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ABSTRACT

The Gender Pay Gap in the Transition from Communism: Some Empirical Evidence^{*}

This short paper investigates the path through the 1990s of the gender pay gap in a number of former communist countries of Eastern Europe and the Soviet Union. The main findings are that the gender pay gap has not exhibited, in general, an upward tendency over the transitional period to which available data relate. Most of the gender pay gap is ascribed to the 'unexplained' component using conventional decompositions and this may partly be attributable to the proxy measure for labour force experience used in this study. Quantile regression analysis indicates that, in all but one country, the ceteris paribus gender pay gap rises as we move up the wage distribution.

JEL Classification: J16, J31, P23

Keywords: Gender, pay, transition, quantile regression

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1. Introduction

The process of economic transition in the formerly socialist countries is reshaping the working lives of women. There are new demands made upon women and these affect their ability to contribute time and money to household activity. Women are re-assessing their lifetime decisions, such as when and if to start a family and whether to participate in tertiary education. These changes feed back into and help create the structure of the transitional labour markets. This paper studies one outcome of all these changes, the gap in pay between men and women.

It may be useful to provide a very brief outline of women and work under communism. Like all generalisations, this does not capture the diversity of experience both between and within countries. Under socialism women spent more time working, on average, than their Western counterparts. The system of wage and price setting more or less ensured that both adults in a nuclear family would have to work in order to maintain a basic living standard. There was no unemployment in the sense of joblessness. This was partly due to the social stigma that attached to unemployment but has also been attributed to the way hard plan targets and soft budget constraints combined to produce continuous excess demand for labour. Women were accorded a wide range of rights and privileges at work, some of which were formal. In the Soviet system for instance, maternity leave was fully paid and women were legally protected from overly physical and dangerous work. Enterprise kindergartens, schools and health care facilities were fairly commonplace in larger enterprises. Other privileges were informal in nature and perhaps a little dubious (e.g., many women workers tell of being allowed time off to queue in shops).

In terms of occupations and industry branch, women and men were at least as segregated as in the West. Relatively few women held senior positions and this was for two reasons. Firstly, women undertook a very large share of domestic duties thus incurring a double burden and leaving them less time to pursue a career than men. Secondly, the revolution in gender relations in the West, which has brought about a slow but fundamental shift in the household division of labour, did not happen in the communist countries.

Historically, in many centrally planned economies, women's labour market participation was high and gender pay gaps low in comparison to many Western economies. For example, in the former Soviet Union (and in the centrally planned economies of Central and Eastern Europe) female participation

rates were regularly over 85 percent, while in the West rates of this kind have only recently been observed in some Scandinavian countries. Since 1989 participation has been falling for women in many transitional economies (see UNICEF (1999)). The scale of the collapse in participation in the era of transition is very large. For example, in Poland, a country roughly the size of one of the larger European Union states, about one and a half million female jobs disappeared between 1989 and 1994.

The labour market remains the conduit through which reform policies impact on a country's standard of living (see Horton, Kanbur and Mazumdar (1991)) and is the market through which many of the rewards of transition are transmitted. An interesting question is how the transitional process has affected the labour market prospects of women. The welfare of both men and women crucially depends on the rate at which the transitional economies recover and develop. Labour market earnings remain an important component of household income in all the transitional economies and provide a direct link between household welfare and economic activity.

In the transitional economies of Central and Eastern Europe (CEE) over the period 1989 to 1995, Allison and Ringold (1996) identified older workers as more vulnerable to unemployment. However, they failed to uncover evidence of a gender dimension to unemployment but noted the decline in female participation in all transitional economies (see also UNICEF (1999)). For the same set of countries covering approximately the same time period, Rutkowski (1996) highlighted the dramatic fall in real wages experienced in all these economies, the increased incidence of low pay, and the steady rise in wage inequality.

There is now a large empirical literature exploring labour market outcomes in the transitional economies. This includes a growing literature investigating the wage position of women in the transitional economies (see, e.g., Orazem and Vodopivec (1995) for Slovenia, Orazem and Vodopivec (2000) for Estonia and Slovenia, Brainerd (1999), Glinskaya and Mroz (1996), Newell and Reilly (1996), Ogloblin (1999) and Reilly (1999) for Russia, Brainerd (2000) and Pailhé (2000) for a selection of central European countries, and Hunt (1997) for a sample of East German workers). The purpose of this paper is to add to this literature and examine patterns and trends in the gender pay gap for a selection of transitional economies with a view to assessing whether, in regard to pay, women are gaining, losing or standing still since the collapse of communism.

The following section presents the methodology to be used and outlines some of its limitations. Section two discusses the empirical results based on the methodology adopted. Section three outlines

a quantile regression approach to computing the *ceteris paribus* gender pay gap. Section four offers some conclusions.

1. Methodologies for Computing the Gender Pay Gap

If we denote W_m and W_f as wages, a common measure used to summarise the female position in the labour market is the ratio of average female pay to average male pay. This is usually expressed as:

$$\frac{\overline{W_f}}{\overline{W_m}} \quad [1]$$

where the bars denote the average.

This gives the fraction of the average male pay earned by women and is known as the gender pay ratio. Viewing the male wage as a ‘mark-up’ on the female wage is generally the dominant approach adopted by economists. The use of natural logarithms allows us to compute an average mark-up (\overline{D}) as follows:

$$\overline{D} = \overline{\ln(W_m)} - \overline{\ln(W_f)} \quad [2]$$

The fundamental problem with either of the expressions [1] or [2] is that, although they may provide an estimate of the gender pay gap, they provide no insight into that part of the gender pay gap attributable to differences in productivity characteristics between the two groups.

In order to inform more usefully on gender wage effects, it is important to control for differences in productivity variables that may exist between gender groups. This requirement demands use of econometric analysis that allows the researcher to control (or hold constant) other factors, whilst exploring the effect of the characteristic of interest (i.e., gender) on the variable of interest (i.e., the wage). The other factors held constant are usually productivity characteristics and thus wage differences mediated through the characteristic of interest are reflecting wage differences for broadly comparable workers. It has become conventional for economists to specify a relationship between the natural logarithm of earnings and the set of characteristics.

If we define \mathbf{w} as the natural logarithm of W and specify a wage equation as:

$$\mathbf{w} = \mathbf{X}'\beta + \delta G + \mathbf{e} \quad [3]$$

where \mathbf{X} is a vector of productivity characteristics (e.g., education and labour force experience), and \mathbf{G} is a qualitative variable for gender adopting a value of one if the worker is male and zero if female. β is a column vector of coefficients representing the effect of the productivity variables on the natural logarithmic wage (\mathbf{w}) and \mathbf{e} is a random error term. The estimation procedure required to provide estimates for the coefficients β and δ is usually Ordinary Least Squares (OLS).

The equation described in [3] can be used to obtain an average estimate for the adjusted gender pay gap. If we denote the OLS estimates for β and δ as $\hat{\beta}$ and $\hat{\delta}$ respectively, we can re-write [3] as:

$$\hat{w} = \mathbf{X}'\hat{\beta} + \hat{\delta}\mathbf{G}$$

The adjusted gender pay gap could be written as:

$$D^A = [\hat{w} | \mathbf{X}, \mathbf{G} = 1] - [\hat{w} | \mathbf{X}, \mathbf{G} = 0] = \hat{\delta} \quad [4]$$

Any difference between \bar{D} and $\hat{\delta}$ (or D^A) is therefore attributable to the productivity characteristics contained in the \mathbf{X} vector. The $\hat{\delta}$ simply reflects a parallel shifting upwards of the regression line. The only part of the estimated relationship allowed to change through the inclusion of the gender variable in this analysis is the estimated intercept term. It may well be the case that the effects of education, labour force experience and other characteristics on earnings differ across gender groups and the pooled regression approach described in [3] fails to capture this.

The constraint imposed by using a pooled relationship has prompted extensive use of separate equations for men and women. In this approach, we estimate a male equation of the following form:

$$\mathbf{w}_m = \mathbf{X}_m'\beta_m + \mathbf{e}_m \quad [5]$$

and a female equation of the following form:

$$\mathbf{w}_f = \mathbf{X}_f' \hat{\beta}_f + \mathbf{e}_f \quad [6]$$

In this case the subscripts m and f denote male and female respectively. It is assumed that the sample of male data points are randomly chosen from the population of males and the sample of female data points are randomly chosen from the population of females.

An important property of the OLS procedure is that the regression plane passes through the means of the data. We could re-write, therefore, the above expressions as:

$$\bar{\mathbf{w}}_m = \bar{\mathbf{X}}_m' \hat{\beta}_m \quad [5']$$

$$\bar{\mathbf{w}}_f = \bar{\mathbf{X}}_f' \hat{\beta}_f \quad [6']$$

If we recall from above, we could re-write [2] as:

$$\begin{aligned} \overline{\ln(\mathbf{W}_m)} - \overline{\ln(\mathbf{W}_f)} &= \bar{\mathbf{w}}_m - \bar{\mathbf{w}}_f \\ &= \bar{\mathbf{X}}_m' \hat{\beta}_m - \bar{\mathbf{X}}_f' \hat{\beta}_f \end{aligned}$$

After some manipulation, we can obtain:

$$= [\bar{\mathbf{X}}_m - \bar{\mathbf{X}}_f]' \hat{\beta}_m + \bar{\mathbf{X}}_f' [\hat{\beta}_m - \hat{\beta}_f] \quad [7]$$

This allows the overall average differential in wages between the two gender groups to be decomposed into a part attributable to differences in productivity characteristics (as evaluated at the male returns) and a part attributable to differences in the estimated relationship between men and women (i.e., the gender differences in returns). These two components have been referred to as the ‘explained’ and ‘unexplained’ components, with the ‘unexplained’ component also referred to as the ‘residual’ component (see Blinder (1973) and Oaxaca (1973)). It is the latter part of expression [7], which provides the average estimate of the gender pay gap adjusted for productivity characteristics. We could express the last part of expression [7] as:

$$\Delta_{UF} = \bar{\mathbf{X}}_f' \Delta \hat{\beta} \quad [8]$$

where $\Delta \hat{\beta} = \hat{\beta}_m - \hat{\beta}_f$

This approach is commonly referred to as the ‘index number’ approach given its similarity to the calculation of index numbers. The use of an ‘index number’ approach is subject to the standard ‘index number’ problem. It is clear that expression [9] could be recomputed using the ‘basket’ of average male characteristics. Under such circumstances, we could re-express [9] as:

$$\overline{\ln(\mathbf{W}_m)} - \overline{\ln(\mathbf{W}_f)} = [\bar{\mathbf{X}} - \bar{\mathbf{X}}_f]' \hat{\beta}_f + \bar{\mathbf{X}}' [\hat{\beta}_m - \hat{\beta}_f]$$

The last part of this expression could be expressed as:

$$\Delta_{UM} = \bar{\mathbf{X}}_m' \Delta \hat{\beta} \quad [9]$$

and provides another estimate for the average adjusted gender pay gap based on average male characteristics. Given the ‘index number problem’, this may be different from expression [8]. The wage structure assumed in the absence of any wage disadvantage, however, need not be the male wage structure. The alternatives suggested by Cotton (1988), Neumark (1988), or Oaxaca and Ransom (1994), for instance, are not pursued here.

Gender pay gap estimates based on $\hat{\delta}$, $\bar{\mathbf{X}}_f' [\hat{\beta}_m - \hat{\beta}_f]$ and $\bar{\mathbf{X}}_m' [\hat{\beta}_m - \hat{\beta}_f]$ are computed using a selection of data sets for transitional economies. The standard error for $\hat{\delta}$ is computed using the OLS formula and expressions for the standard errors of the ‘index number’ quantities are computed in a conventional manner (see Reilly (1987)).

There are a number of potential problems that attach to use of this particular methodology in identifying an accurate empirical measure of female wage disadvantage. These include, *inter alia*, the

measurement of labour force experience (i.e., actual or potential), the legitimacy of including controls that may reflect the outcome of discriminatory processes (e.g., occupations and industries), and the problem of selection bias.

A key explanatory variable in the analysis of the gender pay gap is labour force experience. Since women generally bear the labour market costs of family formation, their level of labour force experience is likely to be below that of men, and this ultimately impacts on the female wage. The most desirable measure to use in such circumstances is actual labour force experience, which accounts for labour force interruptions. Given data constraints this is generally unavailable and many of our data sets contain insufficient household level information to impute labour force experience. Potential labour force experience can be used but it has been shown that its use assigns a greater portion of the actual wage difference to the unexplained or 'residual' component. For the purposes of our analysis we use age and its quadratic as a proxy measure for labour force experience but acknowledge its limitations in this regard.

The methodology outlined relies on the specification of a well defined human capital model augmented by factors designed to capture compensating differentials and monopoly rents associated with an individual's job or sector of attachment. The validity of the methodology is contingent on the estimated equations providing adequate fits to the data and a stable set of parameter estimates that are broadly consistent with economic theory. This should be borne in mind when interpreting the results obtained using this methodology. It remains a matter of debate as to what controls should (and should not) be included in the specifications. The general evidence is that the greater the number of control variables, the greater the explained portion and the smaller the unexplained portion (i.e., the average productivity adjusted pay gap). If there is gender discrimination by occupation or by industry, then such controls may reflect the outcome of a discriminatory process and could not qualify as legitimate exogenous variables in the wage equation. We acknowledge again this potential problem. In order to provide some insight into the effects on the gender pay gap of adding controls, we provide estimates for both an austere specification with only a set of human capital measures and one augmented with additional controls including, in most cases, industry, occupation and region.

In the context of the potential problems associated with sample selection bias, standard correction procedures can be adopted (see Heckman (1979)). However, they have been subject to much

criticism given their sensitivity to identification and their general lack of robustness (see Manski (1989)). The absence of adequate instruments in most of our data sets restricts our ability to address this issue in a satisfactory manner. Thus, the econometric analysis reported in this paper does not deal with the issue of participation selectivity bias.

There are some important methodological issues relating to the measurement of women's wages in the context of transitional economies. The pay measure used in our comparisons relates to earnings received in the main job by employees. The emphasis on main job earnings excludes from the analysis the treatment of secondary earnings. The incidence of secondary job holding has increased in many transitional economies but, to our knowledge, its gender dimension has not been the subject of investigation. Our focus on employees in their main job eliminates measurement errors that would certainly attach to the reporting of earnings on jobs in the secondary or informal sector.

The incidence of payment arrears in a number of the transitional economies (most notably Russia, Ukraine, the Central Asian republics and FR Yugoslavia) may have important gender implications. The dislocations created by the transitional process in both the Russian Federation and the Ukraine triggered inter-enterprise payment arrears as enterprises were faced with hardened budget constraints that ultimately lead to the emergence of employee wage arrears. Enterprise managers might have used wage arrears as a discriminatory tool favouring male workers. The focus on wages paid may be misleading but data limitations confine the analysis to wages paid rather than wages due. As noted by Glinskaya and Mroz (1996), the effect of payment arrears on the gender pay gap is ambiguous but the inclusion of occupational controls may help mitigate such effects.

Finally, our focus on wage data clearly ignores the issue of non-wage benefits. These were of some importance to both gender groups in the pre-transition era where enterprises provided an array of benefits. In particular, women were able to avail of child-care, medical, shopping and recreational facilities. The nature of enterprise reform has largely eliminated these benefits in most transitional economies to such an extent that the gender wage gap may understate the true female income disadvantage. This is not an issue that can be explored with the data available to us.

2. The Empirical Estimates of the Gender Pay Gap

Table 1 contains gender pay ratios based on expression [1] for seventeen countries undergoing transitional change. The data are based on monthly earnings and are obtained from a variety of international and national sources (see notes to table 1). For some of these countries we also report ratios from the period of central planning using gender pay ratio estimates from Atkinson and Micklewright (1992), who used data from the late 1980s. It should be noted that monthly ratios tend to be lower than ratios based on an hourly measure as men, on average, work longer hours than women. Despite this consideration, most of the reported ratios are relatively high compared to Western countries. They also suggest a mild contraction in the unadjusted gender pay gap since the introduction of reforms in Hungary, Poland, the Czech Republic and Slovakia. The ratios for Russia, Romania, Slovenia, the FR Yugoslavia and Kyrgyzstan are characterised by relative stability over the transitional period whilst the Bulgarian ratio registers a fall of five percentage points. The average monthly female earnings for the most recent years for Latvia, Lithuania, Estonia, the Ukraine and Uzbekistan range from two-thirds to just over three quarters of male earnings and suggest little movement in this gap since the period of central planning.

The table provides contrasting country experiences in terms of the size and evolution of the gender pay gap. This type of summary analysis of the gender pay ratio, however, may provide a misleading picture of the female labour market position. Evidence of a stable or contracting pay gap may be due to the influence exerted by measured and/or unmeasured selection effects. For instance, a contraction in the gender pay gap may be partly attributable to the withdrawal from the transitional labour markets of poorly qualified women (see Hunt (1997)). Unfortunately, given the constraints inherent in the data sets available to us, we are unable to pursue this important issue.

Eleven transitional countries provide the basis for our individual-level analysis and these are Bulgaria, Czech Republic, Hungary, Poland, Slovakia, the Federal Republic of Yugoslavia, Latvia, Russia, the Ukraine, Kazakhstan, and Uzbekistan. In the case of the Czech Republic, Hungary, Slovakia, and Uzbekistan the surveys only reported monthly earnings without information on monthly hours worked. All other survey data allowed us to compute our preferred hourly wage gap measure. Table 2 reports a selection of gender pay gaps for these countries over a number of years. The notes to this table provide a description of the sources for the survey data sets used.

The first column of table 2 reports the estimates for the gender pay gap (rather than ratios) unadjusted for controls of any kind (i.e., \bar{D} in expression [2]). If we look at the estimates for the 1990s, the unadjusted hourly pay gaps range from just over 10% for Yugoslav employees to about 35% for Bulgarian employees.

The second column of this table reports the gender pay gaps based on the estimate from a pooled equation that includes only human capital controls (i.e., education and the proxies for labour force experience). The fits of the estimated pooled hourly equations, as measured by the adjusted- R^2 , vary from 0.04 for Russia to 0.19 for FR Yugoslavia – both for 1996. The fits for the pooled monthly equations vary from 0.15 in Uzbekistan to 0.42 for the Czech Republic. All the estimated gender effects are statistically significant at a conventional level. For most countries there is little movement in the gender pay gap once allowance is made for the set of human capital controls. However, in the cases of Latvia, Poland, Russia and the Ukraine, a steep rise in the adjusted gender pay gap is observed relative to the unadjusted measure. This suggests, particularly for these countries, that women's education levels are probably slightly higher than men's, so controlling for this reveals a larger gender pay gap.

The third and fourth columns of table 2 provide estimates for the gender pay gap based on expressions [8] and [9] respectively with standard errors reported in parentheses. \bar{Z} and $\hat{\alpha}$ denote the vectors of human capital characteristics and estimated coefficients respectively. The separate estimation of wage equations offers no new insights with the estimates for [8] and [9] dimensionally comparable to the pooled estimate in almost all cases. All the estimated effects are again significant and the estimates do not appear sensitive to the vector of mean characteristics used.

The fifth column of table 2 reports the estimated gender effect for specifications augmented to include various additional controls. One has to be cautious in interpreting the estimates in this column, however, since the augmented controls are not identical across all countries. In most cases they do include controls for occupations, industry, and region. The inclusion of additional wage determining variables enhances the fits in most cases. The magnitude of the gender wage effect declines in most countries once these controls are entered but actually increases in the case of Uzbekistan. In all cases, however, the estimated gender effects remain statistically significant.

The sixth and seventh columns of table 2 provide estimates for the gender pay gap based on expressions [8] and [9] respectively with standard errors reported in parentheses. \bar{X} and $\hat{\beta}$ denote the vectors of characteristics and estimated coefficients respectively for the augmented specifications. The picture remains broadly the same as column five although there is now some evidence of sensitivity in the estimates depending on which vector of characteristics is used. Nevertheless, one generalization offered by the table is that for all transitional countries, most of the average gender pay gap is attributed to treatment effects rather than differences in endowments. One can only speculate as to how much of the differential assigned to the treatment component is due to our inappropriate measurement of female labour force experience. It is evident that there would be significant differences in the endowment of this particular variable by gender, so it is worth bearing this in mind in interpreting this set of results.

Finally, the implementation of the ‘index number’ decomposition assumes the separation of the various country samples by gender is statistically justified. This is a testable assumption. Table three reports F-test values based on a Chow-test where the null hypothesis is the pooled specification with a binary control for gender. The alternative hypothesis is provided by separate wage equations by gender. In a number of cases (Kazakhstan, Russia, Slovakia, Uzbekistan and FR Yugoslavia) the data are consistent with the null hypothesis suggesting that gender differentials in returns to productivity characteristics are unimportant. Moreover, the results for these particular countries suggest that the specification described in [3] provides an adequate framework for the computation of the *ceteris paribus* gender pay gap.

3. Quantile Regression Analysis

The methodologies outlined up to now rely entirely on the decomposition of a mean regression. An exclusive focus on the average may provide misleading insights into the gender pay gap. The quantile regression approach (see, Chamberlain (1994)) provides a framework within which the *ceteris paribus* gender pay gap can be estimated at a particular quantile of the distribution as opposed to simply the mean. The median regression can be defined by minimizing the sums of the absolute errors rather, as in least squares, minimizing the sum of squares. The estimator is known as the Least Absolute Deviations or LAD estimator. If we return to our pooled specification and introduce the i subscripts for $i= 1, \dots, n$, we could re-express [3] as:

$$w_i = \beta X_i + \delta G_i + e_i \quad [10]$$

The median regression coefficients can be obtained by choosing the values that minimize L

$$L = \sum_{i=1}^n |w_i - \beta X_i - \delta G_i| = \sum_{i=1}^n (w_i - \beta X_i - \delta G_i) \text{sgn}(w_i - \beta X_i - \delta G_i) \quad [11]$$

Where $\text{sgn}(a)$ is the sign of a , and is 1 if a is positive, and -1 if a is negative or zero.

The computation of the estimates is a linear programming problem. In contrast to the OLS approach, the quantile regression procedure is less sensitive to outliers and provides a more robust estimator in the face of departures from normality (see Koenker and Bassett (1978)). Quantile regression models may also have better properties than OLS in the presence of heteroscedasticity. It is desirable to explore quantile regressions other than the median. Using this methodology, the log wage equation is estimated conditional on a given specification and then calculated at various percentiles of the residuals (e.g., the 10th, the 25th, the 75th or the 90th) by minimizing the sum of absolute deviations of the residuals from the conditional specification. In the context of the regression model specified, quantile regression estimation would allow the estimation of the δ parameter at the 10th, 25th, 50th, 75th and 90th percentile. The estimates obtained for δ in this manner allow us to establish the magnitude of the *ceteris paribus* gender pay gap at different points of the wage distribution. This could be important if it was felt that the average provided a misleading picture.

Given our findings in the last section, there appears only minimal loss in modelling earnings for our set of transitional countries using the type of pooled specification described in [10]. Table 4 provides estimates for a set of quantile regressions at the median, the 10th, 25th, 75th and the 90th percentiles. The log wage quantile regression at, for instance, the 10th percentile predicts the value of the log wage at the 10th percentile rather than at the mean and the interpretation of the gender dummy is the percent (or log point) difference, *ceteris paribus*, between the male and female 10th percentile wage. For completeness, the table also reports the estimated gender effect at the mean, which was reported earlier in table 2. Estimates are again reported for two specifications: the basic human capital model and the specification augmented by the inclusion of the additional wage determining controls.

For most countries, there is little material difference between the mean and median regression estimates. One exception is Russia for 1996 where a 0.068 log point difference in the gender effect is reported but in comparison to the estimated standard errors for these estimates this difference is not viewed as sizeable. An examination of the estimated gender effect over the different percentiles provides some insight into the *ceteris paribus* gender pay gap in low paid jobs (10th percentile) and in high paid jobs (90th percentile). For convenience, we only focus on the estimates from the augmented specifications. In most countries there is a steady rise in the estimated gender pay gap as we move from low-paid to high-paid jobs. The most dramatic increase in the gap between the 90th and 10th percentiles is reported for Khazakstan where the difference is of the order of 0.33 log points. Other sizeable differences are reported for Uzbekistan (0.186 log points) and for Russia in 1996 (0.142 log points). In marked contrast to all other countries reported, the estimates for the Ukraine suggest a steady decline as we move across the distribution with the *ceteris paribus* gender pay gap declining by about 0.168 log points between the 10th and the 90th percentile.

4. Conclusions

The review of the evidence and the analysis undertaken suggests that, in general, the gap in pay between men and women in the transitional countries remains low by international standards. It is accepted, however, that there are limitations to the conventional methodology adopted here. For instance, the methodology relies on the existence of well-determined and stable Mincerian earnings equations. These type of equations have provided good fits to data sets from established capitalist economies but the poor fits obtained by some of the estimated models reported here suggests that a degree of caution should be exercised in interpreting our results. In addition, the use of age as a proxy for labour force experience is conceded as imperfect and may result in a higher adjusted gender wage effect. This could be taken to suggest that the actual adjusted pay gap estimates are, if anything, lower than those reported in this paper.

On the basis of the evidence, it could be argued that, although there are marked contrasts in the relative wage position of the average woman across the transitional economies, the adjustment process itself appears, heretofore, to have been approximately neutral to the average pay position of women relative to men. This is perhaps most surprising for Russia and other countries of the Former Soviet Union where there have been large increases in wage inequality. It seems that in these countries, contrary to expectation, the relative pay position of women has not deteriorated.

Hunt (1997), analysing data from former East Germany, highlights the importance of selection effects when interpreting movements in the gender pay gap. In the four years subsequent to unification, the wages of East German women rose by ten percentage points relative to men. Hunt estimates that just over 40% of the rise was attributable to a selection effect induced by the withdrawal from employment of poorly qualified women. This raises the question of whether the apparent stability of the average gender pay gap observed for other transitional economies conceals more important underlying movements in the labour market status of women. Although the data available are limited and prevent a deeper analysis of this issue, the message from other countries is the East German experience is atypical. The departure from the German labour market of poorly qualified women is

facilitated by the existence of a much more generous social safety net in Germany than currently in existence in the other transitional countries. Although selection effects may be present in our data, current labour market conditions and the necessity for households to have dual income earners have prevented a widespread departure from paid work of poorly qualified women.

The quantile regression analysis undertaken provided an insight into the *ceteris paribus* gender pay gap as we move across the wage distribution. In all but one country, there was evidence of larger gaps in the higher paid jobs relative to the lower paid jobs. This general result flags an important issue and clearly justifies further research. In particular, it would be useful to estimate separate quantile regression models by gender and provide decompositions at different quantiles of the distribution (see Mueller (1998)).

The movement of the gender pay gap in the economies undergoing transitional change has been the subject of empirical investigation with researchers adopting a methodology introduced by Juhn, Murphy, and Brooks (1991, 1993) and popularised in the gender field by Blau and Khan (1992, 1994 and 1996). Although Suen (1997) notes a possible conceptual problem with the approach, its primary purpose is to isolate the effects of increased wage dispersion of the gender pay gap. Hunt (1997), Krstic (2000), Orasem and Vodopivec (1995, 1999), and Reilly (1999) used this methodology to provide insights into the evolution of the gender pay for East Germany, FR Yugoslavia, Slovenia, Slovenia and Estonia, and Russia respectively.

Given the limited nature of our data, we are not in a position to report this type of analysis in a systematic way for all of our countries. We are able to review this issue for just two of our countries: Russia (see Reilly (1999) for further details) and FR Yugoslavia (see Krstic (2000) for further details). In both cases the changes in the gender pay gap have been negligible. The contribution of changes in observed characteristics is small suggesting little evidence of a selection effect in observable variables. The effect of wage dispersion on the gender pay gap is most pronounced in Russia but is offset by movements in returns to human capital that appear favourable to women. On the basis of these two transitional economies, and over the time period examined, there is little evidence of anything more than minor movements in the observed gender pay gap. Although these results are broadly in line with the findings of Brainerd (2000) for a selection of transitional economies, and Orasem and Vodopivec (1995, 1999) for Slovenia and Estonia, a more systematic approach across

countries is clearly required to inform more fully on this matter. This remains an important part of the agenda for future research.

In framing the analysis purely in terms of gender pay differentials, it is important to note that pay gaps within gender groups are also important. Waldfogel (1998a) identifies a widening family wage gap for young cohorts of women where young women with children in the United States have higher gender pay gaps than women without children. Waldfogel (1998b) also found that job-protected maternity leave offsets the negative wage effects of having children in both the United States and Great Britain. This is clearly an issue worthy of investigation for the transitional economies where the provision of job-protected maternity leave is unlikely to represent a key priority to employers as private sector employment grows.

Finally, it is salutary to note that the gender pay gap appears to have exhibited a degree of stability over a very volatile period in the post-centralization era in most transitional economies. Nevertheless, the gender dimension of employment and unemployment has not been explored in this paper nor have the effects of enterprise re-structuring, the re-shaping of the social safety net, and the reform of family benefits. Together these have clear implications for the welfare of women. It may be the case that time will show that these factors exert more important influences on the welfare of women in transitional economies than the evolution of the gender pay gap.

Table 1: Monthly Gender Pay Ratios in Transitional Economies for Selected Years

Country	Year	Ratio
Czech Republic ^{a)}	1987	0.661
	1992	0.730
	1996	0.813
Hungary ^{b)}	1986	0.743
	1992	0.808
	1996	0.789
Poland ^{c)}	1985	0.737
	1992	0.790
	1996	0.790
Slovakia ^{d)}	1987	0.661
	1992	0.733
	1996	0.782
Bulgaria ^{e)}	1990	0.740
	1995	0.700
	1997	0.691
Romania ^{f)}	1994	0.786
	1996	0.760
Slovenia ^{g)}	1987	0.870
	1991	0.905
	1996	0.869
FR Yugoslavia ^{h)}	1995	0.899
	1996	0.888
	1997	0.884
Estonia ⁱ⁾	1997	0.750
Latvia ^{j)}	1997	0.799
Lithuania ^{k)}	1997	0.650 (m)
		0.710 (n)
Russia ^{l)}	1989	0.709
	1992	0.685
	1996	0.695
Ukraine ^{m)}	1996	0.777
Kazakhstan ⁿ⁾	1996	0.723
Kyrgyzstan ^{o)}	1995	0.733
	1996	0.730
	1997	0.720
Uzbekistan ^{p)}	1995	0.805

Sources:

a) Atkinson and Micklewright (1992) and Social Stratification Surveys. b) MONEE Database Project. c) Atkinson and Micklewright (1992) and Polish Labour Force Surveys. d) Atkinson and Micklewright (1992) and Social Stratification Surveys. e) MONEE Database Project. f) MONEE Database Project. g) Orasem and Vodopivec (1995) and MONEE Database Project. h) FRY Labour Force Surveys. i) MONEE Database Project. j) MONEE Database Project. k) MONEE Database Project; n and m denotes manual and non-manual workers respectively. l) Atkinson and Micklewright (1992) and Russian Longitudinal Monitoring Surveys. m) Ukraine Living Standards Measurement Survey. n) Kazakhstan Livings Standard Measurement Survey. o) MONEE Database Project. p) The European University Institute and Essex University Survey in Uzbekistan.

Notes to Table 1. The MONEE Database Project is maintained at UNICEF's International Child Development Centre. See UNICEF (1999) for a reference. All the other data sources are either available from the World Bank's LSMS web site or were communicated to the authors, mostly from national statistical offices. The authors will respond to all reasonable data enquiries.

Table 2: Gender Pay Gaps for Selected Transitional Economies

Country	Year	Pay Measure	\bar{D}	$\hat{\delta}_B$	$\bar{Z}_f \hat{\Delta \alpha}$	$\bar{Z}_m \hat{\Delta \alpha}$	$\hat{\delta}_A$	$\bar{X}_f \hat{\Delta \beta}$	$\bar{X}_m \hat{\Delta \beta}$	N_f	N_m
Bulgaria	1995	Hourly	0.301	0.318 (0.031)	0.316 (0.030)	0.313 (0.032)	0.248 (0.034)	0.216 (0.035)	0.293 (0.035)	834	944
		Adjusted-R ²		0.099	0.086	0.036	0.138	0.107	0.106		
Czech Republic	1984	Monthly	0.377	0.371 (0.012)	0.374 (0.012)	0.361 (0.012)	0.327 (0.013)	0.327 (0.014)	0.326 (0.014)	956	1155
		Adjusted-R ²		0.419	0.257	0.141	0.467	0.312	0.233		
Czech Republic	1992	Monthly	0.316	0.312 (0.014)	0.308 (0.014)	0.316 (0.014)	0.282 (0.015)	0.254 (0.021)	0.311 (0.017)	1406	1610
		Adjusted-R ²		0.263	0.182	0.135	0.339	0.278	0.222		
Hungary	1992	Monthly	0.226	0.246 (0.018)	0.248 (0.020)	0.249 (0.019)	0.210 (0.020)	0.186 (0.026)	0.251 (0.023)	833	903
		Adjusted-R ²		0.313	0.290	0.248	0.378	0.366	0.316		
Kazakhstan	1996	Hourly	0.197	0.246 (0.057)	0.263 (0.060)	0.245 (0.057)	0.270 (0.052)	0.303 (0.055)	0.247 (0.056)	572	624
		Adjusted-R ²		0.052	0.054	0.033	0.180	0.133	0.205		
Latvia	1996	Hourly	0.190	0.238 (0.027)	0.244 (0.027)	0.222 (0.027)	0.234 (0.028)	0.249 (0.029)	0.212 (0.036)	853	863
		Adjusted-R ²		0.144	0.129	0.125	0.265	0.272	0.233		
Poland	1992	Hourly	0.095	0.190 (0.007)	0.172 (0.009)	0.205 (0.007)	0.182 (0.007)	0.164 (0.009)	0.196 (0.007)	7003	8364
		Adjusted-R ²		0.266	0.343	0.196	0.274	0.205	0.351		
Poland	1996	Hourly	0.131	0.215 (0.006)	0.202 (0.007)	0.228 (0.006)	0.174 (0.006)	0.162 (0.007)	0.209 (0.009)	7202	8611
		Adjusted-R ²		0.314	0.244	0.355	0.399	0.377	0.445		

Table 2 (cont'd)

Russia	1992	Hourly	0.251	0.302 (0.021)	0.315 (0.021)	0.305 (0.021)	0.200 (0.024)	0.226 (0.030)	0.185 (0.027)	2800	2863
Adjusted-R ²				0.071	0.044	0.052	0.163	0.125	0.165		
Russia	1996	Hourly	0.244	0.299 (0.046)	0.297 (0.049)	0.299 (0.048)	0.230 (0.052)	0.289 (0.060)	0.191 (0.059)	1013	908
Adjusted-R ²				0.041	0.022	0.031	0.127	0.118	0.123		
Slovakia	1984	Monthly	0.332	0.333 (0.016)	0.335 (0.017)	0.329 (0.017)	0.309 (0.018)	0.321 (0.022)	0.294 (0.024)	466	557
Adjusted-R ²				0.400	0.207	0.189	0.427	0.273	0.213		
Slovakia	1992	Monthly	0.310	0.322 (0.017)	0.321 (0.017)	0.325 (0.018)	0.310 (0.020)	0.290 (0.026)	0.363 (0.022)	653	1204
Adjusted-R ²				0.245	0.174	0.105	0.304	0.263	0.178		
Ukraine	1996	Hourly	0.186	0.244 (0.052)	0.236 (0.053)	0.246 (0.053)	0.202 (0.056)	0.180 (0.063)	0.160 (0.069)	496	500
Adjusted-R ²				0.083	0.085	0.069	0.137	0.122	0.148		
Uzbekistan	1995	Monthly	0.216	0.216 (0.032)	0.213 (0.032)	0.218 (0.032)	0.261 (0.030)	0.261 (0.030)	0.260 (0.031)	696	886
Adjusted-R ²				0.155	0.175	0.123	0.297	0.305	0.269		
Yugoslavia	1995	Hourly	0.102	0.120 (0.018)	0.119 (0.018)	0.112 (0.018)	0.113 (0.019)	0.107 (0.021)	0.128 (0.022)	1299	1950
Adjusted-R ²				0.177	0.200	0.152	0.264	0.299	0.231		
Yugoslavia	1996	Hourly	0.116	0.139 (0.020)	0.143 (0.020)	0.132 (0.021)	0.140 (0.021)	0.132 (0.024)	0.150 (0.026)	1266	1930
Adjusted-R ²				0.194	0.197	0.182	0.267	0.293	0.242		

Notes to table 2:

- 1) Data Sources:
 - a) Bulgarian Household Budget Survey
 - b) Social Stratification Surveys
 - c) Social Stratification Surveys
 - d) Polish Labour Force Surveys
 - e) Social stratification Surveys
 - f) FRY Labour Force Surveys
 - g) Latvian Household Budget Survey
 - h) Russian Longitudinal Monitoring Surveys
 - i) Ukraine Living Standards Measurement Survey
 - j) Kazakhstan Labour Force Survey
 - k) The European University Institute and Essex University Survey in Uzbekistan
- 2) $\hat{\delta}_B$ is the estimated OLS coefficient from the pooled equation that includes only human capital controls.
- 3) $\hat{\delta}_A$ is the estimated OLS coefficient from the pooled equation that includes human capital and other controls.
- 4) The Z vector includes only human capital controls; the X vector includes human capital and other controls.
- 5) The Adjusted-R²s reported in columns two to seven refer to the basic pooled equation, the basic female equation, the basic male equation, the augmented pool equation, the augmented female equation, and the augmented male equation respectively.
- 6) N_f and N_m refer to the female and male sample sizes respectively.
- 7) White (1980) standard errors are reported in parentheses.
- 8) The variance for the decompositions are computed as $\bar{Z}_j' \hat{V} \bar{Z}_j$ where j=m or f, and \hat{V} is the sum of the male and female variance-covariance matrices corrected for heteroscedasticity.

Table 3: F-tests for Sample Separation by Gender

Country	Year	F(g, n - k _m - k _f)	F-Value
Bulgaria	1995	F(24,1728)	1.91*
Czech Republic	1984	F(27,2055)	3.47**
Czech Republic	1992	F(35,2944)	1.70*
Hungary	1992	F(27,1697)	1.46
Khazakstan	1996	F(33,1128)	1.07
Latvia	1996	F(21,1672)	1.69*
Poland	1992	F(13,15339)	12.46**
Poland	1996	F(25, 15760)	10.73**
Russia	1992	F(24, 5613)	2.50**
Russia	1996	F(24,1871)	1.48
Slovakia	1984	F(23,975)	1.24
Slovakia	1992	F(31,1793)	1.78*
Uzbekistan	1995	F(19,1542)	1.21
FR Yugoslavia	1995	F(32,3183)	0.68
FR Yugoslavia	1996	F(32,3130)	1.19

Notes to Table 3:

a) The null hypothesis is the pooled equation with a binary control for gender.

b) ** and * denote statistical significance at the 1% and 5% level respectively.

c) g denotes the number of parameter restrictions under test, n is the pooled sample size and k_m and k_f are the number of estimated parameters in the male and female equations respectively.

d) The F-tests are based on the augmented specifications only.

Table 4: Quantile Regression Estimates of the Gender Pay Gaps for Selected Transitional Economies

Country	Year	Pay Measure	Controls	Mean	10 th	25 th	50 th	75 th	90 th	N
Bulgaria	1995	Hourly	Basic	0.318 (0.031)	0.158 (0.031)	0.232 (0.023)	0.294 (0.019)	0.336 (0.023)	0.425 (0.032)	1778
			All	0.248 (0.034)	0.144 (0.032)	0.205 (0.024)	0.235 (0.020)	0.250 (0.024)	0.273 (0.038)	
Czech Republic	1984	Monthly	Basic	0.371 (0.012)	0.366 (0.153)	0.369 (0.011)	0.351 (0.009)	0.365 (0.011)	0.402 (0.015)	2111
			All	0.327 (0.013)	0.322 (0.018)	0.325 (0.012)	0.321 (0.010)	0.341 (0.012)	0.331 (0.017)	
Czech Republic	1992	Monthly	Basic	0.312 (0.014)	0.244 (0.016)	0.292 (0.012)	0.330 (0.010)	0.352 (0.012)	0.348 (0.018)	3016
			All	0.282 (0.015)	0.236 (0.018)	0.269 (0.013)	0.299 (0.011)	0.301 (0.013)	0.320 (0.019)	
Hungary	1992	Monthly	Basic	0.246 (0.018)	0.173 (0.022)	0.239 (0.016)	0.279 (0.014)	0.271 (0.017)	0.345 (0.025)	1736
			All	0.210 (0.020)	0.181 (0.024)	0.211 (0.018)	0.212 (0.015)	0.226 (0.018)	0.253 (0.027)	
Khazakstan	1996	Hourly	Basic	0.246 (0.057)	0.059 (0.072)	0.164 (0.049)	0.276 (0.042)	0.517 (0.051)	0.486 (0.069)	1196
			All	0.270 (0.052)	0.178 (0.063)	0.175 (0.046)	0.272 (0.039)	0.477 (0.046)	0.508 (0.065)	
Latvia	1996	Hourly	Basic	0.238 (0.027)	0.160 (0.035)	0.226 (0.026)	0.242 (0.021)	0.279 (0.026)	0.211 (0.037)	1716
			All	0.234 (0.028)	0.195 (0.037)	0.239 (0.026)	0.250 (0.022)	0.254 (0.026)	0.224 (0.037)	
Poland	1992	Hourly	Basic	0.190 (0.007)	0.141 (0.008)	0.173 (0.006)	0.215 (0.005)	0.235 (0.006)	0.244 (0.008)	15367
			All	0.182 (0.007)	0.133 (0.009)	0.170 (0.006)	0.203 (0.005)	0.227 (0.006)	0.235 (0.009)	
Poland	1996	Hourly	Basic	0.215 (0.006)	0.137 (0.008)	0.184 (0.006)	0.231 (0.005)	0.273 (0.006)	0.296 (0.008)	15813
			All	0.174 (0.006)	0.116 (0.008)	0.152 (0.006)	0.180 (0.005)	0.216 (0.006)	0.231 (0.009)	
Russia	1992	Hourly	Basic	0.302 (0.021)	0.179 (0.027)	0.282 (0.019)	0.355 (0.016)	0.377 (0.028)	0.327 (0.028)	5663
			All	0.200 (0.024)	0.099 (0.030)	0.175 (0.021)	0.194 (0.018)	0.224 (0.021)	0.225 (0.031)	
Russia	1996	Hourly	Basic	0.294 (0.046)	0.225 (0.062)	0.274 (0.041)	0.364 (0.034)	0.299 (0.041)	0.265 (0.059)	1921
			All	0.218 (0.052)	0.117 (0.069)	0.196 (0.047)	0.286 (0.038)	0.293 (0.045)	0.259 (0.061)	
Slovakia	1992	Monthly	Basic	0.333 (0.016)	0.331 (0.021)	0.335 (0.015)	0.336 (0.012)	0.330 (0.015)	0.345 (0.021)	1023
			All	0.309 (0.018)	0.299 (0.025)	0.333 (0.017)	0.333 (0.014)	0.315 (0.016)	0.254 (0.023)	
Slovakia	1992	Monthly	Basic	0.322 (0.017)	0.317 (0.020)	0.281 (0.015)	0.317 (0.012)	0.343 (0.015)	0.395 (0.021)	1857
			All	0.310 (0.020)	0.272 (0.023)	0.289 (0.016)	0.307 (0.013)	0.330 (0.016)	0.363 (0.024)	

Table 4 (Cont'd)

Country	Year	Pay Measure	Controls	Mean	10 th	25 th	50 th	75 th	90 th	N
Ukraine	1996	Hourly	Basic	0.244 (0.052)	0.253 (0.070)	0.218 (0.046)	0.241 (0.046)	0.240 (0.046)	0.291 (0.062)	996
			All	0.210 (0.020)	0.286 (0.074)	0.228 (0.050)	0.255 (0.042)	0.153 (0.051)	0.118 (0.071)	
Uzbekistan	1995	Monthly	Basic	0.216 (0.032)	0.073 (0.050)	0.141 (0.029)	0.151 (0.024)	0.271 (0.028)	0.420 (0.040)	1582
			All	0.261 (0.030)	0.175 (0.041)	0.255 (0.028)	0.227 (0.022)	0.295 (0.026)	0.361 (0.036)	
Yugoslavia	1995	Hourly	Basic	0.120 (0.018)	0.097 (0.024)	0.113 (0.016)	0.101 (0.013)	0.143 (0.016)	0.139 (0.022)	3249
			All	0.113 (0.019)	0.106 (0.025)	0.109 (0.016)	0.131 (0.013)	0.115 (0.016)	0.089 (0.022)	
Yugoslavia	1996	Hourly	Basic	0.139 (0.020)	0.159 (0.028)	0.095 (0.018)	0.122 (0.015)	0.155 (0.017)	0.197 (0.023)	3196
			All	0.140 (0.021)	0.169 (0.029)	0.097 (0.019)	0.100 (0.015)	0.126 (0.018)	0.112 (0.025)	

Notes to table 4:

a) See notes to table 2.

b) The Basic specification contains only human controls.

c) The All specification includes additional controls including region, occupations and industry.

d) Mean denotes the mean regression, 10th, 25th, 50th, 75th, 90th denote the relevant percentiles of the quantile regression.

e) N is the sample size.

f) Quantile regression analysis undertaken using TSP (version 4.3) (See Hall (1996)).

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