WAGE DETERMINATION AND GENDER DISCRIMINATION IN A TRANSITION ECONOMY: THE CASE OF ROMANIA

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Stefano Paternostro and David E. Sahn

ABSTRACT

In this paper we analyze wage determination and gender discrimination in Romania using the 1994 Romanian Household Survey. Wages for both males and females in rural and urban areas are estimated by implementing a Heckman selection model. Gender discrimination analysis is performed on offered wages, addressing the methodological shortcomings found in the literature. The estimations highlight the relevance of human capital and regional labor market segmentation in wage determination. Gender discrimination is found in both urban and rural labor markets**.** While the observed bias in urban areas is comparable to what has been found in other Western countries, in rural settings gender discrimination is much greater.

I. INTRODUCTION

Since the end of the 1980s Eastern Europe and the former Soviet Union have been experiencing a fundamental restructuring of their economic system toward a market economy. In this new phase, one of the many challenges faced by policy makers is to formulate adequate labor market policies and design suitable safety nets. In an attempt to provide some insight into the functioning of the Romanian labor market, this paper focuses on the understanding of wage determination in general and the extent of gender discrimination in specific.

Discerning the differences in wages that stem from endowments such as education, experience and demographic characteristics is in fact critical in assessing the likely outcomes of the ongoing process of economic adjustment in the labor markets. Prior to the current reform period, wages as well as the allocation of labor were heavily regulated in Romania. Wage differentials by skill level were small and compensation by enterprises was primarily determined by the government: remuneration was centrally determined for all categories of employees. Output per unit of time was also decided by the state for those workers whose performance was measurable. Central authorities determined the internal wage structure of the individual enterprise thus depriving the local management of any flexibility regarding wages to suitably attract labor if needed. Furthermore, local management had very limited power in firing workers while central planners controlled labor mobility both across sectors and regions (IMF, 1991).

It is only in 1991 that, within a broad based reform package, the government began to liberalize the labor market by allowing wage scales and hiring and promotion criteria to be determined by collective contracts between workers and managers that are renewed annually. Such contracts however are subject to the national income policy which sets minimum and maximum wages for different categories of workers (World Bank, 1992).

The determination of wages has obvious repercussions on the whole economy. Thus, in this paper, we investigate the determinants of wages as of 1994 by using the first comprehensive household survey ever administered in the country since the beginning of Romania's transition to a market economy. We implicitly apply a neoclassical framework to wage formation and test a variety of hypotheses concerning the returns to higher education and experience, the presence of regional market segmentation, as well as gender and ethnic discrimination. As a whole, labor markets in Romania have received hardly any attention by researchers and little is known even about their basic features. Earle and Pauna (1996) and Kallai and Traistaru (1998) represent two noticeable exceptions providing valuable analysis of labor markets in the country since the transition. Earle and Pauna investigate the incidence and duration of unemployment and show, among others, the prevalence of unemployment among women with a rate that is nearly double of what found in other transition countries. Kallai and Traistaru concentrate on regional labor market trends in Romania during 1990-1995. Regional disparities are found to be of considerable relevance and, similar to the Polish experience, they have remained rather stable throughout the period. Moreover they find a lack of responsiveness in the evolution of regional average real wages to labor market pressures.

Although several studies have focused on the relevance of the gender wage gap in developed as well as developing countries[1], this phenomenon is not very well documented in Eastern Europe and particularly in Romania. The available evidence indicates that gender wage differentials before the transition in Central and Eastern Europe were on average similar to those in Western Europe and non-English speaking countries in general[2] (Fong and Paul, 1992, Atkinson 1992) but significantly better than in countries such as Australia, Canada, the UK and the US (Atkinson 1992). Orazem and Vodopivec, (1995 and 1998) while confirming the above trend for Estonia find, however, that relative female wages in Slovenia were considerably higher than in market economies.[3] More interestingly their study represents perhaps the only significant exception to the lack of formal analysis on gender wage differentials in Central and Eastern Europe after the transition.[4] Their data set in fact enables them to study the two countries up to 1994 and 1992 respectively; they can thus show that in terms of relative wages,[5] women actually gained from the transition.

In this paper we build upon previous analysis of labor markets in Eastern Europe, concentrating on the specific issue of gender-related wage discrimination, and in doing so, addressing the observed effect of productivity differences between women and men on average wage differentials. The identification, and subsequent efforts to eradicate wage discrimination between men and women is not just a socially desirable goal but has direct effects on efficiency and growth. As shown by Becker (1975), if male and female labor are assumed to be perfect substitutes, then economy-wide discriminatory behavior against women will generate not only a gain for men at the expense of women, but will also reduce firms' profits and therefore investments and growth. Moreover, such discrepancy in pay has a direct effect on the level of pensions, unemployment benefits and other means tested benefits paid to workers, contributing to a process of pauperization that, in general, has been severe for women in many Eastern European countries since the outset of economic liberalization[6]. Such issues are particularly relevant in Romania since women represent the majority of both the population and the labor force. Despite the Romanian labor code stipulates equal pay for equal work, little is known on the exact characteristics of pay levels in order to make a clear assessment of discrimination and its economic and social consequences. (Fong 1996).

Our analysis follows recently established econometric techniques as we estimate a Heckman selection model with maximum likelihood techniques. In addition, we are particularly interested

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[1] See among others: Knight and Sabot (1982); Blau and Kahn (1992); Psacharopulos and Tzannatos (1992); Appleton et al (1999).

2 For Yugoslavia however female relative wages were higher than in all other European countries. Note that relative wages do not capture properly the possible presence of discrimination as they do not distinguish it from productivity differences that may be present.

3 As of 1987 female wages were 88 percent of wages for men in Slovenia and 64 percent in Estonia as of 1989.

4 Rutkowski (1996) provides some statistical evidence on the evolution of earnings differentials in some transition countries: The reported female/male ratios are .74 for Bulgaria in 1993, .61 in the Czech Republic in 1992, .84 in Hungary in 1992, .79 in Poland in 1993 and .88 in Slovenia in 1992

[5] More specifically, the increase was to 0.90 in Slovenia and 0.74 in Estonia. This is a consequence of increased labor demand for the more educated in the work force (and that women have on average a higher level of human capital than men in the two countries). The authors however consider such evolution as a transitory one. Women are less mobile and men are gaining an increasing share of the new positions in the expanding sectors thus eroding the early relative gains that accrued to women.

[6] For a discussion on this point see Heinen (1994).

in issues of gender discrimination in transition economies. In addressing this question, we also find that there is still a certain amount of confusion over the proper implementation of the procedures for the estimation of wage discrimination. Consequently, we highlight shortcomings in previous research and offer a correct interpretation of the methods in question.

The paper is organized as follows: Section II discusses the data used in the analysis and the main features of the sample; Section III describes the methodology employed and reviews the literature on similar issues; Section IV presents the results obtained; and, Section V draws the conclusions and presents some possible avenues for future research.

II. THE DATA

The data set used is the 1994 Romania Integrated Household Survey (Government of Romania, 1994) conducted between April 1994 and December 1994. The investigation was conducted on a household sample of 24,560 households randomly selected from all districts of Romania and the city of Bucharest. The survey collected detailed information on household incomes and expenditures, labor market activity, public transfers and a wide range of living standard indicators. Since our study focuses on wage labor, we have excluded household employment as well as self-employment activities, both in agriculture and non-agriculture. Moreover, because we anticipate differences by sector as well as gender, separate models are estimated for men and women for both rural and urban areas.[7]

In analyzing wage labor markets we confine our investigation to individuals between the age of 15 and 65 who are not in school. After deleting observations with missing values, we are left with a sample of 21,297 observations for urban areas (of which 51.63% females) and 20,518 in rural areas (of which 48.99% females). Table 1 reports means and standard deviations of the sample. It is worth noting that while female wage workers in urban areas constitute 46.2% of the total sample, their share drops to 28.4% in rural areas. Moreover, the relatively high averages, and low standard deviations, in the number of hours worked are a clear indication of a very limited amount of part-time workers regardless of gender or location.

With respect to wages, the *observed* log differential of gross hourly wages between men and women is 0.22 in urban areas and 0.16 in rural ones. In other words, women are paid on average 80% and 85% of what men receive in urban and rural settings respectively. Such relative wages are higher than Western European or US equivalents: for example the ratio in Austria and Norway is 73%, while in Germany and United States 68%.[8] With respect to Eastern Europe, Orazem and Vodopivec (1995) in their study of labor markets in Slovenia find, a ratio of 90% in 1991.

As in many other Eastern European countries, the schooling rate is quite high. For each sub-sample at least 98% of the individuals have received some form of education, rural females with a 95.4% rate present the only exception. Also, within each sector, the distribution of individuals

[7] Nevertheless, we do test our assumption and the results obtained confirm its validity. For test results see footnote 13.

[8] Blau and Kahn (1992).

over the different educational levels shows a predominance of females with lower levels of education, particularly in rural areas. Lastly, with regard to the ethnic composition of the population, Romanians account for about 91% of the total while Hungarians are the largest ethnic minority at 7%. Gypsies, who are under-represented, comprise only 1.1% of the sample.

III. METHODOLOGY

We estimate wage equations for men and women, in urban and rural areas. Since labor market participation is not likely to be random, concerns arise over possible sample selection biases in the estimation of the wage function. To account for this potential problem, we estimate an Heckman Selection Model with maximum likelihood techniques.

Formally, let the wage function take the usual Mincerian form:

$$(1)\quad ln(W_i) = \beta X_i + e_i$$

where $1nW_i$ is the natural logarithm of the observed wage for individual *i,* is a vector of observed characteristics, β is a vector of coefficients and e_i is a stochastic error distributed $N(0, \sigma^2_e)$. Individual *i* is included among wage workers if:

$$(2)\quad \gamma\ Z_i + u_i > 0$$

where Z_i is a vector of observed individual characteristics, γ is a vector of coefficients and *u* is a stochastic error distributed *N(0,1)* that has covariance ρ with the error term e_i in the wage equation (1). Mills' ratio estimates are used as starting values for the maximum likelihood estimation. Let *F* be the cumulative probability function for the Normal distribution. The log-likelihood for observation *j* is then:

$$(3)\quad l_j = ln\left(F\left(\frac{I_j + \left(W_j - J_j\right)\rho / \sigma}{\sqrt{1-\rho^2}} \right) \right) - \tfrac{1}{2}\left(\frac{W_j - J_j}{\sigma} \right)^2$$

if W_j is observed and $ln(F(-I_j))$ if W_j is not observed, where $I_j = Z_j\gamma$ from the probit participation equation (2) and $J_j = X_j\beta$ from the wage equation (1).

Next, we proceed to decompose the wage differential between men and women in rural and urban areas. Our technique is based on methods originally developed by Oaxaca (1973) and Blinder (1973) and subsequently refined by Newmark (1988) and Oaxaca and Ransom (1994). Following Oaxaca and Ransom (1994) define the *observed* wage differential G_{mf} as:

$$(4)\quad G_{mf} = W_m / W_f - 1$$

where W_f represents female wages and W_m male wages. In the absence of discrimination, the wage differential between the two groups will reflect pure productivity differences Q_{mf} defined as:

$$(5)\ Q_{mf} = W_{om} / W_{of} - 1,$$

where the 'o' subscript denotes wages that would prevail in the absence of market discrimination. The market discrimination D_{mf} is then defined as the difference, or residual, between the observed wage differential $G_{mf} + 1$ and the portion of it explained by productivity differences $Q_{mf} + 1$. In logarithmic form this can be expressed as:

$$(6)\ ln(G_{mf} + 1) = ln(D_{mf} + 1) + ln(Q_{mf} + 1).$$

The discrimination component[9] can be further decomposed into female underpayment and male overpayment. Thus, we can specify a decomposition equation as follows:

$$(7)\ \ln(G_{mf} + 1) = \ln(d_{of} + 1) + \ln(d_{mo} + 1) + \ln(Q_{mf} + 1);$$

where $d_{of} = W_{of} / W_f - 1$ and $d_{mo} = W_m / W_{om} - 1$.

As shown by Oaxaca and Ransom (1994), within the context of semi-logarithmic wage equations estimated by ordinary least squares (OLS) from cross-section data, (7) can be reformulated as:

$$(8)\ ln(G_{mf} + 1) = ln(\tilde{W}_m / \tilde{W}_f) = \bar{X}'_m(\beta_m - \beta^*) + \bar{X}'_f(\beta^* - \beta_f) + (\bar{X}_f - \bar{X}_m)'\beta^*,$$

where $\tilde{W}$ denotes the geometric mean wage for the respective group, $\bar{X}_m$ and $\bar{X}_f$ are the vectors of mean values of the male and female regressors , β_m and β_f are the vectors of estimated coefficients and β^* is the estimated nondiscriminatory wage structure. Note that each term in equation (8) is the estimated value of the correspondent term of equation (7), i.e., the male advantage, the female disadvantage and the productivity differential.

In this context then, the issue is how to determine the wage structure β^* that would prevail in the absence of discrimination. Such choice poses a well-known index number problem given that we could, for example, use both the male or female wage structure as the non-discriminatory benchmark. While *a priori* there is no preferable alternative, the decomposition can be quite sensitive to the selection made.

If we let:

$$(9)\ \beta^* = \Omega\beta_m + (I - \Omega)\beta_f$$

where Ω is a weighting matrix and I is the identity matrix, then any assumption regarding β^* can be seen as an assumption regarding Ω.

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[9] It should be clear that the term discrimination is used here to identify what is actually a residual component that may actually be generated by other unobserved factors. Such a generalization, which is common to the literature, should be kept in mind when assessing the significance of the results reported.

The literature has proposed different weighting schemes to deal with the underling index problem: first Oaxaca (1973) proposes either the current male wage structure, i.e., Ω=*I*, or the current female wage structure, i.e., Ω=*0* -the null matrix-, as $\beta *$, suggesting that the result would bracket the "true" nondiscriminatory wage structure. Reimers (1983) implements a methodology that is equivalent to Ω=*0.5 I*. In other words identical weights are assigned to both men and women. Cotton (1988) argues that the nondiscriminatory structure should approach the structure that holds for the larger group. In the context of sex discrimination such weighting structure implies an $\Omega = I_m I$, where I_m is the fraction of males in the sample.

A more generalized method is provided by Newmark (1988), who shows that under certain conditions in the underlying utility function[10] the correct non-discriminatory wage structure $\beta *$ can be obtained by OLS estimates on the pooled sample where the model adopted is as in equation (1) i.e., without selection bias correction. As shown by Oaxaca and Ransom (1994), such result is equivalent to a weighting scheme of the form:

$$(10) \quad \Omega = (X^{'} X)^{-1} (X_m^{'} X_m)$$

where X is the observation matrix for the pooled sample and X_m is the observation matrix for the male sample. Such a weighting scheme is not constrained to produce results that are in general a convex, linear combination of the independently estimated male and female wage structures (Oaxaca and Ransom, 1994).[11]

Once these results are extended to a Heckman estimation procedure of offered wages, as we have done in this paper, the decomposition methodology should be carefully tailored to the new estimation setting. Recall that $\hat{\beta} X_i$ from equation (1) is an unbiased estimate of the wage $\hat{W}$ that an individual with characteristics X_i in the population can earn on average (Killingsworth, 1983). Thus, despite the fact that the estimation of (1) is implemented on observations of workers only, we obtain estimates of offered wages for the entire population. In this environment then, we believe that the natural decomposition to be performed is over the wage differential in offered wages for the entire sample. Accordingly, to compute $\beta*$ we apply the weighting scheme as in (10) using the observation matrix X and X_m for the entire sample and all the males respectively. Similarly, the decomposition, as in (8), is performed by taking sample means over the entire sample of males and females. Note further that such a procedure is equivalent to computing $\beta*$ by running an OLS regression on imputed wages, $\hat{W}$, for the pooled sample.

Other authors, having adopted the Heckman model for wage estimation, take a different route in analyzing wage discrimination. For example Reimers (1983), in estimating wage discrimination against Hispanic and African-American men, decomposes offered wages only for people with characteristics equivalent to that of the average worker. In the same fashion, Appleton et al.

[10] Specifically the firm's utility function is homogeneous of degree zero within each type of labor (e.g. skilled and unskilled, or blue-collar and white collar).
[11] Note that this is not true for the Cotton scheme, which is indeed a convex linear combination of the two separate estimates.

(1999) follow the Reimers approach and analyze discrimination between men and women in several African countries computing β^* as in (10) using the observations of workers only, despite the availability of population estimates.

Yet a different approach is taken by Glick and Sahn (1997) where β^* is computed running a Heckman model on the pooled sample of males and females. Consider however that this methodology implicitly assumes that the participation decision is the same for men and women, i.e., from equation (2) $\gamma_m=\gamma_f$. While this is a theoretically admissible case we believe it is not empirically relevant, given that, to our knowledge, men and women have never been assumed or found to have the same participation model. Therefore, unless such an assumption is true, incorrect Mill's Ratios will be used in estimation procedure.

As in Glick and Sahn (1997) we decompose the productivity difference $(\bar{X}_f - \bar{X}_m)'\beta^*$ as in equation (8) into its sub-components: education, experience, etc. Moreover, we extend the same exercise to the male advantage $\bar{X}_m'(\beta_m - \beta^*)$ and female disadvantage $\bar{X}_f'(\beta^* - \beta_f)$ terms, thus enabling us to better appreciate the specific relevance of each set of variables in the determination of the wage differential.

IV. FINDINGS

Wage equations

We discuss first the wage equation results as reported in Table 2.[12] Wald tests results confirm our assumption of differing wage structures between urban and rural areas both for women and men, the equality of wage determinants is rejected at the .001 level in both cases.[13] Overall our regressions seem well specified and yield plausible estimates. The effect of education variables is positive and significant in all four models. Average returns rise consistently with education relative to those with primary or less education for all four estimates. Table 3 reports the marginal returns to education.[14] University degree holders in urban areas have the highest marginal return; moreover, other than that group, rural areas display higher marginal returns than

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[12] The results obtained from the respective probit participation equations are available from the authors. The estimated probit coefficients seem reasonable. Age, education and land holdings are consistently significant in determining the probability of wage labor market participation across all the four sets of estimations. Also as expected the number of children is particularly relevant for women's participation. Other then the above variables the regressions were performed including non earned income, demographic characteristics, marital status, ethnicity, and regional and monthly variables. The principal exclusion restrictions used in the joint maximization are experience and the household demographic variables.

[13] The general form of the test statistic used is $(B_i - B_j)'(V_i + V_j)^{-1}(B_i - B_j)$, where β_i and β_j are the parameters for the two sectors and V_i and V_j are their corresponding variance submatrices. Such a statistic is distributed as a chi-square (j) under the null, where j is the number of restrictions. This test assumes $\text{cov}(B_i - B_j)$ =0 but does not impose equality of the variances of the disturbances for the two groups as a Chow test would do (Glick and Sahn, 1997). The value for the test statistic for males is 55.7 and for females is 45.0, and the degrees of freedom are 22.

[14] Given the semilogarithmic functional form of the wage equation we have computed such returns as exp(c)-1, where c is the untransformed marginal return of an education level. See Halvorsen and Palmquist, 1980.

urban ones. It is also interesting to note that females almost systematically outperform males in both sectors and particularly in rural settings.

Since we have introduced age dummies mainly to control for possible cohort effects, we examine next the coefficients associated with the experience variables. Consistent with our results on education, the coefficients are highly significant and suggest a substantial increase of offered wages with (potential) experience. However this is true only up to about 20 and 27 years of experience for men in urban and rural areas, respectively, and about 23 years for females in both settings; after that returns begin to decline.

Other than for rural women, a small, statistically significant premium is attached to being married, with married men receiving a higher return then women. This result is contrary to the findings in other Eastern European countries where wages (and participation rates) are lower for married women. [15]

Ethnic dummies allow us to investigate possible discrimination along such lines. Note first that the ethnic dummies are significant only for males. Hungarians, by far the biggest ethnic minority represented in the sample,[16] are found to have lower wages in rural and urban areas.[17] The other ethnic groups appear to be treated differently in rural areas, where the sign of the coefficient is positive, than in urban ones where as expected the coefficient is negatively signed. Despite the fact that the result at first may seem counterintuitive, note that such a variable is a composite of all the other ethnic groups, including Germans, present in Romania (about 2.2% of the sample for males). Moreover Gypsies, the other group we would expect to be discriminated against, account for roughly 45% of such residual ethnic groups in the total sample, but drop to 29% and 18% in urban and rural areas, respectively, when only wage workers are accounted for. These results, along with our results from the probit equations,[18] suggest a negative effect of ethnicity for such groups mostly in terms of job access rather then wage offers.

With respect to the land variables, the ploughland variables are negative and statistically significant for males both in rural and urban areas. This corresponds to the theory of employers paying a premium to the permanently available laborer without land because of lower recruitment costs (Bardhan, 1979). The positive sign on pasture land, may be explained by the low labor inputs for owners of this type of land, coupled with the higher reservation wage and wage offers, to landholders (Dasgupta and Ray, 1986).[19] The remaining variables included in the regressions are regional dummies, intended to capture the effects of local labor markets on

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[15] See Heinen (1994). Also, note that in our probit equations, the married variable for both males and females in urban areas and males in rural areas are positive and significant with the male dummies displaying higher coefficients. In light of the concern over the endogeneity of this variable, we tested its exclusion and found that it does not have a large effect on the other parameters.

[16] As indicated earlier, the Gypsy population is likely under-represented in the survey owing to their transient lifestyle.

[17] The difficult relations with ethnic Hungarians and with Hungary itself, which has generated inter-ethnic tensions as recently as 1990, may help explain our findings.

[18] The coefficient for rural males of such ethnic variable is equal to -.5744; with a standard error of .099 and a significance level of 1%.

[19] The inclusion of the land variables in the urban equations is justified by the fact that there are on average 7% of wage workers with positive land holdings in urban areas.

wages.[20] Other than for rural females, such regional dummies are overall statistically significant suggesting that labor market conditions do generate wage differentials among individuals with equal human capital endowments.

Wage Discrimination

We turn next to the comparisons of male and female wage offers in the two sectors. Our decomposition results are reported in Tables 4 and 5. For comparative purposes we also present the decomposition results obtained with weighting schemes other than the one presented in equation (10).

Our results indicate a total log wage differential of .244 in urban areas and of .598 in rural ones; this in turn implies that the average hourly wage for males is 27% higher than that of females in urban areas and 82% higher in rural areas. Thus while in urban areas the wage differential is comparable to that of Poland and, in general, places Romania among the countries with a high female/male wage ratio, in rural areas the difference in wages is considerably higher than in any other country for which figures are available.[21] While the observed log wage differential, as discussed in Section II, is almost equal to the offered one for urban areas, in rural settings the difference between the two is substantial. Such divergence is due to the fact that, only for rural females, observed and offered wages differ, as can be inferred from the significance ρ as reported in Table 2.

As discussed in the previous section we have analyzed wage offers for the entire sample. If one were to replicate the technique implemented by Reimers (1983) and, et al. (1999)(i.e. computing the difference for workers only), the offered log wage differential would be of .216 and .405 in urban and rural areas respectively, thus underestimating the *overall* offered wage differential of the population.

Relative to the relationship between the different weighting schemes, our results confirm those of Oaxaca and Ransom (1994). Consider first the results for urban areas presented in Table 4. The Ω as in equation (10) produces discrimination estimates that are below those generated by the other alternative schemes,[22] and conversely higher values for the productivity differential. Nevertheless, the portion accounted for by the different characteristics between men and women is quite small: 9.6%. In addition, the discrimination component is almost equivalently subdivided between the male advantage and the female disadvantage. With respect to rural areas the chosen weighting scheme produces discrimination estimates that are lower then those generated by the Cotton scheme or by the male weight, but higher then the female weighting scheme. In this case the percentage of the gap explained by differences in productivity is

[20] The inclusion of regional dummy variables is often criticized on grounds of endogeneity. Such criticism is justified only if one is willing to assume a rather high level of labor mobility. We have opted for the inclusion of these variables because in the case of Romania, as well as in many other Eastern European countries we believe that such an assumption on labor mobility is quite unrealistic given that labor market segmentation at the geographic level is a widespread phenomenon (OECD, 1995). For Romania in particular, see Kallai and Traistaru (1998)

[21] Note that the heterogeneity of data sets and time periods makes such cross-country comparisons only indicative.

[22] We do not report the results obtained with the Reimers scheme. The male and female samples are of almost identical size; therefore Reimers estimates and Cotton ones are practically identical. Furthermore note that as claimed the Cotton scheme produces estimated discrimination values that lie between those generated by adopting the female and male wage structure as weight.

sensibly higher, accounting for 20.9%. Still, the male advantage and female disadvantage components of the discrimination portion of the gap are almost equivalent.

We further decompose each term of equation (8) into its main sub-components to gain further insight into the determinants of the wage differential. Results are presented below for each decomposition scheme in Tables 4 and 5. We limit our discussion only to the weighting scheme in equation (10). Furthermore, we concentrate on the most relevant variables: education and experience. Both in rural and urban areas, differences in educational profiles are by far the principal determinants of the explained portion of the wage gap. With respect to the decomposition of the male advantage and female disadvantage care must be taken in interpreting the results. Looking first at the results for rural areas, the contribution of the constant term is actually even greater than the total in both cases. Furthermore, the contributions of educational variables actually have negative signs, thus implying that females are receiving a premium while males are penalized. Note, however, such results are dependent on the choice of dummy variables in our model: following standard procedure we have omitted from the regression the group of people with the lowest level of education. Therefore, our results suggest that for higher educational levels, females receive a (small) premium and males a discount. At the same time, the difference between the male/female intercept coefficients and that of β^* is also generated by a difference in the returns to education for individuals with primary or less education. This would suggest a high level of discrimination may take place among the lowest educated individuals. To further substantiate this inference we have rerun the models omitting the highest educational group from the regression and included the least educated people among the educational variables over which the decomposition is performed. The results obtained confirm our intuition: the male advantage component of the educational variables now becomes equal to .174 and the female disadvantage for education equal to .157, while the difference in the constant terms drops considerably to .106 and to .172, respectively.

In urban areas the results obtained are qualitatively similar. The role of the constant terms is quite high, while education and experience variables have either a negative or almost nonexistent contribution to the computation of the male advantage and female disadvantage. As above, we have rerun the model with the different set of educational dummies, and once again the results confirm the presence of discrimination mostly at the low education level: the new values of the decomposition for education in this case are .056 for the male advantage and .038 for the female disadvantage. The constant terms drop to .038 and .057, respectively.

It is also worth noting that we have not included in our regression any industry specific dummies due to the well-known problem of potential endogeneity that such variables may generate.[23] Nevertheless several studies (OECD, 1995, 1996; World Bank, 1992) have emphasized the prevalence of women in industries and sectors, such as education and health, that pay lower then average salaries, while the opposite is true for sectors like construction and mining where men are predominant. Thus, in our results, the difference in the constant terms may also capture an industry specific premium that, given the male/female distribution across industries, accrues mostly to men.

[23] Moreover if one is willing to make the assumption of exogeneity of the industry variables and thus include them in the regression, then a missing variable problem arises for the non-working individuals.

V. DISCUSSION AND CONCLUSION

Our study of Romania represents one of the first attempts to shed light on the determination of wages since the onset of the transition to a market economy. With the aid of comprehensive microdata we have been able to investigate the structure of labor earnings and gender discrimination both in rural and urban settings.

The relevance of human capital in the determination of wage offers is clearly indicated by our results: increasing returns to education and to experience are consistently significant, for males and females in urban and rural areas. Furthermore, education returns in rural areas, particularly for females, are greater for those in urban ones. Regional segmentation of labor markets is also evident in Romania. Such heterogeneity is likely the result of both economic history, and more specifically, the spatial allocation of resources during the centrally planned system. However, it is only with economic liberalization that the specialization of specific regions translated into differences in regional performance and consequently in the economic situation of local communities (OECD, 1995). Whether market forces alone will redistribute resources across regions in a more or less even way represents an interesting question for future study.

With regard to wage discrimination, we have clarified the conceptual and methodological shortcomings found in previous studies, demonstrating the correct application of estimation techniques. Our results highlight the higher incidence of discriminatory practices in rural areas; also our decomposition of the wage gap into its fundamental components reveals an occurrence of discriminatory behavior mainly at low levels of education, while experience seems to actually overcompensate women. As we have seen, females in Romania have, on average, lower educational attainment then men. Given that the process of adjustment of wages to market forces is not yet complete in Romania (Kallai and Traistaru 1998), and in light of the increasing difficulties faced by less skilled workers elsewhere in the region (Orazem and Vodopivec 1998), we might expect a decrease in the female relative wage as the transition process continues.

Furthermore, we have inferred that our findings may also capture discrimination that leads to high levels of heterogeneity in participation rates across genders in different sectors of the economy. The anecdotal evidence of the presence of discriminatory behavior in sector specific hiring as well as firing practices is actually overwhelming throughout all of Eastern Europe. More research is thus needed to specifically model potential discrimination in the determination of the occupation or sector of employment. As we have discussed, the inclusion of occupation or sector dummies in wage estimation poses well known endogeneity problems, and explains why we, and others, have concentrated on the examination of the wage gap across genders. Modeling discrimination practices that may influence labor market participation *per se*, rather than wage differentials, would therefore be a useful contribution if the endogeneity problems could be overcome.[24]

[24] A first attempt in this direction is offered by Appleton et al (1999). Analogous to the methodology used for wages, they first estimate the participation choice independently for men and women and then obtain the non-discriminatory structure by applying the same weighting matrix Ω used, as in this paper, for the wage model. While this represents an interesting attempt, it in not proven that, as shown by Newmark (1988) for the analysis of wage discrimination, this is the appropriate weighting scheme for the participation model as well.

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TABLE 1

Means and Standard Deviations of Variables

	Females				Males			
	Rural		Urban		Rural		Urban	
	Mean	Dev.	Mean	Dev.	Mean	Dev.	Mean	Dev.
Age(%)								
15-25	0.2121	0.4088	0.1581	0.3649	0.2429	0.4289	0.1684	0.3743
26-35	0.1594	0.3661	0.2491	0.4325	0.1764	0.3812	0.2183	0.4131
36-45	0.1949	0.3962	0.2964	0.4567	0.1810	0.3851	0.3000	0.4583
46-55	0.2071	0.4052	0.1661	0.3722	0.1823	0.3861	0.1748	0.3798
56-65	0.2265	0.4186	0.1302	0.3366	0.2173	0.4124	0.1384	0.3454
Educ. Level Compl. (%)								
None	0.0456	0.2086	0.0155	0.1237	0.0194	0.1379	0.0068	0.0822
Primary	0.2572	0.4371	0.0743	0.2623	0.1708	0.3764	0.0458	0.2091
Secondary I	0.3889	0.4875	0.2406	0.4275	0.3435	0.4749	0.1598	0.3664
Secondary II	0.2872	0.4525	0.5136	0.4998	0.4246	0.4943	0.5444	0.4981
Prof & Tech.	0.0101	0.1002	0.0602	0.2379	0.0233	0.1509	0.1026	0.3035
University I	0.0046	0.0675	0.0131	0.1137	0.0039	0.0625	0.0218	0.1462
University II	0.0065	0.0802	0.0827	0.2754	0.0144	0.1192	0.1187	0.3235
No. Children								
0 -1	0.1017	0.3167	0.0711	0.2721	0.0922	0.3041	0.0707	0.2696
1 - 5	0.2264	0.5039	0.1935	0.4570	0.2073	0.4878	0.1915	0.4546
6 - 14	0.5240	0.9066	0.5733	0.8556	0.5026	0.8850	0.5668	0.8584
No. Adults								
Females 25-65	1.4243	0.8404	1.2477	0.7365	1.7817	0.8862	1.5272	0.7549
Males 15-65	1.5549	0.7287	1.4735	0.7041	1.3624	0.7022	1.3272	0.6482
Elderly >65	0.2035	0.4548	0.0947	0.3212	0.1735	0.4463	0.0674	0.2789
Married(%)	0.7781	0.4155	0.7777	0.4158	0.6910	0.4621	0.7939	0.4045
Land (ha.)								
Plough Orch. Vin.	1.2599	1.5973	0.1142	0.6108	1.2345	1.5696	0.1150	0.6035
Pasture Hay	0.2612	0.6630	0.0259	0.3339	0.2677	0.6851	0.0276	0.3403

Ethnicity (%)								
Romanian	0.9159	0.2775	0.8999	0.3002	0.9113	0.2844	0.8988	0.3016
Hungarian	0.0627	0.2424	0.0761	0.2652	0.0659	0.2481	0.0749	0.2632
German	0.0017	0.0411	0.0036	0.0602	0.0020	0.0447	0.0057	0.0755
Gypsy	0.0106	0.1026	0.0132	0.1141	0.0099	0.0992	0.0127	0.1121
Other	0.0091	0.0947	0.0072	0.0845	0.0109	0.1038	0.0079	0.0883
Regions (%)								
Marramures	0.0662	0.2486	0.0506	0.2191	0.0673	0.2506	0.0484	0.2147
Cristiana-Banat	0.0935	0.2912	0.0852	0.2792	0.0893	0.2852	0.0839	0.2772
Transylvania	0.1581	0.3649	0.2261	0.4183	0.1628	0.3692	0.2303	0.4210
Oltenia	0.1391	0.3461	0.0897	0.2857	0.1376	0.3445	0.0905	0.2869
Muntenia	0.2398	0.4270	0.1624	0.3688	0.2315	0.4218	0.1643	0.3705
Dobrogea	0.0348	0.1833	0.0551	0.2282	0.0351	0.1839	0.0549	0.2277
Moldavia	0.2417	0.4281	0.1703	0.3759	0.2503	0.4332	0.1743	0.3794
R/U Bucharest	0.0268	0.1614	0.1607	0.3673	0.0262	0.1597	0.1535	0.3605
Log real obs. wage	6.3922	0.4944	6.5943	0.4875	6.5598	0.4835	6.8132	0.4966
Hours worked/month	164.22	34.00	163.76	31.69	172.70	32.90	168.41	31.28
Pot. experience (years)	20.00	10.85	20.14	9.45	22.30	12.28	20.11	10.15
# of wage workers	1711		5751		4299		6685	
Total obs.	10050		10997		10468		10300	

TABLE 2

Wage Equations (Dependent Variable: Log of Gross Hourly Wages)

Variable		*Rural* *Males*		*Rural* *Females*		*Urban* *Males*		*Urban* *Females*	
Education dummies	Secondary	0.106	***	0.344	***	0.089	***	0.133	***
(rel. to primary or less)		*0.025*		*0.061*		*0.020*		*0.025*	
	Prof. & Techn.	0.275	***	0.594	***	0.238	***	0.229	***
		0.047		*0.111*		*0.029*		*0.038*	
	University	0.431	***	0.871	***	0.440	***	0.583	***
		0.052		*0.119*		*0.028*		*0.041*	
Experience X 10^{-2}	Experience	0.940	**	2.117	***	1.141	***	2.159	***
		0.386		*0.655*		*0.275*		*0.348*	
	Experience sq.	-1.759	***	-4.573	***	-2.797	***	-4.680	***
		0.678		*1.291*		*0.550*		*0.727*	
Age dummies	26-35	0.052	*	0.096	*	0.087	***	0.026	
(rel. to 15-25)		*0.031*		*0.050*		*0.028*		*0.033*	
	36-45	0.047		0.184	**	0.109	***	0.084	*
		0.043		*0.074*		*0.033*		*0.044*	
	46-55	0.066		0.262	***	0.165	***	0.158	***
		0.048		*0.079*		*0.038*		*0.045*	
	56-65	0.180	***	0.266	*	0.217	***	0.107	
		0.066		*0.149*		*0.058*		*0.096*	
Region dummies	Cristiana-Banat	-0.033		-0.044		-0.093	***	-0.084	***
(rel. to Marramures)		*0.037*		*0.066*		*0.032*		*0.032*	
	Transylvania	-0.115	***	-0.072		-0.085	***	-0.039	
		0.033		*0.060*		*0.028*		*0.028*	
	Oltenia	-0.015		-0.088		-0.085	***	-0.120	***
		0.035		*0.061*		*0.032*		*0.032*	
	Muntenia	-0.077	**	-0.036		-0.131	***	-0.088	***
		0.033		*0.058*		*0.029*		*0.030*	
	Dobrogea	-0.130	***	-0.045		-0.074	**	-0.042	

		0.046	*0.080*	*0.035*	*0.038*
	Moldavia	-0.123 ***	-0.261 ***	-0.129 ***	-0.124 ***
		0.034	*0.060*	*0.029*	*0.030*
	Bucharest	-0.170 ***	-0.084	-0.086 ***	-0.072 **
		0.049	*0.079*	*0.030*	*0.029*
Ethnicity dummies	Hungarian	-0.080 **	-0.025	-0.069 ***	-0.014
(rel. to Romanian)		*0.033*	*0.049*	*0.025*	*0.025*
	Other	0.102 *	0.025	-0.164 ***	-0.078
		0.062	*0.110*	*0.046*	*0.061*
Land X 10^{-1}	Plou+Viny+Orch	-0.129 ***	-0.116	-0.344 ***	0.031
		0.047	*0.081*	*0.114*	*0.120*
	Pasture + Hay	0.005	-0.094	1.140 ***	-0.141
		0.135	*0.233*	*0.215*	*0.392*
Marriage dummy	Married Yes = 1	0.068 **	0.008	0.151 ***	0.034 **
		0.028	*0.030*	*0.025*	*0.015*
Intercept		6.425 ***	5.706 ***	6.448 ***	6.209 ***
		0.082	*0.167*	*0.064*	*0.080*
Sigma		0.465 ***	0.476 ***	0.456	0.442
		0.007	*0.020*	*0.004*	*0.005*
Rho		-0.165	0.356 **	-0.068	-0.062
		0.103	*0.144*	*0.093*	*0.129*
Log Likelihood		-8287.2	-4563.1	-9286	-9339.9
No. of observations		10468	10050	10300	10997

Notes: Standard errors in italics. Significance levels: a)* 10% level, b)** 5% level, c) *** 1% level.

TABLE 3

Marginal Returns to Education

	Rural		*Urban*	
	Males	*Females*	*Males*	*Females*
Secondary	0.111	0.410	0.093	0.142
Prof & Techn	0.184	0.284	0.160	0.100
University	0.168	0.319	0.223	0.424

TABLE 4

Wage Discrimination in Urban Areas

Offered Wage Diff.	*Discrimin.*	*Male Adv.*	*Fem. Disv.*	*Product. Diff.*
$Ln(G_{mf}+1)=.244$	$ln(D_{mf}+1)$	$ln(d_{mo}+1)$	$ln(d_{of}+1)$	$ln(Q_{mf}+1)$
$\Omega=(X'X)^{-1}X_m'X_m$	0.193	0.100	0.093	0.051
Education		-0.045	0.001	0.040
Experience		-0.019	-0.082	0.007
Constant		0.140	0.097	0.000
Ω=I (male)	0.205	-	0.205	0.040
Education		-	-0.035	0.032
Experience		-	-0.100	0.005
Constant		-	0.238	0.000
Ω=0 (female)	0.198	0.198	-	0.046
Education		-0.042	-	0.039
Experience		-0.1	-	0.005
Constant		0.238	-	0.000
$\Omega= l_m I$	0.201	0.102	0.099	0.043
Education		-0.022	-0.0171	0.036
Experience		-0.051	-0.048	0.005
Constant		0.123	0.115	0.000

TABLE 5

Wage Discrimination in Rural Areas

Offered Wage Diff.	*Discrimin.*	*Male Adv.*	*Fem. Disv.*	*Product. Diff.*
$Ln(G_{mf}+1)=.598$	$ln(D_{mf}+1)$	$ln(d_{mo}+1)$	$ln(d_{of}+1)$	$ln(Q_{mf}+1)$
$\Omega=(X'X)^{-1}X_m'X_m$	0.541	0.265	0.276	0.057
Education		-0.092	-0.015	0.052
Experience		0.001	-0.044	0.013
Constant		0.373	0.345	0.000
Ω=I (male)	0.585	-	0.585	0.013
Education		-	-0.076	0.021
Experience		-	-0.03	0.001
Constant		-	0.718	0.000
Ω=0 (female)	0.537	0.537	-	0.061
Education		-0.116	-	0.061
Experience		-0.041	-	0.012
Constant		0.718	-	0.000
$\Omega= l_m I$	0.561	0.263	0.298	0.036
Education		-0.057	-0.039	0.040
Experience		-0.020	-0.015	0.006
Constant		0.351	0.366	0.000