WOMEN'S UNEMPLOYMENT DURING THE TRANSITION: EVIDENCE FROM CZECH AND SLOVAK MICRO DATA[1]

John C. Ham*, Jan Svejnar, Katherine Terrell*****

*Department of Economics
University of Pittsburgh
Pittsburgh, PA 15260
Tel: +(1) 412 648 1762
Fax: +(1) 412 648 1793
E-mail:johnham@vms.cis.pitt.edu

**William Davidson Institute,
University of Michigan
Business School
Department of Economics
University of Michigan
and CERGE-EI, Prague
Tel: +(1) 734 936 5042
Fax: +(1) 734 936 8715
E-mail: svejnar@umich.edu

***William Davidson Institute,
University of Michigan
Business School
Ann Arbor, Michigan 48109-1234
Tel: +(1) 734 647 5844
Fax: +(1) 734 936 8715
E-mail: terrell@umich.edu

Abstract
We analyse women's weekly probabilities of leaving unemployment in the Czech and Slovak Republics (CR and SR) to investigate three questions: a) Why are unemployment rates much lower in the CR than the SR? b) Does the unemployment compensation scheme (UCS) substantially lengthen unemployment spells? c) Why are women's unemployment rates higher than men's? We find that differences in the behaviour of the individuals, employers and institutions in the SR and CR (as measured by differences in coefficients) play a larger role in determining the CR's shorter female unemployment spells than do differences in measured demand and demographic variables. The UCS has only a moderate effect on duration and its impact is greater in the CR. The difference between men's and women's spells (in each republic) are explained more by differences in coefficients than by differences in observed characteristics. (*Keywords*: Czech Republic, Slovak Republic, unemployment duration, unemployment policy; *JEL*: C41, H53, J23, J64, O15, P2)

1. Introduction

The Soviet-type system maintained full employment of labour by centrally setting wages and prices, requiring all able bodied individuals to work, and allocating to enterprises funds to provide the needed jobs. This full employment system was maintained in most Central and East European (CEE) countries until the late 1980s, although the requirement to work (e.g., for housewives) was not always fully enforced. In the early 1990s, the CEE economies started the transition to a market economy and experienced the shocks associated with the elimination of the Council for Mutual Economic Assistance and disintegration of the Soviet market. During this period, all the CEE economies experienced a major decline in output and employment, accompanied by outbursts of high inflation. Real wages also fell dramatically in most of these economies as they devalued their currencies, freed most prices and imposed wage (bill) controls.

As may be seen from Table 1, the transition was accompanied by rapidly rising and persistently high (double-digit) unemployment rates in all CEE countries, except for the Czech Republic (CR) where the unemployment rate remained between 3 and 4 percent throughout the first half of the 1990s. As is evident from Table 1, the principal reason for the rapid rise in the unemployment rate in all the other CEE countries in the early 1990s was the low outflow rate from (long duration of) unemployment in comparison to the CR.

From a policy standpoint, the first important issue we address is why unemployment spells have been so much shorter in the CR, and whether an understanding of this phenomenon could be used in formulating policy for reducing unemployment in the other CEE economies. We investigate the determinants of the unemployment durations of Czech and Slovak women using micro data that we collected specifically for the early transition period of 1991-93. The selection of the Slovak Republic (SR) as a comparison country to the CR is an integral part of our research strategy. Except for an interruption during World War II, the CR and SR formed one country from 1918 to 1993. During this period they shared a common currency, legal system and institutional framework. Yet, as may be seen from Table 1, the Slovak labour market statistics closely resemble those of the other CEE countries and differ markedly from those of the CR. By comparing the CR and SR during the early 1990s, when the two republics formed one country and unemployment diverged so dramatically between the CR and the rest of the CEE countries, we exploit the high variance in the data and yet automatically control for a large number of factors that might otherwise cause bias. Given the similarity of labour market outcomes in Slovakia and the rest of CEE, our Czech-Slovak comparison has important implications for labour market policy in the SR as well as the other CEE countries.

The second important issue addressed in this study is whether the determinants of unemployment duration have been different for women and men. The present study complements in an important way our parallel work on Czech and Slovak men, whose principal results we cite in this paper for comparative purposes.[2] There are several reasons for analysing the unemployment experience of women. First, the communist preoccupation with mobilising the entire potential labour force resulted in the CEE economies having the highest women's labour force participation rates (LFPRs) in the world. Since women comprise nearly half of the labour force, a comparative analysis of women's and men's unemployment is necessary for obtaining a complete understanding of the unemployment phenomenon in the CEE economies.

Another reason for focusing on women is that the experience of women in the transition economies may be different from that of women in the established market economies because women in the transition economies may have been adjusting their LFPRs from artificially high levels. In fact, most observers expected that during the transition period, women would decrease their participation dramatically. However, according to a study by Paukert (1995), the withdrawal of women in the early part of the transition was only slightly greater than that of men.[3] For example, she shows that in the Czech Republic, the labour force participation rates for the population 15 years and older fell from 1989 to 1994 by 6.4 percentage points for men and 9.6 percentage points for women. In Hungary the LFPR declined by about 16 percentage points for men and 20 percentage points for women over the same period. In Slovakia the decline was 5.2 percentage points for men and 7.7 percentage points for women. However in Poland, women's LFPRs declined less than men's (4.4 percentage points for women vs. 7.8 percent for men). Notwithstanding, in 1994 women's LFPRs in the CEE countries were still high, ranging from 66 percent in Poland to 79 percent in the CR, as compared to a rate of 59 percent in the U.S. in 1994 (Ehrenberg and Smith, 1996).[4] Hence, women's labour market experience is not clearly understood in the CEE countries and further analysis of their unemployment experience is necessary to improve our understanding of the principal labour market issues and their differences between the Czech Republic and the other CEE economies.

Third, from the start of the transition observers wondered if women's unemployment rates would be higher than those of men as a result of a gender bias in firing and hiring. An examination of the basic data reveals no consistent relative difference in women's and men's unemployment rates across the CEE countries, although there are systematic differences within individual countries. As may be seen from Table 2, between 1992 and 1996 women's unemployment rates were higher than those of men's in the CR, Poland, and the SR, but substantially lower in Hungary. While the basic unemployment rates by gender are of interest, the discrimination literature clearly points to the importance of controlling for differences in endowments (such as education) when comparing men and women. An important contribution that we make in this paper is to measure differences in the lengths men's and women's unemployment spells within the Czech Republic and Slovakia, controlling for endowments and other explanatory variables.

A fourth reason for focusing on women's unemployment experience relative to that of men is that gender related wage differentials are changing with the transition. Whereas under communism women enjoyed higher rates of return on education than men, the gap has been closing in recent years as the return to education has been rising faster for men than women (see e.g., Flanagan, 1994 and Chase, 1997 for individual studies and Svejnar, 1999 for a survey). There are hence important gender-specific developments in education-related wage differentials that may have important implications for the determinants of labour force participation and unemployment durations of women relative to men.

Finally, the third issue we focus on in this paper is the effect of the unemployment compensation system (UCS) on the female unemployment duration in the Czech and Slovak Republics. We compare these findings with those for men in a parallel study (Ham, Svejnar, Terrell, 1998). In market economies, married women have been found to exhibit substantially higher responsiveness to economic incentives than men or single women in a number of areas, including labour supply. Understanding the responsiveness to the UCS of single and married women in the transition economies is important since the CEE governments face a difficult trade-off in selecting the appropriate level of unemployment compensation. On the one hand, the governments want to provide an adequate social safety net to ensure political stability and support for completing the transition. On the other hand they want to minimise the disincentives of unemployment compensation in order to speed up the transition to an efficient market economy and to reduce the burden on the government budget. We carry out our analysis of the incentive effects of the UCS at two complementary levels. First, we estimate the effects of marginal changes in the UCS on unemployment duration using a sample of recipients of unemployment benefits. Second, we estimate the effect of infra-marginal changes in the UCS by comparing the experience of UCS recipients and non-recipients.[5]

The paper is structured as follows. In Section 2 we briefly describe the salient features of the Czech and Slovak UCSs. In Section 3 we discuss our data and methodology, while in Section 4 we present our estimates of the hazard (probability of exit from unemployment) functions. In Section 5 we compare the experience of recipients of unemployment compensation in each republic to assess the effect of infra-marginal changes in the UCS. In Section 6 we investigate differences in the unemployment durations of women across the two republics, as well as differences between men and women within each republic. We conclude in Section 7.

2. Characteristics of the Czech and Slovak unemployment compensation system

Czechoslovakia introduced a UCS in January 1990. By the time of our study (last quarter of 1991 to the middle of 1993), the system was working fairly smoothly. The implementation of the system in each republic followed the same guidelines. In this section we provide a brief description of the three main features of the UCS: eligibility, entitlement and benefits.[6] The CR and SR were part of the same country until January 1993 and essentially shared the same unemployment system over this period.[7]

<u>Eligibility</u> was quite broad as recently graduating students and anyone who worked for at least twelve months in the preceding three years was eligible for unemployment benefits in 1991-93. The exceptions were individuals who were fired for cause or who quit jobs repeatedly. Until January 1992, individuals out of the labour force were also eligible if they cared for a young child or a sick/disabled relative, or if they were in the military or imprisoned, for twelve months in the previous three years. This January 1992 change in policy is important for this study since it reduced the eligibility of many women.

<u>Entitlement</u>: In 1991, all eligible unemployed individuals were entitled to twelve months of benefits. On January 1, 1992, entitlement was reduced to six months. Since there was no "grandfather clause," those who became unemployed after July 1, 1991 received only six months of benefits. All individuals in our sample fall into this category.

<u>Benefits</u>: For those who worked before entering unemployment, the level of benefits was set in 1991 at 60% of the person's previous wage for the first six months of unemployment. However, individuals who were laid off because of major organisational changes had benefits set at 65% of their previous wage. For both groups, the replacement rate fell to 50% in the second six months of the entitlement period. On January 1, 1992, the replacement rates became 60% for <u>all</u> workers during the first three months and 50% during the second three months of their unemployment spell. Those who had never worked before and graduating students received benefits equal to 60% of the minimum wage in the first half of the entitlement period and 50% in the second half.

Whereas in 1991 there was no upper limit on benefit levels, there was a minimum level set at 1,200 Kcs (60% of the minimum wage).[8] In January 1992 a maximum level equal to 150% of the minimum wage (180% for those in training) was imposed, and the minimum was replaced by the "minimum living standard" (MLS), which is analogous to the household poverty line in the United States.[9] In fact, an unemployed person was eligible for social assistance (welfare) in addition to his/her unemployment benefits if the sum of his/her unemployment benefits and the income of other household members was less than the household MLS.

Once a person's unemployment benefits expired, he/she was eligible for social assistance if his/her household income fell below the MLS. In practice, a single person collected benefits but the amount depended on whether he/she lived at home. A married person only collected benefits if the combined income of other household members was below the household MLS.

A significant number of individuals who were ineligible for unemployment benefits registered as unemployed (the non-recipients). Some registered in order to obtain the assistance of the district labour office in finding a job. Registration was also a prerequisite for receiving welfare. As noted above, those who did not have the necessary work experience in the previous three years (or its equivalent before January 1992) were ineligible for benefits, as were those who were fired for cause or quit repeatedly. Further, if a graduating student started a job and lost it before acquiring twelve months of experience, he/she was not eligible for benefits.

There is a great deal of similarity in the features of the Central and East European UCSs, as they were patterned after the models of the West European countries.[10] Thus, our econometric approach should be useful to those studying the unemployment compensation system in other CEE countries.

3. Data description and econometric model

To carry out our analysis, we collected weekly data on a stratified random sample of men and women who registered at the district labour offices as unemployed between October 1, 1991, and March 31, 1992. We followed these individuals from the onset of their unemployment spell to the end of their unemployment spell or the end of July 1993, whichever came first.

The sample was selected as follows: First, we randomly selected 20 districts in each of the two republics. We then randomly selected 150 individuals in each district who registered at the unemployment office during the last quarter of 1991 or the first quarter of 1992.[11] We eliminated observations for individuals who suffered a prolonged illness, were ineligible for unemployment compensation, had missing values of variables, or took part in training.[12] In the CR, this left us with a sample of 1,269 women, 851 of whom received unemployment benefits (recipients) and 418 who did not receive benefits (non-recipients). In the SR, we analysed a sample with 1,120 women consisting of 902 recipients and 218 non-recipients. The corresponding men's sample consisted of 780 recipients and 482 non-recipients in the CR, and 1063 recipients and 229 non-recipients in the SR. Additional characteristics of the women's sample are presented in appendix Table A1.[13]

We use a duration model to analyse the determinants of unemployment spells. This model is preferred to a regression model because factors such as demand conditions and unemployment benefits change over an individual's unemployment spell, and time changing explanatory variables cannot be captured in a regression framework.[14] Since we have weekly data on the duration of unemployment spells, we denote the hazard function (the probability of leaving unemployment for employment) in week r of the unemployment spell as

$$\lambda(r \mid \theta) = (1 + \exp(-\ y(r \mid \theta)))^{-1} \tag{1}$$

where

$$y(r \mid \theta) = \alpha_0 B(r) + \alpha_1 W + g(E(r)) + Z(r)\ \gamma + h(r) + \theta. \tag{2}$$

In equation (2), B(r) equals unemployment benefits in week r, W is the individual's previous weekly wage, $g(\cdot)$ is a function of remaining entitlement E(r) in week r, Z(r) is a vector of demographic conditions and demand variables, h(r) captures the effect of duration dependence on the hazard and θ denotes an unobserved heterogeneity term. We estimate the model given by equations (1) and (2) by maximum likelihood. The conditional contribution to the likelihood function (conditional on the unobserved heterogeneity component θ) for someone who is still unemployed after r weeks is given by the survivor function

$$S(r \mid \theta) = \prod_{v=1}^{r} (1 - \lambda(v \mid \theta)). \tag{3}$$

The conditional contribution to the likelihood function of someone who completes an unemployment spell of t weeks is given by

$$f(t \mid \theta) = \lambda(t \mid \theta)\, S(t-1 \mid \theta). \tag{4}$$

Let $\phi(\theta)$ represent the density function for the unobserved heterogeneity. The unconditional contribution to the likelihood for the spell that ends in week t is then given by

$$L(t) = \int \lambda(t \mid \theta)\, S(t-1 \mid \theta)\, \phi(\theta)\, d\theta. \tag{5}$$

The contribution of a censored spell is calculated in an analogous manner.

Parallel to HST (1998), in order to identify the benefits and wage coefficients we rely on five sources of independent variation in these variables. First, benefits dropped from 60% to 50% of the previous wage at thirteen weeks of duration.[15] Second, a maximum benefit level was imposed in 1992. Third, a number of individuals had their benefits raised to the minimum level of benefits. Fourth, unemployment benefits were not indexed and hence we discounted benefits by the consumer price index to capture the erosion of the real value of benefits over time. On the other hand, we assumed that the appropriate proxy for the mean of the worker's wage offer distribution is his real wage at the time he began his spell.[16] Fifth, the replacement ratio was 0.65 for those laid off because of a major plant closing prior to January 1992.

In equation (2), we parameterise $g(\cdot)$ as a linear function of (i) remaining weeks of entitlement, (ii) a dummy variable for the last week of entitlement before benefits have been exhausted, and (iii) an exhaustion dummy equal to 1 for all weeks after entitlement has been exhausted. To estimate the impact of remaining entitlement on the hazard rate, we need variation in remaining entitlement that is independent of other determinants of the hazard function, particularly current duration. However, remaining entitlement is a simple linear function of initial entitlement and current duration for women who register for benefits at the district labour office immediately after becoming unemployed, and initial entitlement is constant across individuals. To address this identification problem, we adopt the approach used in HST (1998) and exploit the fact that a significant number of individuals wait to register for unemployment benefits after becoming unemployed. For such individuals, remaining entitlement is not a simple linear function of current duration and initial entitlement.[17] The conditional contribution to the likelihood of an individual who waits T weeks to register and then stays unemployed for t additional periods is

$$L(t \mid \theta) = \lambda(t+T \mid \theta) \prod_{v=1}^{t-1} \left(1 - \lambda(T+v \mid \theta)\right).^{18} \tag{6}$$

As noted above, Z(r) contains variables measuring individual specific demographic characteristics and local demand conditions in week r. The means of these variables are presented in Table A1. All demographic variables except for age are in dummy variable form. We use three dummy variables for educational achievement, where the control group consists of those with only a junior high school education, the basic education required by law. The recent graduate variable is coded 1 if an individual is a graduate within the last year from a university or high school. The marital dummy is equal to 1 for those who are married (formally or by consensual union). We experimented with three variables to account for differences in demand conditions across districts and two republics and a variable to proxy differences in market structure. The first two -- quarterly data on district unemployment and vacancy rates by educational group -- vary quarterly over the duration of a spell and across individuals. The third variable is the real value of per capita industrial production in the district in a given year.[19] It takes on different values across calendar years and across districts. Finally, we used the ratio of employment in agriculture to employment in industry (manufacturing) to proxy differences in market structure across republics.

As mentioned above, h(r) captures the effect of duration dependence on the hazard. Since both remaining entitlement and benefits are a function of duration, it is important to allow for flexible duration dependence and unobserved heterogeneity (θ). However, recent Monte-Carlo evidence (Baker and Melino, 1997) suggests that it is important to be relatively conservative in choosing the number of parameters describing the duration dependence and the heterogeneity distribution. As a result, we use a polynomial in log duration to measure duration dependence and choose the degree of the polynomial using the Schwartz criterion (e.g., Judge et al., 1980), assuming that the degree is less than or equal to five. Further we assume that θ is drawn from a discrete distribution with 2 support points, and thus we use a simplified version of the Heckman and Singer (1984b) approach. We assume that θ is drawn from a discrete distribution with 2 support points. Thus we use a simplified version of the Heckman and Singer (1984b) approach.[20]

We also estimate the model for non-recipients. For these individuals, benefits and entitlement are not defined and we also do not have their data on previous wages. Thus we estimate a smaller version of the hazard function omitting these variables. Moreover, we do not allow for unobserved heterogeneity when estimating the model for nonrecipients. There are two reasons for this. First, the bias from ignoring heterogeneity is likely to be smaller when the explanatory variables do not depend on duration. Second, we have only a limited amount of data for non-recipients, and the likelihood function for male nonrecipients was poorly behaved because of the relatively small samples available for nonrecipients.[21]

Finally, the estimated parameters of the hazard functions are difficult to interpret. To provide results analogous to those produced by regression analysis, we use our parameter estimates to calculate the effect of a change in a given explanatory variable on the expected duration of unemployment. Since we have data on only a relatively short time period after the beginning of the transition, we use a truncated expected duration (ED)

$$\text{ED} = \Big[\ \sum_{t=1}^{K} t\ f(t) + (1 - Pr(t\ < 4yrs))\ *\ 4yrs\ \Big],\ where\ \ K = 4\ \ years\ \ - 1\ week \tag{7}$$

where both f(t) (defined by equation 5), and the term Pr(t < 4) depend on the parameter estimates β and the values of the explanatory variables. We can numerically differentiate this expression with respect to the explanatory variables (or carry out the equivalent exercise for dummy variables) to obtain estimates comparable to regression coefficients.[22]

4. Estimates of the hazard functions for women in the CR and SR

The parameter estimates of the hazard functions for unemployed women in Czech Republic and Slovakia are contained in Table 3. The corresponding results of the expected duration experiments (giving the effect of a change in a dummy variable or a 10 percent increase in a continuous variable on unemployment duration in weeks) are contained in Table 4.

Since models with heterogeneity are better behaved numerically when there are not many insignificant coefficients, we have used the following strategy to choose a parsimonious model. We examined the no-heterogeneity results of the full model (shown in Appendix Table A1) and constrained the coefficients to be the same for married and single women where this seemed appropriate, given the parameter estimates and standard errors. Since there is also multicollinearity in the demand variables, we included only the 'most' significant of these variables in our equations. We then allowed for unobserved heterogeneity in estimating the resulting specification.

4.1 Effects of the UCS

In this section we will be focusing on the results for recipients, reported in columns 1 and 4 of Tables 3 and 4, and the elasticities reported in Table 5.

Benefits: The weekly benefits coefficients, (constrained to be equal for married and single women in each republic based on the results in Table A1) have the expected signs but are insignificant. The point estimates imply an elasticity of 0.21 and .0003 in the CR and SR respectively (columns 1 and 2 of Table 5). Thus the estimated elasticity is larger for women in the CR than in the SR. Comparing the male elasticities from HST (1998) with the female elasticities, in the CR the male elasticity is 0.34 and statistically significant.[23] On the other hand, the male elasticity in the SR is estimated to be 0.06 and is insignificant. Thus men may be somewhat more responsive to benefits in the CR, while men and women in the SR seem quite unresponsive to benefits.

Previous Wage: The wage coefficient in the CR is statistically significantly for married women and implies an elasticity of –0.71, while the coefficient for single women is insignificant. We constrained the wage coefficients to be the same for single and married women in the SR. The coefficient is not statistically significant but does imply an elasticity of –0.41. Thus we find that a higher opportunity cost of time only has a statistically significant effect for married women in the CR, although the true elasticity in the SR may also be substantial. Comparing the CR results for women to those for men, we note that the wage coefficient for men in the CR is insignificant and implies a very small elasticity of –0.01. On the other hand, men in the SR have a statistically significant coefficient but the estimated elasticity is –0.25, which is somewhat smaller than the estimate for men.

Entitlement: The three entitlement coefficients (remaining entitlement, last week of entitlement and entitlement exhausted) generally differ in each republic effect generally differed for married and single women in each republic. Beginning with the CR, the coefficients on weeks of remaining entitlement are statistically significant for both married and single women.[24] There is a statistically significant positive spike in the last week of entitlement for married women, while the last week coefficient is negative and insignificant for single women. Finally the coefficient on the benefits exhausted dummy variable (constrained in Table 3 to be equal for single and married women) is negative and significant. Taken together, these coefficients imply an elasticity of .41 for married women and 0.37 for single women. The CR coefficients for women are qualitatively similar to those for men in the CR; moreover the male elasticity with respect to entitlement of 0.44 is quite similar to those for single and married women.

In Slovakia, the remaining weeks of entitlement coefficients are negative and statistically significant for married and single women. Raising remaining entitlement by one week would raise single women's duration of unemployment by about 0.7 of a week and the spell of married women would be lengthened by about 0.5 of a week. Given the large and different base durations of married and single women, the elasticities implied by these coefficients are 0.28 and 0.17 for married and single women, respectively.[25] It appears that only married women exit at a significantly higher rate in the last week of receiving benefits (coefficient of 0.97 in Table 3). Finally, the benefits exhausted dummy variable has insignificant coefficients for both married and single women. Overall these coefficients imply an estimated elasticity with respect to entitlement of 0.17 for married women and 0.26 for single women. The coefficients on the entitlement variables for men in the SR differ from those for women since the last week and benefits exhausted dummies are statistically significant for both single and married men. Moreover, the entitlement elasticity for men is estimated to be 0.41, somewhat larger than the elasticities estimated for single and married women.

Overall, the results indicate that Czech women are more responsive to the parameters of the UCS than Slovak women. (Czech women have larger benefit, wage and entitlement elasticities, and married women in the CR have a statistically significant wage coefficient.) Married women seem to differ only from single women in their response to the previous wage. Further, the male and female responses to the UCS parameters appear relatively similar in the CR, and both men and women in the CR demonstrate only moderate responsiveness to the UCS when compared to those in other countries (see Devine and Kiefer, 1991, Table 5.2). In our previous work we found that men in the SR were relatively unresponsive to the UCS, and these results suggest that women in the SR are even less responsive to the UCS. The upshot is that policy makers do not have to be overly concerned with the disincentive effects of the UCS based on the results in Tables 3 and 4.

4.2 *Demographic and demand variables*

In this section we will discuss the results for recipients and non-recipients and hence examine the values for the coefficients of the demographic and demand variables in columns (1),(3),(5) and (6) of Tables 3 and 4.[26] We find that a woman's age is not a significant explanatory variable for the probability of leaving unemployment in either republic. We find that education does matter in the sense that in both republics those with only junior high school leave unemployment at a significantly slower rate than those with more education. Moreover, the results in Table 4 indicate that this education effect is substantial. This parallels our result for men in the SR but not in the CR; in the CR we found that there was no difference in unemployment duration across the different educational groups.

The estimated coefficients on Romanies (gypsies) are negative, large and highly significant, confirming the findings in other CEE countries as well that Romanies have much longer spells of unemployment. Whereas handicapped women have on average about fifty percent longer spells than the non-handicapped in the CR, the estimated effect is large but statistically insignificant in the SR.

The base duration results for single and married women in Table 4, which incorporate all the marital status interactions as well as the dummy, suggest that married women recipients stay unemployed longer than single women. However, among non-recipients there is no significant difference between the unemployment durations of married and single women in the CR (using the coefficient for the marital dummy in Table 3), although there is a significant difference among nonrecipients in the SR. The expected duration experiments in Table 4 suggest that these marital status effects are substantial. Women who recently graduated from school do not seem to have longer spells than women who have not recently graduated (with one exception). Women living in Prague leave unemployment more rapidly than women living in other parts of the CR, while recipient women living in Bratislava tend to have longer spells. This result parallels the men's findings and may reflect the fact there was more activity in Prague than in Bratislava during this period. It has also been hypothesised that since Bratislava is very close to Vienna (while Prague is in the centre of the Czech Republic) people in Bratislava had the opportunity to register as unemployed (collecting unemployment or welfare benefits) and work in Vienna.

The coefficients on the demand variables have the expected signs, but in the CR they are only statistically significant for the non-recipients. They indicate that exit rates are positively correlated with higher district level vacancy rates and industrial production and lower unemployment rates. The market structure variable indicates that nonrecipient women in the CR who live in districts with a high ratio of employment in agriculture to employment in production ratio tend to leave unemployment sooner than those who live in districts with lower ratios. However the opposite, more intuitive outcome is found in Slovakia, where the coefficient is marginally significant at standard confidence levels for recipients, as is the vacancy rate variable. The district unemployment rate and industrial production coefficients have the expected sign and are each individually statistically significant for nonrecipients in the SR.

5. Estimating infra-marginal changes in unemployment compensation

The fact that we have data on both recipients and non-recipients enables us to answer the interesting question of how the unemployment duration of a recipient would change if she were deemed ineligible for unemployment compensation at the beginning of her spell.[27] Since we have data on both recipients and non-recipients, we can assess what the expected duration for recipients would look like if their behaviour were governed by the (estimated) hazard function for non-recipients, and compare this value to their expected duration using the (estimated) recipient hazard function.[28]

Since the variables relating to unemployment compensation are not available for the non-recipients, we use a smaller set of explanatory (demographic and demand) variables X^*_{nr} to estimate the hazard rate for the non-recipients. Denote the same smaller set of variables for recipients as X^*_r and the corresponding parameter estimates for non-recipients and recipients as β^*_{nr} and β^*_r, respectively. Using this notation, we calculate

$$Diff\ 1 = ED\left(\beta^*_{nr}, X^*_r\right) - ED\left(\beta^*_r, X^*_r\right), \tag{8}$$

where $ED(\beta, X)$ denotes the expected duration of unemployment at the mean values of the Xs. (For the sake of brevity and clarity of notation, in the rest of the paper we use X to denote mean values in the expected duration calculations.) We describe Diff 1 as "moving someone from being a recipient of unemployment insurance to being a non-recipient." Similarly, we can consider the effect of making a nonrecipient eligible for unemployment compensation at the beginning of her spell

$$Diff\ 2 = ED\left(\beta^*_r, X^*_{nr}\right) - ED\left(\beta^*_{nr}, X^*_{nr}\right). \tag{9}$$

We define Diff 2 as "moving someone from being a non-recipient to being a recipient:"

Table 6 contains the relevant parameter estimates and calculated values. For the CR, we estimate in row 5 that moving a woman from the recipient to the non-recipient category would lower her unemployment spell by 11.9 weeks or by slightly more than 50 percent of the 21.6 weeks expected duration of a recipient (row 1). On the other hand, moving a woman from the non-recipient to the recipient category would raise the length of her unemployment spell by 11.6 weeks (row 6) and hence more than double the length of the 11.5 weeks expected duration of her unemployment spell (row 2).

The results for the SR are sizeable but much less dramatic in percentage terms. In row 5 of Table 6 we estimate that moving a woman from the recipient to the non-recipient category lowers her unemployment spell by 18.6 weeks or less than one-third of the 61.5 week expected duration of a recipient (row 1). Moreover, moving a woman from the non-recipient to the recipient category in the SR only increases duration by 7.7 weeks (row 6) or by less than 10 per cent of her expected duration of 92.6 weeks (row 2).

These findings on the "infra-marginal" effect (Table 6) corroborate those on the marginal effect discussed above in Table 5 in that the women in the Czech Republic appear to be more sensitive to the incentives of the unemployment compensation system than women in Slovakia. Indeed, the results in Table 5 suggest that women in the CR may be quite responsive to the parameters of the UCS. Certainly these estimates of the effect of this "infra-marginal" change in the UCS is substantially larger for women in the CR than for men in the CR. The effect of this "infra-marginal" change appears quite similar across men and women in the SR.

6. Analysing differences in unemployment duration between the republics

6.1 Differences across the republics for recipients and non-recipients

Letting subscripts s and c stand for the Slovak and Czech Republics, respectively, the difference in expected duration of unemployment between the two republics may be expressed as:

$$Ed_s - ED_c = ED(\beta_s, X_s) - ED(\beta_c, X_c). \tag{10}$$

Using the approach developed in HST (1998), we carry out a (nonlinear) Oaxaca-type decomposition to calculate the portion of this difference that arises from differences in parameter estimates and the portion that arises from differences in the explanatory variables. Furthermore, we decompose the contribution arising from differences in the explanatory variables into a portion that is attributable to differences in demand variables and a portion that is due to differences in the other regressors, primarily demographic variables.

Column 1 of Table 7 contains our results for recipients in the two republics. We estimate that recipient women stay unemployed approximately 46 weeks longer in the SR than the CR, and that essentially all of this difference is due to differences in coefficients.[29] Hence, the behaviour of the unemployed women, employers, and labour market institutions is very different in the two markets. Surprisingly, only about three weeks (6 percent) of this difference arises from differences in our measures of demand, which suggests that in future research it will be useful to investigate the effect of other measures of demand than the ones considered in this study. Although demographic variables play a very small role in the difference in expected duration, the decomposition shows an interesting twist. It indicates that the difference in the expected durations of unemployment between Czech and Slovak recipient women would actually shorten by 1.8 weeks if each were to have the others' mean demographic characteristics.

Column 2 of Table 7 contains our results for non-recipients. The 82-week difference in expected duration between non-recipients in the SR and CR is almost double that between recipients in the two republics. Now differences in the coefficients account for 46.2 weeks or slightly over one-half of this difference, while differences in the explanatory variables account for the remainder. Differences in demand conditions account for 13.6 weeks or slightly over one-third of the difference in expected duration that is due to explanatory variables. The remaining two-thirds of the difference that is attributable to differences in the explanatory variables arises from differences in demographics. Examining Table A1 indicates that the much higher proportion of Romanies and those in the lowest education group in the SR among nonrecipients is likely the cause of this difference. The results for women non-recipients assign a significant part of the explanation to the differences in the explanatory variables. In that sense they are much closer to the parallel results obtained by HST (1998) for men than are the results for women recipients.

6.2 Differences between male and female expected unemployment durations

The Oaxaca decomposition of differences in the expected length of unemployment spells of men and women is presented in Table 8. In order to carry this decomposition it is useful to have the same specification for men and women. Hence we re-estimated hazard functions for recipient and non-recipient men in the two republics and present the coefficients from this identical specification in Appendix Table A3. The expected durations for men using this specification are presented in Table A4.[30]

The first finding in Table 8 is that for recipients in the CR, the difference in the unemployment durations between men and women are about 25% of the expected length of a woman's spell. Moreover all of this difference is due to differences in the coefficients and none of it arises because of differences in endowments. However, for nonrecipients there is no difference in the length of unemployment spells between men and women. Second, the decomposition indicates that in both republics, and for both recipients and non-recipients, the differences in men's and women's estimated coefficients are more important than differences in their observed characteristics (demographic and demand variables) in explaining women's longer spells. Among the observed variables, differences in observed demographic characteristics play a more important role than differences in local demand conditions, which should be expected given that men and women are distributed equally among the sampled districts. Hence, we conclude that differences in men's and women's labour supply behaviour, the practices of employers and institutions towards gender are driving the differences between men's and women's exits from unemployment in both these republics.

7. Conclusions

Our analysis of the determinants of women's probabilities of leaving unemployment (and hence their unemployment durations) in the Czech Republic (CR) and Slovak Republic (SR) has been motivated by three important facts. First, throughout most of the 1990s, Slovakia and all other Central and East European (CEE) economies, except for the CR, experienced high unemployment rates and long durations of unemployment spells. Second, the slow rise in the unemployment rate in the CR and its rapid rise in the SR and the other CEE economies in the early 1990s is primarily attributable to higher flows from unemployment to jobs in the CR than in the SR and the rest of CEE. Third, since women represent nearly half of the labour force in these two countries (and in the CEE economies in general) understanding of the unemployment behaviour of women is essential for understanding the unemployment phenomenon. Moreover, relating the analyses of women and men's unemployment duration sheds light on the question of why women have higher unemployment rates in the Czech and Slovak Republics. Since until 1993 the CR and SR shared a common currency, legal system and institutional framework, they represent an excellent laboratory in the CEE region for comparing divergent labour market outcomes while controlling automatically for a number of factors that could otherwise contaminate such a comparison.

We began an analysis of the first two of these facts with an analysis of men's unemployment duration in these two republics in a parallel paper (HST, 1998). Here we continue the analysis by focusing on the determinants of unemployment durations on Czech and Slovak women. For both men and women, we carry out the analysis using data on both recipients and non-recipients of unemployment benefits, thus allowing us to examine the effect of infra-marginal, as well as marginal, changes in unemployment compensation. Furthermore, we decompose the difference in unemployment duration between republics for each gender and recipient group into a portion due to the difference in the explanatory variables and a portion due to differences in the coefficients. Finally, we address the third fact by first comparing our results for men and women across the two papers, and second by carrying out a nonlinear Oaxaca decomposition of the difference in unemployment duration between men and women.

Combining the analysis in this paper with that in the parallel men's paper yields the following results. First, for all but recipient women, differences in the behaviour of the unemployed, the employers and the institutions in these two countries (as measured by differences in their parameters) play a slightly more important role than differences in the measured demand and demographic characteristics in explaining the much shorter unemployment spells in the Czech Republic compared to the Slovak Republic. For recipient women, all of the difference in unemployment duration across the republics is due to differences in the behaviour in individuals, firms, and labour market institutions. Among the explanatory variables, differences in demand conditions were the most important factors for recipient men, while differences in demographics were more important for nonrecipients.

Thus differences in the functioning of the labour markets across republics is very important both for men and women and for recipients and nonrecipients. A full discussion of the factors that are likely to contribute to this labour market difference is contained in HST (1998). Plausible explanations for the difference in the estimated coefficients include the much higher level of foreign investment in the CR as compared to the SR, the relatively greater decline in military production in the SR, the more rapid growth of small-scale service firms in the CR, and differences in the response of the manufacturing sector in each republic arising from differences in the age and location of factories between the two republics.[31]

Our second important finding is that the unemployment compensation system (UCS) is generally having only a moderate effect on the durations of unemployment of both men and women in each of these countries. We find that the responsiveness of both men and women to the UCS is higher in the CR than in the SR. This is true for both the marginal effect (i.e., the elasticities of the unemployment duration with respect to benefits and entitlement) and the "infra-marginal" effect (hypothetically switching an individual from recipient status and non-recipient status and vice-versa) when we experimented with taking away benefit entitlement from a recipient by "making" him/her a non-recipient. In this paper we measured the effect of changes in the UCS parameters on both married and single women in each republic. We found that the married women in the CR are somewhat more responsive to changes in benefits and entitlement than single women, whereas in Slovakia the married and single women are similarly (un)responsive. The finding of a higher responsiveness of married women in the CR parallels the pattern of behaviour of married relative to single women in market economies, while the SR finding diverges from the established market economy pattern. We also found that men and women appeared to be quite similar in each republic in terms of their response to the UCS. However, we note that the one exception to the above discussion is the measured effects of the infra-marginal changes in the UCS on women in the CR.

Third, in terms of the effects of demographic characteristics on unemployment durations, we found in both republics that age is not a factor but that education matters in that the least educated (junior high school graduates) tend to have the longest spells. This effect is stronger in Slovakia than in the Czech Republic (where education is not a significant determinant of the hazard rate for female non-recipients and male recipients). Romanies, both men and women, have much longer spells in both republics. We found that except for nonrecipients in the CR, married women have longer spells than single women.

The Oaxaca decomposition of the differences in the expected length of unemployment spells of men and women indicates that in both republics, and for both recipients and non-recipients, the differences in the estimated coefficients are more important than the differences in observed characteristics (demographic and demand variables) in explaining women's longer unemployment spells. Hence, we conclude that differences in men's and women's behaviour and the practices of employers and institutions towards gender are dominating the differences between men's and women's exit rates from unemployment in both republics.

Endnotes

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[2] See Ham, Svejnar, and Terrell (1998). For studies investigating unemployment duration in other CEE countries, see Abraham and Vodopivec (1995), Gora (1993), Jones and Kato (1993), Mickelwright and Nagy (1994) and Puhani (1996).

[3] In the early years, the expected massive withdrawal of women from the labour force actually took place more among women in older age groups (those who were eligible for old-age pension) than among women in their childbearing years. The labour force withdrawal of workers entitled to old-age pension was encouraged by government measures destined to combat unemployment. The departure of older workers from the labour force affected women more than men in part because female retirement age is five years lower than men's in most countries.

[4] These rates are based on the working age population of women (15-55 in the CEE countries).

[5] As we discuss below, non-recipients are individuals who are unemployed but do not qualify for unemployment benefits.

[6] We refer the reader to HST (1998) for more detail.

[7] Note that we are not trying to explain the difference in unemployment rates between the two republics on the basis of the UCS, and our analysis of the UCS is a separate contribution to paper from our investigation of factors underlying the difference in the unemployment rates between the republics. However, part of the difference in the unemployment rates can be explained by different responses to the UCS in each republic. Interestingly, we find that individuals in the CR are more responsive to increases in the UCS than those in the SR, and thus a reduction in benefits would actually increase the difference in the unemployment rates.

[8] One dollar was equal to 26-30 Kcs (Czechoslovak Crowns) in the 1991-93 period.

[9] See Terrell and Munich (1995) for a detailed description of the MLS.

[10] See HST for a discussion of the similarities between the different UCSs.

[11] There were 78 districts in the Czech Republic and 38 districts in the Slovak Republic at the time of this study.

[12] We do not include these individuals as their unemployment spell can be lengthened by the training period and their behaviour is likely to be different from other recipients. We cannot analyse them separately because we only have a small number in our sample.

[13] The men's sample is described in HST (1998).

[14] Good references on duration models are Flinn and Heckman (1982), Heckman and Singer (1984a), Kiefer (1988), and Lancaster (1990). Devine and Kiefer (1991) provide a comprehensive survey of previous empirical studies.

[15] Benefits also change when unemployment benefits are exhausted after 26 weeks of covered unemployment. As noted above, a single female qualifies for welfare, although her benefits depend on whether she lives at home. Whether a married woman qualifies, and the amount she receives, depends on her spouse's income. We cannot impute welfare benefits for either single or married women and thus we cannot exploit this variation.

[16] Prices and nominal wages rose by approximately 30 percent from the last quarter of 1991 to the second quarter of 1993, the period covered by our data (see Dyba and Svejnar, 1995).

[17] One reason that individuals register late for benefits is that individuals usually exhaust severance pay before collecting benefits. Other individuals simply wait to collect benefits; this phenomenon is similar to the less than full take-up of unemployment benefits in the United States (Anderson and Meyer, 1997).

[18] Using those who register late potentially complicates the econometric framework. See HST (1998).

[19] The industrial production variable is available only at an annual frequency. It is a price-weighted composite of total per-capita industrial production in the district in 1991 prices.

[20] In HST we found that the Schwartz criterion always lead to 2 support points for the male hazard function.

[21] To carry out the comparison of recipients and nonrecipients below, we also estimate the small model for recipients. For comparability with nonrecipients, we do not allow for unobserved heterogeneity when doing so.

[22] We calculate the expected durations at the mean of the explanatory variables, rather than calculating the mean expected duration across the sample (see, e.g., Ham and LaLonde, 1996), since we must use the sample mean values in the decompositions below.

[23] In what follows, we compare the female results to the male results in HST (1998). Since this latter paper has a

somewhat different specification of the hazard, there exists the possibility that our reported differences between men and women may be a result of the different specifications. To investigate this possibility, we re-estimated the male hazards in each republic using exactly the same specification as the hazard used for women in that republic. The results are contained in Appendix Tables A3 and A4 and readers can verify that the male results for the UCS variables do not change if we adopt the female specification of the hazard. We use these estimates in Appendix A3 in the Oaxaca decomposition between men and women below.

[24] Here we are using a one-sided test for single women.

[25] The entitlement elasticity is calculated as follows: Since a one week increase in entitlement would mean a 1/26 or 3.85% increase in entitlement, we divide the values presented in Table 4 (the duration effect from raising entitlement by one week) by 3.85. This gives us the effect of a 1% increase in entitlement on expected duration. We then divide this number by the relevant base duration and multiply by 100 to obtain the relevant elasticity.

[26] Columns (2) and (5) of these tables show the results for the recipients of using the smaller model that we use when we compare recipients to nonrecipients to obtain the infra-marginal effect of the UCS.

[27] One could of course address this question by using the estimates for recipients, and setting benefits, as well as entitlement, equal to zero. However, that approach would entail extrapolating these variables well beyond the experience of anyone in the recipient group, and raises some difficulties in separating duration dependence from entitlement effects.

[28] In this exercise, we are not controlling for possible unobserved differences between recipients and non-recipients. In order to do so, we would need a variable that affects recipient status but does not affect duration (see Ham and LaLonde, 1996). We have not been able to find such a variable.

[29] Note that we are using our sample mean values of the explanatory variables rather than the republic means. The former will differ from the latter since our sample is first stratified across districts. While it would be preferable to use the republic means, they are not available for this time period, and thus we use the sample means. We will address this issue in future work.

[30] We will not discuss these results since they are very similar to those discussed in our parallel study. We present them to enable the reader to compare them with those obtained in the parallel study.

[31] The Czech factories were older than those in the SR, but their locations were determined primarily by market forces, while the location of factories in the SR was determined primarily by central planning.