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Poverty, Labor Force Status and the Social Safety Net

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The first draft of this paper was written as a contractor's report for USAID by the US Bureau of the Census. Nevertheless, the views expressed below are the author's and should not necessarily be attributed to Census.

I. PRELIMINARY REMARKS

Section 1: Overview

The following report culminates the work undertaken in a larger study on income and labor force status commissioned by USAID in 1995. It explores populations at risk in Eastern Europe, and attempts to address three questions: Who are the poor? How much protection do these people receive from the social safety net? and How much might have to be spent to change their economic status? More specifically, a theoretical model linking the incidence of poverty to unemployment and/or being out of the labor force is posited and estimated using logistic regression. The estimates are conditioned on the presence of other known poverty related variables which enter in both confounding and effect modifying capacities. Results from this exercise feed into a quantitative policy assessment of the adequacy of current welfare programs for those populations identified by the model as being at risk.

When interpreting the poverty risk parameters, it is important to recognize that each result is relative to the definition of the poverty threshold and the choice of contrast group.

Throughout the paper, we try to use natural contrasts: for gender, it is female versus male. But there are cases, for example, region, where the contrast is less clearcut. Should one pick a specific locale or use some average which might correspond roughly to the nation? The same problem occurs for the education variable.

The definition of poverty also requires some commentary. We employ both an absolute and relative measures to generate our risk estimates. The first is an index of physical need as calculated and costed out by the relevant statistical agency. The second is not subsistence based *per se*. Rather, it reflects one's position in the income distribution relative to the median. By using both, we test the robustness of our models, and avoid getting into an ongoing methodological squabble between economists and public policy makers over the selection of the most appropriate threshold. Additionally, poverty is calculated where income has been standardized into adult equivalent units according to OECD and LIS practice. People, depending upon their age and living arrangements, are not weighted equally because there may be economies of scale in household consumption.

The report is divided into three chapters plus a brief concluding section. Methodological and data issues are addressed in chapter I, leaving country specific analyses and observations to the remaining three. LIS (Luxembourg Income Study) data sets for Poland (1992) and Hungary (1991) provide the raw material for the bulk of the investigation in chapters II and III. To facilitate cross-national comparisons, generic model structure and variable definition are imposed. Subsequent discussion follows the central theme of identifying/quantifying risks and exploring the impact of government actions aimed at lessening the associated burdens.

Section 2: Review of the Literature

The recent addition of East European household income and expenditure survey data to the Luxembourg Income Study (LIS) data base is a welcome, but long overdue event. Many of the standing questions about material welfare in this region can now be properly addressed. The efforts taken by the LIS team to compile and standardize income data according to accepted international statistical practice open up the prospect for meaningful cross-national comparisons. And, as a result, social scientists can begin to inject some "glasnost" into a sensitive area that has been shrouded in secrecy for almost half a century.

The work presented in the pages below explores this "terra incognito" but doesn't stray from the mainstream of the empirical poverty economics tradition. Precedent for our research focus and choice of methodology are ample and easy to document. A review of that literature over the past 10 years shows that there is growing consensus regarding the root causes/correlates of poverty, and how best to assess their impact. Practically all research is based upon an investigation of some subset of the following variables: age, labor force status, gender, educational attainment, family size and marital status, occupation/sector of employment, race/ethnicity and geographic locale (see: Tsakloglou 1990; Blackburn 1990; Van den Bosch et. al. 1993). These same studies routinely employ multivariate statistical procedures (logit regression, two and three way contingency tables etc.) to make inferences regarding the incidence of poverty (Klugman et. al. 1993; Casper et. al. 1994; Gornick and Pavetti 1990). Many control for measurement flaws by resorting to multiple definitions of the poverty threshold (Geary 1989; McGregor and Borooah 1991; Nolan and Callan 1989). In this respect, our work is not innovative; it travels down a familiar path. Nevertheless, this study contributes to the literature to the extent that its evidence and stylized facts validate the existing paradigm and deepen our understanding of poverty dynamics in economies undergoing wrenching structural adjustments.

Methodological continuity is evident in our decision to employ both absolute and relative specifications of the poverty thresholds. This is a standard response to measurement imprecision.

Given the profession's skepticism that any one metric accurately reflects the point at which poverty ensues, we explore several poverty thresholds in an attempt to bound our risk estimates. We acknowledge methodological criticisms raised by Sen and others and supplement this risk assessment with an investigation of the depth of poverty based on the average cash infusion needed to raise target households above the subsistence minimum.

Thematic continuity with previous studies on poverty in the transition economies of Eastern Europe is immediately apparent. Our assessment of economic risks in Poland and Hungary echoes many of the World Bank's important conclusions. The Bank's February 1995 study on poverty, inequality and social policy (Milanovic 1995) underscores the concern in the West over the high price tag attached to the transition; the particular vulnerability of the unemployed; and the unequal sharing of the costs and benefits across different social groups and regions. Tables 1, 5 and 7 below tell the same story. Similarly, the Bank's June 1995 report on poverty in Russia (Klugman et. al. 1995) reconfirms the risks encountered by those out of work as well as the precarious economic status of women and pensioners. By and large, differences in degree, not kind, separate the approaches. For instance, where quantitative estimates of the cost of eliminating poverty in Poland are provided, the monetary valuation of the poverty line, not the theoretical basis for defining the threshold, is the issue. As social scientists seeking to characterize broad demographic trends, these concerns are secondary; as budget conscious policy makers, they may occupy center stage.

Section 3: Data Issues

The data sets used to test the relationship between poverty and labor force status come from the LIS data base and national statistical agencies. The first two countries covered in this investigation are:

Hungary, as of 1991, and Poland, as of 1992. Micro survey data based on the responses of thousands of households have been processed by the LIS team into consistently defined socio-economic variables which serve as the serve as the regressors and regressand in the statistical analysis below. Our attention focusses on the following variables:

d1..... age of the head of the family
d3..... sex of the head of the family
d4..... number of persons in family
d6..... number of earners in the family
d7..... geographic location of the family
d10.... educational attainment of the family head
d16.... employment by sector of the economy
dpi.... net income after taxes
lfshd.. labor force status of head of the family
lfssp.. labor force status of spouse
foodexp..food expenditures
d18......type (status) of worker
d19......type (status) of worker-spouse

Section 4: Model Specification and Research Strategy

Prior BUCEN work has established the importance of examining the relationship between poverty and labor force status. The following research builds on this insight, and uses a statistical framework, which controls for the impact of other pertinent socioeconomic variables. One model that has been used successfully in similar investigations is the logit model (see: Gornick and Pavetti 1990). It is particularly well-suited to situations in which the dependent variable is dichotomous in nature. Our focus variable, poverty,[1] fits this description since survey respondents are either poor or not. Logit

[1] Household poverty is defined three ways in this investigation: 1) adjusted equivalent income below the adult minimum subsistence level (a physical measure of need specified by each country's relevant statistical agency), 2) adjusted equivalent income below 50 percent of the median adjusted equivalent, and 3) food expenditures greater than 50 percent of income. In the first two instances, adjustment for the number of people in the household is based on the square root of family size. In this way, misleading

models are appealing because they capture the probability of certain events occurring in terms of odds ratios, which are constrained to lie in a range between "zero" and "one." They also produce intuitive "S" shaped curves (ogives) that embody the combined effect of several risk factors on the likelihood of developing a certain condition (becoming poor).

The logit model developed in this paper attempts to establish a given household's odds of being in poverty, given that the head of the household or his/her spouse is not working, (other adult household members are not considered). The latter condition, "not working," in turn, arises from being unemployed or out of the labor force altogether. This risk will depend on a host of conditioning factors such as age, gender, family size, education, sector of employment, region, and so on. Subsets of these variables are included in both confounding and effect modifying (interaction) capacities, but final selection is based upon theory, precedent set in other studies, and limitations of the LIS data sets. The general form of the model being tested is given below in equation 1:

$$LogitP(X) \qquad {}_{i}E_{i} \quad {}_{i}\; {}_{i}V_{i}\; E_{i}\; {}_{j}\; {}_{j}W_{j}$$

where:

P(X) = probability of event X occurring
= baseline odds
$_i$ = coefficients of the exposure effect variables E_i
$_i$ = coefficients of the confounding variables V_i
$_j$ = coefficients of the effect modifying variables W_j

For purposes of estimation, the full model is described by equation 2:

comparisons based on per capita calculations are avoided by substituting a person's income and consumption standardized into adult equivalent units.

$$
\begin{gathered}
LogitP(X) \quad {}_0E_0 \quad {}_1E_1 \quad {}_2V_2 \quad {}_3V_3 \quad {}_4V_4 \quad {}_5V_5 \quad {}_6V_6 \quad {}_7V_7 \\
{}_8V_2V_3 \quad {}_9V_2V_4 \quad {}_{10}V_3V_5 \quad {}_{11}V_4V_6 \quad {}_{12}V_6V_7 \quad {}_{13}E_0V_2 \quad {}_{14}E_1V_2 \\
{}_{15}E_0V_3 \quad {}_{16}E_1V_3 \quad {}_{17}E_0V_4 \quad {}_{18}E_1V_4 \quad {}_{19}E_0V_5 \quad {}_{20}E_1V_5 \quad {}_{21}E_0V_6 \\
{}_{22}E_1V_6 \quad {}_{23}E_0V_7 \quad {}_{24}E_0V_2V_3 \quad {}_{25}E_1V_2V_3 \quad {}_{26}E_0V_2V_4 \\
{}_{27}E_1V_2V_4 \quad {}_{28}E_0V_3V_5 \quad {}_{29}E_1V_3V_5 \\
{}_{30}E_0V_4V_6 \quad {}_{31}E_1V_4V_6 \quad {}_{32}E_0V_6V_7
\end{gathered}
$$

where:

P(X) = probability of event X occurring (being in poverty)

= baseline odds

$_0$ = coefficient of exposure variable E_0 (unemployment)

$_1$ = coefficient of exposure variable E_1 (out of labor force)

$_2$ = coefficient of age variable V_2

$_3$ = coefficient of sex variable V_3

$_4$ = coefficient of education variable V_4

$_5$ = coefficient of family variable V_5

$_6$ = coefficient of region variable V_6

$_7$ = coefficient of occupation variable V_7

$_8 \cdots {}_{12}$ = coefficients of confounding variables which interact multiplicatively

$_{13}$ = coefficient of effect modifying variable E_0V_2 (unemployment x age)

$_{14} \cdots {}_{32}$ = remaining coefficients of effect modification

The model described above is the prototype from which the final equation is developed. To arrive at this destination, we referred to, and adopted where possible, a three part strategy which included:

1) variable specification, 2) assessment of interaction, and 3) assessment of confounding/precision.[2] What follows is a somewhat stylized account of the model refinement process.

Our selection of variables to be included is based upon preliminary analyses of the determinants of poverty found in the general economic literature and the findings contained in the first two Bureau reports to USAID/ENI. We have deliberately chosen the variables to provide the broadest possible meaningful model for initial consideration. Theory and prior research identify six ($V_2..V_7$) confounding variables that should be included as a matter of model validity, and hence should not be removed on the basis of tests of statistical significance since systematic, as opposed to random error, is involved. Further model elaboration allows for interaction between the exposure variables (see equation 2 variable list) and the V_i, and for subsequent testing and removal of such terms if the null hypothesis cannot be rejected at the =.05 level. As a matter of simplification, we do not entertain effect modification more complex than terms of the form EV_iV_j. This restriction facilitates interpretation of coefficients and reduces the likelihood of multicollinearity.

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[2] Neither the underlying theory nor the definition of variables allowed us to adhere to these guidelines 100 percent. The major problems were the lack of clear cut delineation between exposure and confounding variables, and the sacrifice of theoretical detail in light of the estimation problems posed by small data samples.

To illustrate the first problem, take the case of age and unemployment. From one perspective both variables can be treated as exposure variables: losing a job or, past a certain age, living entirely off of pension benefits might separately result in poverty. However, it is also possible to envision instances where the impact of age is indirect and mediated through the chances of finding employment following a layoff. As separate exposure variables, statistical tests can be applied to accept or reject the model's formulation. As an interaction effect, the age*unemployment variable can also be retained or deleted on statistical grounds. But this interaction formulation implies that age also functions as a stand alone confounder, and regardless of significance tests, must be included in the model for the latter to be hierarchically well formulated. Thus, an unresolvable semantic issue can affect the scope of a model and consequently, its representation of the odds ratios .

There were two major issues associated with limited sample size: getting the full model to converge and interpreting results where some parameter estimates were missing or regarded to be infinite. To overcome these hurdles, we did not adhere strictly to the model refinement strategy mentioned in the text. The reader is referred to footnote 5 for a more complete discussion of the tradeoffs and resolution.

We begin with a model that is hierarchically well formulated[3] (HWF), that is, all lower order components of any terms in the model are included. Given that the model is HWF, we refine the prototype using a hierarchical backward elimination procedure. This involves removing variables based on statistical tests if interaction is involved. Higher order terms are screened first, and if retention is warranted, then according to the hierarchical principal,[4] all of the relevant lower order components are kept. With the completion of the assessment of interaction, what remains of the initial model is variously referred to in the literature as the "gold standard" estimate. At this point, the question of precision becomes important. Note that the interaction assessment has been carried out prior to the evaluation of confounding because if there is strong evidence of interaction involving certain variables, then the latter assessment of confounding for these terms becomes irrelevant. However, for those cases in which retention is not mandated, and in which the odds ratios do not change appreciably as specific subsets of the V_i are dropped, a case can be made for further streamlining. The decision depends upon whether there are real gains in the precision of the estimates, for example, narrowing of the confidence intervals. It must be stressed that the final stage of model refinement is a somewhat subjective phase, and unless real gains are apparent, the safest choice is to retain all confounding variables so that validity does not become an issue.

[3] If the model is not HWF, then tests about variables in the model--in particular the highest order terms--will not be independent of the coding.

[4] The rationale behind this principal is the same as that governing HWF formulation, that is, tests about retention of lower order components should be independent of coding.

II. POLAND 1992

Section 1: Model Results

The theoretical model discussed in chapter I, section 4, serves as the blueprint for the empirical analysis below. Early on, it became apparent that certain statistical problems would prove to be intractable and that modifications would be needed to produce meaningful results. Not surprisingly, the full model could not be fit because there were too few sample observations for the number of parameters involved. To get the parameter estimates to converge, we restricted the order of interaction to two-factor product terms of the form $E*V_i$. This simplification solved the issue of size, and allowed us to begin the process of model refinement using Chi squared statistics found in the analysis of variance table. Subsequent tests showed that the vast majority of interaction terms were not significant at the = 0.10 level or better.[5] This occasioned their removal and left a reduced model with two exposure variables and multiple confounders. The final refinements involved the removal of the occupation variable on grounds of logical inconsistency;[6] the substitution of the family

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[5] When choosing between two competing logit models, the decision to reject depends on whether the addition/deletion of some explanatory variable(s) contributes to the model's overall statistical validity. Both the log likelihood function and the Wald test measure this, and are distributed Chi square.

Two of the interaction models identified the following statistically significant cross- product terms: "dispen*d7" and "d1*sex." Tables 1 and 3 do not report these effects because the equations producing them were, in other respects, flawed. First, the inclusion of "dispen*d7" created a situation where sample size became a limiting factor and several of the parameters could not be estimated. The only way around that problem would have been to combine geographic regions and suffer the subsequent loss of important policy detail. It was believed that such a sacrifice was not warranted, given that the noninteracting parameter estimates were substantially the same as those found in the models without the terms "dispen*d7" and "d1*sex." The second concern with these more complete models was the sign reversal associated with the "sex" variable in the presence of interaction between age and gender. There is simply no evidence that being female improves one's chances of staying out of poverty, *ceteris paribus*. All of the data and subsequent analysis suggest the opposite.

[6] Neither the occupation nor industry variables produced the expected results. Regardless of profession, each and every respondent group in the sample was less likely to be poor than the sample taken as a whole. Under normal circumstances, one would expect to see differential rates of poverty across families whose heads were affiliated with industries/occupations, which were

size variable (d4) for the number of earners (d6), again on plausibility grounds,[7] and the selection of d1 to reflect age effects as opposed to dividing the sample into families with heads above or below the age of 60.[8]

Table 1 reports the poverty parameter estimates for models using both an absolute (identified as Model SUBG1) and a relative (median) threshold for poverty (SUBG3). The reader is referred to footnote 1 for definitions of the poverty threshold.

Table 1. Poverty Parameter Estimates, Model SUBG1

Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
UNEMP	0.5537	0.1911	8.39	1	0.0038	1.7397
DISPEN	0.3722	0.1291	8.31	1	0.0039	1.4509
AGE(D1)	0.0237	0.00378	39.10	1	0.0000	1.0240
FAMSIZE(D4)	0.3133	0.0362	74.73	1	0.0000	1.3679
SEX	0.9933	0.1163	73.00	1	0.0000	2.7001
EDCOLL	-2.6018	1.2738	4.17	1	0.0411	0.0741
EDHS	-1.1763	1.2785	0.85	1	0.3575	0.3084
EDGRM	0.5924	1.0591	0.31	1	0.5759	1.8083
REGION(D7)	---	---	87.17	8	0.0000	---

gaining or shedding jobs during the economic transition. It is possible that the coding of industry affiliation precludes those who are unemployed from identifying themselves with their former employers. If that is so, then having an "occupation" is synonymous with being employed and perforce would reduce the risk of poverty, unless wage scales were severely depressed or highly skewed.

[7] Although d6 is statistically significant, its sign is positive. Such a result is counter- intuitive since it implies that as families provide more workers their risk of poverty rises. Causality under these circumstances is unlikely; on the other hand, poor families may opt for higher labor force participation rates to cope with poverty.

[8] The model was estimated with the age of the household head defined as both a continuous and a categorical variable. In the latter instance, a threshold of 60 years of age was used to divide the population into two groups. On statistical grounds, both the continuous and categorical age variables were significant at the =.05 level or better. However, since the model already included a variable "dispen," which was designed to capture the risk of poverty associated with being out of the labor force (that is, retirement, disability, and so on), it was believed that the categorical variant was redundant.

Table 1. Poverty Parameter Estimates, Model SUBG1

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
CAPITAL(1)	-0.4751	0.1447	9.97	1	0.0016	0.6218
NORTH EAST(2)	0.2707	0.1577	2.95	1	0.0861	1.3109
NORTH(3)	0.0144	0.1305	0.01	1	0.9120	1.0145
SOUTH(4)	-0.7688	0.1332	33.31	1	0.0000	0.4636
SOUTH EAST(5)	0.4366	0.1020	18.32	1	0.0000	1.5474
CENTRAL EAST(6)	0.7270	0.1372	28.09	1	0.0000	2.0689
CENTRAL(7)	-0.0546	0.1431	0.15	1	0.7030	0.9469
CENTRAL WEST(8)	0.1362	0.1111	1.50	1	0.2201	1.1459
SOUTH WEST(9)	-0.3044	---	---	-	---	0.7376
CONSTANT	-5.5875	1.0993	25.83	1	0.0000	0.0037

Table 2. Summary Statistics for Model SUBG1

	Chi-Square	d.f.	Significance
-2 Log Likelihood (-2LL)	3516.65	6586	0.0000

Table 3. Poverty Parameter Estimates, Model SUBG3

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
UNEMP	0.5904	0.1964	9.04	1	0.0026	1.8047
DISPEN	0.3785	0.1347	7.89	1	0.0050	1.4601
AGE(D1)	0.0235	0.00395	35.36	1	0.0000	1.0238
FAMSIZE(D4)	0.3319	0.0372	79.51	1	0.0000	1.3936
SEX	1.0063	0.1222	67.85	1	0.0000	2.7355
EDCOLL	-3.2905	1.4578	5.10	1	0.0240	0.0372
EDHS	-1.1664	1.2795	0.83	1	0.3620	0.3115
EDGRM	0.4806	1.0603	0.21	1	0.6504	1.6170
REGION(D7)	---	---	83.39	8	0.0000	---
CAPITAL(1)	-0.4185	0.1511	7.67	1	0.0056	0.6580
NORTH EAST(2)	0.2399	0.1668	2.07	1	0.1504	1.2711

Table 3. Poverty Parameter Estimates, Model SUBG3						
Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
NORTH(3)	-0.0622	0.1410	0.19	1	0.6590	0.9397
SOUTH(4)	-0.8021	0.1433	31.33	1	0.0000	0.4484
SOUTH EAST(5)	0.4978	0.1051	22.45	1	0.0000	1.6451
CENTRAL EAST(6)	0.7187	0.1432	25.18	1	0.0000	2.0518
CENTRAL(7)	-0.0709	0.1518	0.22	1	0.6405	0.9316
CENTRAL WEST(8)	0.1677	0.1155	2.11	1	0.1465	1.1826
SOUTH WEST(9)	-0.2704	---	---	-	---	0.7631
CONSTANT	-5.6612	1.1031	26.34	1	0.0000	0.0035

Table 4. Summary Statistics for Model SUBG3			
	Chi-Square	d.f.	Significance
-2 Log Likelihood (-2LL)	3259.21	6586	0.0000

The statistical validity of the first poverty model is strongly supported by the log likelihood and Wald tests reported in Tables 1 and 2. The log likelihood statistic,[9] comparing the fitted equation with 18 parameters to the baseline case (null hypothesis) where there is only a constant, indicates that the null hypothesis can be rejected at the = 0.0000 level or better. With regard to individual coefficients, the signs are all plausible, and 13 of 18 are significant at the =.05 level or better.[10] The five

[9] SAS, the software estimation package employed in this research, also prints out a likelihood ratio test which compares the fitted model to the so called "saturated" or "perfect" model. Despite the promise of such terminology, the test provides little analytical insight and guidance because the underlying calculations merely refer to a hypothetical situation in which there is one parameter for each cell in the multidimensional contingency table. Unlike classical regression analysis, logistic regression does not produce goodness of fit statistics that are unambiguous and universally accepted. Thus, to avoid convoluted write-ups and other unnecessary confusion, especially when referring to common equations in cross-national perspective, we do not report saturated model results.

[10] These are all two-tailed tests.

variables whose Wald test scores are above the threshold are all retained to preserve the analytical unity of the underlying confounding variables.[11]

All of the Bureau's prior research on poverty in Eastern Europe points to the prominence of labor force status. Those who are unemployed (UNEMP) have poverty risks 74 percent higher than the reference group.[12] For those who are out of the labor force (DISPEN), the risk of poverty is almost 45 percent higher than those who are in.

The interpretations of the "age" and "famsize" odds ratios are consistent with prior expectations: as the household gets larger or as the head matures, poverty risks increase.[13]

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[11] SAS treats the regional variable d7 as categorical, and produces direct estimates for only eight of the nine regional parameters. The ninth parameter is defined as the negative sum of the first eight. Under these circumstances, all regions are bound together as an integral unit. For mechanical reasons, nonsignificant parameters cannot be dropped selectively without first redefining d7 as nine separate variables, and then specifying the appropriate contrast groups for testing the null hypothesis. This latter revision is not obvious. But, since the Chi square test for the regional variable as a unit is highly significant (=0.0000), there is adequate statistical justification for keeping it intact. We believe that retention of the nonsignificant parameters is useful inasmuch as it permits us to identify which regions have average risks when compared to the nation as a whole.

We include and display all three design elements of the educational variable (edcoll--edhs--edgrm), even though two have marginal Wald test scores. To do otherwise would imply that we arbitrarily recoded the covariate. It turns out that purging "edhs" and "edgrm" does produce a significant difference in the log likelihood ratio. It is also noteworthy that parameter estimates are only minimally affected by the choice to retain or delete: there are no sign reversals and the quantitative shifts are all in the second or third decimal place.

[12] The reference group is every household in the sample whose head or spouse was not unemployed.

[13] The logistic regression which included "d27," the number of children under age 18, had a negative coefficient. This result was simply not credible, and subsequent analysis focused solely on family size (d4) without any attempt to isolate the economic burden of young children.

The role of gender and economic status was touched on in previous Bureau reports. Three specific issues were identified: women's unemployment rates were higher and of longer duration than men's; women constituted significantly more than half of the population over the age of 65; single parent families (usually headed by women) had above average rates of poverty. These concerns find expression in the "sex" coefficient: being a household headed by a female raises the risk of poverty by 170 percent over the contrast group.

Education appears to play a decisive role for those on the margins. Households whose heads have schooling beyond high school are only one-thirteenth as likely to be below the poverty line than the population at large when all levels of education are considered. There is also some weak evidence for a similar effect at the secondary level. However, neither the "edhs" nor "edgrm" coefficients had Wald tests which were significant at the = .10 level or better.

Geographic location is the final confounding variable in the model. Because of the manner in which SAS creates and codes dummy instruments, negative parameter estimates indicate decreased peril in the region of interest relative to the average for all regions combined. Thus, we find that poverty risks are highest in the Central East region (107 percent above the contrast group) and lowest in the South (54 percent below the contrast group). These results are broadly consistent with Team Polands estimates for 1993, and also concur substantially with the results found in the first Bureau report to USAID/ENI.[14]

[14] The Polish team and the Bureau logit models agree on the rating of the capital, central eastern, south eastern, and south western regions. There is also common ground between the current work and the November 1994 Bureau report with regard to the central east, central west, and northern regions.

Discussion of the second model SUBG3 (relative poverty threshold) is abbreviated because the results are virtually identical to those found for the absolute measure of poverty (SUBG1). As before, the log likelihood test of goodness of fit indicates that the null hypothesis can be rejected at the =0.0000 level or better. But more importantly, the signs, magnitudes, and significance of the parameter estimates are not materially affected by the redefinition of poverty[15] and this increases our confidence in the robustness of the model and estimation procedures.

Section 2. Policy and the Social Safety Net: a Simulation

a. What the Net Does

Identification of populations in poverty and at risk is logically prior to the evaluation of how well any given society copes with the misfortunes of its most vulnerable citizens. For purposes of analysis, the factors which produce adverse economic selection must be represented in a way which links risk identification quantitatively to remedial policy initiatives. This section makes the connection by highlighting how the social safety net affects economic status. It is an exercise that is counterfactual in nature: poverty rates and odds ratios are computed for Models SUBG1 and SUBG3 where social

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[15] Aside from the fact that parameter estimates differ in the second and occasionally first decimal places, the models are interchangeable.

transfers have been netted out of the income stream, and there has been no offset in taxes. Subsequent comparison of these estimates to those presented in the last section reveals how much protection, as measured by the decline in national poverty rates and odds ratios, has been provided.

As one would expect, shredding the social safety net is a dramatic act. National poverty rates, depending on the threshold selected, jump from 8.56 and 7.71 percent (see Tables 7 and 8) to 44.97 and 44.35 percent, respectively. For the exposure variables, Tables 5 and 6 indicate that in the absence of social transfers, the risk of poverty rises substantially for those out of work and those out of the labor force altogether. Regardless of the definition of the poverty level, the negative consequence of losing one's job as measured by the odds ratios rises by over 200 percent.[16] For those out of the labor force, the repercussions are likely to be harsher, with increases of at least 350 percent. Presumably, the actual provision of unemployment compensation and pension benefits mitigates these worst effects.

By way of contrast to the enhancement of the exposure effects, the impact of benefit recision on the demographic confounding variables is generally towards dilution. Both models present reduced risk in 10 of 15 remaining coefficients (constant term excluded). Our evidence suggests that gender and education impacts are less pronounced. The additional risk of being poor because one is a woman

[16] From 1.7397 to 6.1910 or 445 percent (SUBG1), and from 1.8047 to 4.1099 or 231 percent (SUBG3).

drops from 170 percent to 14 percent or less.[17] At face value, this implies that social transfers go further towards eliminating poverty for males than females.

Education at the post secondary level appears to continue conferring an advantage, albeit a reduced one, judging by the slight shift in the odds ratios. This is somewhat puzzling because there is no organic linkage between higher education and access to, and qualification for, social benefits. It is possible that jobs which require advanced training have better benefit packages, or it may be that the better educated people are more adept at manipulating the bureaucracy. Less can be said about households whose heads left school before college: there are conflicting signals about the direction of the impact and not all of the results are statistically significant.

To gain another perspective on parameter shifts, the following filtering criteria are applied: coefficient revisions are considered meaningful if statistical significance is retained pre- and post-recision, and there is agreement between the models on the direction of the impact. By these standards, age and certain regional risks (found in the capital and south) are enhanced, while diminution occurs in the south east and central east.

Finally, the counterfactual experiment indicates that the existing regional distribution of social benefits may be biased. In the absence of transfers, geographic variations in poverty risks are either equalized

[17]Using the relative definition of poverty (SUBG3), the discrimination effect disappears completely. However, the result is not statistically significant at even the =.10 level. Thus, it seems safer to conclude that gender effect is still present but muted.

(odds ratios move closer to 1 and remain statistically significant) or eliminated outright in the sense that the null hypothesis of uniform regional coefficients cannot be rejected.

Table 5. Poverty Parameter Estimates, Model SUBG1 (no soctrans)

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
UNEMP	1.8231	0.1461	155.77	1	0.0000	6.1910
DISPEN	1.7499	0.0951	338.30	1	0.0000	5.7540
AGE(D1)	0.1080	0.00336	1033.36	1	0.0000	1.1140
FAMSIZE(D4)	0.0928	0.0307	9.17	1	0.0025	1.0972
SEX	0.1344	0.0880	2.33	1	0.1267	1.1439
EDCOLL	-2.4839	0.7499	10.97	1	0.0009	0.0834
EDHS	-1.9765	0.7876	6.30	1	0.0121	0.1386
EDGRM	-1.3272	0.7394	3.22	1	0.0727	0.2652
REGION(D7)	---	---	68.90	8	0.0000	---
CAPITAL(1)	-0.1784	0.0932	3.67	1	0.0555	0.8366
NORTH EAST(2)	0.1067	0.1295	0.68	1	0.4100	1.1126
NORTH(3)	-0.2323	0.1006	5.33	1	0.0210	0.7927
SOUTH(4)	-0.4104	0.0770	28.42	1	0.0000	0.6634
SOUTH EAST(5)	0.4135	0.0876	22.30	1	0.0000	1.5121
CENTRAL EAST(6)	0.4546	0.1327	11.74	1	0.0006	1.5755
CENTRAL(7)	-0.0983	0.1077	0.83	1	0.3615	0.9064
CENTRAL WEST(8)	0.1227	0.0909	1.82	1	0.1770	1.1305
SOUTH WEST(9)	-0.1781	---	---	-	---	0.8369
CONSTANT	-5.9654	0.7646	60.87	1	0.0000	0.0026

Table 6. Poverty Parameter Estimates, Model SUBG3 (no soctrans)						
Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
UNEMP	1.4134	0.1960	52.02	1	0.0000	4.1099
DISPEN	1.6182	0.1170	191.39	1	0.0000	5.0440
AGE(D1)	0.1173	0.00391	900.94	1	0.0000	1.1245
FAMSIZE(D4)	-0.1382	0.0388	12.72	1	0.0004	0.8709
SEX	-0.0389	0.0909	0.18	1	0.6688	0.9618
EDCOLL	-0.5504	1.0514	0.27	1	0.6006	0.5767
EDHS	-0.1901	1.0849	0.03	1	0.8609	0.8269
EDGRM	0.2949	1.0432	0.08	1	0.7774	1.3430
REGION(D7)	---	---	33.19	8	0.0001	---
CAPITAL(1)	-0.2031	0.0979	4.30	1	0.0381	0.8162
NORTH EAST(2)	-0.0438	0.1469	0.09	1	0.7655	0.9571
NORTH(3)	-0.1087	0.1085	1.00	1	0.3167	0.8970
SOUTH(4)	-0.1160	0.0820	2.00	1	0.1570	0.8905
SOUTH EAST(5)	0.3410	0.0953	12.80	1	0.0003	1.4064
CENTRAL EAST(6)	0.4689	0.1464	10.26	1	0.0014	1.5982
CENTRAL(7)	-0.0923	0.1157	0.64	1	0.4248	0.9118
CENTRAL WEST(8)	0.0271	0.1003	0.07	1	0.7873	1.0275
SOUTH WEST(9)	-0.2731	---	---	-	---	0.7610
CONSTANT	-8.0722	1.0856	55.28	1	0.0000	0.0003

b. What Remains to be Done

The above discussion poses the question of what can be done to improve the well being of those who fall through the cracks in the social safety net. Is it possible to quantify how much money needs to be spent to raise the most vulnerable populations out of destitution? The answer is yes, if one has confidence in the representivity of the data.

To generate the estimates, we assume: 1) the absolute and relative poverty measures used in the analysis are robust, 2) sample sizes at the regional level are large enough to make general, but not precise, inferences about rates and magnitudes of poverty,[18] 3) recent trends in average family size follow the pattern prevailing over the decade 1978-1988, and 4) the comparative burden of poverty for those who have lost their jobs, or are out of the labor force altogether, is roughly proportional to their respective number of cases.

Table 7 reports the findings:

[18] Conversations with members of the Polish team indicated that the target sample was randomly drawn and representative of the nation. Our limited efforts to verify their contention are supportive, but there is not enough information regarding the sample standard errors or criteria for selecting precision thresholds to comment about sample size and representativeness at the subnational level.

Table 7. Poverty and Regional Need

Region	Population Weights (1)	Poverty Rate in Households (2)	Average Poverty Gap in Adult Equivalent Units (000 ZL) (3)	Average Number of Adult Equivalent Units (4)	Regional Poverty Weight (need in %) (5)*	Social Transfers by Region (allocation %) (6)	Excess of Need over Allocation in % (5÷6) (7)
CAPITAL	0.1171	0.0590	3335.1	1.3910	0.0716	0.0715	1.0013
NORTH EAST	0.0634	0.1192	2332.9	1.6341	0.0644	0.0694	0.9280
NORTH	0.1006	0.0992	2784.0	1.5667	0.0973	0.1057	0.9205
SOUTH	0.1860	0.0438	2214.5	1.3704	0.0553	0.0948	0.5833
SOUTH EAST	0.1400	0.1488	3295.5	1.5592	0.2393	0.2164	1.1058
CENTRAL EAST	0.0621	0.1923	4091.2	1.5002	0.1638	0.1100	1.4890
CENTRAL	0.0870	0.0913	2801.4	1.4374	0.0715	0.0909	0.7866
CENTRAL WEST	0.1370	0.1142	2961.2	1.6444	0.1703	0.1513	1.1255
SOUTH WEST	0.1067	0.0731	2474.8	1.5416	0.0665	0.0900	0.7389
NATION	1.0000	0.0856	2859.3	1.7723**	1.0000	1.0000	1.0000

* Each poverty weight is formed as the row product of (1)*(2)*(3)*(4) divided by the sum of the products across all nine regions. Given a few caveats, this should produce roughly the same distribution as the calculation based upon the average poverty gap per household summed across all households and regions.

** The average number of equivalent adults per household is higher for the nation than for any of the constituent regions. This is a consequence of the rounding algorithm used to sort the data. It makes no difference for the relative poverty weights in column (5) since the calculation would be invariant to a common scaling factor(1.7723÷1.5053) where the divisor in the latter is the population weighted average of column 4.

Given that the national average poverty gap per household is 5,067,669 zlotys and there are 1,042,603 households,[19] government authorities would have to spend almost 5.3 trillion zlotys to eradicate poverty for 1 year. At an exchange rate of 13,626 zlotys to the dollar, this translates into

[19] Calculated as the estimated number of households (12,182,770) times the national poverty rate reported in Table 7 (0.0856).

$387.8 million, or less than 0.23 percent of GDP.[20] When the median definition of the poverty threshold is used, the cash infusion necessary to eradicate destitution falls slightly to $365 million, and the calculation of regional need, as a percent of the total, is similarly unaffected (column 5). It should be noted that such requirements are not synonymous with the proportion of the population which resides in a given area. For example, while nearly 19 percent of the population lives in the south, its "need" is less than 6 percent of the country total. It is also clear that regional need does not coincide with the regional share of social transfer payments going to the poor (column 6). Again, the south has under 6 percent of the need but over 9 percent of the transfers. Column (7) indicates where potential regional inequities exist, in the sense that an unfair distribution would generate needs-allocation ratios significantly different than "1." By this standard, the south east, central east, and central west are getting shortchanged, under the absolute definition of poverty. For the median standard (Table 8), the list widens to include the north and the capital. Further evidence of this imbalance is found in Tables 1 and 3. With the exception of the capital, all of the regions mentioned have poverty odds ratios above the national average.

The final comments about economic hardship reflect how the burdens vary with labor force status. Preliminary estimates[21] based upon both absolute and median thresholds suggest that less than 10

[20] See the World Factbook 1993, pp. 313-314. Using purchasing power parities, the CIA estimates that the Polish GDP was $167.6 billion in 1992.

[21] Early on, the decision was made to treat the household as the basic unit of analysis. The choice imparted coherence to the treatment of families with dependents and with multiple sources of income. But it also posed problems with the classification of the household depending upon the differing economic circumstances of the head and the spouse. Cross-tabulations of the data showed that poverty in a given household could either be attributed to unemployment or being out of the labor force, depending upon the variables chosen and the order in which the cross-tabulations were performed. Several solutions presented themselves, but rather than make an arbitrary assignment based upon income level or some other equally plausible characteristic, we let the

percent of all poverty cases can be associated with unemployment, while up to 75 percent can be attributed to being out of the labor force. The remaining 15-20 percent is composed largely of working poor (unpaid farm-workers, and so on).

Table 8. Poverty and Regional Need

Region	Population Weights (1)	Poverty Rate in Households (2)	Average Poverty Gap in Adult Equivalent Units (000 ZL) (3)	Average Number of Adult Equivalent Units (4)	Regional Poverty Weight (need in %) (5)*	Social Transfers by Region (allocation %) (6)	Excess of Need over Allocation in % (5÷6) (7)
CAPITAL	0.1171	0.0529	3324.2	1.4286	0.0764	0.0683	1.1186
NORTH EAST	0.0634	0.1006	2328.1	1.5833	0.0610	0.0750	0.8133
NORTH	0.1006	0.0805	3002.7	1.6400	0.1036	0.0945	1.0962
SOUTH	0.1860	0.0366	2248.0	1.3696	0.0544	0.0925	0.5881
SOUTH EAST	0.1400	0.1310	3145.8	1.5714	0.2354	0.2296	1.0253
CENTRAL EAST	0.0621	0.1530	4215.4	1.5370	0.1599	0.1070	1.4944
CENTRAL	0.0870	0.0772	2868.7	1.4762	0.0739	0.0920	0.8033
CENTRAL WEST	0.1370	0.1000	2881.4	1.6627	0.1704	0.1551	1.0986
SOUTH WEST	0.1067	0.0660	2370.5	1.5000	0.0650	0.0861	0.7549
NATION	1.0000	0.0771	2988.0	1.7723**	1.0000	1.0000	1.0000

indeterminacy remain. For this reason, the unrefined numbers do not produce mutually exclusive categories, and the proportions must be considered tentative.

* Each poverty weight is formed as the row product of (1)*(2)*(3)*(4) divided by the sum of the products across all nine regions. Given a few caveats, this should produce roughly the same distribution as the calculation based upon the average poverty gap per household summed across all households and regions.

** The average number of equivalent adults per household is higher for the nation than for any of the constituent regions. This is a consequence of the rounding algorithm used to sort the data. It makes no difference for the relative poverty weights in column (5) since the calculation would be invariant to a common scaling factor(1.7723÷1.5206) where the divisor in the latter is the population weighted average of column 4.

III. HUNGARY 1991

Section 1: Model Results

The comments in the previous section about Polish model refinement fully apply to the Hungarian exercise. With only 2,019 observations, the Hungarian sample is not even a third as large as the Polish, and the task of estimating a complete model with complex interactions becomes practically impossible. Simplification begins with the removal of the education and occupational variables. Both have Chi square test statistics that are not significant at the

= 0.10 level or better. As in the Polish case, D6, the number of wage earners in the household has a counter-intuitive sign and is dropped from further consideration. Even with the model pared down to a reduced set of confounding and exposure effects, the samples are still too small to yield reliable estimates for both the "unemployment" and "dispen" parameters in the same equation.[22] But, in separate logit regressions, finite values can be assigned as long as one of the exposure variables is dropped from consideration. And, as it turns out, the choice of which to omit is simple because the "dispen" variable is either nonsignificant or has the wrong sign *a priori*. Final adjustments produce two acceptable models that vary only in their treatment of the age. We report only one, the continuous age variant, because its important properties are common to its categorical twin. The only statistical difference of note is the magnitude of the age effect. As expected, when age is treated as a continuous variable, poverty risks increase gradually, growing from 1-3 percent for each additional year. When the categorical case is considered, the contrast of elder versus young produces risk increments between 60 and 265 percent.

Table 9. Poverty Parameter Estimates, Model SUBG1

Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
UNEMP	1.4228	0.6090	5.46	1	0.0195	4.1487
D1	0.0327	0.00375	76.17	1	0.0000	1.0332
FAMSIZE(D4)	-0.1891	0.0427	19.58	1	0.0000	0.8277
SEX	1.2444	0.1321	88.76	1	0.0000	3.4709
REGION(D7)	---	---	82.64	4	0.0000	---
FARMSTEAD(1)	0.8669	0.3831	5.12	1	0.0236	2.3795
VILLAGE(2)	0.3966	0.1232	10.36	1	0.0013	1.4868
TOWN/CITY(3)	-0.1473	0.1276	1.33	1	0.2485	0.8630

[22] SAS printouts indicate that both parameter estimates should be considered infinite.

Table 9. Poverty Parameter Estimates, Model SUBG1						
Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
MAJOR CITY(4)	-0.1982	0.1480	1.79	1	0.1806	0.8202
BUDAPEST(5)	-0.9180	---	---	---	---	0.3993
CONSTANT	-1.5730	0.2977	27.92	1	0.0000	0.2074

Table 10. Summary Statistics for Model SUBG1			
	Chi-Square	d.f.	Significance
-2 Log Likelihood (-2LL)	2302.42	1987	0.0000

The statistical validity of the first poverty model is strongly supported by its goodness of fit and Wald tests reported in Tables 9 and 10.[23] The log likelihood statistic comparing the fitted equation with nine parameters to the baseline case where there is only a constant indicates that the null hypothesis can be rejected at the = 0.0000 level or better. With regard to individual coefficients, the signs are all plausible, and seven of nine are significant at the =.025 level or better. The two regional coefficients whose Wald test scores are above the 0.10 threshold are kept in the model with the

[23] Unlike the Polish case, there was enough data to model the connection between poverty and labor force status for all three definitions of subsistence. Unfortunately, the equations where food expenditures were used to set the threshold fit the data very poorly, so they are not reported.

remaining three so that the treatment of geography is unified and consistent with established political subdivisions.

Interpretation of individual coefficients is, for the most part, straightforward. It appears that the nature of the correlations between poverty and specific exposure and confounding variable effects are generally preserved in cross-national comparisons. As before, labor force status is significant. For the unemployed (UNEMP), poverty risks are 315 percent above the reference group. Again, aging enhances the risks of poverty. The majority of the population can expect a modest decline in economic well being as the household head matures. Over the life-cycle, each year raises the risk of poverty by roughly 3 percent.

The social dynamics linking family size to economic well-being appear to operate differently in Hungary than in Poland. The risk of poverty in Hungary decreases by just over 17 percent for each additional member of the household, while in Poland, adding members increases the risk by almost 40 percent. Evidence cited in previous Bureau reports[24] indicates that the demographic composition of the typical household in the two countries varies with regard to age structure, support ratios, and labor force participation rates. These factors, acting separately or in concert with different levels of protection from the social safety net, could account for the sign reversal in the family size coefficient. It may be that the expansion of the household sets in motion two opposing forces: one which improves income prospects as more members work, and the other which tends to dampen living standards because of greater dependency burdens. Perhaps where the

[24]Bureau of the Census, November 9, 1994 and February 23, 1995.

balance is struck determines the sign of the coefficient, but further research is necessary to resolve the issue.

The final two variables cover gender and locale. Again, adverse economic selection appears to be a problem for women. The "sex" coefficient indicates that being female raises the risk of poverty by 247 percent over the contrast group (men). Geographic location also exerts a negative economic effect on farmsteads and villages. These parameter estimates imply poverty risks from 49 to 138 percent above the average for all regions combined.

Review of the second model (SUBG3) illustrates the need to assess the sensitivity of parameter estimates to the definition of the poverty threshold. As before, the log likelihood goodness of fit test indicates that the baseline model should be rejected in favor of the more complete version with exposure and confounding effects. More importantly, labor force status, age, and sex retain their signs and significance. But the two models part company over the role of family size and place of residence.

Table 11. Poverty Parameter Estimates, Model SUBG3

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
UNEMP	1.8719	0.7226	6.71	1	0.0096	6.5006
D1	0.0104	0.00621	2.81	1	0.0935	1.0105
FAMSIZE(D4)	-0.0576	0.0749	0.59	1	0.4414	0.9440
SEX	1.1168	0.1934	33.34	1	0.0000	3.0551
REGION(D7)	---	---	16.39	4	0.0022	---

Table 11. Poverty Parameter Estimates, Model SUBG3						
Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
FARMSTEAD(1)	0.3722	0.6110	0.37	1	0.5425	1.4509
VILLAGE(2)	0.4425	0.1969	5.05	1	0.0246	1.5566
TOWN/CITY(3)	-0.1750	0.2167	0.65	1	0.4193	0.8395
MAJOR CITY(4)	-0.5564	0.2703	4.24	1	0.0396	0.5733
BUDAPEST(5)	-0.0833	---	---	---	---	0.9201
CONSTANT	-3.2864	0.5124	41.13	1	0.0000	0.0374

Table 12. Summary Statistics for Model SUBG3			
	Chi-Square	d.f.	Significance
-2 Log Likelihood (-2LL)	1049.24	1987	0.0000

Section 2. Policy and the Social Safety Net: a Simulation

a. What the Net Does

As in Poland, the provision of social transfers appears to insulate major segments of the population from economic deprivation. Without such benefits, poverty rates of 8.57 and 45.12 percent (see Tables 15 and 16) could jump to 45.86 and 68 percent, respectively.

But the Hungarian experience with supporting its most vulnerable citizens during the transition to a market economy may, in other respects, be unique. For households where one or more of its members is unemployed, simulated cross-national experience does not validate the inference that removing the social safety net increases poverty risks, relative to the contrast group. Tables 13 and 14, report the parameter estimates for models SUBG1 and SUBG3 where age is a continuous variable.[25] Note that both the size and statistical significance of the unemployment coefficient are reduced when social transfer payments are netted out of income: the odds ratios drop from 9 to 28 percentage points depending upon the choice of poverty threshold, and the accompanying Wald tests are no longer significant at the = 0.025 level. While all households would undoubtedly experience greater economic strain under these circumstances, the risks of falling into poverty increase comparatively less for the unemployed than for those who are currently in jobs. Put differently, there may be inherent biases in existing benefit programs which are prospectively revealed by the termination of government assistance. In effect, the incidence of risk burden shifts marginally against people out of work when social benefits are comprehensive and provided universally. Though the unemployed may be helped by the social safety net in a very tangible way, their relative risk position nonetheless deteriorates. It is hard to imagine how this could occur unless unemployment compensation were a minor component of total transfer benefits, while the latter was a large fraction of every household's income stream.[26] The question of programmatic bias is an item that should be

[25] Neither of the models with age as a categorical variable is well estimated.

[26] It is also possible that the results obtained reflect a flaw in data construction. Hungary's household panel survey does not allow one to identify every person who would be classified as unemployed in a western sense. While the labor force status variable singles out those receiving unemployment compensation and first time job seekers, it does not track those who do not receive such income support even though they are out of work. The situation is different in Poland: their labor force status variable follows western conventions in identifying the unemployed.

put on the research agenda for the upcoming country teams report on the scope and administration of welfare.

It is less clear how the mechanism that links the remaining confounding variables in the model to the provision of social welfare benefits produces shifts in poverty risks. On the one hand, a consistent story can be told about gender and age when there is no social safety net. Both SUBG1 and SUBG3 models indicate that the magnitude and statistical significance of the sex coefficient is diluted. This may mean that benefit programs largely or entirely aimed at improving the economic well being of women and their dependents (family allowances, maternity benefits, and so on), do not level the playing field. There is also concurrence regarding the role of age. Left to its own devices, the household becomes less economically secure as its breadwinners mature. However, when it comes to family size and region, variations in the definition of poverty produce contrary motion in the odds ratios. For these variables, no inferences about the safety net are drawn.

Table 13. Poverty Parameter Estimates, Model SUBG1 (no soctrans)

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
UNEMP	1.3148	0.6339	4.30	1	0.0381	3.7240
D1	0.0427	0.00400	114.48	1	0.0000	1.0436
FAMSIZE(D4)	-0.0121	0.0405	0.09	1	0.7658	0.9880
SEX	1.0515	0.1554	45.76	1	0.0000	2.8619
REGION(D7)	---	---	76.45	4	0.0000	---

Table 13. Poverty Parameter Estimates, Model SUBG1 (no soctrans)

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
FARMSTEAD(1)	1.1566	0.5192	4.96	1	0.0259	3.1791
VILLAGE(2)	0.2114	0.1542	1.88	1	0.1703	1.2354
TOWN/CITY(3)	-0.0774	0.1562	0.25	1	0.6201	0.9255
MAJOR CITY(4)	-0.2843	0.1723	2.72	1	0.0989	0.7525
BUDAPEST(5)	-1.0063	---	---	---	---	0.3656
CONSTANT	-1.2864	0.3031	18.02	1	0.0000	0.2763

Table 14. Poverty Parameter Estimates, Model SUBG3 (no soctrans)

Variable	B Coeff.	S.E.	Chi Square	d.f.	Sig.	Exp(B)
UNEMP	1.4589	0.7985	3.34	1	0.0677	4.3012
D1	0.0873	0.00506	297.31	1	0.0000	1.0912
FAMSIZE(D4)	-0.1981	0.0515	14.81	1	0.0001	0.8203
SEX	0.5972	0.1448	17.01	1	0.0000	1.8170
REGION(D7)	---	---	32.49	4	0.0000	---
FARMSTEAD(1)	0.7267	0.4444	2.67	1	0.1020	2.0682
VILLAGE(2)	0.2481	0.1424	3.04	1	0.0815	1.2816
TOWN/CITY(3)	-0.1133	0.1485	0.58	1	0.4454	0.8929
MAJOR CITY(4)	-0.2070	0.1740	1.41	1	0.2343	0.8130
BUDAPEST(5)	-0.6545	---	---	---	---	0.5197

Table 14. Poverty Parameter Estimates, Model SUBG3 (no soctrans)						
Variable	**B Coeff.**	**S.E.**	**Chi Square**	**d.f.**	**Sig.**	**Exp(B)**
UNEMP	1.4589	0.7985	3.34	1	0.0677	4.3012
CONSTANT	-4.8282	0.3893	153.79	1	0.0000	0.0080

b. What Remains to be Done

We employ the same general strategy as before to generate estimates of need at the regional level.[27] However, it needs to be emphasized that regions in this context cannot be understood in a conventional geographic sense. With the exception of Budapest, the Hungarians masked the identity of the sample. The country was stratified into five layers beginning with communities as small as the farmstead and ranging on up in size to major cities. In practical terms this means that the policy implications of what follows are largely limited to rural-urban comparisons.

[27] No attempt is made to disaggregate the incidence of poverty by labor force status. Unlike the Polish case, the Hungarian logit models did not produce statistically significant "dispen" coefficients.

Table 15. Poverty and Regional Need							
Region	Population Weights (1)	Poverty Rate in Households (2)	Average Poverty Gap in Adult Equivalent Units (FORINTS) (3)	Average Number of Adult Equivalent Units (4)	Regional Poverty Weight (need in %) (5)	Social Transfers by Region (allocation %) (6)	Excess of Need over Allocation in % (5÷6) (7)
REGION A (not assigned)	012	.7917	91613.11	1.3158	.0418	.0105	3.9809
FARMSTEAD	.011	.5909	44674.55	1.5385	.0163	.0125	1.304
VILLAGE	.352	.5282	44002.28	1.4347	.4285	.4453	.9623
TOWN/CITY	.281	.4296	42946.27	1.3689	.2591	.2780	.9320
MAJOR CITY	.144	.4433	37317.28	1.3256	.1153	.1304	.8842
BUDAPEST	.200	.3243	48695.19	1.2061	.1391	.1233	1.1281
NATION	1.00	.4512	44450.23	1.6331	1.00	1.00	1.00

Estimates of magnitude of poverty are more sensitive to the definition of the income threshold in the Hungarian case than in the Polish. Whereas worst case to best case scenarios produced cost multiples that were nearly equal to "one" in the previous chapter, we now find multiples of eight. At the upper end, figures from table 15 imply that a program to eliminate poverty could produce expenditures in excess of $1.7 billion dollars for a situation like the one which prevailed in 1991.[28] Full relief would take up resources equal to 2.8 percent of GDP, and this might pose an excessive burden.[29] Table 16

[28] In 1991, there were 3,890,000 households in the country. See 1992 Statistical Yearbook of Hungary, p. 27. Subsidy figure calculated as: 44,450.23 forints*1.6331equiv units*3,890,000 households *.4512poverty rate. When converted at the rate of 74.74 forints to the dollar, subsidy = $1,704,717,385. See The World Factbook 1994 p. 181.

[29] The World Factbook 1992, p.152.GDP was estimated to be $60.1 billion.

indicates that the situation might not be so onerous if the median definition of the poverty threshold is closer to the truth: cash assistance under these circumstances falls to just over $203 million.[30]

Variation in the poverty threshold also complicates the interpretation of regional equity.
Our estimates of the extent to which regional needs are covered by social transfers diverge significantly with respect to Region A (unassigned) and Budapest. This, in turn, skews the overall rural-urban comparison. If rural is defined as farmsteads and villages, then 44.28 percent of the need and 45.78 percent of the allocations are identified under the absolute standard.[31] When the standard shifts to the median, the relevant figures become 38.92 percent and 50.27 percent, respectively. One might expect this growing disparity to be mirrored in the implied urban tally but this does not occur because the needs allocation ratio in Region A is concurrently deteriorating. On further reflection, one might also infer that urban social welfare conditions are static since needs and allocations are roughly in balance (47.53 and 48.75). However, the situation is far from that because residents of Budapest experience an allocation deficit of almost nine percentage points while those living in major cities enjoy a surplus of just under eight. To put this in context, if the policy goal is to eliminate national poverty while preserving virtual parity between needs and allocations at the regional level,

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[30] As the poverty threshold falls, fewer households are adversely selected and the subsidy required to bring those remaining up to some minimum level of subsistence likewise declines.

[31] The UN Demographic Yearbook indicates that the Hungarian population was 37.9% rural in 1990. Our placement of farmsteads and villages in the rural category is based on the correspondence between sample coverage (36.3%) and the UN figure. Region A (unassigned) is omitted from the comparison. See UNDY 1991 p.216.

the conservative strategy would be to adopt the absolute standard. However, the price tag associated with such a move could increase government expenditures by over \$1.5 billion dollars.[32]

Table 16. Poverty and Regional Need							
Region	**Population Weights (1)**	**Poverty Rate in Households (2)**	**Average Poverty Gap in Adult Equivalent Units (FORINTS) (3)**	**Average Number of Adult Equivalent Units (4)**	**Regional Poverty Weight (need in %) (5)**	**Social Transfers by Region (allocation %) (6)**	**Excess of Need over Allocation in % (5÷6) (7)**
REGION A (not assigned)	.012	.4583	63901.14	1.2727	.1354	.0098	13.816
FARMSTEAD	.011	.0909	23177.68	1.5000	.0105	.0098	1.071
VILLAGE	.352	.1056	23580.43	1.4267	.3787	.4929	.7683
TOWN/CITY	.281	.0651	25425.61	1.4054	.1979	.2197	.9008
MAJOR CITY	.144	.0550	11467.27	1.2500	.0343	.1126	.3046
BUDAPEST	.200	.0792	36851.43	1.3750	.2430	.1552	1.566
NATION	1.00	.0857	27868.60	1.6331	1.00	1.00	1.00

[32] This result is hardly surprising. Under the absolute standard over 45 percent of the population are considered poor and prospectively eligible for relief. Given the concentration of need in Budapest, it would be hard to maintain regional equity if the median standard were adopted and less than 9 percent of the households qualified for assistance.

lV. CONCLUDING OBSERVATIONS

The minimal objective of this study was to establish the statistical relationship between poverty and labor force status using the LIS data sets. This has been amply demonstrated. Detailed logit models, covering labor force status and a host of other socio-demographic variables, have been shown to fit the data well. We can now characterize, with some accuracy, how the move from employment to unemployment and from unemployment to exiting the labor force puts the household at economic risk. Because these estimates are conditioned on age, gender, educational attainment, place of residence etc., the results have context that was lacking in the earlier BUCEN reports.

The major findings of this study are organized around three themes: poverty risks; impacts of social assistance and financial assessment of need. Bullet points highlight the results.

Poverty Risks

In general, we conclude that: labor force status and gender are critical socio-economic characteristics in identifying the poor; regional disparities in poverty are large; education plays an important role in reducing poverty risks; and adverse economic selection increases with age. At the more detailed level, we find:

For Polish households, economic risks reach a maximum when the head is out of the labor force, female, lacking a high school education and living in region 6 (Mid East). Under these circumstances, the odds of being poor are 7.3 times as great as the contrast group (median poverty threshold used).

For Hungarian households, the worst case scenario is where the head is female, unemployed and residing in region 1 (Farmsteads). These conditions produce poverty risks which are 8.2 times as great as the contrast group (absolute poverty threshold).

Impacts of Social Assistance

Since the models discussed in Chapters II and III employed variables that were deliberately structured along policy lines, we were able us to use scenario analysis and careful sifting and reclassification of the data to tease out insights regarding the social safety net. We now know roughly how much money is needed to prop up the poor, and where such assistance is likely to have the greatest impact.

To answer the question "does the social safety net protect the most vulnerable members of society?" , national poverty rates and odds ratios are re-calculated under the extreme assumption that all social transfers have been netted out of disposable income, and there has been no offsetting reduction in taxes. For the results which are unambiguous and/or statistically significant, we find:

In Poland, national poverty rates jump by approximately 36 percentage points to just under 45%, regardless of the choice of threshold.

In Hungary, national estimates are sensitive to the threshold. In the median case, the increase is over 37 percentage points (to 45+%). In the absolute case, removing the net raises the rate by 22 percentage points to 68%.

In Poland, individual risks are amplified for those who are unemployed, out of the labor force, or educated at the post-secondary level. The opposite is true with regard to gender and education at the primary level. Contrary signals are given off for secondary education, depending on the poverty threshold chosen.

In Hungary, individual risks are dampened for the unemployed and for women.

Unexpected shifts in the risk amplification ratios may be indicative of programmatic bias in the level and delivery of social welfare benefits.

Financial Assessment of Need

To generate dollar estimates of the magnitude of a poverty eradication program, we assume: 1) the absolute and median poverty measures will produce consistent results in repeated sampling experiments, 2) sample sizes at the regional level are large enough to make general, but not precise, inferences about rates and dimensions of poverty, and 3) regional equity can be meaningfully identified with ratios which compare the size of the poverty gap to the value of social transfer benefits received by the needy.

In Poland, authorities might have to spend between $365 and $388 million to raise households out of poverty for one year.

In Hungary, a program with similar goals could cost between $203 and $1,704 million, depending upon the choice of poverty threshold.

In Poland, the excess of the needs share over allocation share is highest in the Middle East and lowest in the South.

In Hungary, the excess of the needs share over allocation shares is greatest in Budapest or Farmsteads and lowest in Major Cities, depending on the definition of poverty.

In their entirety, the above results are provocative and challenging for future research. There is clearly a need to continue investigating how the social safety net disburses funds to groups at risk. Attention should also be paid to questions of regional equity and recipient category bias. Finally, the parameter estimates raise a number of questions for policy makers: Are gender biases caused by flawed rules/administration, or by impersonal forces at work in the labor market? How do economic incentives, built into the social safety net, affect parental labor force participation rates? What types of human capital investments best reduce poverty risks?

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