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MODELLING MIGRATION ATTEMPTS AND THE ROLE OF GENDER IN ALBANIA

ABSTRACT: *This paper uses the 2002 Albanian Living Standards Measurement Survey to model whether an individual has attempted to migrate conditional on having previously considered migrating. The study addresses the methodological concerns that arise from potential selection bias and empirical issues associated with gender differences. We test for the presence of selection bias using a bivariate probit and apply an Oaxaca-style decomposition technique to analyse gender differentials in the conditional probability of attempted migration. We focus on the roles an individual's living standard, geographic location and local labour market conditions exert on the attempt to migrate. Our empirical findings suggest that there are significant differences in both the conditional probability of attempting to migrate and the relative importance of determining factors across gender.*

KEY WORDS: *Albania, bivariate probit, gender, migration.*

JEL CLASSIFICATION: P27, C25, F22, J16

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INTRODUCTION

The last decade of the twentieth century witnessed one of the largest economic experiments of that century as former communist countries embarked on attempts to transform their economies. The transformation process influenced the direction of economic policies and re-shaped the nature of social policies, business practices and institutions.

The collapse of the central planning system in Europe and the former Soviet Union also provided the erstwhile citizens of communist regimes with opportunities to migrate abroad. Albania experienced a steady increase in migration over the first decade of its transition. By the end of the decade over one-fifth of the Albanian population were estimated to be living abroad, mostly in southern European countries such as Greece and Italy, representing the largest outflow rate of any transitional economy. While early studies of Albanian migration flows have found that migrants tended to be young, male, the better educated (see for example Kule *et al* 2002), there is some evidence of an increasing feminization in international migration. The study of King *et al.* (2005) reports, using data from the 2001 Greek Census, women comprising 41% of Albanians living in Greece.

There is evidence of strong gender differentials in migration decisions in Albania. Castaldo *et al.* (2007) examined the determinants of migration using intentions data and found that men were more willing to consider migration than women.[1] The study also found that women responded differently than men to a set of household and local community conditions. For example, the presence of young dependent children within the household reduced a woman's migration risk but raised a man's. Married women were, on average and *ceteris paribus*, more disposed to consider migration relative to single women but the reverse was the case for married men. Women in poorer households were more inclined to consider migration but the reverse was again the case for men. There were further signs of gender asymmetry in response to local labour market conditions. Local unemployment rates were found to increase the probability of considering migration for both men and women, while local wage levels were found, on average and *ceteris paribus*, to exert no independent effect on a woman's disposition to migrate but a strongly negative effect on that of men.

1 We estimated that men, on average and *ceteris paribus*, were over 22 percentage points more likely to consider migrating than women.

The purpose of the current paper is to explore the factors that determine whether or not an individual has *attempted* to migrate abroad, rather than merely *considered* migrating abroad. In order to investigate this we exploit the 2002 ALSMS, which in addition to living standards data contains detailed information on migration, household characteristics, and important labour market information at the level of the individual and local community. In particular, we focus on the relationship between gender and attempting to migrate and, *inter alia*, household living standards, demographics, labour force status and local labour market conditions.

The structure of the paper is now outlined. Section one discusses a number of conceptual issues relating to the migration measures adopted in this study. A section describing the data source used follows. The econometric methodology is then outlined, followed by a discussion of the empirical results. A final section offers a summary and some policy conclusions.

CONCEPTUAL ISSUES

An understanding of migration behaviour and its determinants is an important objective for researchers and policy-makers but one often hampered by a lack of appropriate data. Economists are increasingly trying to explore migrant behaviour more indirectly through the use of intentions data and cast the empirical analysis in terms of migration willingness.[2] This approach allows the investigator to explore the characteristics of those most at risk of migration and link it explicitly to household and other characteristics.

Intentions data typically provide information on whether an individual has considered or intends to migrate and are increasingly available in household surveys. There is much debate about the extent to which such data can be used to infer behaviour rather than the more traditional reliance within the economics tradition of inference based on actual outcomes. Manski (1990) argues that although there is some informational content in intentions-based survey questions, researchers should not expect too much from such data. In contrast, Louviere, Hensher and Swait (2000) argue that models using stated (rather than revealed) preferences have a good empirical track record, which, they argue, is as impressive as the revealed preference counterparts.

2 For example, see Hughes and McCormick (1985), Papanagos and Sanfey (2002), Ahn *et al* (1999), Ahn *et al* (2002), Drinkwater (2003a), Drinkwater (2003b).

Clearly, there will be a gap between intentions and actual outcomes. There is some evidence that intentions are strongly correlated with actual outcomes (see De Jong (2000)), but there are many reasons why the migration intentions of individuals are not actually realised. The gap between intended and actual migration outcomes is likely to be larger in regard to external rather than internal migration (see Gardner, De Jong, Arnold and Carino (1986)). Intentions may also evolve over time if preferences or, more likely, personal circumstances change. In addition, formulating expectations about the costs and benefits associated with migration is extremely difficult when information on the destination country's labour market is unobservable and/or when individuals are asked to operate in an uncertain environment in the source country.[3] The theoretical work of O'Connell (1997) demonstrates why such uncertainties may induce a 'wait and see' approach to migration and thus partly explains the observed disparity between intentions and actual outcomes.

The analysis undertaken in this paper differs from the literature on migration willingness in a number of key respects. Our study does not explicitly use data derived from an intentions-based question. The analysis draws on responses to the inquiry '*has individual X ever considered moving abroad, even temporarily*?' The question is thus more retrospective and, arguably, may be more immune to criticism than one based on intentions. In our view it captures whether someone could be considered at risk of migrating abroad or not. The analysis is taken one step further by using responses to a supplementary question that was only asked conditional on a positive response to this initial question. Individuals that had considered migration were then asked '*has individual X ever tried to move and failed*?' The responses to the second question provide the core of our empirical analysis, as we are ultimately interested in examining the factors that influence whether an individual has attempted to migrate or not. The responses to this second question are used to inform this issue.

In the context of the Albania Living Standards Measurement Survey (ALSMS) from 2002, there are a number of potential problems that attach to the responses to these questions and to whom the questions were asked. Firstly, the timeframe to which the question relates is broad and it is unclear whether a positive response to either of these questions relates to a more recent period of time or is informed by an individual's longer historical perspective over an earlier five year period. This poses a potential alignment problem in that the factors used to explore the

3 The existence of this type of uncertainty could be taken to characterise many post-socialist economies in the early phase of their transition to a market based economy.

determinants of the conditional probability that someone attempted to migrate relate to the current time period, not past periods. We assume, however, that respondents are more likely to attach a heavier weight to the more recent past in furnishing a response to both these question.[4] Secondly, individuals who have actually migrated abroad in the past five years were not asked these questions for obvious reasons. Thus, the sample that we use is conditioned on those that have not migrated abroad since at least 1997. This poses a potential sample selection problem given the absence of information on those migrants who are currently based abroad. In particular, it could be argued that those left behind have poorer unobservable characteristics that render them less likely to successfully migrate abroad. This issue should perhaps be borne in mind when interpreting the reported results. Finally, given the tendency of those in rural areas to initially engage in internal rather than external migration, such respondents may not be inclined to provide a positive response to the first question. We do not believe that this poses an insurmountable interpretational problem since internal migration is likely to represent the first step towards an ultimate move abroad.

It is readily acknowledged that there are legitimate concerns in making links between whether an individual attempted to migrate at some point in time and actual migration behaviour. These concerns aside, given the absence of data reflecting actual outcomes, we believe an examination of whether an individual attempted to migrate is a useful exercise in its own right. We take the view that the approach provides the basis for a systematic analysis of the profile of individuals considered most prone to migration in Albania and identifies the factors that are most likely to reduce or increase this tendency. This type of analysis is clearly fruitful and has potentially strong policy content. The 'flip-side' of this particular coin is also no less interesting in that we get a sense of the factors and characteristics that do not predispose Albanians to try and attempt to migrate abroad.

DATA

The first round of the Albania Living Standards Measurement Survey (ALSMS), conducted in the spring of 2002, contains a wide range of information on living

4 The issue is also relevant to the use of local labour market variables, which are based on contemporaneous information. However, this should not represent a major problem if there is persistence in the inter-district unemployment and wage rate differentials. Thus, the variation across districts could be argued to be relatively constant over the five-year period to which the question relates. Given data constraints, there is no way to verify this persistence but it does not seem an unreasonable assumption.

conditions of the people of Albania, as well as a module on migration, providing researchers with the opportunity to investigate in more detail Albanian migration and its links with several socio-economic factors.[5]

The ALSMS 2002 was undertaken, as a part of a collaborative five-year project, by the Albanian National Institute of Statistics (INSTAT), with the assistance of the World Bank and the UK's Department for International Development (DfID). The data obtained were used by the World Bank to conduct a poverty assessment (see World Bank (2003)). The ALSMS sample includes 3,599 households comprising a total of 16,521 individuals. The households were selected by using a stratified two-stage cluster design. The sampling frame was divided into four regions (strata), Tirana, Coastal, Central and Mountain. The two-stage clustering involved the selection of 450 primary sampling units (125 in the Coastal, Central and Mountain areas, and 75 in Tirana), and eight households in each of these. The primary sampling units were selected from the 2001 pre-census list of census enumeration areas. The sample is representative at national level, as well as at regional and urban/rural level.[6]

Given the objectives of the present research, the sample of individuals is restricted to include only individuals aged between 15 and 60 who are defined to be in the labour force. These are, by the conventional definition, those who are either employed or unemployed but actively looking for work. The sample is further conditioned on those who have never migrated from Albania since 1997 but expressed an opinion on whether they had ever considered migrating or not. After excluding individuals on whom there are missing values for all variables of interest, we are left with an overall sample of 5,423 individuals that constitute the set of usable data points for our empirical analysis.

The LSMS contains a wide range of information at both the individual and household level. The individual level variables available include age, gender, marital status, educational level achieved, and current labour force status. The set of household level variables that can be constructed from the data comprise household size, presence of dependent children and household consumption expenditure in per capita terms. The LSMS data also enable us to construct measures of the quality of the dwellings in which the individuals reside (e.g.,

5 There are later ALSMS surveys but the 2002 release included a module on household expenditures, thereby allowing us to explore links between poverty and migration decisions.

6 Additional details on the survey methodology and fieldwork are available in the Albanian 2002 LSMS Basic Information Document on www.worldbank.org/lsms.

dwelling size and the age of the dwelling). Furthermore, measures relating to the local labour market conditions, such as the district level unemployment rate and hourly wage rate are also constructed. Finally, from the migration module two key variables for our analysis can be constructed. These are whether or not an individual has ever considered migration abroad and, conditional on having considered it, whether the individual has ever tried (and failed) to migrate abroad.

Table A1 of the appendix describes the variables used in our analysis and table A2 provides corresponding summary statistics using both the whole sample and disaggregated by gender. Comparable summary statistics are provided for the sub-sample of the individuals who have tried to migrate abroad.

The summary statistics in table A2 reveal that almost one-third of the sample considered migration abroad and that men are almost twice as likely as women to have considered this option. Among those who have considered migration abroad, almost half have actually tried to leave the country, with men again being almost twice as likely as women to attempt migration.

METHODOLOGY

The primary concern of this paper is to model whether an individual has attempted to migrate or not. Our analysis is based on questions that were asked in the migration module of the ALSMS. In particular, individuals were asked whether they had ever considered migrating abroad and, if so, had they ever attempted to do so. This sequence of questions and the fact that only a sub-sample of respondents provided information on whether they attempted to migrate warrants caution in the econometric treatment of these responses. An exclusive focus on those individuals who attempt to migrate may overlook the fact that they constitute a self-selected sub-sample. The econometric problem is reminiscent of the more conventional selectivity bias problem within a linear regression framework as popularised by Heckman (1979). In our application 1,785 out of the 5,423 individuals in the sample responded that they had considered migration abroad, and 889 of these responded that they had actually attempted to migrate. The structure of the response sequence suggests use of a bivariate probit model. Meng and Schmidt (1985), building on the original work of Poirier (1980), survey the properties of a number of bivariate probit models with varying degrees of observability.

The decision process in our particular case may be illustrated by the following two linear latent dependent variable equations:

$$y^*_{1i} = \mathbf{x}'_i \boldsymbol{\beta} + v_i \qquad [1]$$

where $y_{1i} = 1$ if the individual has considered migrating ($y^*_{1i} > 0$)
$y_{1i} = 0$ if not ($y^*_{1i} \leq 0$))

$$y^*_{2i} = \mathbf{z}'_i \boldsymbol{\gamma} + u_i \qquad [2]$$

where $y_{2i} = 1$ if individual attempted to migrate ($y^*_{2i} > 0$)
$= 0$ if not ($y^*_{2i} \leq 0$)

The dichotomous y_{2i} variable is only observed in our sample if $y_{1i} = 1$. This is the case of a censored probit model described in Meng and Schmidt (1985) as their case three (*Partial Partial Observability*). Variants of this type of model have been used by Van De Ven and Van Praag (1981) to examine deductibles in private health insurance in the Netherlands, by Farber (1983) to examine the demand for trade unionism in the United States, and by Boyes, Hoffman and Low (1989) to model loan default in the United States.

Equation [1] is fully observed and can be estimated separately. However, separate estimation of the second equation may be subject to selection bias given the potential for correlation between v_i in [1] and u_i in [2]. This possesses obvious consequences for the consistency of the coefficient estimates of [2]. Meng and Schmidt (1985) suggest the joint estimation of equations [1] and [2] by full information maximum likelihood (FIML) methods. The likelihood function (L(·)) for this application is given by:

$$L(\boldsymbol{\beta}, \boldsymbol{\gamma}, \rho; \mathbf{x}_i, \mathbf{z}_i) = \prod_{y1=1, y2=1} F(\mathbf{x}'_i \boldsymbol{\beta}, \mathbf{z}'_i \boldsymbol{\gamma}; \rho) \prod_{y1=1, y2=1} F(\mathbf{x}'_i \boldsymbol{\beta}, -\mathbf{z}'_i \boldsymbol{\gamma}; -\rho)$$
$$\prod_{y1=0} [1 - \Phi(\mathbf{x}'_i \boldsymbol{\beta})] \qquad [3]$$

where:
$\Phi(\cdot)$ = univariate standard normal cumulative distribution function.
ρ = the correlation coefficient between the unobservable errors v_i and u_i.
$F(\cdot\,, \cdot; \rho)$ = bivariate standard normal cumulative distribution function.

The parameters of the equations described in [1] and [2] are estimated jointly by maximizing the log likelihood version of function [3] with respect to the coefficient vectors β, γ and the correlation coefficient ρ. The estimate of ρ provides a test for selectivity bias. If ρ=0, no evidence of selectivity bias is present and no efficiency loss is encountered in the separate estimation of either equation [1] or [2] by a univariate probit.

The identification of such selectivity models is of crucial importance. Identification is achieved by the inclusion of variables in equation [1] that are excluded from equation [2]. Poor identification restrictions can lead to erroneous conclusions regarding the presence of selectivity effects. In the context of our application it would be of some interest to establish if, having controlled for a set of observable characteristics, the sample of respondents possessed unobservable characteristics (for example, motivation) that were in some way different from the sample as a whole. A statistically significant ρ value may provide an insight into this particular issue. However, confidence in the reliability of such a result depends crucially on appropriate identification. The vectors **x** and **z** of equations [1] and [2] contain a set of variables common to both in addition to sets of non-overlapping variables that appear in **x** but not **z**, and conversely (but not crucially for identification) in **z** but not **x.** This issue of identification will be examined in more detailed in the empirical section.

On the assumption that there is some degree of confidence in terms of achieving identification, an important issue is to determine whether the estimated correlation (or selection) coefficient is statistically different from zero or not. In order to ascertain this, several different statistical testing principles are used. The most straightforward one is based on estimation of the unrestricted bivariate probit and use of a Wald test for the estimated correlation coefficient. On the other hand, if the bivariate probit model is treated as the restricted model and the two separate probits comprise the unrestricted model, the Chow-type version of the likelihood ratio test can be used to test the hypothesis of interest. A third test does not require estimation of the bivariate probit model itself and treats the separate probit models as comprising the restricted case. The relevant score (or Lagrange Multiplier) test is based on the work of Kiefer (1982) and is computed using the pseudo-residuals from separate estimation of the two underlying probit models. The test only requires estimation of the restricted outcome (i.e., separate univariate probit models).[7] The nature of this test suggests the possibility for a cruder alternative that could also be implemented using the pseudo-residuals from the separate probit models. If we define the Pearson product moment correlation between the two sets of psuedo-residuals for the n observations where $y_{1i} = 1$ as r, then using the Fisher transformation we could express:

7 It should be noted that the original test developed by Kiefer (1982) was based on bivariate probit models with full rather than partial observability.

$$z_r = \frac{1}{2}\log_e\frac{1+r}{1-r} \sim N\left(\frac{1}{2}\log_e\frac{1+\rho}{1-\rho}, \frac{1}{n-k-3}\right)$$

where k is the total number of coefficients estimated across the two probit models. The z-test could then be computed to test the proposition that the true correlation parameter $\rho=0$.

In estimating the unrestricted bivariate probit it may also prove useful to test whether the effects of covariates common to both the **x** and the **z** vectors have comparable effects on the outcomes of interest. This can be done using simple Wald tests and may allow for the introduction of data admissible cross-equation constraints that render the estimation more efficient econometrically.

The bivariate probit model introduced above allows the calculation of various conditional and unconditional probabilities. For example, $\Phi(\mathbf{x}'_i\boldsymbol{\beta})$ is the unconditional (or marginal) probability that an individual considers migration, $\Phi(\mathbf{z}'_i\boldsymbol{\gamma})$ is the unconditional (or marginal) probability that an individual has attempted to migrate, and (from an application of Bayes' Law) $F(\mathbf{x}'_i\boldsymbol{\beta}, \mathbf{z}'_i\boldsymbol{\gamma};\rho)/\Phi(\mathbf{x}'_i\boldsymbol{\beta})$ is the probability that an individual attempts to migrate conditional on having considered migration as an option. Clearly, if $\rho=0$ the conditional probability $F(\mathbf{x}'_i\boldsymbol{\beta}, \mathbf{z}'_i\boldsymbol{\gamma};\rho)/\Phi(\mathbf{x}'_i\boldsymbol{\beta})$ collapses to $\Phi(\mathbf{z}'_i\boldsymbol{\gamma})$ and the second stage equation is amenable to direct estimation by the more standard univariate probit model.

The interpretation of the estimated effects for the covariates on the probability of the conditional event occurring requires computation of marginal (or impact) effects. Our major interest is to examine the effect of changes in covariates on the conditional probability of the event occurring. Greene (2000) provides the computational details. The marginal effects can be calculated for all individuals in the sample with a mean value that can be computed as a summary statistic. Alternatively, the marginal effect can be computed using the vector of mean characteristics for **x** and **z** in this case. This latter approach facilitates the computation of sampling variances for the marginal effects using the delta method. However, the estimates for the marginal effects can be sensitive to the approach adopted. This follows from the non-linear nature of the joint CDF operator. Train (2003, pp.33-35) provides a nice graphical illustration of the problem in the context of a non-linear probability model and cites empirical evidence where marginal effects based on average characteristics differ from average marginal effects by a factor of two or three. Nevertheless, the approach adopted in this study is to compute the marginal effects evaluated at the average characteristics rather than reporting average marginal effects.[8]

An important objective of the empirical work is to explore differences in outcomes across gender groups. On the assumption that the separation of the sample by gender is supported by the data, an important theme of our work is to quantify the conditional probability differentials by gender and assign any difference to 'treatment' and 'endowment' components. The decomposition of the differential between two continuous dependent variables is relatively straightforward in a linear regression model context using the OLS estimation procedure (see Oaxaca (1973)). The task of decomposing the differential when the dependent variables are binary is complicated by the non-linearity inherent in limited dependent variable models. Gomulka and Stern (1990) outline an approach that is applicable to non-linear binary dependent variable models and we extend it here to the case of the bivariate probit model.

Assume separation of the data is valid and separate bivariate probit models are estimated for men and women. We use the subscripts m and f to denote male and female respectively and N_m and N_f to denote their respective overall sample sizes. The sample average conditional probability of attempting to migrate for men can be expressed as:

$$[N_m]^{-1}[\Sigma F(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_m, \mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_m; \hat{\rho}_m)] \div [N_m]^{-1}[\Sigma\Phi(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_m)]$$
$$\text{where } i=1,\ldots\ldots,N_m \qquad [4]$$

where the circumflexes denote the FIML estimates and the $F(\cdot\ ,\ \cdot\ ;\ .)$ and $\Phi(\cdot)$ operators are defined as above. The corresponding sample average conditional probability of attempting to migrate for women can be expressed as:

$$[N_f]^{-1}[\Sigma F(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_f, \mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_f; \hat{\rho}_f)] \div [N_f]^{-1}[\Sigma\Phi(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_f)]$$
$$\text{where } i=1,\ldots\ldots,N_f \qquad [5]$$

Two counterfactual conditional probabilities are now introduced. The first provides the sample average conditional probability for men if subjected to the female coefficient structure. This can be computed as:

$$[N_m]^{-1}[\Sigma F(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_f, \mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_f; \hat{\rho}_f)] \div [N_m]^{-1}[\Sigma\Phi(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_f)]$$
$$\text{where } i=1,\ldots\ldots,N_m \qquad [6]$$

8 In fact, there are negligible differences in estimated marginal effects between the two computation methods.

The second counterfactual can be constructed for the female sub-sample and provides the sample average conditional probability for women if confronted by the male coefficient structure. This can be computed as:

$$[N_f]^{-1}[\Sigma F(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_m, \mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_m; \hat{\rho}_m)] \div [N_f]^{-1}[\Sigma\Phi(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_m)]$$
$$\text{where } i=1,\ldots\ldots,N_f \qquad [7]$$

The computation of these four measures allows us to decompose the sample average conditional probability between the two gender groups. Using the male coefficients, the total differential = [4] – [5] and the endowment can be computed as [4] – [7]. The treatment effect can then be computed as [7] – [5]. Alternatively, using female coefficients, the total gender differential is [4] – [5], the endowment differential is [6] – [5], and the treatment differential is [4] – [6]. It should be noted that this approach, like the standard index number decomposition, is subject to the standard index number problem and is sensitive to which coefficients (or prices) are used to weight the characteristics. A desirable approach is thus to report both estimates and assess the degree of sensitivity.

Finally, if $\rho = 0$, the decomposition follows the approach of Gomulka and Stern (1990) more closely and the analogous expressions, based on the use of univariate probits, are:

$$[n_m]^{-1}\Sigma\Phi(\mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_m) \qquad \text{where } i=1,\ldots\ldots,n_m \qquad [4']$$

$$[n_f]^{-1}\Sigma\Phi(\mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_f) \qquad \text{where } i=1,\ldots\ldots,n_f \qquad [5']$$

$$[n_m]^{-1}\Sigma\Phi(\mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_f) \qquad \text{where } i=1,\ldots\ldots,n_m \qquad [6']$$

$$[n_f]^{-1}\Sigma\Phi(\mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_m) \qquad \text{where } i=1,\ldots\ldots,n_f \qquad [7']$$

The ‘treatment’ and ‘endowment’ effects can then be derived by substituting expressions [4′] to [7′] for [4] to [7] above. Note in this case the sample size is conditioned on whether the individuals considered migrating, so $n < N$.

EMPIRICAL RESULTS

Table 1 provides the FIML estimates for bivariate probit models fitted to data pooled across gender and also separately for each gender group.[9] The goodness-of-fit measures for the estimated models are reasonably good. The identification of the estimated selection parameter in our application is achieved through the inclusion in equation [1] of three education variables, three dependent children variables, three housing quality variables, and a birth location variable, none of which features in the companion equation [2]. The validity of the identifying restrictions is empirically explored by inserting the excluded variables in a univariate probit model for equation [2]. The set of ten coefficients associated with the identifiers is found to be statistically insignificant from zero on the basis of likelihood ratio tests for all three models.[10] The exclusion restrictions, though somewhat *ad hoc*, thus appear to perform an adequate task in terms of identification.

Table 1: FIML Estimates for Considering to Migrate and Attempting to Migrate – Unrestricted

	Pooled Sample		Male Sample		Female Sample	
	Considered	Tried	Considered	Tried	Considered	Tried
Constant	-1.9640*** (0.6417)	-0.8319 (1.1718)	0.7788 (0.8694)	1.9667 (1.3637)	-4.3312*** (1.0203)	-4.4154* (2.6093)
Aged 15 – 25 years	0.7717*** (0.0857)	0.9561*** (0.1554)	0.8738*** (0.1239)	1.0106*** (0.2088)	0.5544*** (0.1406)	0.9988*** (0.3839)
Aged 26 – 40 years	0.6593*** (0.0699)	0.7166*** (0.1378)	0.6754*** (0.0869)	0.6560*** (0.1667)	0.5168*** (0.1327)	0.8579** (0.3588)
Aged 41 – 50 years	0.5070*** (0.0685)	0.4730*** (0.1311)	0.4915*** (0.0823)	0.3931** (0.1540)	0.3849*** (0.1327)	0.5928* (0.3451)
Aged 50+ years	*f*	*f*	*f*	*f*	*f*	*f*
Male	0.6090*** (0.0396)	0.8483*** (0.0715)	§	§	§	§
Married	0.0604 (0.0633)	0.2300*** (0.0863)	0.0056 (0.0990)	0.2686** (0.1296)	0.1658* (0.0899)	0.2599 (0.1640)

9 The Huber (1967) adjustment to the bivariate probit variance-covariance matrix estimators is feasible for both these models. Its use yielded no material difference to the sampling variances, so the reported standard errors are conventional maximum likelihood estimates. The reservations of Greene (2000) regarding this type of adjustment in the context of maximum likelihood estimators are well taken.

10 The chi-squared values (with prob-values in parentheses) for the joint significance of these coefficients in equation [2] are 5.21 (0.868), 5.835 (0.829) and 0.506 (0.975) for the pooled, male and female models respectively.

	Pooled Sample		Male Sample		Female Sample	
	Considered	Tried	Considered	Tried	Considered	Tried
Less than Secondary Education	*f*	§	*f*	*f*	*f*	§
Secondary Education	0.0786 (0.0509)	§	-0.0429 (0.0659)	§	0.2440*** (0.0853)	§
Vocational Education	0.1793*** (0.0543)	§	0.1056 (0.0699)	§	0.2944*** (0.0922)	§
University Education	-0.0371 (0.0693)	§	-0.0851 (0.0885)	§	0.0715 (0.1238)	§
Born in the Municipality	0.0536 (0.0409)	§	-0.0052 (0.0604)	§	0.0836 (0.0621)	§
Unemployed	*f*	*f*	*f*	*f*	*f*	*f*
Employee	-0.1864*** (0.0628)	-0.2666*** (0.0890)	-0.1801** (0.0836)	-0.3157*** (0.1124)	-0.1598 (0.1026)	-0.1620 (0.1827)
Farmer	-0.2678*** (0.0791)	-0.3482*** (0.1178)	-0.1647 (0.1011)	-0.2443* (0.1411)	-0.2993** (0.1432)	-0.3077 (0.2800)
Self-Employed	-0.2628*** (0.0822)	-0.2076* (0.1223)	-0.1320 (0.1021)	-0.1867 (0.1426)	-0.3430** (0.1534)	0.1817 (0.3086)
Temporary Layoff	-0.2767* (0.1602)	-0.5961** (0.2740)	-0.0432 (0.2473)	-0.2229 (0.3597)	-0.4451** (0.2169)	-1.1286** (0.5751)
Household Size	§	-0.0754*** (0.0202)	§	-0.0901*** (0.0247)	§	-0.0527 (0.0411)
Children in Household: Aged ≤ 4 years	0.0884* (0.0469)	§	0.0553 (0.0626)	§	0.1661** (0.0762)	§
Children in Household: Aged 5 ≤ years ≤ 8	0.0330 (0.0433)	§	0.0854 (0.0576)	§	-0.0389 (0.0706)	§
Children in Household: Aged 9 ≤ years ≤ 14	-0.0521 (0.0400)	§	-0.0553 (0.0531)	§	-0.0283 (0.0661)	§
Log Household Consumption per Capita	0.2038*** (0.0453)	-0.0347 (0.0816)	0.1031* (0.0590)	-0.1769* (0.0947)	0.3259*** (0.0756)	0.2351 (0.1744)
Residence Dwelling Area: ≤ 69 Sq.Metres	*f*	§	*f*	§	*f*	§
Residence Dwelling Area: 70 ≤ Sq.Metres ≤ 130	-0.1289*** (0.0380)	§	-0.1244** (0.0502)	§	-0.1545** (0.0618)	§
Residence Dwelling Area: Sq.Metres > 130	-0.3070*** (0.0939)	§	-0.3511*** (0.1215)	§	-0.2827* (0.1577)	§
Residence Built post-1990	0.1461*** (0.0456)	§	0.2002*** (0.0612)	§	0.1131 (0.0729)	§
Coastal	*f*	*f*	*f*	*f*	*f*	*f*

	Pooled Sample		Male Sample		Female Sample	
	Considered	Tried	Considered	Tried	Considered	Tried
Central	0.1512*** (0.0510)	0.0337 (0.0810)	0.1980*** (0.0691)	0.0192 (0.1035)	0.1004 (0.0790)	0.1415 (0.1572)
Mountain	-0.1049* (0.0572)	0.0728 (0.1006)	-0.0385 (0.0768)	0.1065 (0.1187)	-0.2097** (0.0878)	0.0336 (0.1957)
Tirana	-0.0654 (0.0691)	-0.0394 (0.1055)	-0.0558 (0.0913)	-0.1729 (0.1344)	-0.0314 (0.1132)	0.3122 (0.2042)
Urban	0.0767 (0.0607)	-0.2753*** (0.1042)	-0.0099 (0.0739)	-0.3953*** (0.1182)	0.2037* (0.1186)	0.0337 (0.2365)
District Unemp. Rate (%)	0.0120*** (0.0025)	0.0133*** (0.0042)	0.0180*** (0.0033)	0.0192*** (0.0054)	0.0050 (0.0038)	-0.0021 (0.0078)
District Level Hourly Wage (log)	-0.2966** (0.1245)	-0.0539 (0.2133)	-0.5927*** (0.1673)	-0.2010 (0.2823)	0.0207 (0.1952)	0.2178 (0.3646)
ρ	0.5560** (0.2591)		0.5625* (0.3048)		0.2137 (0.4881)	
Number of Observations	5423	1785	2841	1203	2582	582
Log Likelihood Value	-4300.051		-2586.370		-1622.266	
$[r^2]_{single}$	0.1030	0.0970	0.0780	0.0800	0.0796	0.0572
$[r^2]_{joint}$	0.1167		0.0899		0.0751	
Tests for $\rho = 0$						
Wald Test ~ χ^2_1	4.603** (0.031)		3.404* (0.065)		0.192 (0.661)	
Likelihood Ratio Test ~ χ^2_1	2.132 (0.144)		1.375 (0.241)		0.206 (0.650)	
LM (Score) Test ~ χ^2_1	0.022 (0.637)		0.013 (0.909)		0.005 (0.944)	
Z_ρ	0.408 (0.684)		0.384 (0.535)		0.242 (0.623)	

Notes to table 1:

(a) Asymptotic standard errors for the maximum likelihood estimated coefficients are reported in parentheses.
(b) ***, **, * denote statistical significance at the 1%, 5% and 10% level respectively using two-tailed tests.
(c) § denotes not applicable in estimation and *f* denotes omitted in estimation.
(d) $[r^2]_{single}$ is the squared correlation coefficient between the estimated marginal probabilities using the $\Phi(\cdot)$ operator and the reported maximum likelihood estimates and the single discrete outcome;
(e) $[r^2]_{joint}$ is the squared correlation coefficient between the estimated marginal probabilities using the $F(\cdot)$ operator and the reported maximum likelihood estimates and the joint discrete outcome;
(f) The tests for $\rho = 0$ are described in the text. The values reported in parentheses for these tests refer to the significance level of the tests.

It is possible to ascertain whether the estimated coefficients associated with the variables common to both equations are statistically different from each other. The statistic of choice for this purpose is provided by a Wald test. Table 2 reports the relevant values for testing the different sets of cross-equation restrictions in the three models. In regard to the pooled model the imposition of cross-equation restrictions for the age controls, the labour force status variables, the regional dummies, and the unemployment rate is found to be consistent with the data. Given that the result for the district level wage was marginal it was decided not to impose the cross-equation restriction on this variable's coefficient. Therefore, a total of 11 cross-equation restrictions are imposed in the pooled model. Three fewer cross-equation restrictions are imposed for the male case given that the set of regional effects were found to be statistically different, albeit marginally, across the two equations. In contrast, the data for the female model support the imposition of cross-equation constraints for all the covariates common to both equations.

Table 2: FIML Estimates for Considering to Migrate and Attempting to Migrate – Restricted

	Pooled Sample		Male Sample		Female Sample	
	Considered	**Tried**	**Considered**	**Tried**	**Considered**	**Tried**
Constant	-2.2004*** (0.6312)	0.1596 (1.0437)	0.7482 (0.8627)	2.0547 (1.3364)	-4.3136*** (0.9615)	-4.2018*** (1.0578)
Aged 15 – 25 years	0.8080*** (0.0795)	0.8080*** (0.0795)	0.9047*** (0.1139)	0.9047*** (0.1139)	0.6307*** (0.1381)	0.6307*** (0.1381)
Aged 26 – 40 years	0.6648*** (0.0655)	0.6648*** (0.0655)	0.6582*** (0.0803)	0.6582*** (0.0803)	0.5732*** (0.1320)	0.5732*** (0.1320)
Aged 41 – 50 years	0.4894*** (0.0639)	0.4894*** (0.0639)	0.4569*** (0.0758)	0.4569*** (0.0758)	0.4025*** (0.1321)	0.4025*** (0.1321)
Aged 50+ years	*f*	*f*	*f*	*f*	*f*	*f*
Male	0.6079*** (0.0391)	0.8245*** (0.0759)	§	§	§	§
Married	0.0943 (0.0594)	0.1435* (0.0761)	0.0421 (0.0936)	0.1724 (0.1074)	0.2011*** (0.0747)	0.2011*** (0.0747)
Primary or Less	*f*	§	*f*	§	*f*	§
Secondary Education	0.0791 (0.0513)	§	-0.0404 (0.0667)	§	0.2511*** (0.0830)	§
Vocational Education	0.1853*** (0.0540)	§	0.1169* (0.0703)	§	0.3133*** (0.0886)	§
University Education	-0.0329 (0.0695)	§	-0.0757 (0.0898)	§	0.0921 (0.1144)	§
Born in the Municipality	0.0577 (0.0415)	§	-0.0133 (0.0610)	§	0.0816 (0.0608)	§
Unemployed	*f*	*f*	*f*	*f*	*f*	*f*

	Pooled Sample		Male Sample		Female Sample	
	Considered	Tried	Considered	Tried	Considered	Tried
Employee	-0.2078*** (0.0553)	-0.2078*** (0.0553)	-0.2207*** (0.0732)	-0.2207*** (0.0732)	-0.1649* (0.0921)	-0.1649* (0.0921)
Farmer	-0.2861*** (0.0704)	-0.2861*** (0.0704)	-0.1881** (0.0895)	-0.1881** (0.0895)	-0.3005** (0.1259)	-0.3005** (0.1259)
Self-Employed	-0.2474*** (0.0723)	-0.2474*** (0.0723)	-0.1487* (0.0900)	-0.1487* (0.0900)	-0.2364* (0.1318)	-0.2364* (0.1318)
Temporary Layoff	-0.3486** (0.1563)	-0.3486** (0.1563)	-0.1024 (0.2268)	-0.1024 (0.2268)	-0.5353** (0.2246)	-0.5353** (0.2246)
Household Size	§	-0.0674*** (0.0195)	§	-0.0916*** (0.0237)	§	-0.0328 (0.0359)
Children in Household: ≤ 4 years	0.0777* (0.0467)	§	0.0553 (0.0625)	§	0.1426** (0.0722)	§
Children in Household: 5 ≤ years ≤ 8	0.0268 (0.0435)	§	0.0952 (0.0579)	§	-0.0570 (0.0683)	§
Children in Household: 9 ≤ years ≤ 14	-0.0515 (0.0402)	§	-0.0466 (0.0534)	§	-0.0397 (0.0646)	§
Household Expenditure (Log)	0.2085*** (0.0450)	-0.0675 (0.0719)	0.1086* (0.0586)	-0.2005** (0.0868)	0.2999*** (0.0692)	0.2999*** (0.0692)
Dwelling Area: ≤ 69 Sq.Metres	*f*	§	*f*	§	*f*	§
Dwelling Area: 70 ≤ Sq.Metres ≤ 130	-0.1250*** (0.0383)	§	-0.1197** (0.0508)	§	-0.1519** (0.0608)	§
Dwelling Area: Sq.Metres > 130	-0.3052*** (0.0947)	§	-0.3651*** (0.1230)	§	-0.2758* (0.1546)	§
Residence Built post-1990	0.1523*** (0.0459)	§	0.2024*** (0.0617)	§	0.1061 (0.0718)	§
Coastal	*f*	*f*	*f*	*f*	*f*	*f*
Central	0.1159** (0.0469)	0.1159** (0.0469)	0.1964*** (0.0688)	0.0022 (0.1026)	0.1109 (0.0731)	0.1109 (0.0731)
Mountain	-0.0610 (0.0531)	-0.0610 (0.0531)	-0.0422 (0.0764)	0.1274 (0.1158)	-0.1671** (0.0820)	-0.1671** (0.0820)
Tirana	-0.0576 (0.0626)	-0.0576 (0.0626)	-0.0529 (0.0911)	-0.1884 (0.1383)	0.0491 (0.1014)	0.0491 (0.1014)
Urban	0.0699 (0.0575)	-0.2768*** (0.0831)	-0.0101 (0.0717)	-0.41648*** (0.0995)	0.1698 (0.1036)	0.1698 (0.1036)
District Unemployment. Rate (%)	0.0122*** (0.0023)	0.0122*** (0.0023)	0.0180*** (0.0030)	0.0180*** (0.0030)	0.0035 (0.0035)	0.0035 (0.0035)
District Level Hourly Wage (log)	-0.2560** (0.1215)	-0.1624 (0.1726)	-0.5958*** (0.1665)	-0.1288 (0.2618)	0.0578 (0.1762)	0.0578 (0.1762)
ρ	0.3926** (0.1986)		0.4116* (0.2344)		0.2822 (0.2757)	

	Pooled Sample		**Male Sample**		**Female Sample**	
	Considered	**Tried**	**Considered**	**Tried**	**Considered**	**Tried**
Number of Observations	5423	1785	2841	1203	2582	582
Log -Likelihood Value	-4309.688		-2588.219		-1629.514	
$[r^2]_{single}$	0.0968	0.0972	0.0778	0.0800	0.0786	0.0350
$[r^2]_{joint}$	0.1176		0.0913		0.0777	
Likelihood Ratio Test for Restrictions ~ χ^2_k	19.274* (0.056)		3.698 (0.883)		14.496 (0.488)	

Notes to table 2:

(a) Asymptotic standard errors for the maximum likelihood estimated coefficients are reported in parentheses.
(b) ***, **, * denote statistical significance at the 1%, 5% and 10% level respectively using two-tailed tests.
(c) § denotes not applicable in estimation and *f* denotes omitted in estimation.
(d) $[r^2]_{single}$ is the squared correlation coefficient between the estimated marginal probabilities using the $\Phi(\cdot)$ operator and the reported maximum likelihood estimates and the single discrete outcome;
(e) $[r^2]_{joint}$ is the squared correlation coefficient between the estimated marginal probabilities using the $F(\cdot)$ operator and the reported maximum likelihood estimates and the joint discrete outcome;
(f) The Likelihood Ratio Test reported tests the cross-equation restrictions imposed. The values reported in parentheses for these tests refer to the significance level of the tests.

Given the findings of table 3, the bivariate probit model with sample selection is re-estimated with the relevant cross-equation restrictions imposed. The FIML estimates are reported in table 3.[11] In addition, the separation by gender is also supported by the data in this case. Using the log-likelihood values from table 3, the likelihood ratio test is computed as $\chi^2_{30} = 183.9$.[12] This suggests that a pooled model with an intercept shift is dominated by separate gender models and implies that gender differences in coefficients other than the constant term account for the rejection.

11 Likelihood ratio tests based on a comparison of the log-likelihood values for the unrestricted models in table 1 and the restricted versions in table 3 are reported in the latter table. The validity of the cross-equation restrictions imposed is confirmed for the male and female models but, surprisingly, appears more marginal for the pooled model.

12 The 0.05 cut-off point for a chi-squared statistic with 30 degrees of freedom is 43.77.

Table 3: Wald Tests for Cross-Equation Coefficient Restrictions for Common Variables

	Pooled	Male	Female
Male ~ χ^2_1	14.517*** (0.000)	§	§
Age Controls ~ χ^2_3	3.621 (0.305)	2.184 (0.535)	1.877 (0.598)
Married ~ χ^2_1	5.128** (0.024)	6.264** (0.012)	0.332 (0.565)
Labour Force Status Controls ~ χ^2_4	3.145 (0.336)	1.537 (0.820)	5.231 (0.264)
Household Expenditure (Log) ~ χ^2_1	10.241*** (0.000)	11.069*** (0.000)	0.291 (0.590)
Regional Controls ~ χ^2_3	3.145 (0.336)	6.321* (0.096	3.429 (0.330)
Urban Settlement ~ χ^2_1	10.542*** (0.000)	15.367*** (0.000)	0.517 (0.472)
District Unemployment Rate (%) ~ χ^2_1	0.139 (0.710)	0.064 (0.780)	0.900 (0.343)
District Level Hourly Wage (log) ~ χ^2_1	1.903 (0.168)	2.935* (0.087)	0.321 (0.571)

Notes to table 3:

(a) The Wald tests are computed as: $[db_k]'V^{-1}[db_k]$ where $V = [V1]_k + [V2]_k - 2Cov[1,2]_k$ where db_k is the vector containing the coefficient differences between equations [1] and [2] for the k^{th} category of variables; $[V1]_k$ and $[V2]_k$ are the corresponding sampling variances-covariances for the coefficients for the k^{th} category and $2Cov[1,2]_k$ is the covariance in sampling variances across equations.
(b) ***, **, * denote statistical significance at the 1%,5% and 10% level respectively.
(c) The prob-values for the tests are reported in parentheses.

The evidence from tables 1 and 3 in regard to sample selection is somewhat mixed. The Wald tests for the pooled and the male model suggest evidence of positive selection bias, but there is no indication that this phenomenon is present in the female case. However, and in spite of the large point estimate obtained for the correlation coefficient, none of the other test statistics computed reports a significant result for any of the models. In particular, Kiefer's (1981) score tests and the transformed z-score based on the estimated correlation coefficient between the pseudo-residuals from the separate probit models yield extremely low test values.[13] A cautious approach is thus required for the male sample. On balance, the statistical evidence would appear to support the notion of independence in unobservables between the two processes but the findings for men contain some degree of ambiguity. The econometric implication of this finding is that estimation of separate univariate probits is permissible. Given the magnitude of the estimated correlation coefficient for the men we adopt a cautious approach and retain the bivariate modelling approach. However, for completeness, we also provide the analysis based on the univariate case to determine the degree of sensitivity in regard to this choice.

We now turn to a more detailed examination of the estimates reported in table 3, with a particular emphasis on the male coefficients in the first instance. In general terms, age and labour force status are strong determinants in both equations and are signed as anticipated. The per capita household welfare measure impacts positively on whether men consider migration but (more plausibly) in a negative fashion on trying to migrate. The district level wage is an important determinant of whether a man considers migration but exerts no independent influence on whether migration was attempted. In contrast, and given the cross-equation restriction imposed, the district level unemployment rate yields a well-determined effect for both equations. Both these local labour market proxies, when statistically significant, are signed in the anticipated direction.

In order to get a better feel for these effects, however, it is useful to interpret the FIML estimates in conjunction with the marginal effects reported in tables 4 and 5. Given the findings in regard to selection, we report marginal effects using both the bivariate probit model estimates (table 4) and the corresponding univariate probit model estimates (table 5).[14] The marginal effects reported in both these tables provide the effect of a change in the covariate of interest on the conditional probability of trying to migrate. Thus, on the basis of table 4, a male aged between 15 and 25 years is, on average and *ceteris paribus*, over 26 percentage points more likely to attempt external migration than someone aged over 50 years. An employed male is, on average and *ceteris paribus*, 6.5 percentage points less likely to attempt external migration relative to someone who is unemployed, and comparable effects are reported for farmers and the self-employed. A male residing in an urban settlement is, on average and *ceteris paribus*, 17 percentage

13 The values, however, are consistent with the well known inequality Wald > LRT > Score in finite samples.

14 The maximum likelihood estimates for the univariate probit models are not reported here to conserve space.

points less likely to try and migrate abroad relative to someone who is residing in a rural settlement.

Table 4: Conditional Marginal Effects for Attempting to Migrate – Using Bivariate Probit Models

	Pooled	Male	Female
Aged 15 – 25	0.2435*** (0.3769)	0.2653*** (0.0469)	0.1847*** (0.0525)
Aged 26 – 40	0.2003*** (0.0307)	0.1930*** (0.0324)	0.1679*** (0.0505)
Aged 41 – 50	0.1474*** (0.0269)	0.1340*** (0.0282)	0.1179*** (0.0459)
Aged 50+	*f*	*f*	
Male	0.2748*** (0.0269)	§	§
Married	0.0492* (0.0306)	0.0658* (0.0395)	0.0589*** (0.0226)
Unemployed	*f*	*f*	*f*
Employed	-0.0626*** (0.0183)	-0.0647*** (0.0232)	-0.0483* (0.0275)
Farmer	-0.0862*** (0.0237)	-0.0552** (0.0274)	-0.0880*** (0.0393)
Self-Employed	-0.0745*** (0.0237)	-0.0436* (0.0269)	-0.0692* (0.0398)
Temporary Layoff	-0.1050** (0.0489)	-0.0300 (0.0667)	-0.1568** (0.0713)
Household Size (Numbers)	-0.0285*** (0.0082)	-0.0376*** (0.0093)	-0.0123 (0.1364)
Per Capita Household Expenditure (Log)	-0.0400 (0.0302)	-0.0951*** (0.0357)	0.0878*** (0.0249)
Coastal Region	*f*	*f*	
Central Region	0.0349** (0.0147)	-0.0221 (0.0394)	0.0325 (0.0217)
Mountain Region	-0.0184 (0.0162)	0.0573 (0.0458)	-0.0489** (0.0252)
Tirana	-0.0173 (0.0190)	-0.0711 (0.0554)	0.0144 (0.0296)
Urban	-0.1257*** (0.0294)	-0.1698*** (0.0364)	0.0497* (0.0305)
District Level Unemployment Rate (%)	0.0036*** (0.0269)	0.0053*** (0.0012)	0.0010 (0.0010)
District Level Hourly Wage (log)	-0.0375 (0.0694)	0.0170 (0.0012)	0.0169 (0.0518)

Notes to table 4:

(a) *f* denotes omitted category in estimation.

An examination of the marginal effects for the continuous measures, which proxy for household welfare and local labour market conditions, contains important policy content. The conflict in signs on the per capita household expenditure measure was noted earlier but the conditional marginal effect of a change in this measure is negative for men. The estimated marginal effect suggests that a 5% rise in per capita household expenditure (our proxy for household welfare) reduces the conditional probability of attempting to migrate by, on average and *ceteris paribus*, about one-half of one percentage point. A percentage point reduction in the district level unemployment rate would reduce the conditional probability of attempting to migrate abroad by, on average and *ceteris paribus*, 0.5 of one percentage point. The local labour market hourly wage, a migration opportunity cost measure, exerts no independent influence on the attempt to migrate.

Table 5: Conditional Marginal Effects for Attempting to Migrate – Using Univariate Probit Models

	Pooled	**Male**	**Female**
Aged 15 – 25	0.2846*** (0.0613)	0.2800*** (0.0748)	0.3227*** (0.1202)
Aged 26 – 40	0.1922*** (0.0503)	0.1550*** (0.0548)	0.2737** (0.1130)
Aged 41 – 50	0.1196*** (0.0517)	0.0812 (0.0562)	0.1886* (0.1152)
Aged 50+	*f*	*f*	*f*
Male	0.2732*** (0.0273)	§	§
Married	0.0826** (0.0389)	0.1027* (0.0530)	0.0826 (0.0539)
Unemployed	*f*	*f*	*f*
Employed	-0.0858** (0.0387)	-0.1097** (0.0474)	-0.0477 (0.0600)
Farmer	-0.1049** (0.0505)	-0.0787 (0.0592)	-0.0903 (0.0934)
Self-Employed	-0.0482 (0.0519)	-0.0604 (0.0598)	0.0853 (0.0937)
Temporary Layoff	-0.2194** (0.1118)	-0.0904 (0.1423)	-0.3861** (0.1922)
Household Size (Numbers)	-0.0327*** (0.0080)	-0.0383*** (0.0093)	-0.0188 (0.0140)

	Pooled	**Male**	**Female**
Household Expenditure (Log)	-0.0457 (0.0297)	-0.0863*** (0.0351)	0.0663 (0.0494)
Coastal Region	*f*	*f*	*f*
Central Region	-0.0079 (0.0330)	-0.0200 (0.0398)	0.0450 (0.0518)
Mountain Region	0.0535 (0.0384)	0.0561 (0.0467)	0.0242 (0.0614)
Tirana	-0.0117 (0.0463)	-0.0723 (0.0562)	0.1139* (0.0698)
Urban	-0.1356*** (0.0390)	-0.1692 (0.0434)	-0.0022 (0.0799)
District Level Unemployment Rate (%)	0.0038** (0.0017)	0.0052*** (0.0020)	-0.0011 (0.0026)
District Level Hourly Wage (log)	0.0326 (0.0850)	0.0846 (0.1044)	0.0780 (0.1287)

Notes to table 5:

(a) *f* denotes omitted category in estimation.

In general, the estimated impact effects for the binary measures for women tend to be smaller in magnitude compared to men. In addition, the estimated effect for the household welfare measure is somewhat counter-intuitive given its positive sign, and neither of the marginal effects for the local labour market variables attains statistical significance at a conventional level. Nevertheless, there is a strong gender dimension to the phenomenon investigated here. The magnitude of one aspect of the gender effect can be seen in table 5 using the estimated impact effect for being male in the pooled model. In particular, being male increases the conditional probability of attempting to migrate by over 27 percentage points, on average and *ceteris paribus*.

This finding for gender prompts interest in whether there are other gender differences in coefficients in the reported specifications. In order to explore this in more detail, table 6 reports Wald test values for the difference in estimated coefficients across separately estimated gender models using equation [2] only.[15]

15 These differences are based on using univariate probit models but the findings are congruent with those obtained using the bivariate probit.

There is a strong differential in regard to the estimated constant terms reflecting the finding already alluded to in table 4. In addition, strong gender differences are noted for the household welfare proxy, the local unemployment rate, and more marginally for residing in an urban settlement. These results provide at least *prima facie* evidence that much of the gender differential in average conditional probabilities is attributable to 'treatment' rather than 'endowment' effects.

Table 6: Wald Tests for Gender Differences in Coefficients

	Tried
Constant Terms ~ χ^2_1	6.297** (0.012)
Age Controls ~ χ^2_3	1.653 (0.648)
Married ~ χ^2_1	0.952 (0.329)
Labour Force Status Controls ~ χ^2_4	4.463 (0.347)
Household Expenditure (Log) ~ χ^2_1	6.293** (0.012)
Regional Controls ~ χ^2_3	5.313 (0.150)
Urban Settlement ~ χ^2_1	2.690* (0.100)
District Unemployment Rate (%) ~ χ^2_1	3.697* (0.054)
District Level Hourly Wage (log) ~ χ^2_1	0.191 (0.662)
Household Size ~ χ^2_1	0.952 (0.329)

Notes to table 6:

(a) The Wald tests are computed on the basis of the univariate probit models.
(b) The Wald tests are computed as: $[db_k]'V^{-1}[db_k]$ where $V = [V_{Male}]_k + [V_{Female}]_k$ where db_k is the vector containing the coefficient differences between men and women for the k^{th} category of variables; $[V_{Male}]_k$ and $[V_{Female}]_k$ are the corresponding variance-covariance matrices for the coefficients for the k^{th} category for males and females respectively.
(c) ***, **, * denote statistical significance at the 1%,5% and 10% level respectively.
(d) The prob-values for the tests are reported in parentheses.

The assignment of the gender differential to these two components is now interrogated more thoroughly by using the decomposition methodology outlined in an earlier section. Table 7 reports actual and simulated conditional probabilities for the two gender groups. The sample average conditional probability that an Albanian man attempts to migrate is computed at 0.582. The corresponding figure for an Albanian woman is 0.322.[16] The raw or unadjusted differential in conditional probabilities is thus 0.26. This table also contains two simulated or counterfactual conditional probabilities. The third row of the table contains the average simulated conditional probability for men if the female coefficient structure was imposed. Under this simulation exercise, the estimated average conditional probability would reduce to 0.321 – a reduction for men of 26 percentage points. The fourth row simulates the conditional probabilities for women if the male coefficient structure was imposed. In this case, the average conditional probability would rise to 0.616 – almost a doubling in the conditional probability. It should be noted that the differences between the average simulated conditional probabilities are almost invariant to whether bivariate or univariate probit models are used.

Table 7: Actual and Simulated Conditional Probabilities for Attempting to Migrate

Conditional Probabilities	Expression	Value
Bivariate Probits		
$[N_m]^{-1}[\Sigma F(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_m, \mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_m; \hat{\rho}_m)] \div [N_m]^{-1}[\Sigma\Phi(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_m)]$	[4]	0.582
$[N_f]^{-1}[\Sigma F(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_f, \mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_f; \hat{\rho}_f)] \div [N_f]^{-1}[\Sigma\Phi(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_f)]$	[5]	0.322
$[N_m]^{-1}[\Sigma F(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_f, \mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_f; \hat{\rho}_f)] \div [N_m]^{-1}[\Sigma\Phi(\mathbf{x}'_{im}\hat{\boldsymbol{\beta}}_f)]$	[6]	0.321
$[N_f]^{-1}[\Sigma F(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_m, \mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_m; \hat{\rho}_m)] \div [N_f]^{-1}[\Sigma\Phi(\mathbf{x}'_{if}\hat{\boldsymbol{\beta}}_m)]$	[7]	0.616
Univariate Probits		
$[n_m]^{-1}\Sigma\Phi(\mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_m)$	[4′]	0.582
$[n_f]^{-1}\Sigma\Phi(\mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_f)$	[5′]	0.322
$[n_m]^{-1}\Sigma\Phi(\mathbf{z}'_{im}\hat{\boldsymbol{\gamma}}_f)$	[6′]	0.306
$[n_f]^{-1}\Sigma\Phi(\mathbf{z}'_{if}\hat{\boldsymbol{\gamma}}_m)$	[7′]	0.576

Notes to table 7:

(a) See text for details.

16 These figures can be derived using either the sample average of the conditional probabilities computed with the FIML estimates or they can be directly computed from the raw data. The numerical differences are negligible as one would expect.

Table 8 uses the actual and simulated conditional probabilities and re-casts them in terms of 'treatment' and 'endowment' effects. Using either male or female coefficients, the 'treatment' effect (or coefficient structure) dominates. The geometric average of the two 'treatment' effects yields an estimate of 0.277. Thus, men are 27.7 percentage points more likely to attempt migration than women, on average and *ceteris paribus*. This estimated 'treatment' effect is broadly consistent with the impact effect for gender reported for the pooled models in tables 4 or 5. This analysis confirms that gender differentials in observed characteristics are unimportant in explaining the observed gender differential in our outcome of interest.

Table 8: Decomposition of Gender Differential in Conditional Probabilities For Attempting to Migrate

	Using Male Coefficients	**Using Female Coefficients**
Total Gender Differential	0.260	0.260
Using Bivariate Probits:		
'Endowment' Effects	-0.034	-0.001
'Treatment' Effects	0.294	0.261
Using Univariate Probits:		
'Endowment' Effects	0.006	-0.016
'Treatment' Effects	0.254	0.276

Notes to table 8:

(a) For Bivariate Probits:
(i) Using Male Coefficients: Total = [4] – [5]; Endowment = [4] – [7]; Treatment = [7] – [5];
(ii) Using Female Coefficients: Total = [4] – [5]; Endowment = [6] – [5]; Treatment = [4] – [6];
(see text for further details).
(b) For Univariate Probits:
(i) Using Male Coefficients: Total = [4′] – [5′]; Endowment = [4′] – [7′]; Treatment = [7′] – [5′];
(ii) Using Female Coefficients: Total = [4′] – [5′]; Endowment = [6′] – [5′]; Treatment = [4′] – [6′];
(see text for further details).

CONCLUSIONS

This study examined the factors that determine whether an individual attempts to migrate using data drawn from the 2002 Albania Living Standards Measurement Survey. The analysis used responses to questions from the migration module of this survey. In particular, respondents were asked had they ever considered migrating abroad and, if so, had they ever tried to migrate abroad. The sequence of questions and the fact that only a sub-sample of respondents provided information on whether they attempted to migrate warrants caution in the econometric treatment of these responses. An exclusive focus on those individuals who attempt to migrate neglects the fact that they potentially constitute a self-selected sub-sample. The econometric problem encountered is redolent of the more conventional selectivity bias problem within a linear regression framework and requires a more sophisticated econometric treatment.

In order to deal with the econometric issues generated by this response sequence, we use a bivariate probit model that controls for selectivity bias. This involves estimating an equation for whether an individual considered migrating using the full sample of individuals and, conditional on this, an equation for whether an individual attempted to migrate using the sub-sample of individuals who considered migration. The bivariate probit comprises two standard probits but incorporates a cross-equation correlation in unobervables and requires estimation by the FIML procedure. Our empirical findings are ambiguous for men but suggest, on balance, no selection effects for either gender groups. Thus, those that attempt migration could be taken to comprise a random drawing from the overall sample of Albanian men and women in our study. The independence in the relationship between unobservables across the two equations for both men and women implies that the use of separate univariate probits is permissible.

An important theme of our paper was a focus on gender differences in attempting to migrate. The estimates from the bivariate and univariate probit models were used to decompose the total gender differential in the average conditional probabilities into 'endowment' and 'treatment' components. Our analysis suggests that, regardless of which gender-specific coefficients are used to compute the index number, the 'treatment' effect dominates. Gender differences in characteristics exert little influence in explaining the observed gender disparity in the migration phenomenon of interest to us, with almost all of the difference in conditional probabilities of attempting to migrate being accounted for by treatment effects. Women respond much less strongly to adverse labour market conditions, for example, than men. Thus, the gender gap is largely attributable to 'femaleness' and 'maleness', which in the Albanian context is strongly determined by culture and tradition. This particular gap is thus unlikely to change radically in the near future as these factors tend to evolve in most countries at a glacial pace.

Our estimates also reveal that those that attempt to migrate are largely drawn from rural settlements, which emphasizes that much of the Albanian migration phenomenon has its roots in the disadvantages that exist in rural parts of the country. Our analysis revealed that the rural/urban dichotomy was a considerably stronger determinant of migration attempts than any of the regional controls.

The unemployed were more likely than others to attempt migration and this was true across both gender groups. In addition, per capita household living standards and local labour market conditions provided potent incentives for men to attempt to migrate. In order to get a feel for the magnitude of these two particular relationships, it might prove useful to exploit the econometric estimates for the male equations to provide some simulations. These will help to provide some perspective for the estimated marginal effects for per capita household expenditure (a proxy for household wealth) and the district level unemployment rate. In order to undertake this exercise we note that the Albanian male population aged between 15 and 60 years was estimated to be 903,193 in 2001. This is the relevant age-group for the econometric analysis conducted and this figure is used to undertake our simulations. If per capita household expenditure rises by 10%, according to the bivariate probit model estimates from table 2 and holding all other factors constant, the total number of Albanian males that would attempt to migrate would fall by 8,589. A one-percentage point fall in the district level unemployment rate across all districts, on the other hand, reduces the total number of men who attempt to migrate by 4,787. A doubling of the unemployment rate provides the more pessimistic side of this coin. Under such circumstances, the numbers that would attempt to migrate would rise by over 57,000 – almost seven percent of the total Albanian male population within the relevant age-group.[17]

Our analysis has demonstrated that there is some utility to using more indirect information on migration to inform policy. In our analysis of the factors that influence attempts to migrate, the usual candidate variables emerged. It is worth noting that living standards and local labour market conditions were prominent

17 In using the univariate probit estimates the corresponding numbers are 7,795 for a 10% rise in per capita household expenditure, and 4,696 and 56,359 for the two unemployment rate simulations, and are broadly comparable in magnitude.

drivers of this phenomenon. In order to enhance the policy dimension of our work, some knowledge of the relationship between attempting to migrate and succeeding to migrate would be useful. This would allow a more precise judgment on how the size of migrant flows responds to economic conditions. It is gratifying to note, though, that our econometric results are consistent with the predictions of standard economic behavioural models.

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APPENDIX

Table A1: Description of Variables Used in the Analysis

Variable	Variable Description
Considered	=1 if the individual considered migrating; = 0 otherwise.
Tried	=1 if the individual tried and failed to migrate; = 0 otherwise.
Aged 15 – 25 years	=1 if the individual is in this age-group; = 0 otherwise.
Aged 26 – 40 years	=1 if the individual is in this age-group; = 0 otherwise.
Aged 41 – 50 years	=1 if the individual is in this age-group; = 0 otherwise.
Aged 50+ years	=1 if the individual is in this age-group; = 0 otherwise.
Male	=1 if the individual is male; = 0 otherwise.
Married	=1 if the individual is married; = 0 otherwise.
Primary or Less	=1 if the individual has no education or achieved eight or less primary grades; = 0 otherwise.
Secondary	=1 if the individual achieved secondary level; = 0 otherwise.
Vocational	=1 if the individual achieved vocational level; = 0 otherwise.
University	=1 if the individual achieved university level; = 0 otherwise.
Born in the Municipality	=1 if the individual was born in the municipality; = 0 otherwise
Unemployed	=1 if the individual is unemployed; = 0 otherwise.
Employee	=1 if the individual is an employee; = 0 otherwise.
Farmer	=1 if the individual is a farmer; = 0 otherwise.
Self-Employed	=1 if the individual is self-employed; = 0 otherwise.
Temporary Layoff	=1 if the individual is a temporary layoff; = 0 otherwise.
Household Size	The total number of individuals in the household.
Children in Household: Aged ≤ 4 years	=1 if the individual's household has any children aged between four or less; = 0 otherwise.
Children in Household: Aged 5 ≤ years ≤ 8	=1 if the individual's household has any children aged between five and eight; = 0 otherwise.
Children in Household: Aged 9 ≤ years ≤ 14	=1 if the individual's household has any children of age between nine and fourteen; = 0 otherwise.
Log of Total Household Per Capita Consumption:	The logarithm of the total (monthly) expenditure of the household.

Residence Dwelling Area: ≤ 69 Sq.Metres	=1 if the area of the dwelling is less than 69 square metres; = 0 otherwise.
Residence Dwelling Area:70 ≤ Sq.Metres ≤ 130	=1 if the area of the dwelling is between 70 and 130 square metres; = 0 otherwise.
Residence Dwelling Area:Sq. Metres > 130	=1 if the area of the dwelling is over 130 square metres; = 0 otherwise.
Residence Constructed after 1990	=1 if the dwelling was built after 1990; = 0 otherwise.
Central	=1 if the individual resides in the Central region; = 0 otherwise.
Coastal	=1 if the individual resides in the Coastal region; = 0 otherwise.
Mountain	=1 if the individual resides in the Mountain region; = 0 otherwise.
Tirana	=1 if the individual resides in Tirana; = 0 otherwise.
Urban	=1 if the individual resides in an urban settlement; = 0 otherwise.
District Level Unemployment Rate (%)	The unemployment rate (%) in the district where the individual resides.
District Level Hourly Wage (log)	The average of the logarithm of the hourly wage in the district where the individual resides.

Table A2: Summary Statistics for the Variables Used in the Analysis

	Pooled Sample		Male Sample		Female Sample	
	Considered	**Tried**	**Considered**	**Tried**	**Considered**	**Tried**
Considered	0.3292	§	0.4234	§	0.2254	§
Tried	§	0.4980	§	0.5835	§	0.3213
Age (years)	36.486	34.884	38.061	35.178	34.754	34.277
Male	0.5239	0.6739	1.000	1.000	0.000	0.000
Married	0.7531	0.7401	0.7677	0.7215	0.7370	0.7784
Primary or Less	0.5355	§	0.5012	§	0.5732	§
Secondary Educ.	0.1886	§	0.1925	§	0.1844	§
Vocational Educ.	0.1698	§	0.1873	§	0.1507	§
University Educ.	0.1060	§	0.1190	§	0.0918	§
Born in the Municipality	0.6655	§	0.7339	§	0.5902	§
Unemployed	0.1210	0.1625	0.1214	0.1521	0.1204	0.1838
Employee	0.3487	0.3871	0.4252	0.3982	0.2645	0.3643
Farmer	0.4256	0.3350	0.3182	0.3159	0.5438	0.3746
Self-Employed	0.0889	0.1008	0.1246	0.1222	0.0496	0.0567
Temporary Layoff	0.0159	0.0146	0.0106	0.0116	0.0217	0.0206
Household Size	§	5.1081	§	5.1879	§	4.9433
Children in Household: ≤ 4 years	0.3113	§	0.3157	§	0.3064	§

	Pooled Sample		Male Sample		Female Sample	
	Considered	**Tried**	**Considered**	**Tried**	**Considered**	**Tried**
Children in Household: 5 ≤ years ≤ 8	0.3070	§	0.2964	§	0.3187	§
Children in Household: 9 ≤ years ≤ 14	0.4639	§	0.4579	§	0.4706	§
Log Household Consumption per Capita	8.8429	8.8648	8.8531	8.8244	8.8317	8.9485
Residence Dwelling Area: ≤ 69 Sq.Metres	0.4938	§	0.4924	§	0.4954	§
Residence Dwelling Area: 70 ≤ Sq.Metres ≤ 130	0.4628	§	0.4636	§	0.4620	§
Residence Dwelling Area:Sq. Metres > 130	0.0413	§	0.0422	§	0.0403	§
Residence Built post-1990	0.2261	§	0.2306	§	0.2211	§
Coastal	0.2775	0.2605	0.2728	0.2444	0.2827	0.2938
Central	0.2749	0.3087	0.2703	0.3051	0.2800	0.3162
Mountain	0.3044	0.2756	0.2858	0.3001	0.3249	0.2251
Tirana	0.1431	0.1552	0.1711	0.1505	0.1123	0.1649
Urban	0.4573	0.5249	0.5076	0.5012	0.4020	0.5739
District Unemp. Rate (%)	11.880	12.911	12.118	13.115	11.619	12.490
District Level Hourly Wage (log)	4.2761	4.2825	4.2902	4.2773	4.2607	4.2931
Observations	5423	1785	2841	1203	2582	582

Notes to table A2:

(a) § denotes not used in estimation.