Discussion Paper No. 97–03 E

# All Quiet on the Wage Front?

## Gender, Public–Private Sector Issues, and Rigidities in the Polish Wage Structure

Patrick A. Puhani

# All Quiet on the Wage Front?

Gender, Public–Private Sector Issues, and Rigidities
in the Polish Wage Structure

**by**

**Patrick A. Puhani**
***Zentrum für Europäische Wirtschaftsforschung (ZEW)***
***Centre for European Economic Research***
***SELAPO, University of Munich***

**April 1997**

**JEL classification: J31, J60**
**Keywords:** wage inequality, wage structure, labour market rigidities, Poland.

*Patrick A. Puhani*
*ZEW*
*P.O. Box 10 34 43*
*D–68034 Mannheim*
*Germany*
*e–mail: puhani@zew.de*

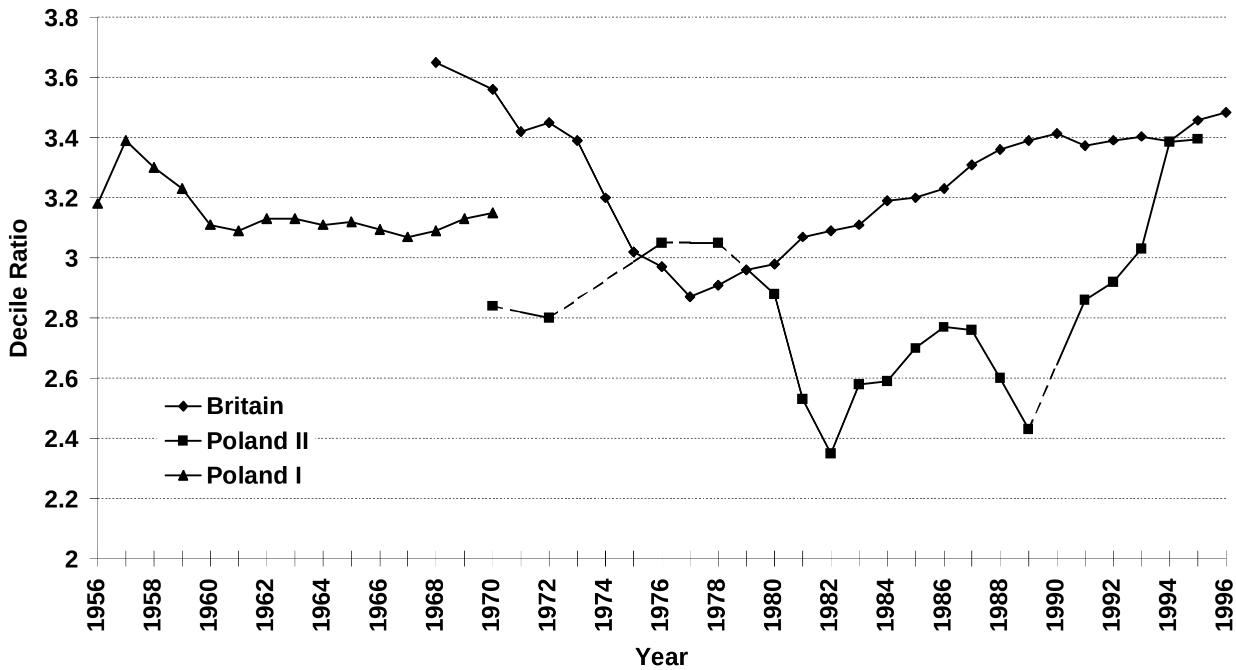
**Acknowledgement**
This research was undertaken with support from the European Commission's Phare ACE Programme 1995 under the project 'Labour Market Policies, Income Support, and Work Incentives in Transition Economies'. Many thanks go to Jan Witkowski, Central Statistical Office (GUS) and Warsaw School of Economics, without whom this research would not have been possible, for his co–operation and advice within this project. I am indebted to Viktor Steiner, ZEW Mannheim, for many helpful comments on this paper. I also very much thank Eugeniusz Kwiatkowski, Czeslaw Domanski, and Leszek Kucharski, University of Lodz, as well as Jan Rutkowski, The World Bank, for helpful suggestions. I am grateful to the participants of the Phare ACE workshop in Ljubljana (21-22 February 1997) for stimulating discussions. Many thanks also go to Anja Triebe for excellent research assistance. All remaining errors are my own.

**Abstract:** We analyse the Polish wage and unemployment structure between 1992 and 1995 on the basis of the Polish Labour Force Survey. It is shown that within this period wage inequality has stabilised. Surprisingly, wage inequality is lower in the private than in the public sector. Our test results show that, contrary to the public sector, there are no gender differences in the valuation of human capital in the private sector. This supports the hypothesis that the transition towards a market economy promotes the equality of the sexes. Although the higher–education wage premium has risen significantly over the observation period, the *ceteris paribus* wage differential between white–collar and blue–collar workers is about zero. Indeed, we find that blue–collar workers as well as workers in the mining, manufacturing, and construction industries have kept or improved their relative wage position despite an increase in unemployment for these groups. There is therefore significant evidence for rigidities in the Polish labour market.

# 1 Introduction

According to widely held opinion, wage inequality has been lower in socialist regimes than in market economies. After all, equality was what socialism was conceived to be all about. Although the statistical evidence suggests that blue–collar workers were comparatively better off in the former command economies than in market economies, inequality levels for all workers were of similar magnitude in the former Eastern and Western blocks (Atkinson and Micklewright, 1992; Redor, 1992). Redor (1992) argues that the former command economies exhibited educational wage premia and a division of labour between managers and workers similar to those of market economies. He even argues that the redistributive effect of benefits in kind in the Visegrád countries[1] has been no greater than in western countries (Redor, 1992, p.182). However, the observed similarities of formerly eastern and western wage distributions also stems from the welfare and minimum wage legislation introduced particularly in western Europe.

**Figure 1: Inequality in Poland and Britain 1956–1996**




*Note:* Here and in the following, the Decile Ratio is the 90$^{th}$ percentile over the 10$^{th}$ percentile.

*Sources:* Polish series I and II up to 1989 and British series up to 1990: Atkinson and Micklewright (1992); Polish series from 1990 up to 1995: Rocznik Statyczny 1996; British series from 1991 up to 1996: New Earnings Survey 1996.

Indeed, as Figure 1 shows, at the end of the 1970s, Britain under Labour had a more equal distribution of earnings (measured by the Decile Ratio) than the People's

[1] The Visegrád countries are the Czech Republic, Hungary, Poland, and the Slovak Republic.

Republic of Poland. However, the Decile Ratios diverged thereafter with Britain undergoing deregulatory pro–market reforms, and Poland seeing the effects of the Gdansk Accord in 1980, when the Solidarity union managed to have a new wage policy implemented. Balcerowicz's 'shock therapy' approach to market reforms, which started at the beginning of 1990, led to a significant increase in inequality, but the change merely closed the gap between Britain and Poland. That is to say, contrary to the fears of some observers, inequality did not soar above the levels observed in western welfare states. At the same time, open unemployment in Poland increased from 0% in 1990 to about 14% in 1992, which is only slightly above western European levels. However, the share of the long–term unemployed in total unemployment continued to increase to 40% in 1995 (Employment Observatory, 1995), which is high by any international comparison.

There are costs and benefits to wage liberalisation. The decline in real wages experienced by most Polish workers during the transition process, as well as the increase in inequality can endanger the transition process through political backlash (Sachs, 1993). On the other hand, wages (prices) are important economic incentives to adjust the economy to changing demands and eliminate the inefficiencies of the socialist period. Economies with appropriate returns to skill will in the long run acquire the right skills to be internationally competitive.

This paper investigates the distributional and structural development of hourly wages in Poland between 1992 and 1995 on the basis of the Polish Labour Force Survey. This period corresponds to the beginning of the second phase of the transition process. Unfortunately, we do not have microeconomic data for the more turbulent first phase of the transition period nor for the period before the transition. However, Rutkowski (1996a) has analysed the Polish wage structure between 1987 and 1992 using individual data from a GUS (Poland's Central Statistical Office) survey of employers and the Household Budget Survey. Based on cross–tabulation evidence, Rutkowski (1996a) concludes that there has been a huge rise in white–collar skills and a considerable increase in the returns to education. He further finds that these premia are much larger in the private than in the public sector. Rutkowski's (1996a) Mincerian (Mincer, 1974) regression results also point to higher returns to schooling and experience in the private sector. In addition, there is slight evidence of a moderate devaluation of work experience gained under the old regime. In sum, these developments have led to an increase in inequality driven mainly by the dispersion of white–collar workers. The distribution of blue–collar workers is found to have become even more equal.

These general tendencies have also been observed for other countries in transition (Rutkowski, 1996b). For the Czech republic Flanagan (1993) and Vecerník (1995) observe falling returns to experience over the transition period, but rising returns to education. For eastern Germany, Krueger and Pischke (1992) and Bird, Schwarze, and Wagner (1994) find decreasing returns to experience, but stable returns to education. Steiner and Puhani (1997) mainly concur with these results but argue that female work experience has not been devalued. Only Orazem and Vodopivec's (1995) results show rising returns to experience with Slovenian data. They attribute this finding to the early retirement schemes which have made experienced labour relatively scarce in Slovenia.

The paper is structured as follows. Section 2 describes the development of the inequality in hourly wages and offers a decomposition of this inequality into the inequality within and between important socio–economic groups. In Section 3 we estimate empirical wage functions, which allow us to decompose inequality changes into the effects of changes in coefficients, observable and unobservable characteristics, respectively. As Poland not only faces inequality in employment, but also great inequality in access to employment, Section 4 tries to identify rigidities in the Polish labour market. This is done by comparing changes in unemployment probabilities with changes in wage premia, holding relevant demographic and socio–economic characteristics constant. Section 5 concludes.

# 2 Wage Inequality 1992–1995

In the following we describe the developments of hourly wage inequality in Poland between November 1992 and November 1995. Our data are from the corresponding waves of the quarterly Polish Labour Force Survey (PLFS), which is carried out by the Central Statistical Office (GUS) of Poland as a representative sample of the Polish population aged 15 and above. During the first four waves (May 1992 to February 1993) the PLFS has been conducted as a pure panel. Since then it has been a rotating panel.[2]

We are interested in gross hourly wages, as we want to analyse the market determinants of wages. However, the PLFS only provides information on net wages. In the face of many people having additional jobs, it is important to notice that we only observe the wage in a person's main job. Part–time employees and self–employed people are excluded from the sample as they do not give information on their wages. To facilitate comparability, we inflate the wages for the years 1992 to 1994 using the Consumer Price Index[3], so that everything is in 1995 old Polish Zlotys (PLZ).

Figure 2 plots selected percentiles of real hourly wages for the years 1992 to 1995. Neither for men nor for women can we observe a remarkable change in inequality. All percentiles saw a fall in the real wage up until 1994 and a subsequent rise thereafter.

---

[2] For more detail on the PLFS, see Szarkowski and Witkowski (1994).

[3] The Polish Consumer Price Index was taken from the Datastream International Data Bank, London.

However, these tendencies were most pronounced for the upper decile. On the other hand, the less well–off amongst Polish workers seem to have a rather stable real wage. The hypothesis of relative stability at the bottom of the wage distribution is also substantiated by Rutkowski's (1996a, p.94) evidence on the period between 1987 and 1992, who argues that the rich have been getting richer, but the poor have not been not getting poorer.

Changes in inequality can also be summarised by standard statistical measures. We report the Decile Ratio (here and in the following the ratio of the ninth over the first decile), the Gini coefficient, and the Mean Logarithmic Deviation (MLD) in Table 1. We report the developments for the public and the private sectors separately.[4] The Decile Ratio has the property that it is not sensitive to errors or real changes at the tails of the distribution, whereas the Gini coefficient and MLD take into account all observations. The Gini coefficient can be given the interpretation that if one randomly draws two people from the population, then the expected wage difference between those two people as a proportion of the average wage is twice the Gini coefficient (Atkinson, 1983). To give an example, the Gini coefficient of 0.238 for men in 1992 of Table 1 says that the expected wage gap between two men chosen at random is 47.6 percent of the average wage.

Table 1 displays the inequality within and between important socio–economic groups (Jenkins, 1995), which were classified into the following categories:

- education: 6 categories as in Table A2 of the appendix.

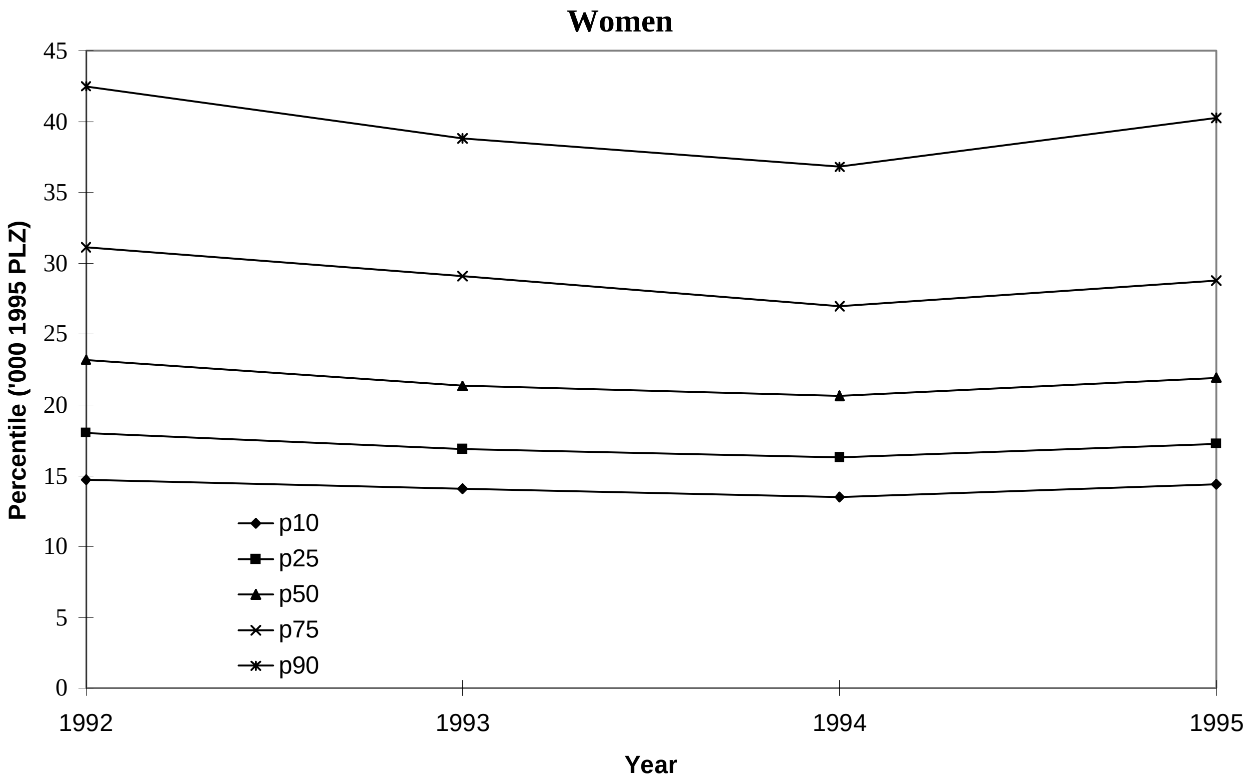
- occupation: 4 categories as in Table A2.

- work experience: 0–9; 10–19; 20–29; 30–39; over 40 years.

- industry: 10 categories as in Table A2 of the appendix.

[4] Co–operatives are also included in the public sector, as they are not profit–maximising.

**Figure 2: Selected Percentiles of Real Hourly Wages 1992–1995**


*Source:* PLFS; own calculations.

**Table 1: Measures of Inequality: Within– and Between–Groups Decomposition**

| | | Decile Ratio | Gini Coeff. | MLD | Educ. | | Occup. | | W. Exp. | | Industry | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | W | B | W | B | W | B | W | B |
| All Men | 1992 | 2.85 | 0.238 | 92 | 72 | 21 | 75 | 17 | 89 | 3 | 84 | 8 |
| | 1993 | 2.79 | 0.238 | 91 | 73 | 18 | 75 | 16 | 89 | 2 | 86 | 5 |
| | 1994 | 2.87 | 0.249 | 101 | 83 | 18 | 84 | 17 | 98 | 2 | 96 | 5 |
| | 1995 | 2.84 | 0.249 | 99 | 77 | 22 | 79 | 20 | 96 | 3 | 95 | 5 |
| Men Public Sector | 1992 | 2.88 | 0.237 | 91 | 70 | 21 | 74 | 17 | 89 | 2 | 80 | 11 |
| | 1993 | 2.82 | 0.236 | 89 | 71 | 18 | 74 | 15 | 87 | 1 | 82 | 6 |
| | 1994 | 2.95 | 0.245 | 97 | 80 | 17 | 81 | 15 | 95 | 2 | 90 | 7 |
| | 1995 | 2.91 | 0.244 | 95 | 77 | 19 | 78 | 17 | 93 | 2 | 89 | 6 |
| Men Private Sector | 1992 | 2.81 | 0.241 | 96 | 77 | 19 | 79 | 17 | 89 | 6 | 92 | 3 |
| | 1993 | 2.81 | 0.243 | 97 | 78 | 19 | 80 | 17 | 93 | 4 | 94 | 3 |
| | 1994 | 2.70 | 0.253 | 108 | 87 | 21 | 87 | 20 | 104 | 3 | 105 | 3 |
| | 1995 | 2.68 | 0.246 | 100 | 73 | 27 | 77 | 24 | 96 | 4 | 98 | 3 |
| All Women | 1992 | 2.89 | 0.243 | 94 | 60 | 35 | 67 | 27 | 92 | 2 | 82 | 13 |
| | 1993 | 2.76 | 0.239 | 91 | 60 | 31 | 67 | 24 | 89 | 2 | 80 | 11 |
| | 1994 | 2.73 | 0.233 | 88 | 58 | 30 | 65 | 23 | 86 | 1 | 80 | 8 |
| | 1995 | 2.80 | 0.236 | 89 | 57 | 32 | 63 | 26 | 87 | 2 | 79 | 10 |
| Women Public Sector | 1992 | 2.88 | 0.243 | 94 | 57 | 37 | 66 | 28 | 92 | 2 | 80 | 14 |
| | 1993 | 2.67 | 0.235 | 88 | 55 | 33 | 64 | 24 | 87 | 1 | 76 | 12 |
| | 1994 | 2.67 | 0.227 | 81 | 52 | 29 | 60 | 21 | 81 | 1 | 74 | 8 |
| | 1995 | 2.73 | 0.232 | 85 | 53 | 33 | 62 | 23 | 85 | 1 | 76 | 9 |
| Women Private Sector | 1992 | 2.77 | 0.240 | 93 | 74 | 19 | 72 | 21 | 89 | 4 | 86 | 7 |
| | 1993 | 2.86 | 0.253 | 104 | 82 | 22 | 79 | 25 | 98 | 6 | 94 | 10 |
| | 1994 | 2.58 | 0.247 | 107 | 77 | 30 | 79 | 27 | 103 | 3 | 98 | 9 |
| | 1995 | 2.57 | 0.233 | 90 | 67 | 23 | 65 | 25 | 85 | 4 | 85 | 5 |

*Note*: $MLD = (1/n)\sum_i \log(\mu/y_i) = \underbrace{\sum_k \nu_k MLD_k}_{within} + \underbrace{\sum_k \nu_k \log(1/\lambda_k)}_{between}$, where

| | | | |
|---|---|---|---|
| n | sample size | y | hourly wage |
| i | index for individual | $\nu_k$ | sample share of $k^{th}$ group |
| k | index for group | $\lambda_k$ | $\mu_k / \mu$ |
| $\mu$ | average hourly wage. | | |

In the table, MLD is multiplied by 1,000. Due to rounding errors the sum of the components of MLD is not always equal to MLD. ‘W’ stands for the within, ‘B’, for the between MLD.
*Source:* PLFS; own calculations.

As we do not observe work experience directly for all waves in the PLFS, we guesstimate this variable by *age minus age at completion of education*. In case of people with a previous unemployment spell, we subtract one year of work experience. For men, we subtract two further years to account for mandatory military service, but only one year for university graduates, as they traditionally serve for a shorter period in Poland. For women, we subtract one year for each child. Unfortunately, years of schooling also have to be inferred from the educational categories in the survey. We have made the following assumptions on age after having completed education (assumed age in brackets): higher (24), post–secondary (21), secondary general (20), secondary vocational (19), basic vocational (17), primary (15), and less than primary (14).

As to the occupation and industry classifications, it is important to note that these have changed in the PLFS over the observation period. Therefore, we have re–classified occupations and industries into broader categories in order to improve the comparability over time (see Tables A3 and A4 in the appendix). In the case of industrial categories, the PLFS fortunately offers both classifications (old and new) in the 1993 wave. Table A5 in the appendix shows that the means of our broader categories are not much affected by whether we use the old or new classification. We can only hope that the same holds for occupations.

Although the MLDs and Gini coefficients in Table 1 show small increases in inequality of male hourly wages between 1992 and 1995, the Decile Ratio exhibits no change. For women, we even observe a small fall in inequality. For men, inequality peaks for all groups in 1994, as also reported in OECD (1997). This relative stability of wage inequality over time is maintained if we look at the whole work force, *i.e.* men and women together. Compared to other countries in transition, wage inequality in Poland is now lower than in Hungary or the Czech republic, but higher than in Slovakia (Pudney, 1993; Rutkowski, 1996b). In general, inequality levels in the Visegrád countries are now within the range of those in western Europe, yet European wage inequality levels are lower than the U.S. one (the U.S. had a Decile Ratio of 4.39 in 1995, see OECD, 1996).

One might expect greater wage inequality in the public sector with some people making a lot of money in modern businesses and others scraping by in simple services. It is interesting, though, to observe that in contrast to Rutkowski's (1996a) results on monthly earnings, this expectation cannot be confirmed by the data, neither for men nor for women. One reason for this finding is that people work longer hours in the private sector (*monthly* wage inequality is higher in the public than in the private sector).

As to the decomposition of MLD into the inequality within and between important socio–economic characteristics, we find there has been little change over the observation period concerning the shares of these two inequalities if one decomposes according to educational categories. However, if one looks at the public and private sectors separately, one observes that for both sexes, the share of the between inequality has increased in the private sector, which is consistent with rising educational premia (*cf.* Rutkowski, 1996a). The same is found if we decompose according to occupational groups in the private sector, which is consistent with rising white–collar premia (*cf.* Rutkowski, 1996a). By contrast, for males, the share of the inequality between work–experience categories has fallen, especially in the private sector. Again this is consistent with falling returns to experience, as found in most studies on eastern Europe (see the references cited in the Introduction). Also, the share of the inequality between industrial sectors has decreased, but here the fall is most pronounced in the public sector rather than the private sector.

To sum up, the inequality measures and their composition gives credence to the view that the Polish transition process is characterised by an increase in educational and occupational premia and a devaluation of work experience. The development seems to be driven mostly by the private sector. Furthermore, the inequality between industrial sectors has decreased, which is what one might expect in well–functioning labour markets after the removal of the socialist regime which had a preference for industrial production over services. In order to make more determinate assertions on the effect of various demographic and socio–economic characteristics on Polish wages, though, an econometric analysis is called for, which we now turn to.

# 3 The Polish Wage Structure 1992–1995

## 3.1 Theoretical and Econometric Issues

Because of the potential effect of industry affiliation as well as regional factors on wages in Poland, we do not consider a pure human capital interpretation (Mincer, 1974) as a suitable explanation of the Polish wage structure in transition. Apart from human capital indicators, we include the industrial sector and the size of the place of residence into our empirical wage equations. In addition, the regional (voivodship) unemployment rate is taken as a proxy for regional aggregate demand factors. As we only have net wages and the Polish income taxation system incorporates joint taxation and child allowances (Bialobreski, 1991), we also have to include relevant household characteristics into the equation.

This is unfortunate as the inclusion of these variables heavily undermines our ability to correct for sample selection (Heckman, 1979). As Leung and Yu (1996) and many others have demonstrated, collinearity problems often arise when effective exclusion restrictions cannot be implemented (see Puhani, 1997, for a short survey). Collinearity can raise the mean square error of the two–step or full–information maximum likelihood estimators way above the one of the OLS estimator. In fact, we cannot find any economically meaningful exclusion restrictions in the data set at hand. Hence, we choose to estimate all wage equations without correcting for selectivity bias.

## 3.2 Estimation Results

Tables 2 and 3 report estimation results of our empirical wage equations for men and women, respectively. Sample means of the reported variables are found in Table A2 of the appendix. We estimate the equations separately for the public and the private sector for the years 1992 and 1995, respectively. Over those three years, the share of private–sector employment in our sample rose from 21 to 34 percent for men and from 16 to 25 percent for women. As the $R^2$ statistics show, the share of the explained variance is lower for men than for women and higher for the public than the private sector. In addition, $R^2$ falls over time. These results are not unexpected. As Table A2 of the appendix shows, women are compared to men far more likely to be employed in industries like health care and education where wages are often set according to measurable characteristics like age, education, work experience, or occupation. Therefore, the higher $R^2$ for women comes as no surprise. The fall of the employment share in these sectors as well as their comparatively low share in the private sector can account for the variation of the $R^2$ over time and between sectors. The estimated variance of hourly wages due to unobserved individual factors shows little variation over sex, time, or sector of employment. However, one can say that unobserved individual factors, like work motivation, are slightly more important in the private than in the public sector. In general, though, the evidence does not point to large changes in the influence of unobserved individual characteristics on the Polish wage structure.

Tables 2 and 3 show the impacts of observed individual characteristics on male and female wages, respectively. In the following discussion, we will focus on the aspects we find most striking. Table 4 gives test results on the equality of coefficients across time, sector of employment, and sex. To give an example, the joint test for the equality of coefficients for the years 1992 and 1995 is carried out by pooling the observations of both years and jointly estimating separate coefficients for each variable and year. We then test for the joint equality of the education, work experience, occupation, industry and place of residence dummies, respectively. The test results will be mentioned in the following discussions of the effects of various socio–economic characteristics on hourly wages in Poland.

**Table 2: Estimated Earnings Functions for Men (see Table A6 for omitted output)**

| Variable | Public Sector | | | | Private Sector | | | |
|---|---|---|---|---|---|---|---|---|
| | 1992 | t | 1995 | t | 1992 | t | 1995 | t |
| *education (basic vocational)* | | | | | | | | |
| higher | 0.351 | 15.40 | 0.295 | 12.21 | 0.297 | 5.61 | 0.441 | 9.34 |
| post–secondary | 0.083 | 2.27 | 0.148 | 3.85 | 0.216 | 2.11 | 0.079 | 0.95 |
| secondary vocational | 0.058 | 4.55 | 0.050 | 3.74 | 0.086 | 3.23 | 0.076 | 4.00 |
| secondary general | 0.065 | 2.25 | 0.057 | 2.03 | –0.041 | –0.48 | 0.136 | 2.63 |
| primary or less | –0.082 | –6.23 | –0.112 | –7.68 | –0.093 | –3.80 | –0.103 | –5.90 |
| work experience | 0.007 | 3.89 | 0.008 | 4.29 | 0.015 | 4.30 | 0.009 | 3.59 |
| work experience^2/100 | –0.014 | –3.25 | –0.019 | –3.94 | –0.035 | –3.71 | –0.022 | –3.45 |
| *occupation (blue collar)* | | | | | | | | |
| manager | 0.168 | 8.85 | 0.276 | 10.30 | 0.282 | 5.87 | 0.265 | 5.79 |
| professional | 0.071 | 4.05 | 0.136 | 7.89 | 0.172 | 3.22 | 0.109 | 3.26 |
| white–collar | –0.045 | –2.59 | –0.035 | –2.00 | –0.034 | –0.79 | –0.002 | –0.07 |
| *industry (mining, manuf.)* | | | | | | | | |
| agriculture, forestry, fishing | –0.248 | –14.79 | –0.258 | –11.20 | –0.165 | –2.00 | –0.156 | –4.41 |
| construction | –0.043 | –2.58 | 0.024 | 1.14 | 0.067 | 3.12 | –0.044 | –0.56 |
| trade, repairs | –0.164 | –6.41 | –0.162 | –7.47 | –0.060* | –1.76 | 0.011 | 0.72 |
| transport, communication | –0.096 | –6.92 | –0.194 | –3.43 | 0.098* | 1.95 | –0.204 | –3.37 |
| financial intermediation | 0.105 | 1.98 | –0.162 | –11.36 | 0.220 | 1.55 | 0.008 | 0.19 |
| health care, social work | –0.140 | –6.08 | –0.169 | –7.38 | 0.120* | 1.79 | 0.075 | 0.54 |
| science, education & arts | –0.074 | –3.40 | –0.048 | –2.77 | –0.125 | –1.31 | –0.044 | –0.29 |
| (public) administration | 0.156 | 8.28 | –0.229 | –7.16 | –0.125 | –0.74 | –0.077 | –1.40 |
| other | -0.029* | –1.82 | –0.232 | –15.56 | 0.050 | 1.46 | –0.099 | –5.00 |
| unemployed before | –0.171 | –7.63 | –0.167 | –11.05 | –0.065 | –2.64 | –0.121 | –8.65 |
| assigned to a disability group | –0.273 | –6.16 | –0.136 | –2.73 | –0.201 | –2.57 | –0.138 | –2.57 |
| *place of residence (rural)* | | | | | | | | |
| 100,000 inhabitants or more | 0.164 | 14.33 | 0.147 | 11.61 | 0.171 | 7.38 | 0.072 | 4.06 |
| 20,000 to 99,999 | 0.079 | 6.32 | 0.112 | 8.94 | 0.078 | 2.95 | 0.052 | 2.93 |
| 19,999 or less | 0.020 | 1.39 | 0.000 | –0.03 | 0.020 | 0.70 | –0.008 | –0.43 |
| voivodship unempl. rate | –0.003 | –2.32 | –0.006 | –4.05 | –0.005 | –2.05 | –0.008 | –4.15 |
| $R^2$ | 0.361 | | 0.357 | | 0.310 | | 0.320 | |
| $\hat{\sigma}^2$ | 0.336 | | 0.342 | | 0.357 | | 0.344 | |
| # observations | 6,143 | | 5,531 | | 1,585 | | 2,873 | |

**Table 3: Estimated Earnings Functions for Women (see Table A7 for omitted output)**

| Variable | Public Sector | | | | Private Sector | | | |
|---|---|---|---|---|---|---|---|---|
| | 1992 | t | 1995 | t | 1992 | t | 1995 | t |
| *education (basic vocational)* | | | | | | | | |
| higher | 0.530 | 25.62 | 0.515 | 25.25 | 0.302 | 4.95 | 0.432 | 9.05 |
| post–secondary | 0.244 | 11.75 | 0.253 | 13.04 | 0.223 | 3.29 | 0.179 | 5.14 |
| secondary vocational | 0.118 | 8.46 | 0.121 | 8.82 | 0.142 | 4.72 | 0.074 | 3.66 |
| secondary general | 0.109 | 6.48 | 0.161 | 9.42 | 0.158 | 3.89 | 0.123 | 4.75 |
| primary or less | –0.096 | –6.53 | –0.084 | –5.52 | –0.086 | –2.40 | –0.035 | –1.32 |
| work experience | 0.014 | 8.02 | 0.009 | 5.05 | 0.010 | 2.10 | 0.010 | 3.40 |
| work experience^2/100 | –0.023 | –5.44 | –0.009 | –2.13 | –0.026 | –2.08 | –0.023 | –2.83 |
| *occupation (blue collar)* | | | | | | | | |
| manager | 0.290 | 12.81 | 0.287 | 11.45 | 0.341 | 4.99 | 0.312 | 4.62 |
| professional | 0.196 | 11.57 | 0.197 | 12.35 | 0.190 | 3.95 | 0.241 | 6.95 |
| white–collar | 0.056 | 4.20 | 0.064 | 4.77 | –0.042 | –1.23 | 0.010 | 0.40 |
| *industry (mining, manuf.)* | | | | | | | | |
| agriculture, forestry, fishing | –0.143 | –5.30 | –0.158 | –4.89 | 0.019 | 0.13 | –0.080 | –1.23 |
| construction | –0.081 | –2.26 | 0.159 | 3.72 | –0.086 | –1.30 | –0.066 | –0.88 |
| trade, repairs | –0.123 | –6.82 | –0.186 | –4.57 | –0.093 | –2.88 | –0.088 | –1.58 |
| transport, communication | 0.000 | –0.02 | –0.113 | –3.58 | 0.041 | 0.55 | –0.153 | –3.84 |
| financial intermediation | 0.158 | 6.03 | –0.014 | –0.79 | 0.239 | 3.12 | 0.043 | 0.63 |
| health care, social work | –0.072 | –5.28 | –0.004 | –0.26 | –0.045 | –0.39 | –0.025 | –0.28 |
| science, education & arts | 0.054 | 3.65 | 0.022 | 1.24 | 0.070 | 0.46 | 0.184 | 2.56 |
| (public) administration | 0.087 | 4.57 | –0.082 | –3.08 | –0.063 | –0.30 | –0.014 | –0.23 |
| other | 0.013 | 0.66 | –0.140 | –11.21 | 0.018 | 0.41 | –0.125 | –5.08 |
| unemployed before | –0.132 | –5.23 | –0.071 | –5.41 | –0.134 | –3.91 | –0.087 | –5.57 |
| assigned to a disability group | –0.089 | –1.32 | –0.024 | –0.48 | –0.195 | –0.92 | –0.177 | –3.02 |
| *place of residence (rural)* | | | | | | | | |
| 100,000 inhabitants or more | 0.080 | 7.00 | 0.034 | 2.93 | 0.125 | 4.16 | 0.086 | 4.16 |
| 20,000 to 99,999 | 0.041 | 3.43 | 0.032 | 2.83 | 0.059* | 1.66 | 0.001 | 0.07 |
| 19,999 or less | 0.032 | 2.26 | 0.014 | 1.16 | –0.008 | –0.21 | –0.014 | –0.57 |
| voivodship unempl. rate | 0.000 | 0.03 | 0.000 | 0.38 | –0.002 | –0.71 | –0.003 | –1.29 |
| $R^2$ | 0.474 | | 0.455 | | 0.341 | | 0.381 | |
| $\hat{\sigma}^2$ | 0.305 | | 0.295 | | 0.345 | | 0.315 | |
| # observations | 5,281 | | 5,338 | | 993 | | 1,825 | |

*Notes to Table 2* and *Table 3:* (1) The dependent variable is the logarithm of the real hourly wage (in 1995 Old PLZ).

(2) The columns beneath the years (*e.g.*, 1992) report the coefficients. The dependent variable is the logarithm of the hourly wage.
(3) Shaded (asterisked) coefficients are significant at the 5 (10) percent level.
*Source:* PLFS; own calculations.

• As can be seen from Tables 2 and 3, there is a high wage premium on completed higher *education* over basic vocational education. This is also found by Jedrzejczak (1994). Because of the semi–logarithmic specification, small coefficients can be interpreted as the approximate *ceteris paribus* wage differential with respect to the base category. The exact interpretation of, say, the higher–education premium for men employed in the private sector in 1995 (*cf.* last column of Table 2) is given by $\exp\left(0.441-1\right)\times 100$, which corresponds to 57.18 percent. For both men and women, the higher–education premium increased significantly (both in the colloquial and in the statistical sense) between 1992 and 1995 in the private sector. However, in contrast to men, the higher–education premium for women in the private sector was still lower than the one in the public sector in 1995.

Table 4 reports test results on the equality of all educational dummies across time, sector of employment, and sex. On the whole, we find no change over time for men, but significant changes for women. By 1995, the differences between the public and the private sector are significant at the 10 percent level for both sexes. Whereas the private sector shows no difference in the educational premia between men and women, the public sector does. If we also look at the tests for equality of the sexes for the other variables in Table 4, we see that the equality of male and female coefficients is a general result for the private sector whereas inequality is a general result for the public sector. This evidence supports the view that the transition process to a market economy improves the equality of the sexes, which is also found by Orazem and Vodopivec (1995) for Slovenia.

**Table 4: Tests of the Equality of Coefficients between Years, Sector of Ownership, and Sex (Tables 2 and 3)**

| Year | Sex | Sector of Empl. | Educ. | Exp | Occ. | Ind. | Pl. of Res. | All |
|---|---|---|---|---|---|---|---|---|
| 1992/95 | Males | Public | – | – | + | + | + | + |
| | | Private | – | – | – | + | + | + |
| | Females | Public | + | (+) | – | + | + | + |
| | | Private | + | – | – | + | – | + |
| 1992 | Males | Public / Private | – | + | – | + | – | + |
| | Females | Public / Private | + | + | (+) | – | – | + |
| 1995 | Males | Public / Private | (+) | – | – | + | + | + |
| | Females | Public / Private | (+) | + | (+) | + | + | + |
| 1992 | Males / Females | Public | + | + | + | + | + | + |
| | Males / Females | Private | – | – | – | – | – | + |
| 1995 | Males / Females | Public | + | + | + | + | + | + |
| | Males / Females | Private | – | – | + | – | – | + |

*Notes:* (1) The tests are for the equality of the coefficients between the categories separated by a slash ‘ / ’ in columns 1 to 3. That is to say, the first, second, and third blocks test the equality of the coefficients between the years, sectors of employment, and sexes, respectively.
(2) ‘All’ means the equality of all coefficients included in the regressions of Table 2 and Table 3 (including the coefficients reported in the appendix) is tested.
(3) A ‘+’ [‘(+)’] sign indicates that the coefficients are significantly different from each other at the 5 [10] percent level.
*Source:* PLFS; own calculations.

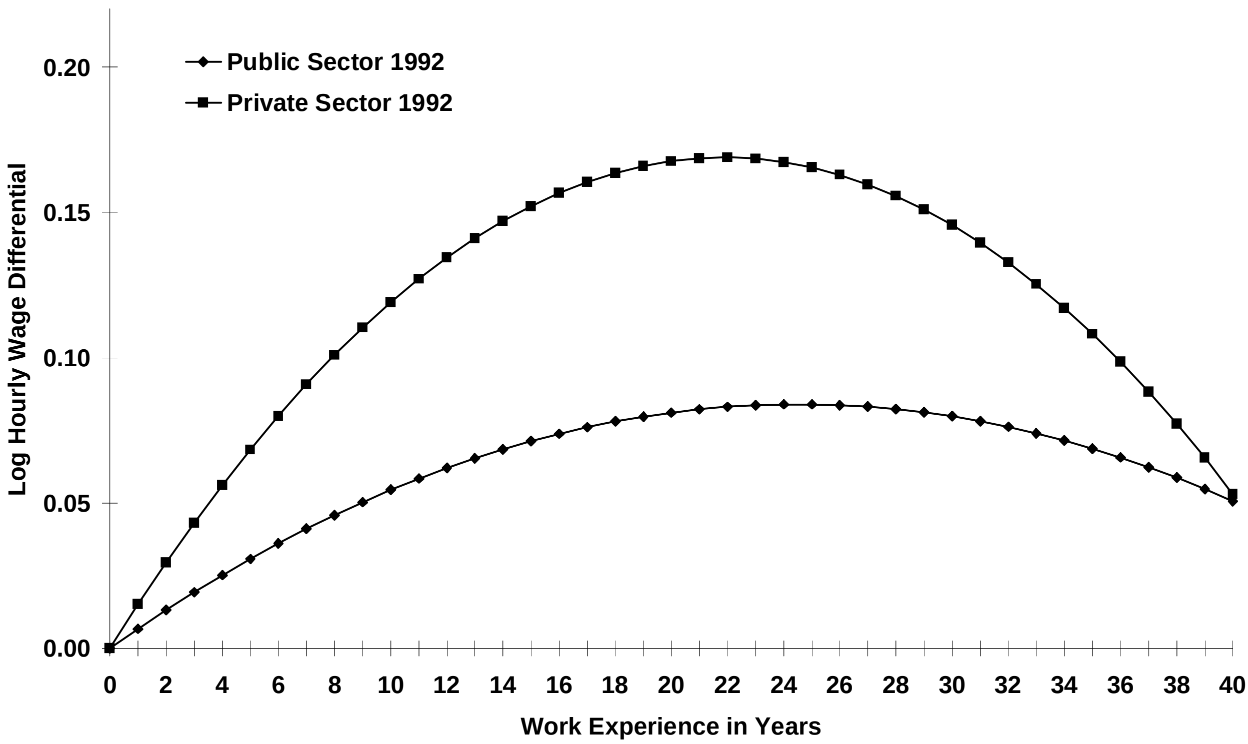
• As the test results of Table 4 show, only for women in the public sector has there been a slight devaluation of *work experience* over time, which is significant only at the 10 percent level. In Figure A in the appendix, we plot the experience–wage profiles estimated by Rutkowski (1996a) for the years 1987 and 1992. It is shown that there has been a slight devaluation of experience in the first phase of transition (the profiles have been estimated jointly for men and women). As can be seen, Rutkowski’s (1996a) profiles for 1992 are much steeper than ours. The reason is probably that we include more variables (*e.g.* occupation, industry) into our wage equation. Such differences have also occurred elsewhere in the empirical literature: for eastern Germany, Bird, Schwarze, and Wagner (1994) estimate flatter profiles than Krueger and Pischke (1992), which they attribute to their more detailed specification of explanatory variables.

As can be seen from the tests in Table 4, experience–wage profiles differ between the public and the private sector (except for males in 1995). As already mentioned above, differences between the sexes can only be observed in the public sector. Taking into account these test results, we plot the relevant experience–wage profiles in Figure 3. The inversely U–shaped form over the life cycle is consistent with human capital theory (see, *e.g.*, Mincer 1974; Franz, 1991; Polachek and Siebert, 1993). Whereas experience–wage profiles for men are steeper in the private sector than in the public sector, we find no change *over time* (see Table 4). A possible explanation for the steeper profiles in the private sector may be a selection effect. The more dynamic private sector is likely to attract the more flexible and able workers away from the public sector by paying them the appropriate rents on their human capital. This explanation is not inconsistent with the fact that, *on average*, the private sector is paying lower wages (see Table A2).

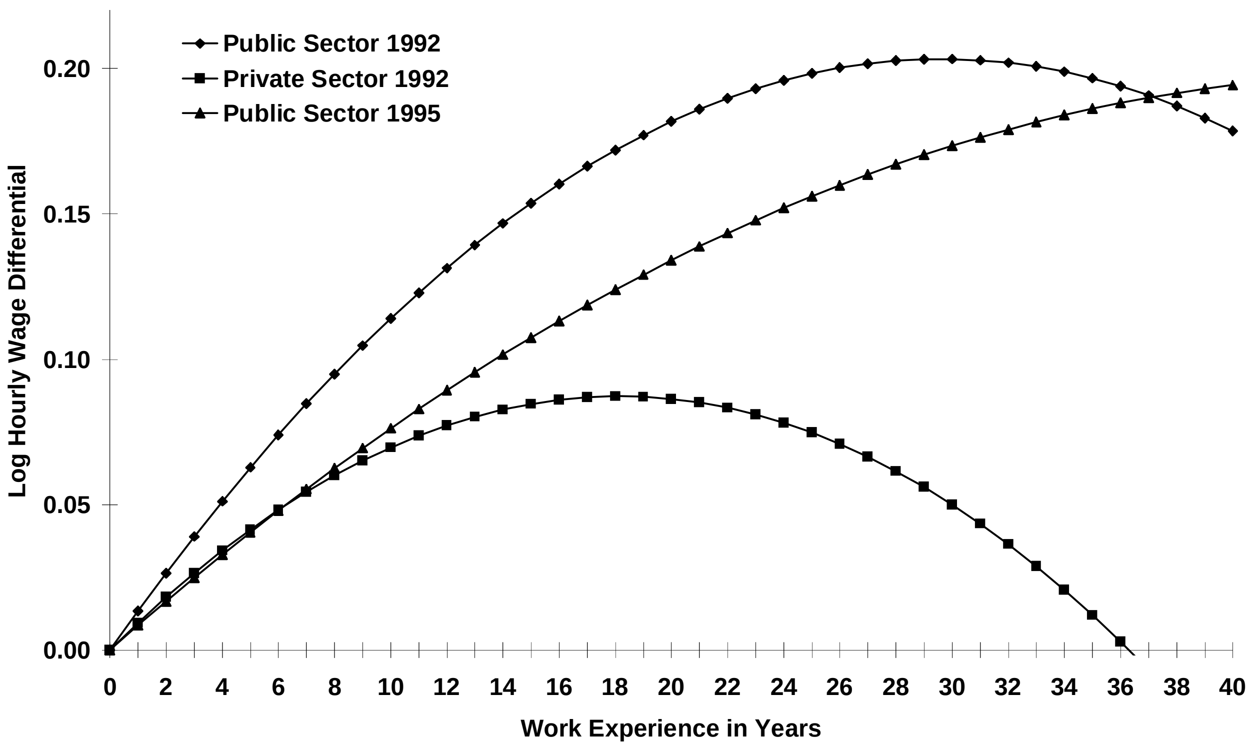
Things are rather different with female experience–wage profiles, though. Here, in contrast to the evidence for men, the profiles are steeper in the public than in the private sector. The reason may be similar to the one mentioned above in this section, namely the high employment share in industries like health care and education with institutionalised wage profiles. The fact that the experience–wage profile in the public sector has flattened somewhat between 1992 and 1995 indicates that the public sector is at least marginally responding to movements in the private market sector. As there is hardly any downturn of the profile with increasing experience, though, one might be suspicious whether the depreciation of human capital over the life cycle is adequately acknowledged in public–sector wage setting.

**Figure 3: Estimated Experience–Wage Profiles**

**Men**




**Women**




*Source:* PLFS; own calculations.

• Male managers and professionals earned higher premia in relation to blue–collar workers in 1995 than in 1992. Otherwise, *occupational* differentials have not changed significantly over the observation period (*cf.* Table 4). We observe no wage premium of white–collar over blue–collar workers, except for women in the public sector. Various factors can account for this. There may be compensating differentials for usually more dangerous and physically more challenging blue–collar work. Union strongholds in typical blue–collar jobs are also possible explanations. Although unions are not that predominant in the newly developing private sector, one should be aware that recently privatised firms are likely to keep many public–sector institutional structures, including union influence. Unfortunately, we are not able to identify employment in recently privatised firms in our data set. Although Rutkowski (1996a) shows that the wage gap between white–collar and blue–collar workers has increased remarkably between 1987 and 1993 by comparing median earnings, we find no such gap in the regression results for the subsequent period we analyse here. However, Rutkowski (1996a) uses a different data set and looks at gross monthly earnings, whereas we analyse net hourly wages. Further, and probably most importantly, we keep the distinction between professionals and white–collar workers as proposed in the PLFS. Rutkowski's (1996a) data set uses a different classification, where professionals and other white–collar workers are grouped into one category.

• In competitive labour markets, persistent non–compensating inter–*industry* wage differentials should be nonexistent. In the transition context, however, we are likely to observe inter–industry wage differentials as disequilibrium phenomena. Growing industries will pay higher wages than shrinking ones thus attracting labour into more productive uses. The estimation results show that wages are amongst the lowest in agriculture, where Poland has an extremely high employment share (27 percent in 1995), which is more than twice as large as in other countries with a comparative advantage in agriculture like New Zealand or Ireland (OECD, 1997). The problems in the Polish housing market (OECD, 1997, Steiner and Kwiatkowski, 1995) probably hamper the movement of workers from rural agricultural areas into cities. Therefore, the negative wage differential for agriculture comes as no surprise. We come back to the results on the inter–industry wage differentials in Section 4.

• The just–mentioned rigidities in the Polish housing market are also likely to explain part of the wage differentials between various *place–of–residence* categories. Hourly wages in big cities are significantly higher than in rural areas, especially for men. The fact that the estimated wage differentials are much smaller for women than for men may well be the result of a selection effect. Women in rural areas with lower average wages have fewer incentives to accept work than women in big cities. Similarly, the zero elasticity of female wages with respect to the voivodship unemployment rate could also be explained by selection effects. For men, this elasticity is negative and significant. In 1995, a 1 percentage point increase in the voivodship unemployment rate led to a reduction of 0.6 and 0.8 percent of the average hourly wage in the public and private sector, respectively.

The issue of wage flexibility will be further discussed in Section 4. Before that, we will in the following subsection use the estimation results of Tables 2 and 3 to gain a deeper insight into the inequality changes discussed in Section 2 above.

## 3.3 Forces Behind Inequality Changes

Following Juhn, Murphy, and Pierce (1993), and Blau and Kahn (1996) we decompose the changes in inequality between 1992 and 1995 into a *characteristics effect*, a *coefficients effect,* and a *residual effect.* The characteristics effect describes the change in inequality resulting from changes in the composition of observed characteristics over time, whereas the coefficients effect describes the change in inequality resulting from changes in the coefficients on observed characteristics. Finally, the residual effect is the change in inequality due to changes in the composition or the effects of unobserved characteristics. These effects can be identified by comparing the distributions of the following variables:

$$y_{92,i} = \hat{\boldsymbol{\beta}}_{92}^{s/p\prime}\, \mathbf{x}_{i,92} + \hat{\sigma}_{92}^{s/p} \hat{\varepsilon}_{92} \qquad [3.1]$$

$$y_{1,i} = \hat{\boldsymbol{\beta}}_{92}^{s/p\prime}\, \mathbf{x}_{i,95} + \hat{\sigma}_{92}^{s/p} \hat{\varepsilon}_{95} \qquad [3.2]$$

$$y_{2,i} = \hat{\boldsymbol{\beta}}_{95}^{s/p\prime}\, \mathbf{x}_{i,95} + \hat{\sigma}_{92}^{s/p} \hat{\varepsilon}_{95} \qquad [3.3]$$

$$y_{95,i} = \hat{\boldsymbol{\beta}}_{95}^{s/p\prime}\, \mathbf{x}_{i,95} + \hat{\sigma}_{95}^{s/p} \hat{\varepsilon}_{95}, \text{ where } E(\varepsilon_t) = 0 \text{ and } \sigma_{\varepsilon_t} \equiv 1 \quad \forall t. \qquad [3.4]$$

The characteristics, coefficients, and residual effects are the differences in the distributions between [3.1] and [3.2], [3.2] and [3.3], and [3.3] and [3.4], respectively. Depending on whether a person is employed in the public or private sector, we use the public– or private–sector coefficient and error variance for the corresponding observation. This way, a move from the public to the private sector is counted as a change in an observed characteristic.

Table 5 shows that although overall changes in inequality as measured by the Decile Ratio were rather modest, the underlying changes in coefficients over time, as well as the mobility between observed characteristics had a larger impact on inequality. In particular, changes in coefficients led to an increase in inequality for both men and women, yet overall inequality changed less, because the change in observed characteristics had a decreasing impact on inequality. This is probably due to the increasing share of private–sector employment, where wage inequality measured by the Decile Ratio is lower than in the public sector (*cf.* Table 1). Steiner and Wagner (1996) report similar results for western Germany in the 1980s: there was hardly a change in overall inequality, but there were underlying changes due to changes in coefficients and observed characteristics, which more or less cancelled each other out.

**Table 5: Decomposition of the Change in Wage Inequality**

| | Men | | Women | |
|---|---|---|---|---|
| | 1992 | 1995 | 1992 | 1995 |
| Decile Ratio (*cf.* Table 1) | 2.85 | 2.84 | 2.89 | 2.80 |
| Δ Decile Ratio | –0.004 | | –0.088 | |
| Characteristics Effect | –0.040 | | –0.114 | |
| Coefficients Effect | 0.049 | | 0.095 | |
| Residual Effect | –0.013 | | –0.068 | |

*Source:* PLFS; own calculations.

The conclusion from this subsection is therefore that the forces working in the Polish labour market have a greater potential effect on inequality than the mere contemplation of inequality changes as carried out in Section 2 suggests. This finding can have important implications for the expectation of future inequality changes. In particular, we have seen that the trend in the revaluation of crucial labour market characteristics (changes in coefficients) works in favour of an increase in inequality. To see whether these changes have been adequate, we compare the changes in wage premia for certain socio–economic characteristics with the changes in the effect of these characteristics on the unemployment probability in the following section. This way, we hope to identify rigidities in the Polish labour market.

# 4 Rigidities in the Polish Labour Market

Labour market rigidities can arise through efficiency wages (Weiss, 1991), the impact of trade unions, incomes policies, or restrictions to mobility due to failures in the housing market.

Poland has two main trade unions, the anti–communist Solidarity (Solidarnosc) and the post–communist OPZZ (Opólnopolskie Porozumienie Zwiazó Zawodowych). While Solidarity is mainly found in large state enterprises in heavy industry, the OPZZ is more concentrated in the service sector. Wage bargaining in Poland is undertaken on a national level by a tripartite commission consisting of the government, trade unions, and employers. However, the agreements of the national commission are only recommendations for the collective bargaining parties at the firm level. Therefore, actual wages are bargained over at the firm level between trade unions and employers. Union power is strengthened, though, by the fact that agreements reached between employers and unions are also binding for non–union members. However, whereas the unions are strongly represented in the private as well as the privatised sector, the unionisation rate in the private sector is very low.[5]

As the OECD (1997) and Steiner and Kwiatkowski (1995) point out, heavy subsidisation of communal rents as well as the poor housing infrastructure are amongst the main obstacles to worker mobility in Poland. In addition, the transport infrastructure cannot compensate these deficiencies.

With the implementation of the stabilisation plan in 1990, Poland introduced a tax–based incomes policy, ‘popiwek’, which helped to keep *average* real wages in check (Hagemejer, 1995). This policy, as efficiency wages and union power, might render the Polish wage structure inefficient by distorting relative wages between socio–economic subgroups. Flanagan (1995) argues for the Czech Republic that the wage rules adopted due to incomes policies there lead to a compression of the wage structure which causes unemployment by preventing *relative* wage adjustment.

In order to identify *ceteris paribus* unemployment probabilities for socio–economic subgroups, we estimate the labour force state (*employed, unemployed, not participating*) by way of multinomial logit models in Tables 6 and 7 for men and women, respectively.[6] To quantify the effects of explanatory variables, we calculate relative odds ratios (RORs), which are reported in the tables. To give an example for the purpose of interpretation, a relative odds ratio of 2.366 in Table 7 in the column ‘Unemployed’ for women aged between 16 and 25 means that

$$\frac{\Pr(\textit{unemployed} \,\&\, 16 \le \textit{age} \le 25) / \Pr(\textit{employed} \,\&\, 16 \le \textit{age} \le 25)}{\Pr(\textit{unemployed} \,\&\, 36 \le \textit{age} \le 45) / \Pr(\textit{employed} \,\&\, 36 \le \textit{age} \le 45)} = 2.366,$$

where $\Pr$ stands for probability. Therefore, if a woman with otherwise the same characteristics as the reference person falls into the age group 16 to 25, her odds for unemployment against employment will be 136.6 percent higher than if she was in the age group 36 to 45.

• For the purpose of comparability with other studies we have chosen to include work experience instead of age into our empirical wage equations. We therefore cannot directly compare the coefficients on the *age* dummies with corresponding coefficients in the estimation of the labour force state. Nevertheless, it is shown that young women have had a very high probability of being unemployed both in 1992 and in

---

[5] I thank Eugeniusz Kwiatkowski, University of Lodz, for helpful discussions on the Polish collective bargaining system.

[6] As people without previous work history do not have information on their occupation nor industry, we group them together with *blue-collar workers*, and the industry *others, unknown*, respectively. An alternative would have been to exclude them from the sample. Both procedures yield the same qualitative results on occupations and industries, though.

1995 (*cf.* Table 7). This is in itself an indication for a rigidity in the labour market. For the market to clear, wages for women with little work experience should fall, however, we observe hardly any change in the experience wage profile for either men or women (see Figure 3 above).

• There is a slight rise in the relative unemployment probability of both men and women with primary or less *education*. Table 2 shows that at least relative wages of men in this educational group have been adjusted downward to account for this development.

• A very clear picture emerges by looking at blue–collar workers. Clearly, they are most likely to be unemployed relative to all other *occupational* categories. This gap has even increased significantly (both in the statistical and the colloquial sense) between 1992 and 1995 for both men and women. However, as can readily be seen from Tables 2 and 3, the small *ceteris paribus* wage gap between female blue–collar and white–collar workers in 1992 has even been closed by 1995. In the case of men, white–collar workers, holding everything else constant, earned even slightly less than blue–collar workers in 1992, but caught up by 1995. A natural explanation for this wage rigidity is that blue–collar workers are more unionised than white–collar workers and can therefore keep relatively high wages, although this hypothesis cannot be tested using our data. As blue–collar workers had a special status in communist ideology, blue–collar workers in Poland did not have to fight for a *change* (increase) in their relative wages in order to reach the current status quo, but only had to *resist a change*, namely falls in relative wages. The social and political difficulty of agreeing on a negative change can explain why this wage rigidity can be so persistent.

• *Industries* for which the unemployment probability has increased over the observation period are construction, administration (for both sexes), transport and communication (especially for men), as well as trade and repairs for women.[7] By 1995, though, both male and female construction workers earn – *ceteris paribus* – amongst the highest wages of all industries, both in the public and the private sector (only women in science, education and arts in the private sector earn more). Here increasing unemployment has not led wages to adjust. Things are different with wages in administration, transport and communication, and to some degree in trade and repairs (for women), as the relative wage positions of these industries have fallen. As far as mining and manufacturing (the base category) is concerned, wages in this

[7] The extremely high estimated relative odds ratio (ROR) for women in the industry *trade and repairs* (31.83), comes as no surprise: in the total female sample, about 9% are unemployed, 51% are employed and 40% are not participating. However, in *trade and repairs*, the corresponding figures are 30%, 13%, and 57%, respectively. It follows that the odds ratio for unemployment over employment is about 0.16 for the total sample and about 2.13 for trade and repairs, respectively. Hence we get a relative odds ratio of 2.13 / 0.16 = 14.44 from cross tabulation. This increases to our estimate of 31.83 under the *ceteris paribus* condition.

category are high relative to other industries, even more so in 1995 than in 1992. As mining and manufacturing is also among the industries with comparatively high unemployment rates, we may conclude that there is some wage rigidity in this sector, which may be explained by union influence, although this cannot be tested here.

• *Place–of–residence* wage differentials have disappeared by 1995, which suggests that there might have been an improvement in worker mobility between 1992 and 1995. This improvement in worker mobility could be the result of a more efficient housing market, or cuts in the generosity of the unemployment benefit system (Puhani, 1996).

To sum up, there is significant evidence for labour market rigidities in Poland. In particular, blue–collar workers as well as workers in the industries mining, manufacturing, and construction have kept or improved their relative wage position between 1992 and 1995, although unemployment has risen for these people. From an efficiency point of view, wages seem too high for these groups.

**Table 6: Estimates of the Labour Force State for Men (see Table A8 for omitted output)**

| Variable | 1992 ROR U | t | 1992 ROR N | t | 1995 ROR U | t | 1995 ROR N | t |
|---|---|---|---|---|---|---|---|---|
| *age between (36 and 45)* | | | | | | | | |
| 16 and 25 | 1.099 | 1.03 | 3.965 | 12.85 | 1.435 | 3.91 | 2.664 | 9.71 |
| 26 and 35 | 0.893 | –1.44 | 0.791 | –2.02 | 1.041 | 0.50 | 0.552 | –5.56 |
| 46 and 55 | 0.972 | –0.27 | 2.221 | 7.92 | 1.155 | 1.49 | 1.739 | 6.36 |
| 56 and 65 | 1.190 | 1.23 | 12.637 | 23.25 | 1.071 | 0.45 | 9.970 | 23.78 |
| *education (basic vocational)* | | | | | | | | |
| higher | 0.688 | –2.37 | 0.895 | –0.75 | 0.591 | –2.83 | 0.677 | –2.86 |
| post–secondary | 0.826 | –0.70 | 1.319 | 1.07 | 1.044 | 0.17 | 0.830 | –0.69 |
| secondary vocational | 0.953 | –0.60 | 1.274 | 2.73 | 0.835 | –2.25 | 1.354 | 3.95 |
| secondary general | 0.939 | –0.38 | 3.003 | 8.20 | 0.980 | –0.12 | 2.567 | 7.87 |
| primary or less | 0.989 | –0.15 | 2.840 | 15.69 | 1.141 | 1.96 | 3.532 | 21.01 |
| *occupation (blue–collar, unkn.)* | | | | | | | | |
| manager | 0.671 | –2.88 | 0.889 | –0.91 | 0.372 | –5.92 | 0.394 | –7.21 |
| professional | 0.677 | –2.81 | 0.704 | –2.70 | 0.369 | –7.00 | 0.652 | –4.14 |
| white–collar | 0.696 | –2.78 | 0.434 | –5.57 | 0.459 | –7.34 | 0.263 | –12.17 |
| *industry (mining, manuf.)* | | | | | | | | |
| agriculture, forestry, fishing | 0.455 | –7.59 | 0.180 | 16.17 | 0.183 | –14.55 | 0.084 | –26.46 |
| construction | 1.824 | 7.04 | 1.238 | 2.00 | 8.179 | 21.87 | 5.457 | 15.74 |
| trade, repairs | 1.366 | 2.55 | 0.837 | –1.16 | 1.192* | 1.87 | 0.923 | –0.75 |
| transport, communication | 0.862 | –1.17 | 1.264* | 1.95 | 12.432 | 14.84 | 9.763 | 13.12 |
| financial intermediation | 0.206 | –1.55 | 0.285 | –1.48 | 0.054 | –7.59 | 0.050 | –9.69 |
| health care, social work | 0.832 | –0.76 | 0.331 | –3.60 | 0.485 | –2.96 | 0.257 | –6.21 |
| science, education & arts | 0.883 | –0.61 | 0.981 | –0.11 | 0.564 | –2.86 | 0.576 | –3.21 |
| (public) administration | 1.027 | 0.15 | 2.141 | 4.79 | 3.501 | 9.32 | 2.691 | 7.21 |
| other, unknown | 3.575 | 15.87 | 10.173 | 31.03 | 1.778 | 7.51 | 2.877 | 15.38 |
| unemployed before | 4.755 | 22.23 | 0.526 | –4.63 | 2.429 | 14.48 | 0.584 | –6.55 |
| assigned to a disability group | 2.122 | 6.58 | 18.463 | 37.73 | 1.917 | 5.85 | 18.457 | 41.91 |
| *place of residence (rural)* | | | | | | | | |
| 100,000 inhabitants or more | 1.109 | 1.38 | 1.244 | 2.98 | 0.962 | –0.52 | 1.259 | 3.48 |
| 20,000 to 99,999 | 1.402 | 4.21 | 1.492 | 4.99 | 1.071 | 0.93 | 1.277 | 3.55 |
| 19,999 or less | 1.289 | 2.90 | 1.535 | 4.88 | 1.096 | 1.11 | 1.228 | 2.57 |
| voivodship unemployment rate | 1.085 | 11.26 | 1.040 | 5.23 | 1.096 | 12.29 | 1.046 | 6.55 |
| log likelihood | –9,596.69 | | | | –11,009.66 | | | |
| Pseudo $R^2$ | 0.393 | | | | 0.417 | | | |
| # observations | 18,982 | | | | 22,064 | | | |

**Table 7: Estimates of the Labour Force State for Women (see Table A9 for omitted output)**

| Variable | 1992 ROR U | t | 1992 ROR N | t | 1995 ROR U | t | 1995 ROR N | t |
|---|---|---|---|---|---|---|---|---|
| *age between (36 and 45)* | | | | | | | | |
| 16 and 25 | 2.366 | 9.04 | 7.672 | 24.63 | 2.460 | 9.33 | 8.884 | 27.64 |
| 26 and 35 | 1.556 | 5.91 | 3.138 | 16.33 | 1.466 | 5.16 | 2.509 | 14.44 |
| 46 and 55 | 1.221* | 1.85 | 2.736 | 12.25 | 0.932 | –0.70 | 1.901 | 9.03 |
| 56 and 65 | 1.067 | 0.37 | 19.406 | 31.66 | 0.503 | –3.36 | 12.794 | 29.89 |
| *education (basic vocational)* | | | | | | | | |
| higher | 0.518 | –3.90 | 0.612 | –3.99 | 0.440 | –4.51 | 0.572 | –5.18 |
| post–secondary | 0.719 | –2.19 | 0.727 | –2.59 | 0.898 | –0.73 | 0.672 | –3.32 |
| secondary vocational | 0.935 | –0.84 | 0.921 | –1.11 | 1.066 | 0.84 | 0.947 | –0.84 |
| secondary general | 0.924 | –0.78 | 1.373 | 3.71 | 1.106 | 1.04 | 1.611 | 6.43 |
| primary or less | 0.705 | –4.29 | 1.854 | 9.39 | 0.934 | –0.88 | 2.409 | 15.09 |
| *occupation (blue–collar, unkn.)* | | | | | | | | |
| manager | 0.269 | –7.28 | 0.463 | –6.08 | 0.097 | –10.30 | 0.090 | –16.31 |
| professional | 0.460 | –6.76 | 0.648 | –4.79 | 0.162 | –17.36 | 0.175 | –23.57 |
| white–collar | 0.584 | –5.99 | 0.541 | –8.28 | 0.259 | –17.13 | 0.216 | –25.27 |
| *industry (mining, manuf.)* | | | | | | | | |
| agriculture, forestry, fishing | 0.176 | 13.29 | 0.118 | –23.13 | 0.094 | –18.79 | 0.037 | –36.57 |
| construction | 1.995 | 3.73 | 1.351* | 1.75 | 6.116 | 9.26 | 4.080 | 7.85 |
| trade, repairs | 1.328 | 2.80 | 1.132 | 1.43 | 31.830 | 28.19 | 37.029 | 32.48 |
| transport, communication | 0.801 | –1.29 | 0.872 | –1.00 | 1.548 | 2.35 | 2.133 | 5.43 |
| financial intermediation | 0.429 | –3.23 | 0.583 | –2.77 | 0.326 | –4.62 | 0.309 | –6.76 |
| health care, social work | 0.449 | –5.51 | 0.664 | –4.09 | 0.763 | –2.15 | 0.798 | –2.48 |
| science, education & arts | 0.457 | –5.65 | 0.709 | –3.54 | 1.282* | 1.67 | 1.760 | 5.33 |
| (public) administration | 0.552 | –3.04 | 0.699 | –2.40 | 1.663 | 2.86 | 1.480 | 2.63 |
| other, unknown | 5.587 | 20.22 | 13.040 | 36.36 | 0.925 | –0.99 | 1.391 | 5.66 |
| unemployed before | 5.935 | 21.79 | 0.744 | –2.63 | 2.074 | 11.38 | 0.584 | –8.42 |
| assigned to a disability group | 2.394 | 6.70 | 10.957 | 30.67 | 2.066 | 5.79 | 11.692 | 34.71 |
| *place of residence (rural)* | | | | | | | | |
| 100,000 inhabitants or more | 1.119 | 1.49 | 1.077 | 1.22 | 0.886 | –1.62 | 1.241 | 3.88 |
| 20,000 to 99,999 | 1.316 | 3.39 | 1.207 | 2.83 | 1.021 | 0.28 | 1.071 | 1.18 |
| 19,999 or less | 1.112 | 1.17 | 0.983 | –0.23 | 1.074 | 0.86 | 0.976 | –0.36 |
| voivodship unemployment rate | 1.080 | 10.26 | 1.034 | 5.36 | 1.060 | 7.86 | 1.016 | 2.75 |
| log likelihood | –11,583.41 | | | | –13,624.44 | | | |
| Pseudo $R^2$ | 0.369 | | | | 0.372 | | | |
| # observations | 19,762 | | | | 23,452 | | | |

*Note to Tables 6 and 7*: ROR U and ROR N stand for the relative odds ratio for unemployed and not participating, respectively.

*Source:* PLFS; own calculations.

## 5 Conclusions

We have analysed the Polish wage and unemployment structure between 1992 and 1995 on the basis of the Polish Labour Force Survey. In particular, we looked at the development of inequality measures, considering men and women, and the public and private sector separately. We also decomposed the inequality measures into inequality within and between important socio–economic groups. The following estimates of empirical wage equations allowed us to track the development of the impact of important socio–economic characteristics on hourly wages in Poland. Additionally, by estimating unemployment probabilities by way of multinomial logit models of the labour force state, we were able to compare changes in *ceteris paribus* unemployment probabilities with changes in wage differentials. This way we managed to identify some rigidities in the Polish labour market. In the following, we summarise the most striking results of our analysis.

(i) While Rutkowski (1996a) has shown that there has been a significant increase in inequality during the first phase of transition (1990–1992), our data on the second phase (1992–1995) show that the level of inequality has stabilised within the observation period. The current level of inequality in Poland is comparable to the ones of other European countries, but much lower than in the U.S. A surprising result is that hourly wage inequality measured by the Decile Ratio is lower in the public sector than in the private sector. As a consequence, fears that further inequality increases are expected when the share of private sector employment rises further are unwarranted by looking at the overall distribution of hourly wages. However, the decomposition of the total inequality into the inequality within and between educational and occupational groups shows that the private sector is leading the trend towards a larger inequality between these groups. This can be explained by a rising demand for highly educated people and a falling demand for less well–educated people.

(ii) Our test results show that, contrary to the public sector, there are no gender differences in the returns to human capital in the private sector. This is also found by Orazem and Vodopivec (1995) for Slovenia. Hence, we can also give credence to the hypothesis that the transition towards a market economy promotes the equality of the sexes.

(iii) Our empirical wage equations confirm that there has been a significant rise in the higher–education wage premium as in other transition economies except Eastern Germany (Steiner and Puhani, 1997). Whereas Rutkowski (1996a) finds that there has been a devaluation of work experience in Poland between 1987 and 1992, our results point to rather stable profiles between 1992 and 1995. An exception is the profile for women in the public sector, which has flattened. The public–sector experience–wage profile is steeper than the private–sector one for women, but the reverse is true for men. We explain the result for women by the comparatively high share of female employment in industries like health care or education, where wages are set according to fixed regulations and therefore less responsive to market forces. The steeper private–sector profiles for men can be accounted for by a selection effect, *i.e.* the private sector is able to attract workers with more valuable work experience.

(iv) It is shown that blue–collar workers have by far the highest *ceteris paribus* unemployment probability of all occupational groups. Their relative unemployment probability has even increased between 1992 and 1995, but wage differentials have not responded to that. Similar rigidities are found for the industries mining, manufacturing, and construction. Unions may play a decisive factor for these rigidities, although we could not test the impacts of unions with our data.

(v) Although wage inequality has stabilised within our observation period, inequality in access to employment has risen significantly through high unemployment and rising long–term unemployment. Therefore, the costs of the transition seem to be distributed very unequally across the population (Puhani, 1996; 1995). According to Dolado and Jimeno (1995), the high unemployment rate of Spain, a country Poland is often compared with (Sachs, 1993), can *inter alia* be traced back to the failure to make labour market institutions more flexible. The Spanish example shows how inflexible collective bargaining institutions can paralyse a labour market for decades after the transition period.

(vi) The fact that we have found considerable evidence for wage rigidities in the Polish labour market has important implications for the potential effectiveness of labour market policies. These policies have become ever more prominent in Poland since 1992 as an alternative to mere passive payments of unemployment benefits or social welfare. The aims of these policies are to promote labour market attachment (public works), to place workers temporarily into jobs (public works and subsidised employment) and to improve and to update skills (training). For those measures to have a long–run payoff in the labour market, a well–functioning labour market is needed in the first place. Obstacles to wage flexibility are harmful to the re–employment chances of the unemployed. If these obstacles are reduced or removed, job chances are likely to improve for outsiders, and active labour programmes will more easily be able to play their part in supporting the disadvantaged.

(vii) Many authors who analyse the wage structures of economies in transition conclude that by now, the wage structures of the transition economies have more or less converged to the ones of market economies. We believe that it would be a great mistake to see this as a sign of a successful transformation. Indeed, western European labour markets are currently facing unemployment problems similar to the ones of central and eastern Europe. Unemployment as well as the need to restructure the economy seem to be general problems of European labour markets. Bertola and Ichino (1995) emphasise that the changes in labour demand since the 1970s have increased wage inequality in the flexible U.S. labour market, but have led to a rise in unemployment in the more rigid European labour markets. Of course, a move towards U.S.–style labour market institutions would increase inequality in Europe significantly and also lead directly to the question of the sustainability of the social safety nets that Europeans, East and West, have been accustomed to since the end of WWII. Hence, the pressure to bring down unemployment by letting wages being set more efficiently is likely to cause social and political battles in Poland and elsewhere. Just looking at the moderate inequality changes in the recent period therefore blurs the true picture of the Polish labour market.

## References


Atkinson, A. B., and Micklewright, J. (1992): Economic Transformation in Eastern Europe and the Distribution of Income, Cambridge University Press, Cambridge.

Atkinson, A. B. (1983): The Economics of Inequality, Clarendon Press, Oxford.

Bertola, G., and Ichino, A. (1995): Wage Inequality and Unemployment: US Versus Europe, CEPR Discussion Paper No. 1186, London.

Bialobrzeski, J. (1991): The Reform of Personal Income Taxation in Poland, in: The Role of Tax Reform in Central and Eastern European Economies, Centre for Co–operation with European Economies in Transition, OECD Publications, pp. 65–70, Paris.

Bird, E. J., Schwarze, J., and Wagner, G. G. (1994): Wage effects of the Move toward Free Markets in East Germany, *Industrial and Labor Relations Review*, Vol. 47, No. 3,pp. 390–400.

Blau, F. D., and Kahn, L. M. (1996): International Differences in Male Wage Inequality: Institutions versus Market Forces, *Journal of Political Economy*, Vol.104, No.4, pp. 791–837.

Commander, S., Coricelli, F., and Staehr, K. (1991): Wages and Employment in the Transition to a Market Economy, Policy, Research, and External Affairs – National Economic Management, Working Papers No.736, Economic Development Institute, The World Bank.

Dolado, J. J., and Jimeno, J. F. (1995): Why is Spanish Unemployment so High? CEPR Discussion Paper No. 1184, London.

Employment Observatory (1995): Central and Eastern Europe, Employment Trends and Developments, No.8, European Commission, Brussels.

Flanagan, R. J. (1995): Wage Structures in the Transition of the Czech Economy, *IMF Staff Papers*, Vol. 42, No. 4, pp. 836–854.

Hagemejer, K. (1995): What Role for the Minimum Wage in the New Polish Labour Market?, in: Standing, G., Vaughan–Whitehead, D.: Minimum Wages in Central and Eastern Europe: from Protection to Destitution, Central European University Press, Budapest, pp. 68–84.

Heckman, J. J. (1979): Sample Selection Bias as a Specification Error, *Econometrica*, Vol.47, No.1, pp. 153–161.

Jedrzejczak, A. (1994): Inequality of Wage Distributions in Poland on the Basis of the ‘Labour Force Survey’, Chair of Statistical Methods, University of Lodz, mimeo.

Jenkins, S. P. (1995): Accounting for Inequality Trends: Decomposition Analyses for the UK, 1971–86, *Economica*, Vol. 62, pp. 29–63.

Juhn, C., Murphy, K. M., and Pierce, B. (1993): Wage Inequality and the Rise in Returns to Skill, *Journal of Political Economy,* Vol. 101, No. 3.

Krueger, A. B., and Pischke, J.–S. (1992): A Comparative Analysis of East and West German Labor Markets Before and After Unification, ZEW Discussion Paper No. 92–11, Mannheim.

Krueger, A. B., and Summers, L. H. (1988): Efficiency Wages and the Inter–Industry Wage Structure, *Econometrica,* Vol. 56, No. 2, pp. 259–293.

Mincer, J. (1974): Schooling, Experience, and Earnings, Studies in Human Behavior and Social Institutions No. 2, N.Y., Columbia University Press.

OECD (1996): Employment Outlook, Paris.

OECD (1997): Economic Surveys, Poland 1996–1997, Paris.

Orazem, P. F., and Vodopivec, M. (1995): Winners and Losers in Transition: Returns to Education, Experience, and Gender in Slovenia, *The World Bank Economic Review,* Vol. 9, No. 2, pp. 201–230, The World Bank, Washington.

Polachek, S. W., and Siebert, W. S. (1993): The Economics of Earnings, Cambridge University Press.

Pudney, S. (1993): Earnings Inequality in Hungary: A Comparative Analysis of Household and Enterprise Survey Data, University of Leicester and DAE Cambridge, mimeo.

Puhani, P. A. (1995): Labour Supply of Married Women in Poland, A Microeconometric Approach Based on the Polish Labour Force Survey, ZEW Discussion Paper No.95–12, Mannheim.

Puhani, P. A. (1996): Poland on the Dole. Unemployment Benefits, Training, and Long–Term Unemployment during Transition, ZEW Discussion Paper No.96–30, Mannheim.

Puhani, P. A. (1997): Foul or Fair? The Heckman Correction for Sample Selection and Its Critique, A Short Survey, ZEW Discussion Paper No.97–07E, Mannheim.

Puhani, P. A., and Steiner, V. (1996): Public Works for Poland? Active Labour Market Policies During Transition, ZEW Discussion Paper No.96–01, Mannheim.

Redor, D. (1992): Wage Inequalities in East and West, Cambridge, University Press.

Rutkowski, J. J. (1996a): High Skills Pay Off: The Changing Wage Structure during Economic Transition in Poland, *Economics of Transition,* Vol. 4 (1), pp. 89–112.

Rutkowski, J. J. (1996b): Changes in the Wage Structure during Economic Transition in Central and Eastern Europe, World Bank Technical Paper No.340, Social Challenges of Transition Series, The World Bank, Washington.

Sachs, J. (1993): Poland's Jump to the Market Economy, MIT Press, Cambridge, (MA).

Steiner, V., and Kwiatkowski, E. (1995): The Polish Labour Market in Transition, ZEW Discussion Paper No.95–03, Mannheim.

Steiner, V., and Puhani, P. A. (1997): Economic Restructuring, the Value of Human Capital, and the Distribution of Hourly Wages in Eastern Germany 1990–1994, in: Dunn, Th. A., and Schwarze, J. (ed.): Proceedings of the 1996 2nd International Conference of German Socio–Economic Panel Study Users, *DIW Vierteljahresheft*, Vol. 66, No.1, Berlin.

Steiner, V., and Wagner, K. (1996): Has Earnings Inequality in Germany Changed in the 1980's ?, ZEW Discussion Paper No.96–32, Mannheim.

Szarkowski, A. and Witkowski, J. (1994): The Polish Labour Force Survey, Statistics in Transition, *Journal of the Polish Statistical Association*, Vol.1, No. 4, pp. 467–483.

Vecernik, J. (1995): Changing Earnings Distribution in the Czech Republic: Survey Evidence from 1988 – 19941, *Economics of Transition*, Vol. 3 (3), pp. 355–371.

Weiss, A. (1991): Efficiency Wages, Models of Unemployment, Layoffs, and Wage Dispersion, Clarendon Press, Oxford.

# Appendix

**Table A1: Selection of the Sample for the Wage Functions**

| | 1992 | 1995 |
|---|---|---|
| total PLFS sample | 45,739 | 54,469 |
| age between 16 and 65 | 38,763 | 45,727 |
| full–time employees | 15,535 | 17,381 |
| monthly wage given | 14,995 | 16,523 |
| hourly wage given | 14,294 | 15,813 |
| hours between 20 and 100 | 14,002 | 15,567 |
| men public sector | 6,143 | 5,531 |
| men private sector | 1,585 | 2,873 |
| women public sector | 5,281 | 5,338 |
| women private sector | 993 | 1,825 |

*Source:* PLFS (Polish Labour Force Survey); own calculations.

**Table A2: Sample Means of Variables Reported in Table 2 and Table 3**

| Variable | Men | | | | Women | | | |
|---|---|---|---|---|---|---|---|---|
| | Public Sector | | Private Sector | | Public Sector | | Private Sector | |
| | 1992 | 1995 | 1992 | 1995 | 1992 | 1995 | 1992 | 1995 |
| hourly wage (‘000 ‘95 old PLZ) | 27.98 | 28.00 | 27.68 | 24.06 | 24.46 | 24.11 | 22.61 | 20.33 |
| *education* | | | | | | | | |
| higher | 0.115 | 0.131 | 0.073 | 0.068 | 0.132 | 0.158 | 0.057 | 0.061 |
| post–secondary | 0.016 | 0.018 | 0.015 | 0.011 | 0.097 | 0.094 | 0.051 | 0.046 |
| secondary vocational | 0.233 | 0.254 | 0.206 | 0.209 | 0.296 | 0.321 | 0.256 | 0.271 |
| secondary general | 0.028 | 0.029 | 0.026 | 0.021 | 0.118 | 0.111 | 0.116 | 0.105 |
| basic vocational | 0.438 | 0.433 | 0.507 | 0.538 | 0.200 | 0.187 | 0.368 | 0.392 |
| primary or less | 0.170 | 0.134 | 0.174 | 0.152 | 0.157 | 0.128 | 0.152 | 0.126 |
| work experience | 17.822 | 18.204 | 13.811 | 14.245 | 18.090 | 18.854 | 13.958 | 14.172 |
| work experience^2/100 | 4.217 | 4.314 | 2.890 | 3.071 | 4.203 | 4.458 | 2.903 | 3.015 |
| *occupation* | | | | | | | | |
| manager | 0.106 | 0.058 | 0.070 | 0.054 | 0.084 | 0.048 | 0.050 | 0.031 |
| professional | 0.156 | 0.224 | 0.067 | 0.087 | 0.377 | 0.436 | 0.137 | 0.182 |
| white–collar | 0.063 | 0.081 | 0.085 | 0.106 | 0.277 | 0.266 | 0.456 | 0.403 |
| blue collar | 0.675 | 0.636 | 0.779 | 0.753 | 0.261 | 0.250 | 0.356 | 0.385 |
| *industry* | | | | | | | | |
| agriculture, forestry, fishing | 0.091 | 0.051 | 0.018 | 0.033 | 0.034 | 0.017 | 0.011 | 0.012 |
| mining, manufacturing | 0.428 | 0.403 | 0.396 | 0.428 | 0.243 | 0.208 | 0.368 | 0.445 |
| construction | 0.082 | 0.052 | 0.295 | 0.010 | 0.017 | 0.012 | 0.047 | 0.005 |
| trade, repairs | 0.035 | 0.054 | 0.146 | 0.238 | 0.093 | 0.012 | 0.433 | 0.030 |
| transport, communication | 0.103 | 0.004 | 0.043 | 0.010 | 0.061 | 0.016 | 0.014 | 0.044 |
| financial intermediation | 0.007 | 0.117 | 0.005 | 0.044 | 0.037 | 0.062 | 0.013 | 0.013 |
| health care, social work | 0.036 | 0.061 | 0.001 | 0.001 | 0.193 | 0.184 | 0.009 | 0.008 |
| science, education & arts | 0.061 | 0.104 | 0.006 | 0.002 | 0.201 | 0.088 | 0.012 | 0.003 |
| (public) administration | 0.073 | 0.028 | 0.006 | 0.022 | 0.069 | 0.024 | 0.006 | 0.025 |
| other | 0.083 | 0.127 | 0.084 | 0.210 | 0.052 | 0.377 | 0.087 | 0.414 |
| assigned to a disability group | 0.010 | 0.011 | 0.011 | 0.013 | 0.006 | 0.008 | 0.006 | 0.011 |
| *place of residence* | | | | | | | | |
| 100,000 inhabitants or more | 0.313 | 0.283 | 0.382 | 0.311 | 0.347 | 0.325 | 0.445 | 0.323 |
| 20,000 to 99,999 | 0.213 | 0.262 | 0.181 | 0.204 | 0.236 | 0.276 | 0.171 | 0.238 |
| 19,999 or less | 0.132 | 0.137 | 0.117 | 0.140 | 0.148 | 0.155 | 0.142 | 0.138 |
| rural | 0.343 | 0.318 | 0.320 | 0.345 | 0.269 | 0.244 | 0.242 | 0.301 |
| voivodship unemployment rate | 13.690 | 13.240 | 13.912 | 13.228 | 13.722 | 13.344 | 13.872 | 13.133 |
| # observations | 6,143 | 5,531 | 1,585 | 2,873 | 5,281 | 5,338 | 993 | 1,825 |

*Source:* PLFS; own calculations.

**Table A3: Classification of Occupations**

| Combined Classification | Old GUS Classification | New GUS Classification |
|---|---|---|
| manager | top manager<br>middle manager<br>lower manager | manager |
| professional | professional | professional<br>technician |
| white–collar | simple white–collar | white collar<br>personal services |
| blue–collar | blue–collar | farmer<br>industrial worker<br>simple blue–collar<br>other simple jobs |

**Table A4: Classification of Economic Sectors**

| Combined Classification | Old GUS Classification | New GUS Classification |
|---|---|---|
| agriculture, forestry, fishing | agriculture, forestry | agriculture, forestry, fishing |
| mining, manufacturing | mining, manufacturing | mining, manufacturing |
| construction | construction | construction |
| trade, repairs | trade | trade, repairs |
| transport, communication | transport, communication | transport, communication |
| financial intermediation | finance, insurance | financial intermediation |
| health care, social work | health care, social aid | health, social work |
| science, education, arts | science, education, arts | education |
| (public) administration | public administration, justice, political and social organisations | (public) administration |
| other | other in the material sphere<br>housing, community services<br>tourism, leisure, sport | electricity, gas, water<br>real estates, renting<br>other services, none, not known |

**Table A5: Comparison of Means of Industry Dummies in 1993**

| *Industry* | Means calculated from | |
|---|---|---|
| | Old Classific. | New Classific. |
| mining, manufacturing | 24.1 | 23.89 |
| construction | 5.68 | 5.5 |
| trade, repairs | 10.17 | 9.57 |
| transport, communication | 4.75 | 4.2 |
| financial intermediation | 1.13 | 1.65 |
| health care, social work | 4.43 | 4.53 |
| science, education & arts | 6.56 | 5.59 |
| (public) administration | 3.48 | 3.25 |
| other | 4.84 | 5.86 |
| unknown | 16.69 | 18.22 |

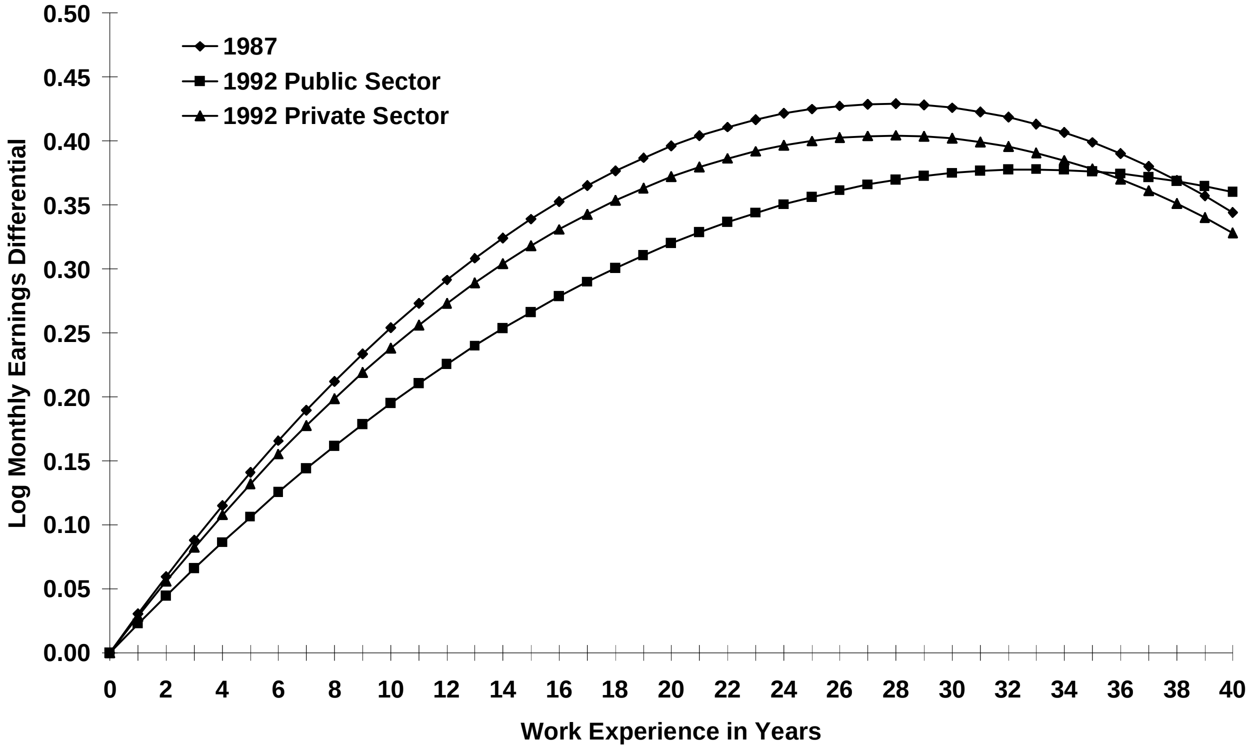
**Table A6: Estimated Coefficients on Household Characteristics in Table 2 (Men)**

| Variable | Public Sector | | | | Private Sector | | | |
|---|---|---|---|---|---|---|---|---|
| | 1992 | t | 1995 | t | 1992 | t | 1995 | t |
| single | 0.017 | 0.85 | –0.125 | –4.75 | –0.162 | –3.63 | –0.039 | –0.95 |
| children | 0.017 | 3.70 | –0.001 | –0.20 | 0.010 | 0.82 | 0.002 | 0.23 |
| no information on children | –0.145 | –6.41 | –0.054 | –2.00 | 0.036 | 0.82 | –0.065 | –1.58 |
| net monthly earnings of wife | 0.005 | 2.06 | 0.000 | –0.17 | 0.001 | 0.24 | 0.007* | 1.68 |
| no wife found | 0.078 | 2.89 | –0.041 | –1.32 | –0.057 | –1.08 | –0.003 | –0.06 |
| constant | 10.121 | 373.63 | 10.242 | 360.04 | 10.079 | 189.36 | 10.136 | 271.38 |

**Table A7: Estimated Coefficients on Household Characteristics in Table 3 (Women)**

| Variable | Public Sector | | | | Private Sector | | | |
|---|---|---|---|---|---|---|---|---|
| | 1992 | t | 1995 | t | 1992 | t | 1995 | t |
| single | 0.025* | 1.84 | 0.040 | 3.02 | 0.017 | 0.39 | 0.099 | 3.11 |
| children | 0.006 | 1.15 | 0.007 | 1.39 | –0.002 | –0.14 | –0.020 | –2.05 |
| no information on children | –0.072 | –3.95 | –0.097 | –5.32 | –0.113 | –2.21 | –0.149 | –4.28 |
| net monthly earnings of husb. | 0.005 | 3.26 | 0.006 | 4.23 | 0.004 | 0.75 | 0.011 | 3.67 |
| no husband found | 0.047 | 2.26 | 0.069 | 3.39 | 0.006 | 0.11 | 0.059 | 1.44 |
| constant | 9.667 | 322.22 | 9.717 | 329.48 | 9.924 | 126.20 | 9.836 | 204.03 |

**Figure A: Estimated Experience–Earnings Profiles from Rutkowski (1996a)**




*Note:* The dependent variable in Rutkowski's (1996a) estimations is the logarithm of gross monthly earnings.
*Source:* Rutkowski (1996a), p.99; regressions (1) and (4).

**Table A8: Estimated Coefficients on Unreported Characteristics in Table 6 (Men)**

| Variable | 1992 ROR U | t | 1992 ROR N | t | 1995 ROR U | t | 1995 ROR N | t |
|---|---|---|---|---|---|---|---|---|
| single | 1.242 | 1.60 | 2.226 | 7.82 | 1.720 | 3.92 | 1.228* | 1.87 |
| children | 0.906 | –0.98 | 0.580 | –6.20 | 0.994 | –0.06 | 0.536 | –7.89 |
| children not found | 2.397 | 7.31 | 0.754 | –2.98 | 1.756 | 4.29 | 2.437 | 7.27 |
| gross earnings of wife | 0.994 | –0.40 | 0.980 | –1.20 | 0.972 | –1.62 | 0.974* | –1.69 |
| wife not found | 0.761* | –1.70 | 0.732 | –2.10 | 0.917 | –0.52 | 0.301 | –7.85 |
| private sector | 1.137* | 1.82 | 1.105 | 1.23 | 0.563 | –8.73 | 0.184 | –21.16 |

**Table A9: Estimated Coefficients on Unreported Characteristics in Table 7 (Women)**

| Variable | 1992 ROR U | t | 1992 ROR N | t | 1995 ROR U | t | 1995 ROR N | t |
|---|---|---|---|---|---|---|---|---|
| single | 0.832* | –1.83 | 0.796 | –3.21 | 1.310 | 2.77 | 0.674 | –5.81 |
| children | 1.519 | 4.59 | 1.078 | 1.11 | 1.516 | 4.59 | 1.215 | 3.07 |
| children not found | 1.325 | 2.35 | 0.710 | –3.87 | 1.751 | 4.41 | 1.624 | 5.30 |
| gross earnings of husband | 1.006 | 0.55 | 1.049 | 5.94 | 1.019 | 1.74 | 1.041 | 5.37 |
| husband not found | 0.999 | –0.01 | 1.445 | 3.57 | 0.967 | –0.26 | 0.812 | –2.11 |
| private sector | 0.833 | –2.21 | 1.073 | 1.06 | 0.320 | –15.00 | 0.188 | –24.84 |