The unemployment rate of East Germans had increased from 2 percent before unification to 12 percent by the end of 1993 (ibid.). The considerable shock reunification visited on the East German labor force was tempered by training programs specifically targeted at them to smooth the transition from a planned to a competitive labor market. Such programs apparently worked. Between 1990 and 1991 women, low-wage workers, and low-skilled workers, in particular, gained most from the unification of the labor market (Hunt 1999). East German females, in particular, appear to have gained from political and economic unification, illustrated by Hunt's (1997) finding that between 1990 and 1994 the monthly wage of this group increased 10 percent relative to males. This trend, however, hides the growing problem of female unemployment in East Germany, which according to Hunt (1997) and Lange (1996, 19) has become "disproportionately higher than those of their male counterparts." Lange (1996) argues the increasing female unemployment problem is due to gender occupational discrimination.

The next section reviews the literature on gender discrimination and gender earnings differentials using both decomposition and stochastic frontier analysis, with specific attention paid to unified Germany. Section 3 then presents the stochastic panel frontier method that is used. Section 4 discusses the data. In Section 5 we present and interpret results. We end with a conclusion.

2. Previous Literature

It has been standard for several years to decompose gender or racial wage gaps into a portion explained by observed differences in individual characteristics and a residual that remains unexplained (Blinder 1973; Oaxaca 1973). The explained and unexplained wage gaps are typically estimated with Mincerian wage equations, separately for each group. The wage differential accounted for by the model is labeled a productivity difference and the residual or unexplained wage difference is often attributed to gender or racial discrimination. The decomposition technique has been refined and applied in numerous subsequent studies (e.g., Neumark 1988; Hinks, Atkins, and Allanson 2000). Applying these standard techniques, Gerlach (1987), Hubler (1991) and Black, Trainor and Spencer (1999) find evidence that gender discrimination

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exists in the German labor market. The most comprehensive and comparable study to this paper though comes from Brookes, Hinks, and Watson (2000). Using panel estimates for 1991 to 1993, they find that German females were on average paid 26 percent less than observationally similar males, whereas in the UK the figure was 34 percent. When decomposed into discrimination components, it was found that differences in returns to observed characteristics dominated these differentials, explaining over 50 percent of the female-male wage differential in both countries.

Decomposition analysis does, however, suffer limitations. In particular, the residual wage gap includes omitted variables, such as worker's ability, that will contaminate its interpretation as an estimate of the extent of discrimination in a given labor market. Whilst this bias can be mitigated if information on omitted variables is available, such as IQ tests, such data are still relatively rare. Another criticism of the decomposition procedure is the generally ad hoc nature of estimating a competitive wage structure. Probably the most damaging criticism, though, is the very interpretation of the unexplained component as evidence of labor market discrimination. Because it is a residual, the unexplained wage difference might just as reasonably be interpreted as resulting from differences in unexplained personal characteristics.

These criticisms have led to the application of alternative techniques to identify the extent of discrimination and underpayment. One such alternative is stochastic frontier analysis, an econometric technique that adapts the wage equation to include a one-sided error term to capture the effect of underpayment. This technique has been used to estimate non-white and female wage underpayments (Robinson and Wunnava 1989). It assumes that, by estimating a wage frontier for females dependent on human capital characteristics, the extent of labor market discrimination can be represented by the difference between observed female earnings and the earnings that would be paid at the efficient wage frontier. This approach is superior to the decomposition method because it relaxes the assumption that males and females have identical unmeasured qualities and because it allows for the estimation of an individual discrimination term rather than a sample average. The principal problem with stochastic frontier analysis is the interpretation of the error term remains arbitrary. The majority of papers interpret the one-sided error term as the underpayment caused by labor market frictions. Such analyses, for example by Hofler and Murphy (1992) and Watson (2000), adopt a job search method whereby workers set a reservation wage below their productivity. When individuals maximize lifetime income and firms maximize profits, firms derive monopsonistic power and, given the individual's reservation wage strategy, workers cannot be overpaid. In contrast, Robinson and Wunnava (1989) interpret the one-sided underpayment error term as evidence of direct discrimination. This paper undertakes this standard analysis but acknowledges that underpayment can be caused by both discriminatory and labor market friction factors. This means both males and females can be underpaid due to labor market frictions in the search for jobs, but that females may also encounter wage discrimination and hence suffer even greater underpayment.

3. The Model

We adopt the standard Mincerian (log) earnings function:

$$Ln(w_{it}) = x_{it}\, b + v_{it} \tag{1}$$

where x_{it} is a vector of worker characteristics and b is a vector of parameters to be estimated. The vector of worker characteristics includes schooling, experience, and age. Adding a one-sided inefficiency term to (1) produces the stochastic earnings frontier:

$$Ln(w_{it}) = x_{it}\, b + (v_{it} - u_{it}) \tag{2}$$

The earnings function is perceived as a frontier. The frontier represents the potential (maximum) wage that a worker could earn given his or her characteristics. The inefficiency score represents the extent to which the actual earnings fall short of potential earnings. We classify this difference as an underpayment index.

In order to estimate u_{it} using maximum likelihood (ML) methods, it is necessary to specify the distribution of the residual components:

$$v_{it} \sim N(0, s_v^2) \tag{3}$$

and

$$u \sim N(m, s_u^2) \tag{4}$$

where u_{it} follows the truncated-normal distribution[1].

To simplify expressions that follow we drop the subscripts i and t. The density function that corresponds to equation (3) is defined as:

$$f(v) = \frac{\exp - \frac{1}{2}\left(\frac{v}{s_v}\right)^2}{s_V \sqrt{2p}} \qquad \text{for all } v \tag{3a}$$

$$g(u) = \frac{\exp - \frac{1}{2}\left(\frac{u-m}{s_u}\right)^2}{s_u\left(1 - \Phi\left(-m / s_u\right)\right)\sqrt{2p}} \qquad \text{for } u \geq 0 \tag{4a}$$

$= 0$ otherwise,

[1] The one-sided error term implies overpayment cannot occur. While this may appear to be a strong assumption, it is consistent with both the job-search model and discrimination model that are analyzed in this paper.

where $\Phi\ (\cdot)$ represents the distribution function for the standard normal random variable. These are normal and truncated-normal distributions respectively[2].

The log-likelihood function, assuming $u_i\ (= u_{it})$ follows the truncated-normal distribution and v_{it} is normal, is given by

$$L(J^*;Y) = -\frac{n}{2}T\ln s^2 - \frac{n}{2}T\ln 2p - \frac{n}{2}(T-1)\ln(1-g) - \frac{n}{2}\ln[1+(T-1)g]$$
$$-n\ln[1-\Phi(-z)] - \frac{n}{2}z^2 + n\ln[1-\Phi(-z_*)] + \frac{n}{2}z_*^2$$
$$-\frac{nT}{2}\frac{(Ln(w_{it}) - x_{it}b + m)^2}{(1-g)s^2} \qquad (5)$$

where $z = \frac{m}{(gs^2)^{1/2}}$, $z_* = \frac{m(1-g) - gT(y - x_{it}b)}{\{g(1-g)s^2[1+(T-1)g]\}^{1/2}}$ and $J^* = (b_0, b_1, b_2, m, s^2, g)$.

Individual efficiency is calculated from the conditional probability of u_i given e_{it} (i.e., $e_{it} = v_{it} - u_i$):

$$E[\exp(-u_i|e_i)] = \exp(-m_* + s_*^2/2)\left[\frac{1-\Phi(s_* - (m_*/s_*))}{1-\Phi(-m_*/s_*)}\right] \qquad (6)$$

where $m_* = \frac{(-s_u^2\bar{e}_i + ms_v^2T^{-1})}{s_u^2 + s_v^2T^{-1}}$ and $s_* = \left(\frac{s_u^2 s_v^2}{Ts_u^2 + s_v^2}\right)$ such that $\bar{e}_i$ is defined as $\bar{e}_i = T^{-1}\sum_{t=1}^{T}\hat{e}_{it}$.

An alternative technique used in this paper is a time-variance panel model, based on the formulation offered by Battese and Coelli (1992), that assumes u_{it} is a deterministic function of time:

$$u_{it} = (\exp[-h(t-T)])u_i \qquad (7)$$

where h is the time-varying parameter that either increases, decreases, or remains constant as t increases. If h is constant, then efficiency is time-invariant. If h is positive, efficiency monotonically increases over time (inefficiency falls), whereas if h is negative, efficiency monotonically decreases over time (inefficiency rises). T denotes the last period in the panel and the time period under consideration is denoted by t. Therefore, if a particular worker is observed in the last season $t = T$ and $u_{iT} = u_i$ (since $exp(0) = 1$). In this case efficiency will monotonically rise to this level if h is positive ($u_{iT} > u_i$) and monotonically fall to this level if h is negative ($u_{iT} < u_i$). Because h is the same for each employee, the ordering of efficiency is preserved in all time periods.

For this model the log-likelihood function is[3]:

$$L(J^*;Y) = -\frac{n}{2}T\ln s^2 - \frac{n}{2}T\ln 2p - \frac{n}{2}(T-1)\ln(1-g) - \frac{n}{2}\ln[1+(h^2-1)g]$$
$$-n\ln[1-\Phi(-z)] - \frac{n}{2}z^2 + n\ln[1-\Phi(-z_*)] + \frac{n}{2}z_*^2$$
$$-\frac{nT}{2}\frac{(Ln(w_{it}) - x_{it}b)^2}{(1-g)s^2} \qquad (8)$$

where z is as defined previously,

$z_* = \frac{m(1-g) - gh(y - x_{it}b)}{\{g(1-g)s^2[1+(h^2-1)g]\}^{1/2}}$ and

$J^* = (b_0, b_1, b_2, m, h, s^2, g)$.

Individual efficiency is found by applying the formula:

$$E[\exp(-u_i|e_i)] = \exp(-hm_* + h^2s_*^2/2)\left[\frac{1-\Phi(hs_* - (m_*/s_*))}{1-\Phi(-m_*/s_*)}\right] \qquad (9)$$

where $m_* = \frac{(-hs_u^2\bar{e}_i + ms_v^2)}{h^2s_u^2 + s_v^2}$ and $s_* = \left(\frac{s_u^2 s_v^2}{h^2s_u^2 + s_v^2}\right)$.

4. Data

The paper uses data from the German Socio-Economic Panel (GSOEP) for the period 1991 to 1993. The panel sample was reduced from 18,041 to 11,658 to restrict the sample to employed persons aged 18 to 65. Of this sample 7,541 were male and 4,117 were female. The wage frontier is assumed to depend on human capital variables (age, level of education, on-the-job experience), personal characteristics (married, number of children), industry dummies based on ISIC codes, year dummies, and an East-West Germany dummy. The reference industry consists of Finance, Insurance, Real Estate and Business Services. On-the-job experience is defined in terms of length of employment on the current job. When we have data on how long a person has been employed in a particular job, we code a person as a “known insider” if he has been employed in the current job for over 5 years. A “known outsider” is an individual who has remained in their current job for less than 2 years. The reference group consists of “initiated” workers with 2 to 5 years of job-specific experience. It is expected that workers with greater job experience will earn more than less-experienced workers[4].

5. Results

Initially, time-invariant wage frontiers are estimated for male and female full-time workers. Table 1 illustrates that sign on most of the coefficients are as expected and sta-

[2] The term u_{it} can also be estimated using half-normal distributions, but it was found that the half-normal distribution was rejected in favor of the truncated normal distribution, by the log-likelihood test. These results are available upon request from the authors.

[3] Full derivation of the log-likelihood can be found in the appendix of Battese and Coelli (1992).

[4] The returns to being an experienced worker (insider) may be biased downward because individuals who failed to indicate how long they had been employed were designated new workers (outsider). Those individuals were retained in our sample to maintain a sufficiently large sample to conduct the analysis.

tistically different from zero at conventional levels of significance. Wages follow a quadratic relationship with respect to age, consistent with human capital theory, and the coefficients on the age terms are highly significant. Married women earn significantly less than their unmarried counterparts, but earnings of married and unmarried males do not statistically differ. This finding is consistent with empirical evidence from the United Kingdom and the United States, where Hill (1985) and Juster and Stafford (1991) find that married women workers spend more time on domestic household duties than single females. They hypothesize that women thus gain less labor market experience and, therefore, command a lower wage. The coefficient on number of children is statistically significant and positive for males, but negative and insignificant for females. This finding is consistent with the conjecture that males with families work harder and therefore, earn more to support their offspring, and, that females with children participate less in the labor market and command lower earnings because they enjoy a comparative advantage working at home (Polachek and Siebert 1993). Such an argument has been criticized by Bergmann (1981, 84) for not recognizing that the role of the female in a household is dominated by the role of the male, and that essentially 'caste' discrimination against females dictates this apparent comparative advantage. The one surprising result is that relatively new male and female workers (known outsiders) earn more than newly initiated insiders, though the estimates are insignificant for females. Known insiders predictably earn more than either outsiders or the newly initiated, for both male and female workers. As predicted by human capital theory, coefficients on education are positive and highly significant. Finally, East German workers command significantly lower wages relative to West German workers.

The findings indicate that both East and West German female workers are paid 27 percent less than their potential earnings (100–73 percent). Robinson and Wunnava (1989) interpret such a shortfall as evidence of direct labor market discrimination.

Comparisons with the labor market in the United Kingdom in the same period have been undertaken, with the finding that females earn 47 percent less than their full earnings potential. The difference in the earnings shortfall of women in Germany and the United Kingdom can be explained by differences between the social and political institutions of Germany and the United Kingdom, including more powerful trade unions in Germany and minimum wage legislation that protects the employed worker. Note

Table 1

Time-Invariant Efficiency, Full-Time Sample Only (German Data)

Variable	Half-Normal		Truncated-Normal	
	Male	Female	Male	Female
Constant	6.756 (0.074)	6.620 (0.104)	6.774 (0.071)	6.587 (0.095)
1992 dummy variable	0.010 (0.007)	0.024 (0.010)	0.008 (0.007)	0.027 (0.010)
1993 dummy variable	0.025 (0.007)	0.051 (0.010)	0.023 (0.007)	0.053 (0.010)
Agriculture and Fishing	–0.262 (0.037)	–0.335 (0.049)	–0.271 (0.036)	–0.336 (0.047)
Mining and Quarrying	–0.113 (0.042)	–0.023 (0.081)	–0.121 (0.040)	–0.034 (0.078)
Manufacturing	–0.083 (0.025)	–0.189 (0.027)	–0.084 (0.025)	–0.185 (0.025)
Electricity, Gas and Water	–0.027 (0.037)	0.052 (0.067)	–0.029 (0.037)	0.044 (0.061)
Construction	–0.056 (0.027)	–0.113 (0.050)	–0.055 (0.026)	–0.104 (0.048)
Retail and Sales	–0.146 (0.029)	–0.251 (0.029)	–0.146 (0.028)	–0.246 (0.027)
Transport and Communication	–0.134 (0.030)	–0.075 (0.040)	–0.144 (0.029)	–0.090 (0.037)
Community and Social Services	–0.150 (0.026)	–0.087 (0.027)	–0.159 (0.025)	–0.094 (0.025)
Age	0.049 (0.004)	0.051 (0.006)	0.041 (0.003)	0.049 (0.005)
Age^2	–0.0005 (0.00004)	–0.0006 (0.00007)	–0.0004 (0.00004)	–0.0005 (0.00007)
Known insider	0.039 (0.015)	0.014 (0.020)	0.041 (0.015)	0.015 (0.019)
Known outsider	0.057 (0.014)	0.055 (0.020)	0.049 (0.014)	0.053 (0.019)
Married	0.011 (0.013)	–0.049 (0.016)	0.003 (0.012)	–0.043 (0.015)
Number of Children	0.016 (0.005)	–0.0004 (0.009)	0.019 (0.005)	–0.001 (0.009)
Years of Education	0.075 (0.002)	0.067 (0.003)	0.075 (0.002)	0.067 (0.003)
East German resident	–0.552 (0.015)	–0.488 (0.022)	–0.554 (0.015)	–0.482 (0.020)
σ^2	0.279 (0.008)	0.376 (0.014)	0.762 (0.047)	1.079 (0.073)
γ	0.815 (0.007)	0.848 (0.008)	0.930 (0.005)	0.945 (0.004)
μ	0	0	–1.684 (0.144)	–2.020 (0.149)
LL	–1880.271	–1435.122	–1836.138	–1372.587
Iterations	25	25	37	32
ME	**0.713**	**0.684**	**0.759**	**0.727**
N (CSO)	7541 (3573)	4117 (2073)	7541 (3573)	4117 (2073)

Source: Authors' calculations.

that the 27 percent underpayment term estimated using the stochastic frontier analysis is exactly half of the German female underpayment term calculated using decomposition analysis in Brookes, Hinks, and Watson (2000), and in this regard is consistent with the findings of Robinson and Wunnava (1989).

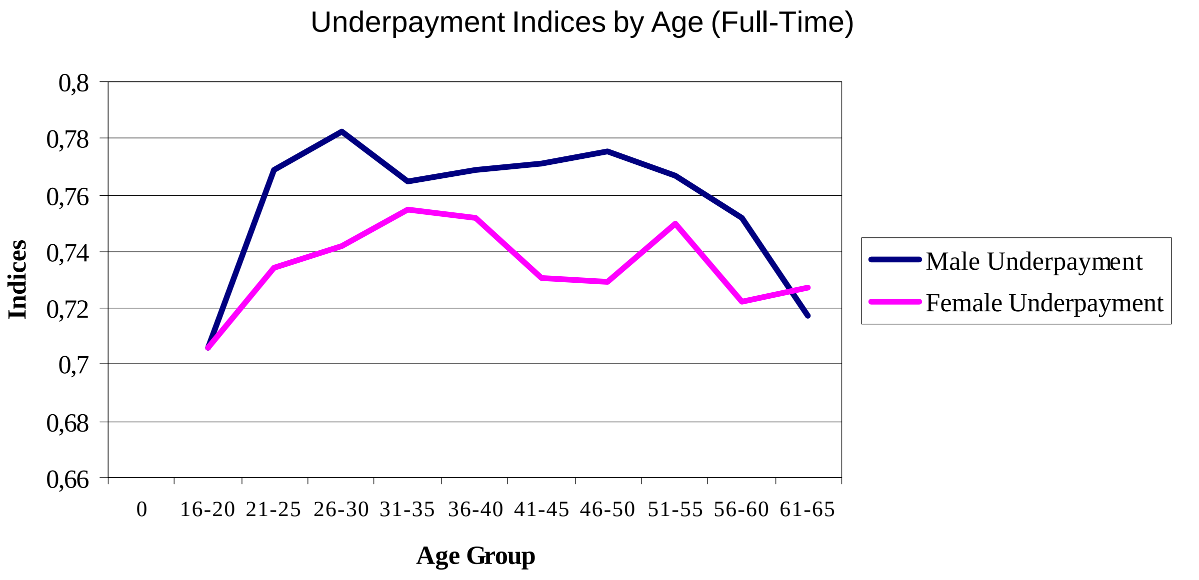
The setting of reservation wages, however, suggests that male workers will also be underpaid. This implication reflects the assumption that workers must account for the marginal benefits and marginal costs of job search in the presence of labor market frictions. Unlike Robinson and Wunnava (1989), we therefore also estimate an earnings frontier for male workers.[5] Table 1 indicates that German males receive 76 percent of their potential wage. It then appears that gender wage discrimination can be inferred from the differential between the male and female average underpayment. Table 1 shows that this differential is approximately 3 percent. However, it is also problematic to refer to this differential as discrimination. Dynamic monopsony indicates that expected underpayment depends on the level at which the reservation wage is set. Any gender differences in the marginal benefits and marginal costs of job search would imply non-discriminatory underpayment differences. For instance, the period that an accepted job will be kept until break-up[what does break-up mean? until the worker quits?] may differ according to gender. While prime-age males may intend that the next job will last for a prolonged period, females will consider the impact that family commitments have on labor market participation and, therefore, expected duration employed in the next job. Such differences imply the marginal benefits of additional search are, ceteris paribus, higher for prime-age males than for females. Consequently, the reservation wage of prime-age males will be higher than reservation wages of females and should result in larger underpayment to women. Further analysis into the relevance of the job search framework is required to evaluate whether gender underpayment differences reflect discrimination.

A comparison of age groups indicates underpayment follows a quadratic relationship in age for both males and females, with underpayment more acute for the young and old. The principal finding illustrated by Figure 1 is the difference in underpayment between prime-age males and females. Between 31 and 50 years of age, male underpayment stabilizes while female underpayment increases. An explanation for this finding is the decline in female participation in the labor market after bearing children/getting married. Lower female participation reduces the returns to search and therefore reservation wages. The lack of search time (i.e., high job search cost) means women can suffer greater underpayment than men because of the social structure of the Western family, where the male is usually still the principal breadwinner.

Watson (2000) also tests the relevance of the dynamic monopsony approach by comparing average underpay-

[5] By assuming that male workers are also underpaid we are not following traditional neo-classical discrimination theory. Traditionally, it is argued that the minority worker (female) will be discriminated against and the majority worker (male) will possibly gain from nepotism on the employers' behalf.

Figure 1



Source: German Socio-Economic Panel (GSOEP), 1991 to 1993 ment rates for workers with different levels of education. This approach relies on the observation that job search behavior is associated with human capital characteristics. In particular, workers with high levels of education are more likely to engage in on-the-job search. One assumes that if higher educated people have lower marginal search costs, they set a reservation wage closer to their highest potential wage. Using data from the United Kingdom, Watson (2000) finds that as number of years of education increases, average underpayment declines. We do not find a similar association in our analysis. Indeed, there is no significant correlation between education and underpayment. It can be argued, however, that this lack of association is consistent with the German labor market due to the nature of the German education system. As described by Nickell and Bell (1996), compared to the United Kingdom, the German education system produces a much more compressed distribution of human capital. Basic education variables will, therefore, capture less of the relevant variation in human capital differences in German data.

The time variance stochastic panel frontier analysis also highlights some of the difficulties one encounters in applying underpayment analysis to discrimination. Using the Battese and Coelli (1992) approach described above, Table 2 indicates that the underpayment levels of males and females are rising. Moreover, it also indicates that the difference between male and female underpayment is increasing. While the time period covered is insufficient to undertake a robust analysis into the trend in underpayment, using it as part of discrimination analysis suggests that gender discrimination increased over the period 1991 and 1993. This result is at odds with the declining pattern of gender discrimination found by Brookes, Hinks and Watson (2000) and may be due to weaknesses with the econometric approach used. For instance, (as described earlier) the stochastic panel frontier approach assumes that underpayment is a deterministic function of time, implying that underpayment is increasing, decreasing, or stable for all workers. However, the increase in underpayment could reflect fundamental changes in the German wage distribution. As shown by Van Dijk et al.'s (1998) comparison of underpayment in the United States and Netherlands, estimated underpayment will reflect the extent of wage compression. This can be important, since between 1991 and 1993 there is evidence of change in the nature of the wage distribution. While reunification has been accompanied by substantial increases in East German wages, there is also evidence that the ratio between average earnings in the lowest and highest earning deciles of the female earnings distribution declined by 5 percent. The deterioration in the wage distribution will increase the estimated underpayment and the differential between males and females.

6. Conclusion

This paper used a stochastic panel wage frontier model to measure male and female underpayment. Unlike in standard decomposition work, the authors found that underpayment is increasing in Germany. While this result could be interpreted as evidence of increasing gender discrimination, it is also consistent with labor markets operating monopsonistically. The paper finds evidence consistent with monopsony theory. For instance, when age cohorts are taken into consideration, prime-age males are underpaid by a stable amount, while equivalent females encounter increased underpayment. Since reunification, the wage distribution of females has been less compressed. This increasing earnings inequality may explain the apparent increase in relative female underpayment.

Table 2

Time-Varying Efficiency, Full-Time Sample Only (German Data)

Variable	Half-Normal Male	Half-Normal Female	Truncated-Normal Male	Truncated-Normal Female
Constant	6.767 (0.073)	6.607 (0.010)	6.782 (0.071)	6.591 (0.096)
1992 dummy variable	0.041 (0.009)	0.107 (0.011)	0.045 (0.008)	0.109 (0.010)
1993 dummy variable	0.087 (0.013)	0.217 (0.017)	0.093 (0.010)	0.213 (0.016)
Agriculture and Fishing	–0.259 (0.036)	–0.344 (0.046)	–0.269 (0.035)	–0.334 (0.044)
Mining and Quarrying	–0.109 (0.041)	–0.020 (0.075)	–0.117 (0.040)	–0.023 (0.073)
Manufacturing	–0.083 (0.025)	–0.183 (0.026)	–0.084 (0.024)	–0.174 (0.025)
Electricity, Gas and Water	–0.029 (0.037)	0.042 (0.063)	–0.032 (0.036)	0.034 (0.062)
Construction	–0.052 (0.027)	–0.108 (0.048)	–0.049 (0.026)	–0.093 (0.046)
Retail and Sales	–0.145 (0.029)	–0.228 (0.028)	–0.149 (0.028)	–0.221 (0.027)
Transport and Communication	–0.140 (0.029)	–0.087 (0.038)	–0.150 (0.029)	–0.090 (0.037)
Community and Social Services	–0.151 (0.026)	–0.083 (0.026)	–0.156 (0.025)	–0.081 (0.024)
Age	0.043 (0.004)	0.048 (0.005)	0.039 (0.004)	0.045 (0.005)
Age^2	–0.0005 (0.00004)	–0.0005 (0.00007)	–0.0004 (0.00004)	–0.0005 (0.00006)
NEWBIE	0.041 (0.015)	0.019 (0.020)	0.040 (0.015)	0.017 (0.019)
OLDHAND	0.059 (0.014)	0.061 (0.020)	0.055 (0.013)	0.061 (0.019)
Married	0.011 (0.013)	–0.038 (0.015)	0.008 (0.012)	–0.030 (0.015)
Number of Children	0.016 (0.005)	–0.003 (0.008)	0.019 (0.005)	–0.002 (0.008)
Years of Education	0.075 (0.002)	0.067 (0.003)	0.074 (0.002)	0.067 (0.003)
East German resident	–0.565 (0.015)	–0.525 (0.021)	–0.569 (0.015)	–0.510 (0.020)
σ^2	0.330 (0.014)	0.597 (0.033)	0.966 (0.059)	2.047 (0.158)
γ	0.846 (0.008)	0.914 (0.006)	0.946 (0.004)	0.975 (0.003)
μ	0	0	–1.911 (0.154)	–2.826 (0.187)
η	–0.092 (0.016)	–0.236 (0.020)	–0.123 (0.014)	–0.283 (0.020)
LL	–1865.036	–1370.551	–1811.290	–1297.528
Iterations	28	30	39	37
ME T1	**0.732**	**0.733**	**0.782**	**0.776**
ME T2	**0.720**	**0.696**	**0.768**	**0.739**
ME T3	**0.708**	**0.660**	**0.754**	**0.703**
N (CSO)	7541 (3573)	4117 (2073)	7541 (3573)	4117 (2073)

Source: Authors' calculations.

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