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to my parents, who have put up with my

ups and downs throughout this journey

to all my relatives and significant ones

who have contributed in shaping who I am

to all the people I have met during this experience

and who have contributed to

this final result or my personal growth

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SUMMARY

In developing countries, poverty remains one of the most significant social problems the policy maker has to deal with. The lack of economic resources has a greater impact on the overall level of material well being than on the subjective one. In fact, capturing individual happiness through subjective well being (SWB), it has been consistently found that the process of increasing material well being is largely decoupled from the one of becoming happier (Easterlin, 1974). Several studies have demonstrated that in developing countries wage receipts are the most important sources of income the poor can rely on to escape poverty (see references in Fields, 2004). This is largely due to their own labour being the poor's almost unique asset. In turn, the human capital literature (Mincer, 1974) has ascertained the strong connection between the salary obtained in adulthood and the amount of education received at a younger age. Therefore, it seems interesting to investigate the variables associated to the participation of youngsters in schooling or labour (i.e., the activity which most often competes for pupils' time.

The present thesis is a collection of three contributions which investigate children's schooling and labour participation, the correlates of subjective well being and the differences in wage levels in Morocco and Albania. The papers try to carry out empirical analyses of these issues based on the most recent developments in the relevant literature and to contribute to the empirical methodology commonly used.

The first paper deals with the schooling opportunities and labour engagement of Moroccan children in the period 1998/1999. This essay employs the theoretical model in Glick (2008) to identify the variables, which might be affected by policy actions, correlated in a statistically significant manner to the probability of a child being involved in any labour activity or schooling initiative.

Employing the Moroccan Living Standard Measurement Study (LSMS) survey, this work estimates two probit models with the same explanatory covariates to compare and contrast the different empirical patterns which characterise the two binary decisions. Two different models seem appropriate since the sample proportion of pupils being either idle or engaged in both activities is negligible. The distinctive contribution of this work includes the elaboration, using dedicated software, of anthropometric data available in the LSMS to obtain the child's Body Mass Index

(BMI) (Ayalew, 2000). Moreover, the marital status of the household head is employed to verify whether an unsettled family situation limits the pupil's achievements and favours her involvement in work activities (Francesconi *et al.*, 2005). Finally, since females frequently face harsher living conditions in developing countries, the results are presented separately for males and females.

The main findings report that the accumulation of education appears the highest at young age while being adequately fed increases the probability of being out of school and into work. It is surprising that adults endowed with high quantities of human capital have smaller impact on pupils' attendance compared to the one associated with adults who enjoy lower qualifications. Females' school attendance plummets, compared to the men's, after the maximum probability of being in school has been secured. Somewhat surprisingly, the age of the household head (i.e., a proxy for the declining propensity to send the youngsters to school the older the adult) is instead associated with a higher probability of a pupil being out of work. Similarly, although maybe biased by reverse causality, the higher the yearly household expenditure the higher the probability of being working. The apparent transmission of the present working status of the old generation to the current condition of the young one is a worrying result. The gender specific child labour models seem to highlight that the time away from work is a luxury good and females are more likely to be stuck at work.

The second paper addresses the broader issue of characterising the SWB levels of a sample of Albanians in the year 2005. This work, which constitutes an extended and preliminary version of a paper written with Dr. Barry Reilly and Dr. Julie Litchfield and currently under consideration for the publication on a special issue of the Journal of Economic Behaviour and Organisation, has a twofold purpose. On the one hand, it contributes to the literature on Albania since it is, to the best of my knowledge, the first attempt to assess SWB in Albania. Moreover, it tries to build on a compendium of the SWB literature since it selects the set of explanatory variables out of the covariates most commonly employed in similar studies. On the other, the paper improves on the ordered probabilistic modelling of SWB that the literature has implemented so far.

Making use of the 2005, and most recent, wave of the Albanian LSMS (ALSMS) the established variables in SWB equations are selected alongside some for religious affiliation, the government's involvement in welfare provision, the collapse of the pyramid schemes, the number of close friends and the increase in local population. Treating the responses to a SWB question as ordered in nature, the paper notes how the contributions employing ordered models do not properly assess their theoretical underlying hypotheses. To fill this gap, the paper focuses on ascertaining whether the model is mis-specified in its functional form, the error term is not normal, its variance is not constant and the thresholds are heterogeneous for different characteristics controlled for (Machin and Stewart, 1990). Among the possible violation of these hypotheses, heteroscedasticity is the most worrying one since, in an ordered probit, it leads to the estimated betas being biased and inconsistent. Therefore, the commonly applied Huber (1967) correction for heteroscedasticity is rather meaningless. The present work contributes to modelling, rather than curing, heteroscedasticity drawing on the insights of Stewart (1983) as further developed by Caudill and Jackson (1993). Moreover, particular attention is dedicated to keep the model parsimonious according to statistical tests.

The statistical testing carried out on a traditional ordered probit model verifies that only the hypothesis for the correctness of the pseudo functional form is upheld. Applying the proposed modelling, the validity of all of them is restored. The downside of this methodology is the diminished statistical significance of the estimates due to the standard errors being larger. The measure of per capita household consumption is found to be large and highly significant. This is in contrast with much of the literature on SWB which finds small effects of such a variable but this divergence from the norm can be easily explained by the low national standards of living. Moreover, the psychological scars of the collapse of the Albanian pyramid schemes are still vivid in the mind of the people and Muslims enjoy a higher satisfaction with life due to their religion being the majority one in the country. An individual living in a community which benefits from the presence of a collective organisation will have higher satisfaction with life while the occurrence of thefts induces an opposite change.

The third paper remains focused on Albania and analyses the differences in the wage levels, enjoyed by a sample of adults involved in their primary job, as deriving from a number of individual gaps. Using a human capital model (Mincer, 1974) as the main interpretational framework it tries to disentangle the contribution of differences in gender, human capital endowments, occupation characteristics, geographical location, migration possibilities and industry affiliation to explaining the level of wage inequality.

After having reviewed the recent literature on the correlates or determinants of inequality, the paper provides some background information on the Albanian post-communist development path and the resulting current state of the economy. This constitutes a necessary piece of knowledge which might help in commenting the results of the two Mincer equations, estimated using OLS, for the year 2002 and 2005. The explanatory variables considered include the most popular ones in the literature and some capturing peculiar national phenomena. The paper implements a very granular classification of the individual's industrial affiliation instrumental to the identification of which productive sector might endow its employees with the monetary resources granting higher material well being.

The operating sample for 2002 is characterised by a Gini coefficient, calculated on the level of real hourly wages, equal to 0.34 while the one for the year 2005 expresses a level of inequality of 0.28. The heteroscedasticity robust estimates detect a persistent wage premium accruing to males which is larger in 2005 than in 2002. Moreover, in both years higher wages are paid to individuals who hold a university degree or higher. International migration has affected Albania significantly. This experience seems to have endowed those who have moved abroad, and then returned, with skills and education which appear to grant higher wages – compared to the ones of non migrant – on the domestic market. Yet, the estimated premium is lower in 2005 compared to 2002. Quite surprisingly, the public sector's workers are paid less than those in the private one. This gap might reflect genuine differences in the levels of productivity in these positions rather than the existence of position rents accruing to those in more protected jobs. The recently achieved higher productivity in the public sectors might be associated with the negative wage gap being smaller in 2005 than in 2002. The most surprising and somewhat puzzling result, in face of the established literature on the national economic conditions, is the consistent evidence, across the two years, of almost every sector paying lower salaries compared to agriculture. This seems to contrast with the perception that the Albanian agriculture is underdeveloped and a refuge for the poor. On the contrary it might be enjoying a sort of within country comparative advantage with respect to sectors which might be more dependent upon a well developed infrastructural system. The latter having improved in recent years, also in concurrence with several trade liberalisation initiatives, might have helped in reducing the shortfalls between 2002 and 2005.

SCHOOL ATTENDANCE, CHILD LABOUR AND GENDER BIAS IN MOROCCO

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Abstract. *Using cross-sectional data and employing probit models, the paper investigates child school attendance and labour in Morocco. Results indicate that education is a luxury good which can be purchased more easily by the non-poor. Moreover, a child is more likely to be in school if he/she is residing in an urban area and is the son/daughter of the head of the household. However, the Body Mass Index (BMI) - used to account for the impact of proper nutrition status on attendance probability - deters school attendance. More detailed empirical evidence is obtained by estimating the same type of models on gender based subsamples. Poverty status, household size and the provincial average of time taken to travel to school in minutes lessen the probability of female school attendance. If a female child lives in a household headed by a female, the pupil should have higher likelihood of school attendance. Besides being a quite common result, this occurrence can be employed to devise a policy initiative of disbursing to the female head a monetary or in-kind transfer devoted to pupils' education which can be most cost-effective in closing the gender gap as required by the second Millennium Development Goal (MDG). The probability of a male going to school is limited by the significant role of some regional location dummies and living with an aged household head. The variables which are significant in both single-genders models generate higher probabilistic effects for females compared to males.*

As expected, the child labour models estimate coefficients which are most of the time opposite in sign to the ones in the educational. The variable which is consistently associated with a negative influence on the dependent variable is the age of the household head. The significance of the yearly per capita expenditure in the child labour investigations might be driven by some endogeneity or the inclusion of asset variables. Besides the econometric problem, it undermines the confidence attached to the implementation of a cash/in-kind transfers devoted to curb child labour and promote school attendance. Industry localization effects and complementarity in within-household labour choices are present. The investigation carried out on gender-specific samples finds a inverted U shaped female trajectory in age probably due to early pregnancy. Moreover, poverty condition improves males probability of working while it diminishes female's school participation.

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Introduction

Child labour and pupils' school attendance are crucial policy issues particularly in developing countries. According to ILO (2006), in 2004 166 million children aged 5-14 years were labourers while 74 million were involved in hazardous work. Despite the "large numbers" the two stocks have fallen in size by 11 and 33 percent respectively since the last global survey in 2000 (ILO, 2006:6-7). In 2001/02 115 million children, in primary age, were out of school (UIS, 2005) while UNICEF (2004) reports that 121 million children[1] were denied the right to universal education.

It could be argued that at any point in time the stock of child labourers and out of school pupils is determined by the dynamics of the factors which can increase or reduce the likelihood child labour. These incentives and mitigating forces occur at the international or national and household or family level[2].

At the "macro" level, a higher incidence of child labour can be due to perverse incentives arising from economic policies centred on national growth. The attraction of foreign direct investments or multinational corporations in developing countries might lead to subsidiary branches employing locally sourced child labour. Hence, greater global integration may have mixed effects, at least on the poor through the rise in national production and labour demand (Ravallion and Wodon, 2000). Nonetheless, Cigno *et al.* (2002) in a cross section study do not find any support for increased trade between countries raising, *per se*, the incidence of child labour. The norms which are likely to deter child labour at the "macro" level are the 1990 Convention on the Rights of the Child, international trade sanctions toward countries where child labour is a relevant phenomenon, a ban on child labour and a minimum wage policy.

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[1] Despite the absence of a clear indication of the geographic reference for this figure (i.e. worldwide, in Africa or in Sub-Saharan Africa), it is believed that this count corresponds to a worldwide scale. The quoted figures are likely to underestimate the true extent of the child labour phenomenon heavily. Admassie (2003) attributes the measurement error to official surveys being unable to record those pupils who work at home, in subsistence agriculture and attend unpaid jobs in the informal sector. Additional distortion arises from child labour being illegal in several countries such that parents underreport on their children's work.

[2] For expositional purposes the former two aggregation levels will be considered as the "macro" level of the problem while the latter two the "micro" level.

The Convention on the Rights of the Child is probably the most authoritative piece of international jurisprudence which protects the children's human rights to spend their youth playing and socializing[3], avoiding dangerous work practices[4] (Admassie, 2003). International trade sanctions might be ineffective or even counterproductive. Bhalotra (2003) and Admassie (2003) suggest that "the proportion of output [from child labour] exported is rather minimal" (Admassie, 2003:169) in several subsistence economies of the African continent to make any significant difference. Jafarey and Lahiri (2002) identify a set of conditions, which might be common to several countries, that could lead to a rise in the amount of time spent working at the expense of schooling time. On the other hand, Jafarey and Lahiri (2002) demonstrate that the existence of a national credit market can mitigate the extent of child labour increasing schooling attendance.

A national ban on child labour is likely to hurt those household which are extremely poor or use child labour as an essential coping strategy in presence of market failures (Beegle *et al.*, 2006). In these conditions, a timely government transfer compensating for forgone income could preserve household consumption levels while inducing higher child schooling attendance (Ravallion and Wodon, 2000). Grimm (2005) suggests that proactive education policies may be a substitute for monetary transfers in that making the knowledge distribution more even could diminish the country's poverty and inequality incidence.

Finally, a minimum wage policy would be welcomed by the adults participating in a labour market largely dominated by young participants. Due to the children's limited marginal productivity of labour (Patrinos and Psacharopoulos, 1997) and the displacement of adults' work (Admassie, 2003), the wage rate of the latter is likely to be lower[5] and their unemployment rate higher than the ones that would occur in an adult's only labour market. Therefore, the probability of adults' poverty is enhanced (Patrinos and Psacharopoulos, 1997). The effectiveness of a minimum wage policy in

[3] Patrinos and Psacharopoulos (1997) promote the controversial idea that it is indeed work that "… may help in the process of socialization, in building self-esteem and for training" (Patrinos and Psacharopoulos, 1997:388).

[4] The Convention states that a pupil should not undertake "[…] any work that is likely to be hazardous […]" (Article 32) while should be involved in "primary education compulsory and available free to all" (Article 28) (UN, 1989 cited in Gibbons *et al.*, 2003:1).

[5] Admassie (2003) reports that Bequelle and Boyden (1988) note that children earn less than adults even if both perform the same tasks and the former do not enjoy fringe benefits, insurance or social security payments. The latter income streams might constitute a sizeable percentage of the remuneration of the adults' labour and provide significant protection in case of unexpected unfavourable events.

limiting the extent of child labour is tied to the incidence of wage child labour. It is well known that children in developing countries are more likely to attend to household and farm chores (i.e., childminding, fetching water and wood or herding cattle, ploughing and seeding) than to work for a monetary or in kind wage.

The household economics literature frequently explains the children's participation to income generating activities using the luxury axiom. According to it if the non-child labour income falls significantly the household head will send the children in the household to work[6]. The extent of child labour can be contained by national governments deploying transfers to poor households to relax their budget constrain allowing more pupils to attend school without having to combine it with work. Additional schooling enhancing policies are those intended to reduce the opportunity and additional cost of education. They include the provision of good infrastructures, high quality schools and even child care initiatives. The latter are particularly important given that childminding is likely to compete directly with schooling time (i.e., to be carried out at home at a time when school is on) while chores like fetching water and wood compete for children's spare time (i.e., can be carried out on the way to or from school) (Admassie, 2003).
Moreover, the preference for child labour compared to schooling can be due to old parents failing to perceive correctly the importance of education for the pupils' development and the connected future returns (Mincer, 1974). In this vein, child labour can be seen as a practise promoting human capital dis-investment (Admassie, 2003) and perpetuating poverty in the future generations (Moser, 1996 cited in Ersado, 2005).

Gender bias is an additional problem which hampers many development countries. Because of its relevance, the United Nations places the provision of equal opportunities for both genders at the second place in the list of the MDGs. Despite recent progress in fostering gender equality, in several countries of Africa, South Asia and the Middle East, girls' access to education is still inferior compared to males' (Lewis and Lockheed, 2006 cited in Glick, 2008; Glick, 2008). This happens despite the common finding that female education guarantees, simultaneously, higher private

[6] The same axiom can be proposed for the schooling decision. In this case the child is going to school only if the household income is above the poverty threshold.

and social returns. Educated females will tend to have lower fertility, communicate health and sanitation knowledge to their children, and be keener on sending them to school (among others Schultz, 2002 cited in Glick, 2008). Likewise, the reduction in gender bias will raise the cohesion of the national population inducing higher female well-being, aspirations of increasing future income, lower poverty persistence and, possibly, the achievement of broader political rights. Gibbons *et al.* (2003) suggest choosing appropriate empirical measures so that the true extent of gender discrimination is captured in survey data and the policy recommendations are therefore more adequate. Culture and norms, religion and more "economic" reasons may have caused and perpetuated gender bias. The "economics of gender bias" can be largely due to parents evaluating the pupils' future transfers to them, the chances of each child finding a suitable job or to marry a desirable spouse (Kamaruddin, 2006)[7]. The World Food Programme (WFP, 2001) attributes gender bias to the distance of schools from students' homes, lack of drinkable water and/or single gender toilets, the girls' contribution to domestic chores and the fathers' religious/social concerns that young females at school will be interacting with males, possibly of an older age.
The form of gender bias considered here supposes that the biased pupil takes part in economic activities that will sustain household income while the preferred one attends school.

This paper aims to identify policy variables which are correlated, in a statistically significant way, with schooling attendance and child labour in Morocco using cross-sectional data[8]. It will do so exploiting the theoretical model in Glick (2008) which provides an interesting neoclassic framework to deal with the adults' allocation of children's time to a combination of school, work and leisure. Because it explicitly allows for the differential treatment of the males and females' condition, it is believed to be particularly suitable to the investigation of gender bias. Empirically, it estimates two probabilistic models for schooling and child labour participation using a

[7] For a formalization of the elderly insurance component of gender bias see the model in Glick (2008) or the discussion proposed in Section 2.
[8] The paper relies on some of the covariates most commonly used in studies of child labour and schooling. Part of the literature considers some of them as determinants of true causal relationships (Ersado, 2005). The paper will make clear which could and have been considered endogenous to the dependent variables. Because of the limited access to instrumental variables to solve all the econometric problems highlighted, the paper is "downgraded" to an analysis of statistically significant correlations.

compendium of the RHS variables more commonly employed in the established literature. Moreover, data from the anthropometric section of the household survey are elaborated using a specialist software and dummies for the marital status of the household head are included to account for the presence of fragile household environments. The use of the same set of covariates in modelling school attendance and work participation aims at highlighting which variables affect each probability and determining the difference in magnitude for those significant in both. A statistical test supports the case for exploring the impact of the covariates across single gender models.

The paper is organised as follows. Section 1 summarizes the Glick (2008) model and discusses recent empirical probabilistic models; Section 2 reports the econometric approach implemented here; Section 3 describes the data and the covariates employed in the models; Section 4 interprets the results discussing the marginal/impact effects (henceforth ME/IEs, respectively); the penultimate section summarizes the findings while the last one draws some policy implications.

1. Some insights into the issues

When setting out to investigate the characteristics of children's schooling attendance and working participation two major issues come to the fore and require careful handling. In the first instance, children's allocation of "disposable time" (i.e., all the hours in a day which are not reserved to sleep) is not completely in their hands. It is likely that either their parents or the head of the household, being pupils' main source of economic means and all-round care, play a decisive part in determining the final "optimal" allocation of time (Becker, 1974). Secondly, in achieving the "optimum"[9] the family (household) guided by its head is likely to behave as a sole entity composed of different individuals with distinctive needs.

Both issues are present in the model Glick (2008) compiles and which is believed to be proficiently employable in the present study.

[9] The equilibrium is optimal in the sense that it arises from the maximization of a particular utility function (see below) according to the traditional neoclassic founding hypotheses.

This model is preferred to the family/household ones proposed by Becker (among others 1974; 1976; 1981) and McElroy and Horney (1981) and Chiappori (1988).

The models in Becker (1974; 1976; 1981) appear to be unsuitable to be employed here because they largely resort on altruism. In Becker (1974) altruism is the behaviour by virtue of which the reduction in the transfers from the household head to one of the member does not change the distribution of resources among all the household members. A compensatory mechanism within the household will make sure that nobody experiences unpleasant feelings. This very feature of Becker's models seems to rule out the possibility of some household members experiencing an inferior condition.

Likewise, the intra household allocation of resources achieved through a bargaining process is the main feature of the models in McElroy and Horney (1981) and Chiappori (1988). In these models, the final household equilibrium is obtained comparing the utility achieved in the household equilibrium and living alone outside the household[10]. These models would require the researcher to model the utility levels of children outside the household. Since the present paper does not model and define the "threat" points in a model *a la* McElroy and Horney (1981) and Chiappori (1988), the intra household bargaining models are disregarded to use the Glick (2008) model instead.

In fact, Glick (2008) seems particularly suitable to be used in empirical studies which want to estimate child schooling and labour participation accounting for the role of parents, the distinction between genders and the allocation of time to different activities. It is a two-period model of the parents investing in their sons' and daughters' education. The adults work in the first period to finance household consumption. In the second period they expect to live out of the children's direct transfers and wealth. Therefore, the parents' utility can be expressed as:

$$U = U_1(C_1) + U_2(C_2, W_g, W_b) \quad (1)$$

where C_t is the consumption at time *t* and W_g and W_b are the second period wealth of the female and male children respectively. First period parents' utility depends on

[10] Within models for marriage, the utility of the husband and wife living together as a family is contrasted with the utility both of them would experience in case they were divorced.

their contemporary consumption while the second period one is an indirect function of consumption through the transfers from their dependents and the genuine desire of having them enjoying higher wealth. In turn, children's wealth in period two depends on the investment in schooling early in life (S_i) and the experience accumulated in productive work (L_i) according to the following general formula:

$$W_i = W(S_i, L_i) \qquad i = \text{b, g} \tag{2}$$

According to equation (2) labour is a source of wealth and remunerable skills like schooling is. For this matter, the Glick (2008) model embodies the relevance of "child labour" postulated in Patrinos and Psacharopoulos (1997).

Supposing that parents do not have access to consumer credit, the first period consumption can be expressed as the contemporary income net of the schooling expenses:

$$C_1 = Y_p + Y_g\left(L_g\right) + Y_b(L_b) - P_s\,S_g - \;P_b S_b \tag{3}$$

where Y_p is the income of the parents, Y_g and Y_b are the contribution of the male and female pupils in the household to first period consumtion through their labour earnings (L_i) and P_i is the direct (monetary) cost of one unit of schooling which is allowed to vary across gender. Second period consumption depends on children's wealth (W_i) and the shares of "remittances" each pupil, according to her gender, pays back to the parents (r_i):

$$C_2 = r_b W_b + r_g\,W_g \tag{4}$$

The optimal quantity of education can be derived applying the usual maximization of (1), subject to the (2), (3), (4) and individual overall time constraint, with respect to S_i and L_i. Incidentally, the equilibrium will be characterized by the equality between the marginal benefits from education and its costs. A change in the relative prices of both consumption and investment goods, in the first period, will determine the actual opportunity cost of education allowing for a change in the optimal demand levels in equilibrium.

The model postulates an intrinsic trade-off between additional hours spent in class or at work. It could be argued that estimating two "single activity" (i.e., schooling and working probability only, respectively) probits (see section 2) and judging the incidence of opposite signs, across models and for the same variables, allows for gauging the presence or absence of the aforementioned trade-off.

Following Amin *et al.* (2006) it could be useful to remind that both child labour and school attendance can be measured as dichotomous variables expressing the probability of either conditions or as the number of hours dedicated to each activity. Dichotomous variables give rise to probabilistic models of the like presented here[11]. Using the number of hours spent in child labour and/or schooling would lead to OLS or Tobit based models. Finally, note that schooling participation lends itself to be analysed using the completed educational levels or years of attendance. This type of analysis can be carried out using econometric techniques for censored variables (i.e., survival analysis) or devising dedicated variables compatible with, for instance, OLS (e.g., Patrinos and Psacharopoulos, 1997).

Works on Nepal (Dancer and Rammohan, 2006), Peru and Pakistan (Ray, 2000) and rural India (Drèze and Kingdon, 2001) are the leading examples of probabilistic models for schooling attendance considered here. Dancer and Rammohan (2006) estimate pooled and gender-specific models controlling for the effect of, among others, religion, wealth, income and whether the parents are earning their livelihood from being employed in the formal or informal sector. They find that household size and the age of both parents are the most relevant insignificant variables. Ray (2000) employs logit models[12] and finds that age and its square value, gender, working status, language, years of education of the most educated woman and community characteristics are relevant in both countries. In none the gender of the household head is significant. In Pakistan, poverty status, household size and a few wage variables are significant too. In Ray (2000), household size is captured by both the number of children and adults. The effect of an additional child is the reduction in the likelihood of attending school by the pupil in schooling age. On the contrary, one

[11] Amin *et al.* (2006) provides additional references to education and child labour probabilistic models.
[12] For a better appreciation of what follows it might be useful reminding that to compare logit and probit coefficients it is necessary to divide former by $\frac{\pi}{\sqrt{3}}$.

more adult in the household releases children's time so that they might allocate some of it to human capital accumulation. The effect of increasing household size by one child is compatible with gender bias against female pupils who are forced to childminding instead of going to school. Dancer and Rammohan (2006) and Ray (2000) both find that females are consistently are worse off than males in their pooled models.

The empirical literature has developed the probabilistic models to investigate school attendance across genders, localization areas or both of them. The present paper is concerned with the difference across gender but, for instance, Ray (2003) finds different coefficients across gender and settlement area in a child labour model for Ghana. A previous paper for Peru (Ray, 2000) highlighted the presence of gender bias only. Grimm (2002) estimates a probit for the permanence in school of both male and female children in Cote d'Ivoire. The localization dummies all display significant negative effects for both genders. Likewise, all but one (for the mother in the male model) of the coefficients for the completed schooling career are significant and are rising in magnitude. Surprisingly, the child of the household head suffers from a negative IE. In both Dancer and Rammohan (2006) gender specific models neither the father's age nor the mother's are significant. The father's education, the mother dummy for the higher one and all, but the second, wealth quintiles (in the male model) are significant for both genders and foster attendance. The religion dummy and the mother being employed in paid work are the variables significant only in the male model. Likewise, the dummy for the mother possessing land, the mother dummy for primary education, the second wealth quintile, the number of children under five and the size of the household are significant in the female one. Note that the latter two effects have contrasting signs. This might confirm the gender bias against women in education as it happened in a very similar manner in Ray (2000). In Drèze and Kingdon (2001) the asset indicator raises female participation.

Psacharopoulos (1997) in a study on Bolivia and Venezuela and Ray (2000) on Peru and Pakistan estimate logit models for the probability of the child being employed. In both countries in Ray (2000), alongside the significance of several individual and household level demographical characteristics, the poverty status dummy, the gender and age of the household head are not significant. The number of adults in the household suggests the substitution of adults' work for children's. Child labour in

Peru seems tightly linked to the variables reporting the household's financial situation (i.e., expenditure per equivalent adult, the maximum wage earned by a male member and dummies for sewage treatment) through the luxury axiom. The first and second degree term in the wage earned by a female member and the dummy for bad water storage are the variables associated with the luxury axiom in Pakistan. Psacharopoulos (1997) is not looking for a non-linear relationship with age but includes dummies for gender, being indigenous, having a female household head and a continuous measure for household income. It is interesting to note that all the covariates favour child labour. In the Venezuelan model, the urban dummy is a substitute for the racial one and is the only variable which curbs child labour. The marginal effects in this country are much smaller than those recorded for Bolivia.

At least to the author's knowledge, single gender probabilistic models for child labour are rare but Cockburn (2000) on Ethiopia is a useful reference. It originally employs the log of age to account for non-linear relationship in this variable; dummies for land ownership, fertility and slope of the plot; dummies for permanent crops and other assets plus the minutes required to fetch water. The model estimated on the female subsample achieves the significance of the log of age and the number of younger boys only. Both increase girls' involvement in working activities. In the male subsample, the number of infants and of older boys; the number of oxen/bulls/hoes and the presence of permanent crops on the farm are significant in addition to those significant for females.

2. Econometric models for schooling and working

The present work employs the Glick (2008) model to define the theoretical framework for the analysis of child participation in education. Empirically, it estimates discrete response models for both probabilities and for both genders. The theoretical foundation of probabilistic models in terms of random utility functions can be found in the seminal work of Domencich and McFadden (1975). Despite the equilibrium originating from the Glick (2008) model is by definition optimal[13], the comments provided in this paper will imply that child school attendance is morally and ethically

[13] It is impossible to find a combination of schooling attendance, child labour and leisure time different from the optimum yielding a Pareto improvement in utility given that every equilibrium is Pareto efficient.

preferable to child labour. This will allow suggesting policies which diminish the probability of child labour while enhancing that of schooling[14].

Let $d_{i,1}$ be a dummy variable taking value one if a child goes to school and zero otherwise, $d_{i,2}$ a categorical variable expressing whether the pupil is involved in child labour, $\boldsymbol{w}_{i,k}^{'}$ $k = \{1,2\}$ the set of explanatory variables described in section 3 and $\varepsilon_{i,k}$ the error terms independent and identically distributed according to the normal distribution $N(0; 1)$. The probit models for schooling and working attendance may be formalised as

$$\begin{aligned} d_{i,1}^{*} &= \boldsymbol{w}_{i,1}^{'}\theta_1 + \varepsilon_{i,1} \\ d_{i,2}^{*} &= \boldsymbol{w}_{i,2}^{'}\theta_2 + \varepsilon_{i,2} \end{aligned} \tag{5}$$

where $d_{i,1}^{*} \geq 0,\ d_{i,2}^{*} \geq 0$ are the latent variables associated to $d_{i,1} = 1$ and $d_{i,2} = 1$. According to Filmer (1999), it will be assumed that $d_{i,k}^{*}$ is the underlying demand for either schooling or working. Whenever either of them exceeds zero by any positive quantity, the respective dichotomous variables will be equal to 1. This approach reconciles probit models (i.e. discrete models) with the continuous nature of the levels of demand for both schooling and child labour attainable from developing the Glick (2008) model.

The system (5) does not consider the working and schooling decisions as being simultaneously determined. Likewise, it does not consider the pupil's working condition as completely exogenous. In the present paper, 13.38% of 5,357 children aged 7 to 15 not head of the household is idle, 9.46% is working while not attending school, 76.27% is studying while not working and a mere 0.88% is doing both activities. The negligible percentage associated to the last combination leads to the irrelevance of estimating a model where the two conditions are endogenous to each other. Nonetheless, to preserve the relevance of the decision process which allocates the pupil's time to schooling and/or work, the models are estimated using the same set of RHS covariates. This practise originates from the child labour literature (Grootaert and Patrinos, 1999). Since child labour is the activity which is likely to compete the

[14] Although it possible to record idle pupils (i.e., neither going to school nor going to work) the whole paper – for expositional purposes – will assume that a child not going to school would work and the opposite.

most with education for children's time, the models in (5) are expected to yield numerous variables with opposite sign across the two.

The set of RHS variables reported in section 3 is quite comprehensive but the extent of the omitted variable bias might still be relevant. For instance, failing to account for asset variables which are likely to have an effect on the marginal productivity of child work might prevent us from obtaining a statistically significant coefficient for the luxury hypothesis (Basu and Van, 1998 cited in Ray, 2000). This issue is highly relevant since an asset with intrinsic high productivity will raise household income (hence reducing poverty) which, in turn, is frequently a covariate which enhances enrolment. Cockburn (2000) considers the assets' contribution to the productivity of child labour a variable informing policy making. In fact, Cockburn (2000) recommends the implementation of a policy increasing the poor's access to productive assets to tackle the connected issues of child labour, school attendance and poverty when a market for child labour is missing.

Because of the relevant number of unobservable individual characteristics cross-sectional data cannot account for (e.g., the effect of talent (Ashenfelter and Rouse, 1998)), Beegle *et al.* (2004) warn that the findings of such econometric investigations should be interpreted as suggesting statistically significant correlations and not causal relationships. In the present paper, being unable to deal thoroughly with the possible endogenous nature of the household size, some within-household joint decisions, cost of schooling and the Body Mass Index (BMI) suggests embracing the aforementioned position.

Model (5) evaluates the impact of a consumption measure on the probability that the children go to school or to work. It is likely that per capita consumption depends on the household members' allocation of time to productive or leisure activities. Moreover, expenditure - as well as income - is often subject to measurement errors arising from, for instance, self-employment activities or auto-consumption of the produce of the family farm (Bhalotra and Tzannatos, 2003). The use of an Instrumental Variable (IV) technique (Sudhanshu, 1999) to limit the impact of joint determination and measurement error bias is investigated in Appendix A. Since the tests upheld the exogeneity of consumption, Table 3, 5, 7 and 8 rely on the non-instrumented version of this variable.

Household size might be endogenous to the children's allocation of time through fertility choices (Bhalotra and Tzannatos, 2003). The latter can potentially give rise to a trade-off between quantity and quality of the pupils (Becker and Tomes, 1976 cited in Bhalotra and Tzannatos, 2003). It might be argued that the standard argument of endogenous fertility decisions does not apply to this model since the model does not consider parents but household heads only. Yet, the head's decision about the number of people belonging to the household might be jointly determined with, for instance, the square metres. In fact, it is possible that the household head wants to ensure a minimum level of per capita floor space and, in turn, decent living conditions. Econometrically, this variable would require IV which instead is not implemented due to the possible absence of suitable instruments and the need for demanding identifying restrictions. Therefore, the variable is supposed exogenous (Ray, 2003) but a "sensitivity analysis" is performed amending the present variable from the model. This practise does not induce significant changes in the schooling results while in the child labour model the coefficients related to wealth and income variables are affected, though seldom significantly.

The occupational status of the household head is potentially jointly determined with the probability of the children going to school or to work. It would be correct to exclude these variables from the specification only after having tested for the separability, within the household, of the decision-making process and having found that the independence of behaviours is upheld[15] (Bhalotra and Tzannatos, 2003). Neither the test for independence nor the IV estimation is deployed here but excluding these three dummies from the educational pooled-model did not have major effects on the significance and size of the coefficients. The age of the household head and the dummy for her having a Secondaire education became insignificant in the probit model for child work while the female household head became significant.

The expenditure for schooling purposes and the time that takes the child to get to school are variables which are endogenous to the pupil going to school. Both individual level variables are averaged over all the observations which belong to the same province to make them exogenous (Grootaert and Patrinos, 1999; Ray, 2003). The averaging procedure removes endogeneity, maintains the analysis at a reasonable micro level and avoids collinearity with the regional dummies.

[15] For hints on separability refer to Appendix A.

It is well known that the probit estimates are interpretable solely in terms of standardized probit index (Reilly, 2007). To view the results in probabilistic terms it is necessary to turn to the ME/IEs which may be quantified in probability points (pob.p.) or in percentage points (per.p.)[16]. Besides the difference between a probit index and a probabilistic interpretation of the results, it is necessary to choose between reporting on ME/IEs associated to significant *βs* or on statistically significant ME/IEs themselves. Conflicting evidence between the two approaches might arise[17]. In the current paper ME/IEs for significant *βs* are presented[18].

This work aims at identifying the ME for the variable age taking into account its first and second degree terms whenever both of them are significant. This effort requires the use of the following expression

$$\frac{\partial P_i}{\partial Age} = \left(\gamma_{age} + 2\gamma_{age^2}\,\overline{Age}\right)\cdot\phi(\overline{\boldsymbol{w}}) \qquad (6)$$

where $\overline{\boldsymbol{w}}$ is the vector of the linear combination of the mean characteristics in the estimating sample and the probit coefficients, $\phi(\cdot)$ is the normal probability density function and $\overline{Age}$ is the mean value for the corresponding variable[19].

The need for two gender-specific models is tested econometrically transforming model (5) in a fully interacted specification (Mukamel *et al.,* 2002) as follows:

$$\begin{aligned} d_{i,1}^{*} &= \boldsymbol{x}_{i,1}^{'}\beta_1 + \varepsilon_{i,1} \\ d_{i,2}^{*} &= \boldsymbol{x}_{i,2}^{'}\beta_2 + \varepsilon_{i,2} \end{aligned} \qquad (7)$$

[16] Henceforth only the abbreviation for percentage point(s) (per.p.) and probability point(s) (prob.p.) will be used.

[17] An example appears in Table 7 and Table 8 in Cockburn (2000).

[18] Moreover, the statement *"on average and ceteris paribus"* should complement each effect. For expositional convenience, it is always omitted.

[19] The author credits Dr. Barry Reilly for signalling this formulation. This expression reflects the traditional formulation of the ME ($\beta\phi(\overline{\boldsymbol{w}})$)**.** Nonetheless, given the original estimated model is of second degree in age, the combined ME is a linear combination of the "traditional" ME associated to a first degree variable and $2\,\beta_{age^2}\phi(\overline{\boldsymbol{w}})$ which corresponds to the second degree term evaluated at the mean of the age variable ($\overline{Age}$) as it is usually done in STATA.

where the covariates' vector $\boldsymbol{x}_{i,k}^{'}$ is the partition of vector $\boldsymbol{w}_{i,k}^{'}$ obtained by removing the gender dummy and multiplying what is found by the gender dummy. A Wald Test[20] on the vector of interaction terms yielded significant results denoting the statistical difference of the results across genders. According to Glick *et al.* (2004) (footnote 25), because of the standardization of the variance of the error term carried out in probits and logits, model (7) yields the same results of estimating model (5) for the two genders separately. To make the presentation of the results more convenient, the present paper will estimate and present single-gender models.

3. Data

The Moroccan Enquete Nationale sur les Niveaux de Vie des Menages[21] (ENNVM) 1998/1999, carried out by the National Statistical Office assisted by the World Bank, is a nationally representative household survey which employs a two stages stratified sampling procedure. The Primary Sampling Units (PSU) are representative across rural and urban areas and reflect the different social classes in the society. Cartographic information and the National Census of the Population and Homes held in 1994 provided an universe of 1,500 PSUs from which 432 were drawn respecting stratification and base parameters criteria. The Secondary Sampling Units (SSU) are the households. For each PSU, 12 households are selected according to equal probabilities, giving rise to an expected final sample of 5,184 observations. Non-responsiveness and the inconsistency in some records drove the total number of usable household records down to 5,129. On the other hand, Vecchi (2001) warns that at the beginning of 2001 several sections of the survey are being checked for consistency to determine the soundness of the data.

The organization of the survey respects the traditional guidelines elaborated by Grosh and Glewwe (1996) and provides information regarding social mobility, housing conditions, energy supply and expenditure, education, health, employment status, transportation means, migration, fertility and anthropometric measures.

The sample for the schooling attendance model is composed by 4,957 observations while the one for child labour by 4,910.

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[20] The traditional F test becomes a Wald test because of the correction for heteroscedasticity implemented throughout the work.

[21] An appropriate translation might be National Survey of Household Living Standards.

The data for pooled and gender-specific models are defined and described in Table 2 using averages and standard deviations in parentheses when appropriate. These variables were selected to minimize the extent of multicollinearity in the RHS using a cut-off point of 0.3 for the Spearman Rank Correlation[22].

TABLE 2 ABOUT HERE

The dependent variable is a dummy for individuals 7 to 15 years old currently at school, for the educational model, or currently at work, for the working status model. Due to the survey's limitations these models could not be casted employing, for instance, the number of hours spent in either school or work like in tobits or two stage Heckman OLS models. The definition of child employed here is consistent with the classification employed in World Bank (2001a) and largely overlaps with the one commonly used for child labour studies (Gibbons *et al.*, 2003). Since the model uses the head of the household's characteristics as explanatory variables, the only one occurrence of a child who is also head of the household is excluded from the sample. This practise is believed to introduce a negligible bias.

The section of the household questionnaire reporting the biometric data for the individual is employed to calculate the children's BMI. This characteristic is included since once in school, "[a] healthy child is more able to attend class attentively, is less likely to miss class due to illness and as a result […] to repeat grade" (Ayalew, 2000:2). Moreover, it is a proxy for the household per capita wealth and long term well-being.

Despite the specialized literature prefers the height to age z-scores (HAZ) and the weight to age ratios (WAZ) to the BMI to describe the pupils' nutritional condition (WHO, 1986 cited in Gorstein *et al.*, 1994:273), the latter measure is employed here. The preference for BMI stands with the age of the individual in month, necessary to calculate HAZ and WAZ but not BMI using the ANTHRO software[23] (Gorstein and

[22] Principal Component Analysis (PCA) (see Ruel *et al.* (1999) for an application and StatSoft (2007) for technical insights) is an alternative method for purging collinearity yielding equivalent, or even superior results.

[23] ANTHRO software, [on line], available at <http://www.cdc.gov/nccdphp/dnpa/growthcharts/anthro.htm>.

Sullivan, 1999), being unavailable in the biometric section. This choice should avoid the occurrence of a "substantial systematic bias" in the HAZ and WAZ measures (Gorstein, 1989 cited in Gorstein *et al.*, 1994:275).

The paper relies on describing children's relationship with the head of the household since tracing parents in a consistent way throughout the survey proved impossible (Vecchi, 2001) [24].

The proportion of female children in schooling age out of the total number of pupils in the household is included to capture the competition - within the household - over scarce resources devoted to education. Moreover, according to Dancer and Rammohan (2006) it captures the extent of gender bias when

> " […] for a given family size, it must be the case that a male child, growing up in a household with only brothers, may have fewer resources than if he were to grow up with sisters only" (Dancer and Rammohan, 2006:12).

Instead, the household size denotes the competition over "multiple ends" resources (Grootaert and Patrinos, 1999).

The gender of the household head is used to test whether, as reported for instance by Unni (1998) cited in Dancer and Rammohan (2006), female household heads are keener on sending their children to school when compared to male household heads. The estimation of gender-specific models will provide results useful to test whether female heads, by favouring pupils of their same sex, contribute to closing down the gender gap against women (Jayachandran, 1997 cited in Drèze and Kingdon, 1999).

The household head's highest diploma is a proxy for the educational attainment of the parents which the specialized literature has proved to be one of the major determinants of children's enrolment (among others Drèze and Kingdon (2001)). The statistical significance of a positive coefficient for this variable could initiate an intergenerational valuing and maximization of the private rate of return from pupils' education (Francavilla and Lyon, 2002) while providing the society with some of the public returns associated with it (e.g., adequate health and social consciousness).

[24] Throughout the whole work, those of the household head will substitute the variables describing parents' characteristics.

Instead, Parsons (1975) considers this same variable a determinant of the willingness to subsidize children's education transferring to them part of the adults' wealth.
The occupational status of the household head could affect the other household members' allocation of time. A priori there might be complementarity or substitution between the adults' and children working conditions. Complementarity would occur if the household head and the children are both working. Substitution would mean that if the household head is employed the children will have higher probability of schooling. Basu and Van (1998), as cited in Ray (2000), test this proposition evaluating the significance of these dummies.

Francesconi *et al.* (2005), who analysed the consequences of growing up in a fragile family environment on educational attainment in Germany, inspired the inclusion of the marital status of the household head. The age of the household head is included to proxy her inclination towards education. In Morocco 9.30% of the non-attendance cases are associated to parental disbelief in the importance of schooling. Chernichovsky (1985) reports that Ashton (1945) believed the same phenomenon occurred in Botswana. Likewise, Admassie (2003) notes that in Ethiopia parents are very uncertain about whether formal education grants higher rates of return compared to on the job training. Moreover, they are reticent to provide equal educational opportunities to boys and girls because the former will remain in the family lineage and are more likely to assist the elderly while the latter will abandon their parents' house. The assumption for this variable is that the older the household head the lower the probability of school attendance, the higher the one of working.

The indirect cost of education is captured, in the present model, by the average, at the provincial level, of the distance from school in minutes. Time distance is preferred to kilometres since it captures the nature of the terrain encountered *en route* and is a proxy for the development of public infrastructures such as roads, transportation and traffic management systems. The direct cost of education is the log of the mean, at the provincial level, of the total expenditure for schooling purposes. The total expenditure includes schooling fees[25], clothing, material, accommodation, transportation, extra

[25] Both for public and private schools.

courses and unexpected expenses in the last twelve months[26]. Pecuniary costs of schooling are relevant in Morocco, according to World Bank (2001b), because "[a]lthough education is free, the 1998/1999 LSMS confirms that families incur non-negligible out of pocket costs" (World Bank, 2001b:47). In this condition the poor's access to education can be heavily undermined. In Ray (2003) this variable was used as a proxy for school quality. This interpretation is debatable since it is difficult to distinguish, from the data, whether higher expenditure was induced by a better institution rather than an idiosyncratic household head's willingness to spend on children's schooling equipment or quality.

Household expenditure, the number of square metres the household occupies and the type of possession right on their dwelling are the variables which denote the impact of financial adequacy on children's schooling and working probabilities. Since in Morocco 8.87% of the households share living spaces the question on the number of square metres excluded common areas such as balconies and courtyards (Royaume du Maroc, 1997). According to the Cockburn (2000) analysis of the asset variables, it might be argued that dwelling characteristics might capture the child involvement in housekeeping chores.

Note that the literature usually employs parental characteristics while this paper uses the household head ones. This will require an approximation procedure when commenting on the IE for the presence of a female household head and making cross-country comparisons with the surveyed literature. The difference in size of the *βs/*ME(IE)s for the education of the mother in the available single gender models will achieve this goal.

Table 1 describes the expected effects of the model's covariates on the children's probability of going to school or work. They are to be intended as the foreseeable changes in the probabilities associated to a rise in the explanatory variables.

TABLE 1 ABOUT HERE

[26] The aggregate measure was created considering the missing values in each component as zeros.

4. On the interpretation of the results

The interpretation of the results attempts at comparing the Moroccan findings with the reference literature and at highlighting instances of conflicts with the established "priors" presented in Table 1.

Across all the estimated models, the descending ranking of the IEs for the dummies for the relationship with the household head is consistent with its prior and reads as follows: child, grandchild, other relationships with the household head, sibling and being a son/daughter-in-law of the head. Due to this regularity, this variable will not appear in the comments for any model.

4.1. School attendance

TABLE 3 ABOUT HERE

The ME/IEs for all the significant covariates, but the primary education level, in Dancer and Rammohan (2006) are bigger (at least in absolute terms) than the ones for Morocco.

The relationship between $d_{i,1}$ and age is a concave downward parabola. The maxima, and the age spans on which the current and reference models were estimated appear in Table 4. The Moroccan MEs (for the first and second degree terms) lie between those in Drèze and Kingdon (2001) and Dancer and Rammohan (2006). Equation (6) yields a negative ME of -0.0157. Therefore, inducing an infinitesimal change in age above the sample average causes the probability of attending school to decline by 1.57 per.p. Consistently with less developed and Muslim countries' evidence, being a male grants a higher probability of going to school by 13.92 per.p., compared to a female. In Dancer and Rammohan (2006) gender bias against females is three times the one in Morocco while in Drèze and Kingdon (2001) is two per.p. lower than in Morocco. This evidence supports the development of a gender specific model which follows.

A 5% rise in the BMI induces a decline in schooling participation by 0.032 of a per.p. According to this result, the school feeding programme that Morocco has been employing “since the early ‘70s” (WFP, 2001:6), which includes take–home ration[27] targeted to rural females, seems to have yielded an undesirable results. A tentative interpretation for this surprising finding is that a healthy body is crucial in working activities rather than cognitive ones. This hypothesis will be verified looking at the significance and sign of this same variable in the child labour model.

Pupils living in a household whose head is female have a higher probability of schooling, compared to a male head, by 4.70 per.p. This outcome conforms to, among others, Unni (1998). Since male household heads are deemed to support child labour, further comments on the coefficient linking head–pupil of the same gender are postponed after the estimation of the child labour model. In Drèze and Kingdon (2001), the effect of mother’s literacy on children’s school attendance is more than ten times the father’s. In Dancer and Rammohan (2006), the father records larger positive IEs than the mother’s.

Residing with a household head that has a diploma Fondamental grants the pupil a higher probability of going to school by 9.30 per.p., compared to the condition he or she would experience under an illiterate head. It can be argued that “decreasing marginal returns from the head’s education” are present and are due to the head’s diploma Secondaire increasing the child schooling probability less than the Fondamental. This occurrence is consistent with most of the literature that underlines the importance – in underdeveloped countries – of the basic literacy enjoyed by authoritative people in the house. Nonetheless, it conflicts with the international evidence taken into account here. In fact, in Nepal “increasing returns” to both fathers and mothers’ education are present. Moreover, the IEs are larger for the father than the mother. In Drèze and Kingdon (2001) household head’s education is measured by attended years and MEs are roughly ten times smaller than the Moroccan IEs[28]. In Ray (2000) a similar effect is measured by the number of years of education of the most cultured female. Considering the different nature of the variables, the Peruvian β equates circa half the Moroccan’s while the Pakistani is more than a tenth.

[27] Girls attending school regularly used to “receive a ration of 100 kilograms of wheat and 10 litres of vegetable oil per year, distributed in two instalments. This is equivalent to approximately one or two monthly incomes of a typical beneficiary family, or between 1.5 and four monthly salaries of a girl working as a domestic helper in a Moroccan city” (WFP, 2001:6).

[28] This result may be explained by relying on the different nature of the continuous and dummy variables.

A 5% increase in both expenditure and square metres triggers an improvement in the attendance probability by 0.19% and 0.105 of a per.p., respectively. These two MEs are in agreement and complement the information provided by the indigence IE in supporting the luxury hypothesis. Interpreting the Moroccan results for the asset variable according to Cockburn (2000), it displays decreasing child-labour productivity in household chores. In Dancer and Rammohan (2006), a twofold phenomenon occurs: the four wealth quintiles display a positive impact on pupil's attendance while their interaction with the dummy for the son has negative effects. In Drèze and Kingdon (2001), the ME on the asset index[29] induces a rise in attendance, which is ten times smaller than the ME on square metres in Morocco.

4.2. Gender-specific models for educational attendance[30]

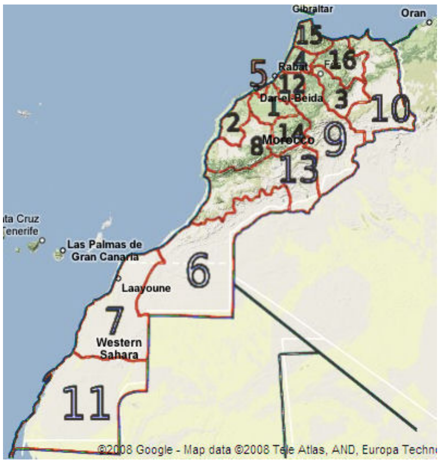
TABLE 5 ABOUT HERE

Among the variables significant uniquely in the male model, being located in the regions 4, 8, 3 and 15 in Figure 1 is associated with a lower probability of being at school compared to a pupil of the same gender living in Oued Ed-Dahab-Lagouira (region 11 in the map).

[29] Interpreting the asset index as a continuous variable does not seem particularly worrying.

[30] The presentation of the gender specific results hereafter (and the related policy implication drawn in Section 6) are likely to appear as implying that if a variable is significant in the female model and not in the male one, inducing a change in that variable might reduce gender bias in the probability of going to school/work. Glick (2008) (footnote 12) notes that despite it being a common practise in applied research it is not a statistically sound one (i.e., methods like difference-in-difference regression or the comparison in the probabilistic effect carried out in Glick *et al.* (2004) (Appendix 2.2) should be used). This paper as well as Glick (2008) follows the established convention. The use of the Glick *et al.* (2004) method of evaluating the relative gender responsiveness of child labour and schooling participation to the change in the models' RHS variables is left as a future research effort.

Figure 1 Moroccan regions on a terrain map



Source: author's elaboration on Google (2008) and Wikipedia (2008)

All the significant regions except number 3 have access to the sea. Therefore, it is possible to imagine that the pupils might be involved in some activities related to the fishing or commercial industry. Despite child labour not being a perfect substitute for school attendance, it can be the most likely and preferred outside option. The dummy for region 3, is likely to pick up characteristics of the difficult terrain encountered on the way to school which are not fully captured by the variable expressing the duration of the journey.

An infinitesimal increase in the age of the household head reduces a male schoolchild chance of attendance by 0.0013 of a prob.p. The negative attitude of an old household head is likely to be strongly determined by her lack of any completed education, being out of the reach of modern media and left out of government literacy programmes.

Instead, raising by 5% the yearly per capita expenditure increases the male attendance probability by 0.235%.

The dummies for poverty condition, female leadership of the household, the household head holding a Secondaire diploma, the variable for the household size and the provincial average time to school are the variables which affect only female pupils in Morocco.

The present evidence respects the findings in Dancer and Rammohan (2006) and Drèze and Kingdon (2001) where maternal education compensates for gender bias against female pupils. If a woman leads the household, female students have a higher probability of being in school - compared to their condition under a male head - by 8.76 per.p. In Dancer and Rammohan (2006) the IE on girls' school attendance of the mother's completion of secondary education is almost 2.5 times the effect of the same variable on males' attendance. In Drèze and Kingdon (2001), the father's education improves boys' chances while girls benefit from the one of both parents. Nonetheless, the ME for mothers exceeds the father's by 0.5 of a per.p. In Grimm (2002) mothers' education induces *βs* for female students which are almost twice as big as those for males while fathers' does not play any significant role. Surprisingly, the dummy for mothers' secondary education or above induces higher schooling attendance probability for males compared to females.

Despite the Secondaire diploma is significant, "declining marginal returns to education" arise since the IE associated to this diploma for a female pupil is 3.6 per.p. lower than the one associated to the Fondamental.

The household size supports the existence of a significant gender bias since, according to the Cameron and Trivedi (2005:488-9) methodology, the increase of household size by one unit diminishes female probability of schooling by 0.76 of a per.p.[31]. In Dancer and Rammohan (2006) the size of the household fosters female attendance.

The physical effort required to cover long distances between home and school in a tough natural environment is likely to explain the finding that increasing the travelling time by 5%, the probability of a female girl being at school falls by 0.05 of a per.p.

The remainder of this subsection will concern variables which are significant for both genders.

The coefficient for age and age square are larger, in absolute value, for females. The maxima in the parabolic relationships between schooling probability and age appear in Table 4. The result for Morocco is consistent with both the findings of the respective pooled-model and the presence of gender bias against women due to the steeper decline in probability of attending school for females. According to equation (6), an infinitesimal change in a male's age, above the sample average, causes a fall in the attendance probability by 0.35 of a per.p. For females, the reduction in schooling probability is 3.46 per.p. The MEs in Dancer and Rammohan (2006) for both age and age square are similar across gender and in line with the pooled-model's ones. The negligible difference in the single-gender MEs that Drèze and Kingdon (2001) estimate for India is particularly surprising due to the country's reputation of opposing

[31] Because this variable varies by one unit at time the correct effect is the "exact unit change" and happens to be quite close to corresponding ME.

female education. This should be particularly the case in states, as Bihar, Madhya Pradesh, Rajasthan and Uttar Pradesh, which record particularly low attendance rates and are the main focus in Drèze and Kingdon (2001).

Improvements in nutritional status are more detrimental to women's probability of going to school than to men's. A 5% increase in the BMI reduces the schooling participation probability by 0.015 and 0.0575 of a per.p. for males and females, respectively. These results reinforce the findings of the pooled model for schooling.

A 5% increase in the square metres available to the household causes females to have higher probability of school attendance by 0.187 of a per.p. The same improvement enhances males' chances by 0.047 of a per.p. In Drèze and Kingdon (2001) the asset indicator raises female participation, only.

The strongest effect for the highest diploma held by the household head corresponds always to the Fondamental one. A household head having this certificate is associated with a female pupil having a higher probability of schooling by 0.1677 prob.p. compared to living with a household head without any diploma. In the male case, the IE denotes a higher schooling likelihood by 0.0403 prob.p. The trend involving the marginal returns from education is maintained when moving from pooled to gender specific models both in Morocco and in Dancer and Rammohan (2006). In Grimm (2002), all but one (for the mother in the male model) coefficients for the completed schooling career are significant and are rising in magnitude. In Drèze and Kingdon (2001), the father's education ME in the girls probit is slightly more than three times the one in the male model. Urban settlement and the relationships with the household head in Morocco favour more female school attendance than males' one.

4.3. Estimating a model for child labour

TABLE 7 ABOUT HERE[32]

The comments to the ME/IEs will attempt to highlight the differences with the previous pooled and gender-specific models. The statistically significant dummies for the household head's education are many more in the working condition rather than in the school attendance model. The IEs they originate are fairly homogeneous across

[32] Please refer to the appropriate section of the appendix for information on the estimation problems and their solutions.

the head's diplomas such that the returns from the head's education might be deemed constant.

The relationship between the dummy for being at work and age is linear. An infinitesimal change in age raises the probability of being employed by 1.43 per.p. The Moroccan results lie between those calculated for the two models in Psacharopoulos (1997) where both probability curves are downward parabolas, though. The Peruvian peaks at 17.87 years while the Pakistani at 17.85. Note that the two maxima happen to be very close to each other despite the likely differences in economic development, legislations, attitude toward child labour in the countries and the definition of children employed in these two studies.

Being male entails a lower probability of employment compared to a female *"other things being equal"* by 1.26 per.p. This is a result compatible with the presence of gender bias in education against women. Likewise, the reviewed empirical evidence reports that males appear to be more involved in work because domestic chores are commonly excluded from the surveyed declinations of child work. Living in an urban – rather than rural – settlement diminishes the probability of working by 14.63 per.p. The present coefficient provides complementary information to the very same one in Table 3 and is consistent with the international evidence.

The significant coefficients for regional settlement suggest a reduction in the working probability compared to the omitted region. This phenomenon might be explained assuming that these very same regions suffer from particularly depressed economic conditions. This argument might hold for region number 13, 10 and 9 since, according to Figure 1 (page 18), they seem widely dominated by natural conditions likely to prevent the fishing or agricultural activity. The former is prohibited in landlocked regions while the latter activity is distressed by the presence of the Atlantis mountains and the desert. On the other hand, it might be argued that the better natural conditions in region 2 and 12, compared to the omitted one, raise the level of per-capita household consumption relaxing the household budget constraint and reducing the luxurious nature of children's spare or educational time.

A 5% rise in the BMI boosts the probability of working by 0.01 of a per.p. This result confirms the hypothesis developed, after estimating the model for education, that substantial physical capabilities are needed in activities that are likely to be manual and highly effort-demanding. This occurrence seems to be a Moroccan peculiarity since a similar result is absent in Ray (2000).

A unit change in family size increases the probability of going to work, for a youngster aged seven to fifteen, by 0.36 of a per.p. This probabilistic effect denotes the presence of constraints and competition over available resources among the household members. The Moroccan β and the "standardized" logit coefficient for the Venezuelan model in Psacharopoulos (1997) are close in magnitude and of the same sign. The two coefficients in Ray (2000) look remarkably similar and, once transformed, are in absolute value 33% larger than the Moroccan β for the household size. The education of the household head curbs child labour in Morocco and the associated βs are 10 times bigger than those converted, for both countries, in Ray (2000).

Living with a household head who is employed boosts a pupil's probability of employment by 1.70 per.p. compared to the excluded category[33]. Therefore, the "less than optimal" test for the separability, within the household, of the working decision[34] supports complementarity.

An infinitesimal increase in the age of the household head diminishes the probability of going to work by 0.27 of a per.p. and generates an unexpected result. A 5% rise in the provincial average time taken to go to school raises the opportunity cost of attendance by 0.002 of a per.p. This is consistent with the expectation that a cost surge, above economic convenience, is likely to induce a substitution of other activities for school. Among these, work is the most likely. A 5% rise in the square metres available diminishes the children's working probability by 0.019 of a per.p., confirms the luxury axiom and suggests that house-based asset variables report on the decreasing marginal productivity of a child involved in housekeeping.

All the dummies for the possession right of the dwelling are associated with a higher probability of child labour. Nonetheless, the ownership induces the smallest increase in probability since it surges by 3.23 per.p. while renting or occupying the house "for

[33] It is possible that this result is genuinely spurious. The significance of the dummy for the household being employed may arise from it being the only variable for working condition left in the model after the one for unemployment was dropped due to perfect classification.

[34] This attempt is defined "less than optimal" because it neglects several other relevant issues which should be taken into account in a test like this. A dedicated example is Bhalotra (2006).

free” raises it by 5.91 and 5.55 per.p., respectively, compared to “other rights”. Dwelling ownership is the right which is the likeliest to originate a wealth effect which, in turn, could relax the household budget constraint. The stronger the wealth effect, the smaller the marginal returns from the contribution of the children to the household budget and, in turn, from child labour. According to this interpretation, ownership shows a strong wealth effect by means of increasing the probability of child work very little. On the other hand, renting is the tenancy agreement which is associated with the highest increase in the probability of child labour.

The weak wealth effect arising from occupying a dwelling for free produces a rise in child labour probability. Therefore, there are incentives towards higher working engagement which exceed the curbing effect brought about by the availability of spare money freed up by this peculiar tenancy agreement and, allegedly, devoted to conspicuous consumption which need to be funded[35].

Inducing a 5% rise in the per capita yearly expenditure boosts the probability of being employed by 0.08%. It is possible that the higher the yearly expenditure, the higher the share devoted to conspicuous consumption[36] and, in turn, the higher the likelihood of children working to contribute to the household budget.

4.4. Gender-specific working condition models

TABLE 8 ABOUT HERE[37]

Among the variables significant only in the male sample, being settled in Fes-Boulemane raises the probability of child work by 5.93 per.p., compared to living in Oued Ed-Dahab-Lagouira. Region 3 is a landlocked, mountainous region located in

[35] To avoid that this argument brings about reverse causality it is necessary to postulate that the consumption effects captured by the tenancy agreements are not fully captured by the consumption variable.

[36] The positive sign of this ME may be interpreted as suggesting reverse causality i.e. the level of per capita consumption being affected by the number of active people in the household, hence by the probability of children being involved in child labour. The determination of the “true”, if any, causality direction is not addressed in this paper.

[37] Problems in estimating these models are reported in the appropriate section of Appendix A.

the north of the country but might be enjoying a local economy more thriving than the one in the reference region, which is deserted and geopolitically unstable[38].

Poverty is associated with child work through the luxury hypothesis. Similarly, inducing a 5% increase in the BMI, a boy has higher probability of working by 0.009 of a per.p.

An infinitesimal increase in the proportion of pupils in schooling age reduces the probability of a male being working by 0.63 of a per.p. This might require the female schoolchildren in the same household to work more to compensate for the lower masculine effort. This is somewhat confirmed by noting that the female predicted probability of working is higher than male. In Cockburn (2000), the probabilistic effects of a higher number of infants and older boys are contrasting in sign. The former fosters work participation by 2.3 per.p. while the latter reduces it by 0.4. This figure is quite comparable with the Moroccan one.

Recall that the effect of a female household head was a surge in the attendance chances for youngsters and girls, in particular. Consequently, it is quite surprising that this covariate influences males' opportunities only when estimating child labour gender-specific models. In fact, living in a household whose head is female diminishes the male probability of working by 2.21 per.p. According to the Cameron and Trivedi (2005:488-9) formulation of the unit change probabilistic effect, increasing the household size by one member raises the probability of a young man being employed by 0.39 of a per.p.

If the household head is married, boys have a higher probability of being at work by 1.24 per.p. compared to when the head is single. If she were a widow(er), the difference in chances of being working would be 0.3804 prob.p., compared to the base category.

Raising the "time-distance from school" by 5% boosts the probability of child work by 0.006 of a per.p. Finally, living in a free house increases the chances, for a male aged seven to fifteen, of being at work by 0.0408 prob.p. while a 5% increase in household living space diminishes it by 0.021 of a per.p. The IE for living in a house for free is consistent with the tenancy agreement inducing higher consumption of

[38] Region number 6 (Guelmim-Es Semara), 7 (Laâyoune-Boujdour) and 11 (Oued Ed Dahab-Lagouira) are territories contended between the Kingdom of Morocco and the Polisario front which claims this territory as part of the Sahrawi Arab Democratic Republic (Wikipedia, 2008).

goods and, in turn, requiring more people to help financing the budget rather than letting pupils enjoy their spare time or attend formal education.

The female probability of working is concave with respect to age and the minimum is located at 17.55 years. Employing equation (6), an infinitesimal increase in individual age, above the sample average, raises the probability of working by 3.50 per.p. The positive ME is due to the average age, in the female sample, being smaller than the value recorded at the turning point.

Additional variables significant only for females are the ones for living in region 6, 13, 2 and 9 in Figure 1 (page 23). They cause the probability of being at work for a female child to be lower by 0.0250, 0.0252, 0.0255 and 0.0305 of a prob.p. compared to living in the omitted region, respectively. The numeric impacts are all quite close to each other. The reduction in the working probability recorded in region 6, 13 and 9, compared to region 11, can be associated to the presence of the desert and absence of access to the sea (region 9). Region 2 produces a counterintuitive IE since it should be a region with low land and access to the sea (i.e., providing good business conditions). In turn, children could end up supplying some of the labour required locally in business prone regions. According to this argument, it is possible that this counterintuitive result is due to under-reporting the extent of child work.

A female pupil living with a household head with a Secondaire Diploma has a lower probability of going to work by 10.99 per.p., compared to living with an illiterate head. A 5% increase in the yearly per capita expenditure raises the probability of being working by 0.1065%.

The ME/IEs associated to variables significant in both genders are greater for the female model. For instance, the IE for a female living in an urban area is more than double the males' and associated to a smaller incidence of child labour. A female pupil settled in the Oriental region has a lower probability of being at work by 3.29 per.p. compared to residing in the omitted region. Men's differential is minus 1.78 per.p.

The IEs associated to the Diploma Fondamental are quite similar across genders. In fact, a female aged seven to fifteen living with a head with such a diploma has a lower likelihood of going to work by 0.0168 prob.p. compared to the reference dummy. For a male with the same characteristics, the differential is 0.0152 prob.p.

In both gender models, it is evident that there is complementarity between the adults and children's choices of participating in the labour market. In fact if the head is employed, a female pupil has a higher working probability by 1.81 per.p. compared to living with an inactive head. A male has a higher likelihood by 1.44 per.p., instead.
Owning the place where the household lives causes an increase in the working probability, compared to "other" possession rights, by 3.57 per.p. in the female model while in the male model the surge is 2.13 per.p.

Summary

The maxima in the models for the probability of going to school are consistent with the agents' optimizing behaviour postulated in the two period models for the optimal accumulation of human capital. According to Azzi and Ehrenberg (1975), concentrating human capital accumulation in the early stages of an individual's life is a rational behaviour which will grant the highest returns for a longer time span compared to studying later in life. The better nourished a pupil is (i.e., the higher the BMI) the more likely he/she is to be out of school. In fact, a well nourished body seems to be better employed at work than at school. The relationship between the household head and pupil, the urban or rural localization and gender are the variables causing the most significant changes in schooling probabilities. The sizeable link between adults and pupils testifies the importance of the intergenerational benefits of policies which promote education even when school attendance is heavily impeded. Nonetheless, it is striking that this impulse gets transmitted with decreasing marginal returns from the education of the household head. In fact, in countries with low average education levels, it would be expected that the heads of the household who are highly literate would urge the pupils they care for to "consume" ever increasing quantities of formal education.
The squared term in age in all the models presented above denotes that – beyond the probability maximizing point – the females' enrolment chances decline at a pace ten times the males'. A female-specific cause for low schooling attendance, beside child work (which is likely to take the form of childminding), is the occurrence of an early pregnancy and the associated stigma of being "a bad example for the students" (Fentiman *et al.*, 1999:345) when going back to school.

The comparison across genders and countries of the age level which maximizes the probability of schooling attendance provides the surprising result that India records the smallest distance between males and females. In this special ranking, Nepal is second - since male maximise their probability 1.72 years after females - and Morocco is third due to the difference in favour of men stretching to 1.93. Additional covariates, which stress the female inferior condition, are the poverty status (both in terms of monetary and nutritional adequacy), the family size and the distance from school. Surprisingly, the other ME/IEs associated to significant *βs* seem to favour markedly female – rather than male – prospective students.

As expected the model for child labour displays most of the coefficients with opposite sign compared to the schooling one. This occurrence could be interpreted as a sign of a degree of substitution – between the time being allocated to schooling and work – close to one. The tentative nature of the explanation is justified by the models presented here not explicitly accounting for any "simultaneity" in the pupils' time allocation between school and work. The decision of approaching "simultaneity" using an indirect technique is due to more complex and appropriate models not fitting the fragile nature of the Moroccan dataset. The age of the household head and the per capita yearly expenditure stand out for the peculiar estimated betas. Characteristics proxied by the household head's age are negatively correlated with both children's schooling and working. The significance of the per capita yearly expenditure might be due to reverse causality, especially in the child work model. The allocation of pupil's time (i.e., going or not to school/work) might have a remarkable impact on the household expenditure budget or wealth of households which face asset or income constraints. The occurrence of a linear, as opposed to concave, relationship between the dummy for working status and age may be due to the 7-15 age range used in these models. In this span, a child has already experienced some schooling such that a non-monotonic probability of school attendance is conceivable. If a pupil is marginally or not engaged in working, a concave parabolic curve for the probability of being employed should be unlikely to occur.

The significance of some regional dummies in the working model as opposed to none for school attendance might be caused by different role played by localization incentives across the two models. On the one hand, production activities are likely to be highly "location intensive". On the other, sound education policies may want to scatter the education facilities efficiently throughout the country. The regions which come up significant are largely coastal or “agriculture-friendly”. Nonetheless, these two characteristics seem associated with an unexpected decline in the probability of child labour. A variable, which is relevant solely in the child labour model, is the household head being employed. Because of the complementarity between the pupils’ and the heads’ working status, the children in the household are likely to suffer from a detrimental “intergenerational” effect which limits their human capital accumulation. Unfortunately, the insignificance of this same variable in the attendance model makes the interpretation of this phenomenon somewhat incomplete.

Female gender-specific ME/IEs in the child labour model are only double the male ones. In the schooling attendance model, they averaged roughly three times the male. The concavity in age is verified for the female model while in the male case no co-variation employment–age is present. Juvenile pregnancies occurring at the start of a female’s working career are likely to undermine her chances of long term job tenure. The lack of any connection between a male’s age and his probability of being working is an unexpected and fairly inexplicable result. The regional dummies are more frequently significant in the female working condition model vis-à-vis any male gender-specific model. Poverty is connected to male child labour while it used to be linked to females in the educational model. This might be a confirmation of the luxury hypothesis and gender bias. It reaffirms that provided the household is poor every member’s contribution to the common budget is required. When the poverty threshold has been overcome, the male pupils are likely to drop out of work and go back in class. Likewise, females have fewer opportunities and might be stuck at work. Among the relationships with the household head, being his/her son is the condition associated to the largest reduction in child labour. The second largest similar effect accrues to being a widow of the household head. This result could be somewhat spurious given the high number of widows in the survey. The latter phenomenon is caused by the respondents’ preference to be classified in this way instead of being recorded as divorced (Royaume du Maroc, 2007). The wealth and income variables present opposite signs in the model for male child labour. In the female specification, both variables are connected with a rise in the working probability. A female household head promotes the education of pupils of her same gender but in the child labour model she removes male children from work increasing their chances of schooling attendance. It appears that both gender-specific models find some complementarity between household heads and children's working conditions.

Future research on this topic would select the covariates employed in such econometric estimations more carefully to minimize the possibility of inducing endogeneity. The investigation of separability and the bargaining power in the household decision making process could be two extensions, to the current framework, which could be undertaken both theoretically (McElroy and Horney, 1981) and empirically. In the latter case, a robust and suitable dataset could provide the basis for the estimation of a bivariate or simultaneous model.
Probably the fiercest criticism of the present work concerns the extent of perfect classification which is due to the high number of dummies employed in estimation. A more pragmatic approach would have reduced their number. This would have preserved sample size but yielded a less straightforward and precise interpretation of the results.

Policy Implications

To achieve both micro (i.e., according to the literature on human capital accumulation of Mincer (1974)) and macro-level (i.e., according to the endogenous growth theory of Romer (1990)) growth and limit migration, the national government could raise school presence and quality, expand the teaching staff and improve infrastructures to facilitate children's commute to school. The focus on developing Moroccan infrastructures is mainly driven by the empirical finding that the distance from school (and the personal security issues connected with it) suggests that opportunity costs hinder school attendance more than pecuniary ones do. According to Glick (2008), this policy measure is particularly important given the common finding that female enrolment reacts much more to a change in variables denoting school distance or availability, compared to male.
A combined effort has to be exerted in bundling together educational and labour market reforms. The latter should aim to create a dynamic labour market which will facilitate literate youngsters to find a satisfactory job soon after leaving the educational system. Minimizing the time the educated youth will spend in unemployment will improve the chances that, once they will have become parents or

household heads, they will promote dependants' participation in schooling. Their experience will be the best example of the benefits of education. National policies devoted to providing child care facilities should be taken into consideration to reduce the extent of gender bias against female schoolchildren given their likely involvement in childminding within the household.

The recommendation for a monetary transfer devoted to relax the household budget constraint and facilitate the allocation of resources to education should be formulated after an analysis of schooling participation which exceeds the consideration of the sole luxury axiom. A thorough analysis of the causes of child labour and of the actual allocation of the funds could determine whether disbursing monetary or in-kind transfers to families with children in schooling age maximizes the number of those who will enrol. If child labour is employed to generate resources devoted to hedonistic consumption enjoyed by the adults in the household, then a restriction or ban on child labour may be conducive to a better children's development. Therefore, conditions for receiving school subsidies have to be attentively devised to limit the extent of both under coverage and leakage. The estimates for Morocco respect the established literature that a female adult is the most appropriate household member to manage the transfers when the policy focus on education is at stake.

The Moroccan evidence identifies gender bias against females since their school attendance possibilities are fewer while their probability of being child labourers increases. Changes in the adults' attitude toward young girls in the household and their human capital accumulation opportunities are likely to provide better results in curbing gender bias against females than relying on economic incentives only. This is likely to imply a radical cultural change in a deeply masculine society and a more egalitarian bearing, across the two genders, of favourable and unfavourable household-level shocks. Exceptional care has to be devoted to the selection of suitable means of monitoring the accomplishement of these reforms. For instance, the age of the household head – used in the Moroccan model as a proxy for the adult's attitude towards pupils' allocation of time – does not seem appropriate since it displayed a negative correlation to both schooling and working participation.

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Appendix A

Issues concerning the covariates of the parsimonious pooled-model for education

Instrumentation of the natural logarithm of the per capita expenditure

Greene (2008) warns that having an endogenous right hand side variable when estimating a binary choice model is a more serious problem than when using OLS. This happens because the IV procedure is centred on the moments of the data, embedded into the least square estimation procedure, while binary models usually employ maximum likelihood estimation (MLE). The generalized method of moments (GMM) might be preferable to the IV but, in both cases, "nothing is gained in

simplicity or robustness of this approach to full information maximum likelihood estimation [FIML][39]" (Greene, 2008:20).
Attempting the IV estimation, the procedure required an appropriate set of instruments granting relevance and over-identification. Suitable variables were looked for among those for durable goods belonging to the household and dummies for sanitation facilities in the accommodation. The instruments, which achieved relevance[40] and exogeneity[41], were the dummies for the presence of toilet and bath, the number of dishwashers and personal computers. At the end of the procedure, the test for the exogeneity of the log of the per capita expenditure is insignificant at conventional levels. Therefore, the need for instrumentation is statistically rejected.

About the separability of individual decisions in a household

According to Bowlus and Sicular (2003) a simple *t* test on the significance of explanatory variables which affect only one of multiple joint decisions is sufficient. Identifying these covariates confidently is always challenging since it might be difficult to establish a theoretical and testable independence. For example, Bowlus and Sicular (2003) argue that household size and composition should affect household consumption demand but not the amount of labour it requires to work its farm. Notwithstanding the popularity of this approach,[42] it might be argued that both characteristics impact on both consumption and working choices. Nonetheless, it is possible that the empirical test drives the theoretical conclusion.

Heteroscedasticity in a probit model with survey data

In STATA, it is not possible to test for heteroscedasticity after having estimated a probit model. Nevertheless, the latter suffers inherently from this problem (Reilly, 2007). Beside this theoretical aspect, Dancer and Rammohan (2006) point out that:

[39] Because of the absence of a procedure for both FIML and LIML (limited information maximum likelihood) in STATA 9.2, the IV estimation relies on a two-step procedure.
[40] This test F(4, 431) = 20.55 was significant a 1% confidence level.
[41] The over-identification Sargan test $\chi^2(3) = 5.704$ was not significant at conventional levels granting instrument exogeneity.
[42] Bowlus and Sicular (2003) develop the Benjamin (1992) model using panel data estimation techniques.

> "In using cross-section data […] households within the cluster may have similar characteristics, and ignoring these cluster fixed effects is likely to give us biased estimates. If these similarities are not specified in the model […] the intra-cluster correlations lead to heteroscedasticity, which will bias the estimated standard errors." (Dancer and Rammohan, 2006:6)

One solution to this econometric problem would be to include community level covariates that denote characteristics common to groups of households. Unfortunately, employing such variables here is impossible since the relevant questionnaire is available for the rural (Douar) communities only. Therefore, the STATA correction for cluster fixed effects was implemented using the PSU[43] identifier as the cluster variable.

Estimation issues for the models on education

The schooling equation in model (5) suffered from collinearity between the dummies for father/mother of the household head and father/mother in law. Likewise, the dummies for the head of the household holding the "diplomes de technicien" and the "diplomes professionnelle" were perfect classifiers of the favourable case. They were dropped in order to maximise the number of observations used in estimation.

The fully interacted model suffered from perfect predictors in excess of those affecting the pooled one. In the first place, the interacted covariates for the Laayoune-Boujdour-Sakia El Hamra region, the household head with Secondaire and Technicien Superieur diploma required exclusion. The latter variable was itself a perfect classifier of the absence from school. Moreover, the dummy for being a son/daughter-in-law of the household head was collinear to other covariates. All of them were dropped. Despite this refinement, 62 successes completely determined were a source of additional concern. This problem was eased estimating the model on observations yielding less than unitary predicted probability[44]. Consequently, the interacted dummy for living in the Guelmime Es-Sema region was collinear and was amended from the final model. This specification was the basis for evaluating the joint statistical significance of the interaction terms.

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[43] PSU denotes the variable taken into account as reference when preparing a representative sample.

[44] This customized procedure is incidentally equivalent to a more elegant, but cumbersome, official one (STATA Corp., 2005)

The male model reported problems very similar to those arising for the fully interacted one while the female added the dummy of the household head holding a Technical-Superieur diploma to the set of perfect predictors.

Estimation issues for the working status models

The working status equation in model (5) suffered from the dummies for the household head having a technical diploma and professional qualification, him/her being unemployed and divorced; the condition of being a son/daughter-in-law of the household head being all perfect predictors of the individual's inactive condition. Once these covariates were dropped from the model, 30 failures were still completely determined. A procedure similar to the one above was employed. Correcting the existing problems and estimating the model on the same sample for the educational one, the dummy for living in the Taza-Al Hoceima-Taounate region was excluded due to collinearity.

The fully interacted version of the model for working condition is, once again, affected by perfect prediction and collinearity problems involving the interacted dummy for living in Laayoune-Boujdour-Sakia El Hamra; for the household head holding a diploma Secondaire, diploma de Faculte Superieur and Technicien Superieur. The dichotomous variable for diploma Technicien Superieur was itself a perfect predictor of working. All the problematic variables were dropped.

Estimating the model on the male sub–sample, the perfect prediction problem affected the following variables: the dummy for Laayoune-Boujdour-Sakia El Hamra; the household head's diploma Secondaire, diploma Superieur and of Technicien Superieur. All of them were dropped. In the female case, the only indeterminate variable is the household head with the technical superior diploma.

Appendix B (tables)

Table 1 Summary of the expected effects of the RHS variables on the probabilities

Variables	Effect on the probability of going to school	Effect on the probability of going to work
Age	(0): no clear expectation on the combined ME but a concave parabola is expected (Dancer and Rammohan, 2006; Ray, 2000; Drèze and Kingdon, 2001)	(0): even less clear expectation due to the very fragmented nature of child labour. Probably, absence of non linear relationships.
Dummy for gender	(+): relying on the "usual" direction of the gender bias	(-): the "usual" direction of bias will preserve men from working
Dummy for urban	(+): due to higher concentration of educational facilities	(0): due to failing to account for the sector of employment
Regional dummies	(0): no expectations	(0): no expectations
Poverty dummy	(-): luxury hypothesis (Basu and Van, 1998)	(+): luxury hypothesis (Basu and Van, 1998)
BMI	(+): due to the contribution of a well nourished body to attentiveness etc. (Ayalew, 2000)	(+): due to the need of physical energies to complete demanding tasks (Ayalew, 2000)
Expenditure measure	(+): luxury hypothesis on education (Basu and Van, 1998)	(-): luxury hypothesis on education (Basu and Van, 1998)
Relationship with the household head dummies	(+):the closer is the tie, the higher is the probabilistic effect	(-):the shorter is the relationship the stronger is the IE.
Proportion of female pupils	(0): it is possible to make tentative expectations in the gender-specific models only (Dancer and Rammohan, 2006)	(0): it is possible to make tentative expectations in the gender-specific models only (Dancer and Rammohan, 2006)
Female household head	(+): a female head values pupils' human capital accumulation more than a male (Unni, 1998; Jayachandran, 1997)	(-): a female household head induces a stronger substitution effect (schooling for labour time) than a male head (Unni, 1998; Jayachandran, 1997)
Household size	(0): the possible lack of resources on a per capita terms would yield (Walters and Briggs, 1993) (-) but the presence of more people available for child minding will free up school-age children (+)	(0): larger households require far higher resources (Walters and Briggs, 1993) (+) but the large number of "employable" people in the household will free pupils' time enough to yield (-)
Household head's diploma	(+): a more literate head will try to perpetuate his/her condition in the next generation too. Moreover, the higher the qualification the larger the effect on probability (among others Drèze and Kingdon, 2001 and Francavilla and Lyon, 2002)	(-): a literate head will value education a lot, hence will try to move his/her children out of the field and into the class (among others Drèze and Kingdon, 2001 and Francavilla and Lyon, 2002)

Table 1 Summary of the expected effects of the RHS variables on the probabilities *(continued)*

Variables	Effect on the probability of going to school	Effect on the probability of going to work
Occupational status of the head	(+): substitution hypothesis based on the violation of the separability test (Basu and Van, 1998)	(-): substitution hypothesis based on the violation of the separability test (Basu and Van, 1998)
Marital status of the head	(+): living with a married head provides a better environment for human capital accumulation (Francesconi *et al*, 2005)	(-): living with an additional adult, likely to work, might induce a substitution effect for the children's time (Basu and Van, 1998)
Age of the head	(-): assuming that age is inversely related to the positive perception of education (Ashton, 1945). It is possible that benevolence inverts the sign (+)	(+): an old head is likely to have experienced child labour him/herself (Ashton, 1945). Benevolence might reverse the direction of the effect (-)
Distance from school	(-): according to the opportunity cost of schooling argument	(+): this sign might be particularly appropriate when considering "local" jobs: housekeeping, child minding, herding and other tasks on the household farm
Expenditure for schooling	(-): according to the opportunity cost argument (World Bank, 2001b:47)	(+): besides the standard argument for the opportunity cost (World Bank, 2001b:47), it is even stronger when a wage job is taken into account
Wealth variables	(+): standard substitution effect but Cockburn (2000) warns that according to the type of assets the effects vary quite dramatically and yield, on aggregate, (0)	(-): standard substitution effect but it is possible that the net effect of several Cockburn (2000) assets is (0)

Source: author's compilation on cited literature

Table 2 Descriptive statistics for the education models

Variable	Definition	Education Model Pooled model	Education Model Female Model	Education Model Male Model	Working Model Pooled model	Working Model Female Model	Working Model Male Model
P_s	=1 if the pupil goes to school, 0 otherwise	0.7753 N/A	N/A	N/A	N/A	N/A	N/A
P_s^F	=1 if the female pupil goes to school, 0 otherwise	N/A	0.6924 N/A	N/A	N/A	N/A	N/A
P_s^M	=1 if the male pupil goes to school, 0 otherwise	N/A	N/A	0.8530 N/A	N/A	N/A	N/A
P_w	=1 if the pupils is working, 0 otherwise	N/A	N/A	N/A	0.1049 N/A	N/A	N/A
P_w^F	=1 if the female pupil is working, 0 otherwise	N/A	N/A	N/A	N/A	0.1297 N/A	N/A
P_w^M	=1 if the male pupil is working, 0 otherwise	N/A	N/A	N/A	N/A	N/A	0.0804 N/A
Age	Age of the pupil in years	10.7809 (2.5123)	10.8531 (2.5528)	10.7121 (2.4708)	10.7825 (2.5132)	10.8539 (2.5550)	10.7122 (2.4698)
Age squared	Age of the pupil in years squared	122.5384 (55.0516)	124.3047 (55.9843)	120.8517 (54.0890)	122.5768 (55.0744)	124.3317 (56.0294)	120.8488 (54.0733)
Male	=1 if the child is male, 0 otherwise	0.5041 N/A	N/A	N/A	0.5039 (0.5000)	N/A	N/A
Urban	=1 if the child is living in an urban area, 0 otherwise	0.5348 N/A	0.5281 N/A	0.5294 N/A	0.5303 (0.4991)	0.5238 (0.4995)	0.5368 (0.4987)
Oued Ed-Dahab–Lagouira	=1 if the child lives in the Oued Ed-Dahab–Lagouira region, 0 otherwise. Base category	0.0089 N/A	0.0085 N/A	0 N/A	N/A	N/A	N/A
Laayoune-Boujdour-Sakia El Hamra	=1 if the child lives in the Laayoune-Boujdour-Sakia El Hamra region, 0 otherwise	0.0153 N/A	0.0142 N/A	0 N/A	0.0155 N/A	0.0144 N/A	0.0166 N/A
Guelmime Es-Sema	=1 if the child lives in the Guelmime Es-Sema region, 0 otherwise	0.0339 N/A	0.0317 N/A	0.0370 N/A	0.0342 N/A	0.0320 N/A	0.0364 N/A
Souss-Massa-Daraa	=1 if the child lives in the Souss-Massa-Daraa region, 0 otherwise	0.0970 N/A	0.0980 N/A	0.0986 N/A	0.0980 N/A	0.0989 N/A	0.0970 N/A
Gharb-Chrarda-Beni Hssen	=1 if the child lives in the Gharb-Chrarda-Beni Hssen region, 0 otherwise	0.0720 N/A	0.0712 N/A	0.0747 N/A	0.0727 N/A	0.0718 N/A	0.0736 N/A
Chaouia-Ouardigha	=1 if the child lives in the Chaouia-Ouardigha region, 0 otherwise	0.0525 N/A	0.0488 N/A	0.0575 N/A	0.0530 N/A	0.0493 N/A	0.0566 N/A

Table 2 Descriptive statistics for the education models ***(continued)***

		Education Model			Working Model		
Variable	**Definition**	**Pooled model**	**Female Model**	**Male Model**	**Pooled model**	**Female Model**	**Male Model**
Tensift Al Haouz	=1 if the child lives in the Tensift Al Haouz region, 0 otherwise	0.1009 N/A	0.1074 N/A	0.0969 N/A	0.1018 N/A	0.1084 N/A	0.0954 N/A
Oriental	=1 if the child lives in the Oriental region, 0 otherwise	0.0565 N/A	0.0606 N/A	0.0538 N/A	0.0568 N/A	0.0612 N/A	0.0525 N/A
G.Casablanca	=1 if the child lives in the G.Casablanca region, 0 otherwise	0.1017 N/A	0.1009 N/A	0.1051 N/A	0.1024 N/A	0.1014 N/A	0.1035 N/A
Rabat-Salé-Zemmour-Zaer	=1 if the child lives in the Rabat-Salé-Zemmour-Zaer region, 0 otherwise	0.0694 N/A	0.0627 N/A	0.0780 N/A	0.0701 N/A	0.0632 N/A	0.0768 N/A
Doukala Abda	=1 if the child lives in the Doukala Abda region, 0 otherwise	0.0664 N/A	0.0651 N/A	0.0694 N/A	0.0670 N/A	0.0657 N/A	0.0683 N/A
Tadla Azilal	=1 if the child lives in the Tadla Azilal region, 0 otherwise	0.0537 N/A	0.0533 N/A	0.0554 N/A	0.0542 N/A	0.0538 N/A	0.0546 N/A
Meknes Tafil	=1 if the child lives in the Meknes Tafil region, 0 otherwise	0.0637 N/A	0.0627 N/A	0.0665 N/A	0.0644 N/A	0.0632 N/A	0.0655 N/A
Fes-Boulemane	=1 if the child lives in the Fes-Boulemane region, 0 otherwise	0.0520 N/A	0.0541 N/A	0.0513 N/A	0.0525 N/A	0.0546 N/A	0.0505 N/A
Taza-Al Hoceima-Taounate	=1 if the child lives in the Taza-Al Hoceima-Taounate region, 0 otherwise	0.0664 N/A	0.0740 N/A	0.0604 N/A	0.0670 N/A	0.0747 N/A	0.0594 N/A
Tanger-Tetouan	=1 if the child lives in the Tanger-Tetouan region, 0 otherwise	0.0898 N/A	0.0867 N/A	0.0953 N/A	0.0904 N/A	0.0874 N/A	0.0934 N/A
Indigence	=1 if the children is poor according to the poverty line form the World Bank, 0 otherwise	0.2080 N/A	0.2136 N/A	0.2078 N/A	0.2098 N/A	0.2151 N/A	0.2045 N/A
Body mass index	Children's BMI	17.2568 (3.3502)	17.4397 (3.2026)	17.0733 (3.4847)	17.2579 (3.3443)	17.4311 (3.1863)	17.0873 (3.4851)
No relationship with the H.H.[45]	=1 if the child has no relationship with the household head, 0 otherwise	0.0079 N/A	0.0134 N/A	0.0025 N/A	0.0079 N/A	0.0135 N/A	0.0024 N/A
Sibling of the H.H.	=1 if the child is the sibling of the household head, 0 otherwise	0.0129 N/A	0.0146 N/A	0.0115 N/A	0.0130 N/A	0.0148 N/A	0.0113 N/A
Grandchild of the H.H.	=1 if the child is the grandchild of the household head, 0 otherwise	0.0801 N/A	0.0797 N/A	0.0813 N/A	0.0804 N/A	0.0800 N/A	0.0808 N/A

[45] This category comprises those without a link to the household head and those who were working for the household.

Table 2 Descriptive statistics for the education models *(continued)*

		Education Model			Working Model		
Variable	**Definition**	**Pooled model**	**Female Model**	**Male Model**	**Pooled model**	**Pooled model**	**Female Model**
Sons of the H.H.[46]	=1 if the child is the son of the household head, 0 otherwise	0.8691 N/A	0.8605 N/A	0.8760 N/A	0.8686 N/A	0.8600 N/A	0.8771 N/A
Son-in-law of the H.H.	=1 if the child is the son-in- law of the household head, 0 otherwise	0.0006 N/A	0.0012 N/A	0 N/A	0.0006 N/A	0.0012 N/A	0 N/A
Other relationships with the H.H.	=1 if the child has other relationships with the household head, 0 otherwise	0.0295 N/A	0.0305 N/A	0.0287 N/A	0.0293 N/A	0.0304 N/A	0.0283 N/A
Proportion of female children	Proportion of female children in schooling age out of the household size	0.3665 (0.2860)	0.5424 (0.2480)	0.1939 (0.2035)	0.3669 (0.2861)	0.5428 (0.2479)	0.1937 (0.2037)
Female household head	=1 if the household head is female, 0 otherwise	0.1158 N/A	0.1131 N/A	0.1162 N/A	0.1143 N/A	0.1125 N/A	0.1160 N/A
Family size	Number of people in the household	8.7503 (3.0461)	7.8832 (3.1303)	7.6267 (2.9743)	8.7466 (3.0468)	8.8752 (3.1295)	8.6200 (2.9582)
No Diploma	=1 if the household head has no diploma, 0 otherwise. Base category	0.8303 N/A	0.8360 N/A	0.8279 N/A	0.8303 N/A	0.8366 N/A	0.8242 N/A
Diploma Fondamental	=1 if the household head has the diploma Fondamental, 0 otherwise	0.1283 N/A	0.1225 N/A	0.1322 N/A	0.1281 N/A	0.1215 N/A	0.1346 N/A
Diploma Secondaire	=1 if the household head has the diploma Secondaire, 0 otherwise	0.0202 N/A	0.0216 N/A	0.0181 N/A	0.0202 N/A	0.0218 N/A	0.0186 N/A
Diploma Superieur	=1 if the household head has the diploma Superieur, 0 otherwise	0.0157 N/A	0.0151 N/A	0.0156 N/A	0.0159 N/A	0.0152 N/A	0.0166 N/A
Diploma Technicien Superieur	=1 if the household head has the diploma Technicien Superieur, 0 otherwise	0.0008 N/A	0.0004 N/A	0.0012 N/A	0.0008 N/A	0.0004 N/A	0.0012 N/A
Diploma de Technicien	=1 if the household head has the diploma de Technicien, 0 otherwise	0.0044 N/A	0.0041 N/A	0.0049 N/A	0.0045 N/A	0.0041 N/A	0.0049 N/A
Diploma Professionelle	=1 if the household head has the diploma Professionelle, 0 otherwise	0.0002 N/A	0.0004 N/A	0 N/A	0.0002 N/A	0.0004 N/A	0 N/A
Inactive	=1 if the household head is inactive, 0 otherwise. Base category	0.1840 N/A	0.1855 N/A	0.1836 N/A	0.1839 N/A	0.1851 N/A	0.1827 N/A
Unemployed	=1 if the household head is unemployed, 0 otherwise	0.0226 N/A	0.0183 N/A	0.0263 N/A	0.0218 N/A	0.0181 N/A	0.0255 N/A

[46] This category comprises pupils born in the household and adopted.

Table 2 **Descriptive statistics for the education models** ***(continued)***

		Education Model			Working Model		
Variable	**Definition**	**Pooled model**	**Female Model**	**Male Model**	**Pooled model**	**Pooled model**	**Female Model**
Employed	=1 if the household head is employed, 0 otherwise	0.7934 N/A	0.7962 N/A	0.7901 N/A	0.7943 N/A	0.7968 N/A	0.7918 N/A
Single	=1 if the household head is single, 0 otherwise. Base category	0.0105 N/A	0.0134 N/A	0.0078 N/A	0.0106 N/A	0.0135 N/A	0.0077 N/A
Married	=1 if the household head is married, 0 otherwise	0.9250 N/A	0.9219 N/A	0.9310 N/A	0.9269 N/A	0.9228 N/A	0.9309 N/A
Divorced	=1 if the household head is single, 0 otherwise	0.0069 N/A	0.0049 N/A	0.0074 N/A	0.0061 N/A	0.0049 N/A	0.0073 N/A
Widow/Widower	=1 if the household head is a widow or a widower, 0 otherwise	0.0577 N/A	0.0598 N/A	0.0538 N/A	0.0564 N/A	0.0587 N/A	0.0542 N/A
Age of the Household Head	Age of the household head in years	47.8759 (11.1738)	47.9361 (11.1863)	47.8628 (11.2071)	47.8456 (11.1636)	47.8682 (11.1488)	47.8234 (11.1805)
Time to go to school	Provincial average time to school in minutes	17.6024 (4.6805)	17.7683 (4.7983)	17.6538 (4.4150)	17.6680 (4.6514)	17.8336 (4.7693)	17.5049 (4.5274)
Education expenditure	Natural logarithm of the provincial average total educational expenditure	6.0169 (0.3972)	6.0162 (0.4029)	6.0060 (0.3897)	6.0123 (0.3960)	6.0116 (0.4017)	6.0130 (0.3904)
Square metres	Squares metres the household lives on	3.2455 (1.5497)	3.2705 (1.5350)	3.2193 (1.5762)	3.2473 (1.5543)	3.2713 (1.5398)	3.2235 (1.5684)
Other ownership rights	=1 if the household lives in a dwelling with other arrangements, 0 otherwise. Base category	0.0484 N/A	0.0480 N/A	0.0485 N/A	0.0487 N/A	0.0480 N/A	0.0493 N/A
Owned house	=1 if the household lives in an owned dwelling, 0 otherwise.	0.7260 N/A	0.7457 N/A	0.7170 N/A	0.7318 N/A	0.7512 N/A	0.7126 N/A
Rented house	=1 if the household lives in a rented dwelling, 0 otherwise.	0.1442 N/A	0.1277 N/A	0.1573 N/A	0.1440 N/A	0.1281 N/A	0.1597 N/A
Free house	=1 if the household lives in a dwelling occupied for free, 0 otherwise.	0.0813 N/A	0.0785 N/A	0.0772 N/A	0.0756 N/A	0.0727 N/A	0.0784 N/A
Expenditure	Natural logarithm of per capita yearly household expenditure	8.6457 (0.6704)	8.6428 (0.6731)	8.6269 (0.6591)	8.6373 (0.6663)	8.6339 (0.6681)	8.6405 (0.6646)
Obs.		4,957	2,458	2,435	4,910	2,436	2,474

Source: author's compilation on ENNVM 1998/1999 data

Table 3 Pooled-model for school attendance of pupils

Variable	Probit coefficients	ME/IE
Age	0.61779***[47]	0.1211
Age squared	-0.0324***	-0.0063
Gender	0.7022***	0.1392#
Urban	1.3259***	0.2792#
Indigence	-0.2644***	-0.0567#
Body mass index	-0.0326***	-0.0064
Sibling of the household head	2.4052***	0.1226#
Grandchild of the household head	3.4277***	0.1793#
Child of the household head	3.1562***	0.8853#
Son/daughter-in-law of the household head	2.1479***	0.1164#
Other relationships with the household head	2.8936***	0.1341#
Female household head	0.2724*	0.0470#
Household Head with Diploma Fondamental	0.6362***	0.0930#
Household Head with Diploma Secondaire	0.6998***	0.0892#
Age of the Household Head	-0.0067*	-0.0013
Provincial average total minutes to go to school	-0.0300**	-0.0059
Square metres	0.1083***	0.0212
Ln of per capita yearly expenditure	0.1935**	0.0379
$\hat{y}_i$		0.8834
Number of observations		4,957
Joint Significance		Wald $\chi 2$ (46) = 695.37***
Pseudo R^2		0.3558
***Log* pseudo likelihood**		-1701.4633

Source: author's estimation on ENNVM 1998/1999 data **Note:** ***, **, * denote variables significant at 1%,5% and 10% respectively; # denotes an IE

Table 4 Maxima in the probabilities of schooling on aggregate and in gender-specific models

Country	Maximum at			Ages span
	Overall	Male	Female	
Morocco	9.53	10.39	8.46	7 – 15
Nepal	10.05	10.87	9.15	6 – 17
India	11.34	11.58	11.34	6 – 14
Peru	8.34	-	-	6 – 17
Pakistan[48]	9.22	-	-	10 – 17

Source: author's calculation on displayed estimates, Dancer and Rammohan (2006), Drèze and Kingdon (2001) and Ray (2000)

[47] Statistical significance based on robust standard errors originating from 432 clusters in PSU.
[48] This is the first instance, of many more, of a maximum falling outside the age range considered in estimation. According to Monge Zegarra (private conversation) this event does not denote any major mistake.

Table 5 Single gender models for educational attendance and ME/IE

Variable	**Male probit**	**Male** ME/IE	**Female probit**	**Female** ME/IE
Age	0.8996***[49]	0.1131	0.4349***[50]	0.1223
Age squared	-.04329***	-0.0054	-0.0257***	-0.0072
Urban	1.1340***	0.1590#	1.4654***	0.4117#
Gharb-Chrarda-Beni Hssen	-0.6359*	-0.1173#	0.0229	0.0064#
Tensift El Haouz	-0.6854*	-0.1275#	0.1421	0.0381#
Fes-Boulemane	-0.7787**	-0.1580#	-0.1598	-0.0476#
Tanger-Tetouan	-1.1375***	-0.2601#	-0.1077	-0.0314#
Indigence	-0.1493	-0.0201#	-0.3555***	-0.1080#
Body mass index	-0.0239**	-0.0030	-0.0409***	-0.0115
Sibling of the household head	1.1616*	0.0622#	3.0360***	0.2140#
Grandchild of the household head	2.3656***	0.0921#	3.8188***	0.2973#
Child of the household head	2.2499***	0.6381#	3.4463***	0.8895#
Daughter-in-law of the H.H.	N/A	N/A	2.8008***	0.2022#
Other relationships with the H.H.	2.0436***	0.0718#	3.0561***	0.2285#
Female household head	0.3289	0.0341#	0.3528*	0.0876#
Household size	-0.0030	-0.0004	-0.0344*	-0.0097
H.H. with Diploma Fondamental	0.4006**	0.0403#	0.7970***	0.1677#
H.H. with Diploma Secondaire	N/A	N/A	0.6259*	0.1315#
Age of the Household Head	-0.0102**	-0.0013	-0.0028	-0.0008
Prov. average total minutes to school	-0.0266	-0.0033	-0.0355**	-0.0100
Square metres	0.0740**	0.0093	0.1332***	0.0374
Ln of per capita yearly expenditure	0.3769***	0.0474	0.0950	0.0267
$\hat{y}_i$		0.9357		0.7986
Number of observations		2,435		2,458
Joint Significance		Wald χ^2 (40) = 291.81***		Wald χ^2(44) = 559.90***
Pseudo R^2		0.3115		0.3726
***Log* pseudo likelihood**		-699.9158		-951.7132

Source: author's estimation on ENNVM 1998/1999 data **Note:** ***, **, * denote variables significant at 1%,5% and 10% respectively; # denotes an IE

Table 6 "Conversion" table to be used with Figure 1

Region Number	**Region Name**
1	Chaouia-Ouardigha
2	Doukala-Abda
3	Fès-Boulemane
4	Gharb-Chrarda-Béni Hsen
5	Grand Casablanca
6	Guelmim-Es Semara
7	Laâyoune-Boujdour
8	Marrakesch-Tensift-El Haouz
9	Meknès-Tafilalet
10	Oriental
11	Oued ed Dahab-Lagouira
12	Rabat-Salé-Zemmour-Zaer
13	Souss-Massa-Daraâ
14	Tadla-Azilal

[49]Statistical significance based on robust standard errors originating from 414 clusters in PSU.
[50]Statistical significance based on robust standard errors originating from 423 clusters in PSU.

15	Tanger-Tétouan
16	Taza-Al Hoceïma-Taounate

Source: Wikipedia (2008)

Table 7 Probit model for the children being employed

Variable	**Probit**	ME/IE
Age	0.2449**[51]	0.0143
Gender	-0.2155***	-0.0126#
Urban	-1.7367***	-0.1463#
Souss-Massa-Daraa	-0.5521*	-0.0213#
Oriental	-1.3102***	-0.0289#
Rabat-Salé-Zemmour-Zaer	-0.4559*	-0.0183#
Doukala Abda	-0.4891*	-0.0191#
Meknes Tafil	-0.9967***	-0.0270#
Indigence	0.2335**	0.0156#
Body mass index	0.0358***	0.0021
Sibling of the household head	-2.9017***	-0.0272#
Grandchild of the household head	-2.8671***	-0.0417#
Child of the household head	-2.9390***	-0.7134#
Other relationships with the household head	-2.9295***	-0.0304#
Household size	0.0294*	0.0017
Household Head with Diploma Fondamental	-0.4632***	-0.0196#
Household Head with Diploma Secondaire	-0.5063*	-0.0185#
Household Head with Diploma Superieur	-0.5839**	-0.0199#
Household Head employed	0.3539***	0.0170#
Age of the Household Head	-0.0046*	-0.0003
Provincial average total minutes to go to school	0.0075**	0.0004
Square metres	-0.0641*	-0.0037
Owned house	0.7268***	0.0323#
Rented house	0.6431**	0.0591#
Free house	0.5868*	0.0555#
Ln of per capita yearly expenditure	0.2596**	0.0151
$\hat{y}_i$		0.0249
Number of observations		4,910
Joint Significance		Wald χ^2 (42) =487.63***
Pseudo R^2		0.3482
Log pseudo likelihood		-1074.2983

Source: author's estimation on ENNVM 1998/1999 data **Note:** ***, **, * denote variables significant at 1%,5% and 10% respectively; # denotes an IE

[51] Statistical significance based on robust standard errors originating from 427 clusters in *PSU*.

Table 8 Gender-specific models for working condition

Variable	**Male Probit**	**Male** ME/IE	**Female Probit**	**Female** ME/IE
Age	0.0144[52]	0.0006	0.4668***[53]	0.0288
Age squared	0.0057	0.0002	-0.0133*	-0.0008
Urban	-1.4637***	-0.0843#	-2.0370***	-0.1917#
Guelmime Es-Sema	-0.4213	-0.0112#	-0.8111*	-0.0250#
Souss-Massa-Daraa	-0.4587	-0.0125#	-0.6580*	-0.0252#
Oriental	-1.1301**	-0.0178#	-1.5848***	-0.0329#
Doukala Abda	-0.2175	-0.0071#	-0.7402*	-0.0255#
Meknes Tafil	-0.6875	-0.0151#	-1.1840***	-0.0305#
Fes-Boulemane	0.7357**	0.0593#	-0.1502	-0.0081#
Poor	0.2687*	0.0128#	0.2067	0.0143#
Body mass index	0.0445***	0.0018	0.0249	0.0015
Sibling of the household head	-2.0978***	-0.0169#	3.2707***	-0.0298#
Grandchild of the household head	-2.6960***	-0.0269#	-2.9186***	-0.0447#
Sons of the household head	-2.6026***	-0.5472#	-3.1124***	-0.7628#
Other relationships with the household head	-2.4086***	-0.0188#	-3.2340***	-0.0334#
Prop. of female children in schooling age	-0.1573**	-0.0063	0.0020	0.0001#
Female household head	-1.2519***	-0.0221#	0.1074	0.0072#
Family size	0.0378*	0.0015	0.0298	0.0018
Household Head with Diploma Fondamental	-0.5793***	-0.0152#	-0.3489*	-0.0169#
Household Head with Diploma Secondaire	N/A	N/A	-0.2175**	-0.0110#
Household Head employed	0.4709***	0.0144#	0.3546**	0.0181#
Household Head married	0.4720*	0.0124#	0.0131	0.0008#
Household Head widow	1.9815***	0.3804#	-0.3471	-0.0160#
Prov. average total minutes to school	0.0307*	0.0012	-0.0062	-0.0004
Square metres	-0.1048**	-0.0042	-0.0510	-0.0031
Owned house	0.6818**	0.0213#	0.7947***	0.0357#
Free house	0.5949**	0.0408#	0.5979	0.0601#
Ln. of per capita yearly expenditure	0.1368	0.0054	0.3444***	0.0213
$\hat{y}_i$		0.0159		0.0267
Number of observations		2,474		2,436
Joint Significance		Wald $\chi^2(37) = 274.19$***		Wald $\chi^2(40) = 330.70$***
Pseudo R^2		0.3386		0.3766
***Log* Pseudo Likelihood**		-457.8721		-585.9348

Source: author's estimation on ENNVM 1998/1999 data **Note:** ***, **, * denote variables significant at 1%,5% and 10% respectively; # denotes an IE.

52 Statistical significance based on robust standard errors in 420 PSUs.
53 Statistical significance based on robust standard errors in 419 PSUs.

HOW HAPPY ARE THE ALBANIANS: AN EMPIRICAL ANALYSIS OF LIFE SATISFACTION

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Abstract. *This paper uses the nationally representative Albanian Living Standards Measurement Survey from 2005 to investigate the determinants of life satisfaction. In common with much of the existing empirical economics literature that models life satisfaction (or subjective well-being), this paper exploits an ordered probit model. In contrast to this literature, however, the current study places an important emphasis on regression model evaluation. Diagnostic testing revealed a number of econometric model deficiencies but the explicit incorporation of a heteroscedastic function into the ordered probit model resolved all detected problems. The tenor of the key findings generally reflects that found in the literature on the determinants of life satisfaction for both advanced capitalist and transitional economies. However, a number of additional themes with a strong Albanian flavour were interrogated. In particular, our study revealed evidence of long memories among Albanian respondents with respect to the collapse of that country's notorious pyramid scheme and the scarring effects of this episode continue to impact on life satisfaction even with the passage of almost eight years. In addition, a sizeable effect for communal level crime activity on life satisfaction was also detected.*

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Introduction

Subjective well-being (SWB) data have always attracted the attention of psychologists and philosophers interested, most of the times, in defining happiness or well-being (WB) (Diener, 1984). Diener (1984) refers to the pioneering work of the psychologist Wilson (1967) as the first example of a study which reports the characteristics of a happy individual. Blanchflower and Oswald (2000) report that, among others, Warr (1980) and Chen and Spector (1991) suggest that self-reported measures might be shaped by a combination of circumstances (i.e., life domains *a la* Winter *et al.* (1999)), aspirations, comparison with others and individual inherent happiness (i.e., sociability or optimism).

The present paper aims to analyse happiness in Albania, as recorded in the satisfaction with life, taking into account as many of the aforementioned components as possible. Moreover, it dedicates a great deal of attention to a few technical aspects which appear to have been neglected in the most common analysis of SWB data.

Probably, the most interesting use of SWB data outside psychology is the investigation of poverty according to its multidimensional definition provided by Sen (1999). In fact, before his contribution, economists mainly dealt with the analysis of poverty, as the primary condition expressing WB deprivation, using the monetary value of income or consumption. Sen (1999), formalizing a broader concept of deprivation which includes several different domains (i.e., housing, expenditure, health, job security and political freedom), lays the foundation for employing a broader yet synthetic measure of WB. Many, if not all, of the domains that Sen (1999) includes in his multidimensional measure of deprivation are taken into account when an individual is asked to report her SWB (see Winter *et al.* (1999) for Poland).

Despite the important role SWB data might have in quantitative economic analysis, practitioners have been worried primarily that different definitions of SWB and personal characteristics unobservable or mis-reported in survey data could undermine the soundness of their findings (e.g., Bertrand and Mullainathan, 2001). Indeed, there are instances where the results obtained using objective and subjective measures are different (Pudney and Francavilla, 2006) but further psychological research (e.g., Larsen *et al.*,1985 and Pavot and Diener, 1993) has demonstrated that "self-reported SWB is a stable concept that can be measured reliably over time" (Winter *et al.*, 1999:4). Additional trust in SWB data has arisen verifying the occurrence of coefficients with stable signs and size when estimating the determinants of WB in different countries (among others Blanchflower and Freeman, 1997; Blanchflower and Oswald, 2001; Di Tella *et al.*, 2001 and Alesina *et al.*, 2004).

These reassuring and consistent results have allowed researchers to exploit the full potential of SWB data and tackle issues beyond the mere investigation of the determinants of WB or subjective poverty.

In particular, the capacity of SWB data to capture national welfare in presence of social turmoil and ongoing political reforms has generated a consistent literature on SWB in transition countries. Sanfey and Teksoz (2005) distinguish between transition and non transition countries, Hayo and Seifert (2003) focus on Eastern Europe while several studies investigate countries which constituted the, now dissolved, communist bloc (Namazie and Sanfey, 2001 for Kyrgyzstan; Winter *et al.*, 1999 for Poland; Ravallion and Lokshin, 2000, Eggers *et al.*, 2006 and Senik, 2004 for Russia).

SWB data have been used to investigate the role institutions, ideology, psychological traits, macroeconomic variables and different economic paradigms have in shaping individual WB. Frey and Stutzer (2000) deal with the impact of Swiss institutions of direct democracy on WB while Blanchflower and Freeman (1997) is probably the first example of testing ideological convictions using SWB data. They find that living in former communist countries meant having different perceptions regarding equality, state intervention, support for trade unions and satisfaction with jobs compared to living in western countries. Blanchflower and Oswald (2001) highlight the differences between Eastern and Western Europe estimating microeconomic “wage curves”. They find that wages, unemployment and the psychological damage from joblessness are similar across Europe. Alesina *et al.* (2004) compare the ideological attitude of Europeans and Americans towards inequality as captured by the income mobility in the society. According to them, ideology is reflected in the individuals’ behaviour and political orientation. They find that inequality concerns Europeans and, among them, the leftists in particular. It is argued that the lack of influence on Americans is due to equality being associated to the lack of stimuli and rewards for a consistent working effort. Hayo and Seifert (2003) incorporate in their model the perception of the communist regime and the individual’s religion and service attendance. It might be argued that the latter two characteristics capture spiritual beliefs which might be as important as politics in shaping individual satisfaction with life. They found that there was a mild appreciation of and positive expectation toward the capitalistic system, religion influenced heavily the levels of WB but church attendance did not.

The present paper investigates the determinants of SWB in Albania using an ordered probabilistic model and the latest wave of the Albanian Living Standard Measurement Survey (ALSMS). Despite focusing on the "traditional" investigation of WB determinants, it will try to incorporate the impact of religion, health, government benefits provision, distinctive features of the national economic history and of local communities in a parsimonious manner. Besides exploiting the new features of the most recent wave of the ALSMS, the paper innovates on the application of probabilistic models to the investigation of SWB. It does so after having tested for the fulfilment of some of the hypotheses which underpin the model. The ultimate aim is to provide a more precise and reliable set of estimates which might avoid the misleading effect that Eggers *et al.* (2006) attribute to econometrics when it is used to inform policy makers' development programmes.

The remainder of the paper is organized as follows. The next section reviews the literature on the most common variables, focusing on the effect of age, monetary measures and working status. Moreover, it surveys the econometric techniques used to investigate SWB. Section three describes the ALSMS data with particular reference to the more complex variables and those uniquely available in the wave we exploit. It also spells out how we define them to suit our specification needs. Section four details the testing we implement to ascertain the adequacy of traditional ordered probabilistic models and reports on our alternative approach. Section five provides the economic interpretation of the results, highlighting how the Albanian evidence compares with the reviewed literature. Finally, it reports how the estimation procedure proposed here delivers a more complete understanding of the Albanian SWB and, most importantly, solves the specification problems the testing procedure highlighted. The last section summarises and concludes.

1. Literature Review

This section discusses the findings of the established literature on SWB with respect to its dependence on age, education, marital and health status, gender, household size, localisation, ethnicity, measures of the incidence of criminality, money metric welfare variables and labour market status[1].

The dependence of SWB on age has been investigated more frequently empirically than theoretically[2]. This might be due to the finding that early theoretical models have not resisted to their extensive empirical testing. According to Charles *et al.* (2001), the simplest and oldest hypotheses available on the link between these two variables appear in Banham (1951), Buhler (1935) and Frenkel-Brunswik (1968) who postulate the decline of emotional WB as physical functioning diminishes with age. The theory of the existence of a "mid-life crisis", which would raise the question of the purpose in life as death approaches, (Levinson *et al.*, 1978 cited in Charles *et al.,* 2001) might be seen as an evolution of the one which predicts a linear declining trajectory. More recent contributions include action theories, cohort effects and socioemotional selectivity. Charles *et al.* (2001) describe the action theories, developed by Brandtstadter (1999) and Brandtstadter and Greve (1994), as centred on an intentional overall behaviour which is instrumental to maintaining an all-around individual functioning. Actions are supposed to shape both the biological and environmental personal sphere. Charles *et al.* (2001) stress that the cohorts effects postulated by Felton (1987) might include the geopolitical, social and – in general – socio-cultural changes which can deeply affect individual perception. Because some of these changes happen once in a lifetime they constitute distinctive traits of a subset of the analysed sample yielding age-specific dependence. Carstensen *et al*. (1999) propose a model of the influence of the passing of time on individual behaviour which they name socioemotional selectivity theory. In particular, they argue that young age people engage in social interactions because of their informational and educational components. When old age, as a proxy of the end of the life, (or any other future constraining condition) raises the relevance of the present with respect to the future, relationships which yield higher positive affect and diminish the negative one are

[1] Dr. Jeffery Round pointed out that discussing "relationships" without a motivating theory is incorrect. We certainly agree with this point and, since the present work is mainly empirical, we would call on the comment that Blanchflower and Oswald (2000) made on the conflict between analysing associations and cause-effect relationships to excuse our approximation in terms: "The pragmatic response, here and elsewhere, is that at this point in the history of economic research it is necessary to document patterns and to be circumspect about causality" (Blanchflower and Oswald, 2000:15).

[2] For a very comprehensive list of references on some of the most recent contribution to the multidisciplinary investigation of SWB data see Blanchflower and Oswald (2007). In particular, it is a precious source on the controversial issue of the shape of the SWB-age relationship.

actually preferred. The shift from relationships with high educational content to ones with high emotional assonance induces age-related behaviours. Despite Carstensen *et al*. (1999) report a significant body of empirical supportive evidence, the socioemotional selectivity theory has not been exploited outside psychology.

Lastly, Blanchflower and Oswald (2007) argue that the economic theory of permanent income, behind the maximization of life cycle happiness (the latter being a measure of utility), does not substantiate the common empirical result of a U-shaped trajectory but rather can predict the independence of WB from age.

Empirical investigation has preferred specifying a higher order polynomial expression in the variable age than using cohort dummies (Lelkes, 2006 and Frey and Stutzer, 2000). Ravallion and Lokshin (2000), using a difference-in-difference ordered probit, find that age is not a significant explanatory variable of the changes in welfare rungs occurred during the '90s in Russia. Graham and Pettinato (2002) estimate a monotonically decreasing linear trajectory in Peru using an ordered logit. Blanchflower and Oswald (2000), using the same technique, record a monotonically increasing trajectory only when variables for age, age squared, gender, race and a time trend are used as explanatory ones. When a more complete set of controls is used, the more common result of a U shaped trajectory is restored for both the US and Britain. Surprisingly, Winter *et al.* (1999) estimating an OLS model for the impact of exogenous variables on life domains and another for the impact of the latter on reported WB find that the estimates suggest an inverted U shape.

The bulk of the literature has reported consistent evidence of U shaped curves whose minima range between 27 (Senik, 2004) and 63 (Namazie and Sanfey, 2001) years. The turning point is commonly located in the mid-thirties (e.g., Hayo and Seifert, 2003 and Graham and Pettinato, 2002 for Russia). Sanfey and Teksoz (2005) establish that the minimum in the aggregated sample is at 46.9 while it occurs at 52.2 in a subsample of transition countries. In the subsample of non-transition countries the minimum is at 44.8 instead.

Blanchflower and Oswald (2000; 2007) stress that the existence and, in turn, the location of a "point" at which SWB is at its minimum might be the result of the continuous tension between adaptation to present circumstances and expectation of more enjoyable living conditions. Striving for better life domains is largely driven by the so called Hirschman tunnel effect[3] (Hirschman, 1973 cited in Graham and Pettinato, 2002). Once adaptation has won the battle, the individual might fully enjoy her present circumstances such that a rise in SWB can be recorded. Yet, failing to control for some positive or negative characteristics common to individuals growing up in exceptionally good or bad times might have caused the U-shaped age dependence to be spurious (Blanchflower and Oswald, 2007)[4]. Blanchflower and Oswald (2007) address this concern and, using datasets of repeated cross-sections and decadal dummies for being born between 1900 and 1980, demonstrate that the relationship is robust. The only appreciable difference is a shift of the minimum towards a more mature age. Blanchflower and Oswald (2007) verify that the same finding holds when the higher order polynomial in age is replaced by a set of age dummies (i.e., a non parametric investigation is carried out) but suggest that the change in culture and behaviours might hinder the comparability over time of SWB data. Blanchflower and Oswald (2007) explain that the positive attitude towards life developed by those cohort members who have survived their peers' demise might cause some of the cohort unobservables behind the common U-shape in age. Moreover, individuals with higher inherent happiness are likely to self-select into the oldest survivors determining the upward sloping part of the U. Clark (2007) performs an analysis similar to the one in Blanchflower and Oswald (2007) but, because of the availability of a long panel data for Britain, includes individual fixed effects. The results remark the presence of a U shaped relationship with age and of a U shaped trajectory in the fixed effects which, in turn, might be interpreted as cohort ones. Clark (2007) notes that cohort fixed effects do vary by gender and level of education, does not investigate their origins yet suggests that reference group theory might be an explanation.

Theoretically, the pursuit of knowledge in all its dimensions has been deemed important to the development of several individual domains (Carstensen *et al.*, 1999).

[3] The "tunnel effect" is the expectation of higher satisfaction levels, formulated by an unsatisfied individual, that arises from appreciating the advancements enjoyed by individuals who have characteristics similar to hers (Graham and Pettinato, 2002). Easterlin (2003) notes that (hedonic) adaptation and (social) comparison are two terms drawn from psychology. Their counterparts in economics are habit formation and interdependent preferences. Yet, it appears that the terminology from psychology has become the standard for other disciplines as well.

[4] We are grateful to Dr. Jeffery Round for his determination in asking us to evaluate the dependence of the relationship SWB-age on life-expectancy. The latter is likely to be somewhat common to the members of one cohort and is indeed unobservable to the researcher.

Easterlin (2003) argues that education is crucial in shaping more informed preferences which, arguably, translate into a more efficient and effective optimisation of individual utility. Within a framework of revealed preferences, the latter should yield higher reported WB. Empirically, education has been found to impact positively on SWB with increasing marginal benefits the higher the level of the title acquired or the more years an individual has been at school (e.g., Hayo and Seifert, 2003; Alesina *et al.*, 2004). Education seems to increase SWB because it enhances income prospects rather than because it generates positive affects *per se*. In fact, its positive impact on WB might completely disappear when previously accumulated knowledge is tested against major structural breaks and mutated economic conditions (i.e., transition from communism to capitalism) (Namazie and Sanfey, 2001; Ravallion and Lokshin, 2000). Blanchflower and Oswald (2000) forcefully disagree with the previous interpretation noting that, in an analysis of successive cross-sections for the US, education plays an autonomous role from income.

When marital status variables are significant, married people appear often vastly more satisfied than singles (Senik, 2004; Eggers *et al.*, 2006; Namazie and Sanfey, 2001 and Ravallion and Lokshin, 2000). Graham and Pettinato (2002) estimate that the psychological benefit of marriage in Peru is comparable to the increase in SWB originating from the log of expenditure. The inferior condition of widowers and divorced, compared either to married or singles, is a consistent result throughout this literature. Lelkes (2006) constitutes an exception since she found that being a widower in transient Hungary meant having higher WB compared to being single.

Males often enjoy higher SWB than women (Senik, 2004; Graham and Pettinato, 2002; Hayo and Seifert, 2003 and Ravallion and Lokshin, 2000) and perceived good health might raise WB even more than income (Senik, 2004). Ravallion and Lokshin (2000) find that actual health is not significant while a dramatic degradation in personal conditions during the '90s in Russia could almost neutralize the positive effect of income. Ravallion and Lokshin (2000) and Senik (2004) include variables for household size and a set of localization dummies. Household size is significant only in Senik (2004) but in both it reduces WB. Likewise, the significant localization variables in both studies diminish SWB. Winter *et al.* (1999) control for household size using the number of workers and find its positive effect on satisfaction. This result might be driven by the significant correlation between material and subjective measures of WB and the strong dependence of the former from the number of providers.

Ethnicity (Lelkes, 2006; Namazie and Sanfey, 2001); settlement characteristics (Namazie and Sanfey, 2001); auto thefts, murder rates and race (Alesina *et al.*, 2004) have been used as additional explanatory variables to estimate SWB equations. Hayo and Seifert (2003) control for an objective measure of community size which, although insignificant, informs that the higher the number of people the lower the WB. Graham and Pettinato (2002), Winter *et al.* (1999) and Lelkes (2006) find that urban living impacts negatively on SWB. Besides this limited set of community variables, other measures of local infrastructure and characteristics appear to have been neglected by the SWB literature.

Eggers *et al.* (2006) highlight that mis-specifying the relationship between SWB and money metric measures of the standard of living (i.e., using income rather than the life cycle funding of consumption levels) might lead for instance to a spurious U shaped relationship between SWB and age. On a similar note, Clark *et al.* (2008b) argue that constraining the effect of income on SWB to be picked up by a single coefficient, rather than allowing it to vary according to age, might lead to an overestimation of the effect of age on happiness (or other SWB measures).

Surprisingly, the estimated impact of the absolute level of either consumption or income on WB in cross-section analyses frequently appears to be moderate in the SWB literature. Coefficients between 0.1 (Namazie and Sanfey, 2001) and 0.8 (Graham and Pettinato, 2002) have been estimated using the whole distribution of income or consumption and ordered probit and logit, respectively. Hayo and Seifert (2003) obtain 0.7 for the highest income quartile in an ordered logit specification deployed for 8 countries in Eastern Europe. An interpretation of this phenomenon suggests that the relationship between SWB and an absolute measure of money metric WB, at any point in time and using cross-section data, might hold only for low income countries while it might vanish for an average income above a critical satiation point of US$ 15,000 (Layard, 2005). On the other hand, Easterlin (1974) observes that, over time, an objective measure of WB displays a clear upward sloping trend while the subjective one appears flat. This divergent behaviour in two variables measuring WB, appearing only in time-series data, was named the Easterlin paradox and has attracted a lot of research. The most agreed upon explanations for the insignificant impact of monetary measures of WB on SWB include the hedonic treadmill (Brickman and Campbell, 1971 in Clark *et al.*, 2008b) and adaptation to present conditions. Clark *et al.* (2008b) provide a detailed analysis - based on psychological, economic, recent and established evidence - of the role of these two hypotheses in explaining the Easterlin paradox. Ravallion and Lokshin (2000) argue that it might be due to the structure of measurement errors in SWB and income or consumption.

Due to the mild influence of either absolute consumption or income levels on SWB, the psychological and economic literature have explored the evaluation of individual SWB on relative rather than absolute terms. The economic preference for the dependence of SWB on relative measures, and of monetary adequacy in particular, is largely determined by using SWB as a proxy for utility[5]. Relative rather than absolute money metric measures of WB consider the monetary value of income or consumption enjoyed by the individual in relation to a reference group. Therefore, a significant body of literature has investigated the definition, existence and formation of reference groups.

Clark and Oswald (1996) estimate the "comparison income" as the income earned by people who have the same education, job and marital status of the person under investigation. Brown *et al.* (2008) report that Bygren (2004), Festinger (1954) and Law and Wong (1998) choose neighbours, individuals of the same age, wage and in similar occupations to form the reference group in their studies[6]. Clark *et al.* (2008b) in discussing the strengths and weaknesses of the most popular approaches to the determination of comparison income warn that the average income by geographical area might measure the consumption of public goods rather than the private financial resources in that area. Similarly, the predicted values from an instrumental income equation might be contaminated by the individual's expected future income. Brown *et al.* (2008) cite theoretical (Kahneman, 1992) and empirical evidence (Ordoñez *et al.*, 2000 for judged salary fairness) that more than one reference group can be used at any time by an individual asked about different aspects of work satisfaction. Once the

[5] Clark (2008) suggests that the usefulness of SWB's explanatory variables cast in relative terms might exceed that of income or consumption only and, potentially, involve a large number of covariates. Yet, Easterlin (2003) is sceptical about the possibility of forming adequate estimates, useful for social comparison, in the domains of family life and health. He argues that this process is far easier in the material domains. As far as we are concerned, this paper is the first significant attempt at drawing a picture of SWB in Albania such that we would like to keep it simple. Such an improvement is left to further research effort.

[6] Clark *et al.* (2008a) provides a more comprehensive survey of the different definitions of reference group used in this literature.

reference group(s) has (have) been chosen, it is still uncertain whether the evaluation is made according to mean levels or extreme values and the variance of the relevant characteristic (Janiszewski and Lichtenstein, 1999; Volkmann, 1951 cited in Brown *et al.*, 2008). Several issues regarding relative measures and interpersonal comparison are still rather nebulous. For instance, there is little agreement on how the process of electing a reference group unfolds (Brown *et al.*, 2008). In fact, Brown *et al.* (2008) suggest that it might occur through a sort of "similarity-based" matching[7] (Brown *et al.*, 2008:378), the norm theory postulated by Kahneman and Miller (1986) or the "natural" functioning of the memory (e.g., Brown *et al.*, 2007; Hintzman, 1986; Nosofsky, 1986). Moreover, although data from non-controlled experiments seem to suggest that individuals are able to produce a reasonable estimate of the relevant objective money metric measure of WB enjoyed by their reference group, there is no actual guarantee of such a knowledge. Furthermore, Clark *et al.* (2008b) review contributions which suggest that endogenously determined reference groups might induce the associated levels of living standards to be increasing in individual ability (Falk and Knell, 2004) and that the domains considered when building interpersonal comparisons are individual specific (i.e., the researcher cannot assume some of them being relevant for the whole sample used) (Oxoby, 2004). Therefore, some caution in interpreting these findings is suggested (Brown *et al.*, 2008) especially since the empirical literature on these issues is not fully established yet (Clark *et al.*, 2008b).

Lastly, relative measures can be created by ranking people (or asking individuals to rank themselves) on an income (Namazie and Sanfey, 2001), wealth or wage ladder. Brown *et al.* (2008), in a laboratory experiment, asked undergraduates to place their appreciation for every level of initial salary, drawn from six different distributions, they were offered knowing that the other ten levels accrued to their peers.

Alesina *et al.* (2004), Graham and Pettinato (2002) and Senik (2004) use relative measures of income to establish the presence of the Hirschman "tunnel effect" (Hirschman, 1973 cited in Graham and Pettinato, 2002). Senik (2004) and Brown *et al.* (2008) are two examples of papers which evaluate whether relative money metric measures of WB explain SWB better than absolute ones. The former, employing both the "comparison income" and various other absolute income measures finds that in Russia the relative measure is significant in all specifications. Nonetheless, its

[7] Empirically this could be modelled through a propensity score matching procedure.

estimated coefficient is not markedly larger than the one for the logarithm of income. The latter, provides results for workers being more concerned about their pay relative to those of the co-workers rather than its absolute value. Moreover, it informs that the comparison is more relevant when carried out according to ordinal rank rather than simple share of the individual salary to the average one. Namazie and Sanfey (2001) succeed in estimating ordered probit coefficients for the dummies denoting being located at the very top of the income ladder, compared to being at the very bottom, which ranged between 1 and 1.5.

Personal labour market status may influence SWB significantly due to the variability in financial and psychological rewards associated with each condition. Ravallion and Lokshin (2000) do not find that holding a job means experiencing higher SWB but ascertain that the transition to unemployment is associated with a negative probit coefficient. Hayo and Seifert (2003), Eggers *et al.* (2006) and Namazie and Sanfey (2001) using probabilistic models find that being unemployed entails lower SWB.

Andrén and Martinsson (2006) estimate a high negative impact of unemployment on the happiness of Romanians after the transition from communism. They attribute it to unemployment being a fairly recent phenomenon in post transition Romania such that individuals have not yet adapted to it (*a la* Blanchflower and Oswald, 2000) and are still suffering from its welfare reducing effect. Clark (2008) confirms that incomplete adaptation to unemployment might be the main explanation for the significance and negative impact of being unemployed on SWB pointing to corroborative evidence for several different datasets in Clark (2006) and for long run German panel data in Clark *et al.* (2008b).

Furthermore, Sanfey and Teksoz (2005) using an ordered probit find that any condition different from full time employment causes a decline in SWB[8]. Eggers *et al.* (2006) and Lelkes (2006), in particular, constitute an exception to the above finding because they estimate that self-employment is associated with increases in SWB even compared to full time work.

[8] Dr. Simon Fauser signalled how, according to this finding, developed countries where part time work is highly developed (i.e., the Netherlands) would be characterized by low levels of SWB. We would respond to this remark stressing that the result in Sanfey and Teksoz (2005) is not to be generalised due to their sample being restricted to transition and non transition European low income countries. In general, the purpose of this survey of the literature is to compare and contrast previous studies, to report on empirical regularities but not to establish general findings.

Clark and Oswald (1994) and Winkelmann and Winkelmann (1998) estimate the impact of unemployment over and above the mere disappointment arising from the forgone income that this status entails. Clark and Oswald (1994) employ the '90s WB data for Britain and demonstrate that having lost the job causes worse psychological damage than getting divorced. Nonetheless, Winkelmann and Winkelmann (1998) suggest that the drop in social rewards associated to losing the main occupation can be identified consistently only if the econometric specification includes the income variable as well. Ravallion and Lokshin (2000) note that these arguments are in stark contrast with the standard literature on utility functions which postulates the individual's preference for spare time.

The econometric approach to the investigation of SWB issues has employed OLS (Di Tella *et al.*, 2001; Blanchflower and Oswald, 2001 and Winter *et al.*, 1999), simple (Winkelman and Winkelman, 1998) and ordered probabilistic models (e.g., Hayo and Seifert, 2003; Senik, 2004). The preference for OLS lies with its results being qualitatively comparable to those of an ordered probabilistic model and the availability of established fixed effect estimators for this technique. The latter are particularly appealing to practitioners who want to tackle the effect of personality traits and country or year-specific unobservables. Using OLS, though, implies that SWB is treated as a cardinal rather than ordinal phenomenon. There might not be difference between the results of either approach (Ferrer-i-Carbonell and Frijters, 2004), yet it may not be possible to build a general methodological rule and prescribe the use of OLS rather than ordered probit or logit.

The conflation of multiple levels of recorded SWB into two is instrumental to applying established fixed effect estimators, like Chamberlain's (1984), to a simple probabilistic model (Winkelman and Winkelman, 1998).

In our opinion an ordered model is the preferred econometric framework since it respects the discrete and ordinal coding of SWB levels. Moreover, as Blanchflower and Oswald (2000) note, treating SWB as an ordinal measure helps in using SWB as a proxy for utility and interpreting the latter using indifference curves. The actual specification can be either logit (e.g., Hayo and Seifert, 2003; Graham and Pettinato, 2002 and Alesina *et al.*, 2004) or probit (e.g., Frey and Stutzer, 2000; Senik, 2004 and Ravallion and Lokshin, 2000). Eggers *et al.* (2006) are the only ones who justify choosing the ordered logit because it displays a higher goodness-of-fit as recorded by the pseudo R^2.

The heteroscedastic nature of probabilistic models is the most common econometric issue tackled in this literature. Usually, this treatment is applied without relying on proper preliminary testing. Eggers *et al.* (2006), Senik (2004) and Sanfey and Teksoz (2005) correct the estimated standard errors implementing a "cluster" specification. Alesina *et al.* (2004) achieve the same goal following Moulton (1986) and Liang and Zeger (1986). Instead, Lelkes (2006) relies on the White (1980)/Huber (1967) "sandwich" estimator for the robust variance.

Broader econometric issues like ruling out instances of reverse causality, the endogeneity of income and collinearity have been addressed within the SWB literature. Diener (1984), Kenny (1999), Graham and Pettinato (2002) and Diener and Biswas-Diener (2002) are concerned with the possibility that some of the variables commonly used as determinants of WB are affected by the level of WB, instead (i.e., suffer from reverse causality). The endogeneity of income is attributed to the role of measurement errors and unobservable individual characteristics such that Ravallion and Lokshin (2000) and Senik (2004) develop fixed effect estimators to remove them and obtain consistent results. In particular, Senik (2004) exploits the lagged values of individual income, year and location dummies to estimate an original version of the fixed effect estimator. Instead, Hayo and Seifert (2003) are the only ones who investigate and remove collinearity using the variance inflating factors.

While it is possible to argue that measurement errors might be eliminated using the fixed effect correction for unobservable traits, the literature on ordered response models seems to have underestimated the importance of ensuring that the error term respects the underlying assumed distribution (e.g., normal for the probit and logistic for the logit). Likewise, due to the necessary standardization carried out when setting up the likelihood function it is crucial to verify that the variance is constant. Finally, the estimation of ordered models provides a set of thresholds which usually are only assumed constant across the subsamples created by the different characteristics the researcher is controlling for. To our knowledge, a battery of tests for non mis-specification, normality and homoscedasticity of the error term as well as threshold homogeneity has never been applied to an ordered model for SWB. We aim to fill in this gap implementing a version of the tests presented in Machin and Stewart (1990) and to provide remedial treatment if any econometric assumptions are violated. In particular, we propose a treatment of the possible heteroscedasticity of the error term which provides information on its determinants and sheds light on the direction of the conditioning effects. This approach goes beyond treating heteroscedasticity of an unknown form as a mere hypothesis violation which requires correction.

2. Data

This paper employs the 2005 wave of the ALSMS which collects answers to a household, community and price questionnaire. A diary for recording household consumption provides valuable information on expenditure patterns.

The 2005 ALSMS is the result of a two stages stratified sampling. The Primary Sampling Units (PSUs) are the Enumeration Areas (EAs) used in the 2001 Population Census. The EAs were selected according to their geographic location (mountain, central, coastal and Tirana), being in areas characterised by big/small towns and rural environment. Tirana is subject to oversampling using non-randomly selected EAs because of the large share of population living there but we ignore this additional piece of information and rely only on the 455 EAs randomly sampled. The Second Stage Sampling Units (SSUs) were the households. Within each EA, 12 households were selected to be involved in the survey. Eight of them constituted the main interview sample while the other four were used as replacement in case of non response. As a whole the ALSMS provides 16,387 usable individual observations distributed in 3,638 households.

The household section is based on recommendations from Grosh and Glewwe (2000) and includes novel data on migration, fertility, subjective poverty, agriculture, non-farm enterprises and social capital. The present work exploits valuable information from some of these distinctive sections.

We restrict the individuals' age range between 15 and 65 and exclude students from the sample to model the individuals' labour market status more easily. Within the active population, the unemployed are people who during the last four weeks have tried to find an occupation or to start an independent business but have failed to do so. It is possible to classify the main occupation of employed people as work outside the household, in a household related business and self employment.

Dummies for religious belief should highlight whether belonging to different religions has implication for the subjective measurement of WB. Clark (2008) submits that the finding that religious people report higher SWB than the atheist is well established in the research literature. Besides the "automatic" influence of religion on personal spiritual life, there might be other ways confessions shape SWB. Among these, the main channel may be the impact of religion on the growth of income at the micro and macro-level.

For instance, Protestants might record higher SWB due to their historical appreciation of and supposed contribution in creating modern capitalism (Weber, 1905 cited in Guiso *et al.*, 2003). McCleary and Barro (2006) note that in Weber (1905) religious beliefs, rather than simple belonging, are crucial for enhancing individual economic success through the development of, among others, work ethic, honesty, trust and thrift[9]. The Calvinist flavour of Protestantism, despite putting predestination at the centre of the salvation process, regards economic success and religious faith as signals of having been chosen (Calvin, 1585 cited in McCleary and Barro, 2006). Economic success being a source of happiness is likely to induce higher levels of SWB in a subset of the Protestant group. Guiso *et al.* (2003) find that Christian religions are more positively associated with attitudes conducive to economic growth. Yet, Guiso *et al.* (2003) report that Putnam (1993) attributes the lack of trust of the others in the South of Italy to the Catholic teaching of the superiority of the "vertical" bond with the Church to the "horizontal" one with the peers. This theory has been confirmed by cross country studies (La Porta *et al.*, 1997 and Inglehart, 1999 cited in Guiso *et al.*, 2003). Similarly, Landes (1998) cited in Guiso *et al*. (2003) highlight the negative influence that the Catholic Church had on the Spanish development during the Inquisition. McCleary and Barro (2006) report that Kuran (2004) supports the idea that countries of Muslim majority adopt legal structures which generate a lot of red tape affecting contract enforcement, credit and insurance provision and corporate ownership. All of this overhead might discourage the economic activity and is likely to diminish SWB. According to the evidence above we anticipate the OTHERREL

[9] In the present paper it is not possible to control for beliefs. Yet we have the chance to compare the impact of belonging on a SWB measure with the one on the growth rate of real GDP (McCleary and Barro, 2006) or on other economic, political and legal outcomes (Barro, 1997 and La Porta *et al.*, 1999 cited in McCleary and Barro, 2006).

dummy (accounting for the Protestants in our specification) to display a positive coefficient while the MUSLIM one could come up with a negative sign.
Lastly, different religious affiliation in an unstable political and social framework might be a cause of segregation and discrimination and, in turn, of lower SWB. It is expected that belonging to the most represented religious groups will protect the individual from these welfare reducing instances. Clark and Lekles (2009) sustain our hypothesis by finding that Catholics believers were more satisfied when their creed was dominant in the region both in absolute and relative terms. The result for Protestants was qualitatively similar yet insignificant. Moreover, the change in the ranking of religions is associated with a fall in the WB experienced by the followers of the former leading one.

We use expenditure as our absolute money metric measure of WB. Household wealth is captured by the dummies for living in a single household dwelling, its area[10] and having an inside toilet.
The monthly household expenditure used in this specification excludes housing and health expenses due to their distortive impact on the aggregate measure. The housing outlays are skewed by the little relevance of the rental market in Albania which, in turn, produces unreliable values of imputed rents for people who own their dwelling. Likewise, health expenditure is excluded because it is dominated by the disbursement for medicines and because, occurring after one or more negative shock[11], it is highly non-representative of the typical monthly outlays. Moreover, its inclusion in a multidimensional measure of expenditure can alter the welfare ranking previously obtained using an aggregate which excludes the health expenditure. Owning a health licence in Albania grants discounts on the cost of medicines or allows getting them for free. World Bank (2003) affirms that it is difficult to single out the household member entitled to this benefit. Yet there is evidence that those who buy medicines at a discount, after having visited the public health provider, are not in absolute poverty, live in urban areas and concentrate in Tirana. The significance and positive contribution to WB of a measure of per capita household expenditure might lead to the policy prescription of raising the resources available to the individual. If this

[10] We were prompted with the usefulness of including a measure of per capita floor space of the dwelling. Despite the attractiveness of this measure, it cannot be implemented in this paper because the survey's questionnaire does not collect any continuous measure of the dwelling size.
[11] The visits to the public health provider are instead free of charge (World Bank, 2003).

happened through a higher provision of free or subsidised healthcare, according to the above, the Albanian poor might experience a worsening in their condition. Similarly, incurring in significant and unexpected health expenditures under a tight budget constraint might force household members to sell off assets or borrow money to finance the associated surge in the outflow of money. Since it is difficult to distinguish between such an occurrence and the outlays associated to a long term chronic illness, this type of expenditure is neglected to reduce measurement errors.

We do not make assumptions on the economies of scale and size in consumption when we transform the data in per capita terms. World Bank (2003) supports this choice verifying that different assumptions for the economies of scales in consumption do not alter the poverty estimates significantly. Due to the relevant differences in the national distribution of prices we correct nominal expenditure using the Paasche index calculated at the PSU level and standardized with respect to the capital. We deflate expenditure by the national rate of inflation between 2002 and 2005 to make expenditure consistent with the poverty profile based on the 2002 national poverty line.

Finally, we follow Senik (2004) and take the natural logarithm of the real per capita expenditure to account for the non-linear relationship with SWB and to estimate, contrary to Alesina *et al.* (2004), a beta of a meaningful size. The non-linear relationship most commonly assumed in analysis of SWB data is one of decreasing marginal returns (see Easterlin, 2005 and Oswald, 2005 for references to supportive theoretical and applied works). Nonetheless, Easterlin (2005) and Oswald (2005) provide empirical and theoretical evidence which confutes the most recent findings suggesting caution and qualified interpretations. In particular, the former notes how the finding of a dependence from some logarithmic transformation of income (or other measures of monetary WB) in cross-section studies is commonly extended to a time-series framework. Easterlin (2005) is worried that, under the above assumption, cross section results are used to justify policy actions which favour the poor (for whom the gains in SWB would be the largest). The passing of time implied by such a policy effort would require a time-series study in which case he demonstrates that the SWB's decreasing marginal returns from income disappear in favour of an inexistent relationship between the two variables. We believe the policy implications drawn from this study (see below) comply with the recommendation in Easterlin (2005). On the other hand, Oswald (2005) requires us to refine our wording and state that we are investigating the impact of per capita expenditure on *reported* SWB since the concerned literature does not know anything about the relationship between reported and actual happiness.

A set of mutually exclusive dummies which identifies whether the household receives zero, one, two or more transfers is employed to capture the impact on SWB of the government's involvement in welfare provision. The Albanian transfers' portfolio includes subsidies for economic hardship; pensions for the elderly, the disabled and war survivors as well as benefits for unemployment, war veterans, maternity, illness and social care. Following Winkelman and Winkelman (1998), the coefficients for this set of dummies will denote the psychological implication of integrating the autonomous funding of consumption with government transfers. It is expected that the higher the share of consumption funded through welfare provision the lower the level of SWB[12].

The recent Albanian history has been dominated by a few negative events. The internal political conflict, the collapse of the pyramid schemes, the rampant corruption and the problems in Kosovo might be singled out as the most relevant ones (Jarvis, 1999). Among them, the schemes attracted a lot of research interest because of their essence, causes and consequences.

The Albanian pyramid schemes were fostered by depositors putting money into lending companies which, at the peak of their popularity and of the need for fresh funds to keep the system running (November 1996), were promising rates as high as 47% a month (i.e., trebling depositors' money in three months) (Jarvis, 1999)[13]. Due to the high promised returns, the schemes attracted funds which cumulatively accounted for a maximum of almost 50% of the GDP. The high returns paid out to the older investors leaving the schemes were financed either through the money raised from more recent entrants or some sort of illegal activity (e.g., arms, drugs and people's trafficking) but not from the schemes' real investments. They happened to be

[12] Following a comment from Dr. Jeffery Round, it has been verified that a higher number of transfers corresponds to a larger mean value for the amount received. In turn, we do not see complications in using the former instead of the latter.

[13] Jarvis (1999) is our main source for this paragraph on the pyramid scheme and provides additional interesting knowledge about Albanian history and development. Therefore, we refer the interest reader to it for a more complete picture.

of negligible value and return. Besides the attractive rates of return, additional causes for the schemes' unprecedented popularity include:

- The transition from a very poor, isolated and underdeveloped centrally planned country to a modern market economy generated several business opportunities especially in the agricultural and newly privatised manufacturing sector. As a consequence, several strata of the population enjoyed higher wealth but struggled to re-invest their cash in both the real and financial markets largely because of the inadequacy of the formal, not yet completely privatised, financial system. The latter was suffering from the inability of proficiently mobilising the stream of national and international (coming from migrants' remittances) savings to the healthy borrowers in the economy.
- Consequently, the economy was flooded with large amounts of unproductive cash waiting for alternative and profitable investment opportunities. The imbalance of available funds and investment opportunities induced the rise of a flourishing informal credit market. At its outset it was largely tolerated by national authorities, external observers, the World Bank and the IMF. In fact, it was perceived as a low-cost solution to a market failure and a provider of funds to viable national businesses. Unfortunately, those informal lenders which invested the collected savings on their own account developed into pyramid schemes.
- The complacency of some politicians, from the then ruling party, for the schemes provided, to the eyes of prospective contributors, some endorsement of these investment practices. In fact, at times there was the perception that the government would have bailed out the contributors to these schemes if they ever went bust. Moreover, because the Bank of Albania was alone in voicing its own, the World Bank's and the IMF's concerns regarding the spiralling instability of the Albanian financial system, new schemes were constantly created and the more established ones grew to astonishing dimensions.

The schemes collapsed at the turn of the year 1997 and, as far as the economy is concerned, this caused the currency to devaluate by more than 40%, the GDP to fall by 7%, the inflation rate to skyrocket to 42%, the budget deficit to soar to 12% of GDP and the unemployment to rise by 3% (ETF, 2006). On the social front, several riots exploded especially in the richer south of the country, the police and the army spiralled out of control, arms were stolen and the clashes caused around 2,000 victims (Jarvis, 1999).

We exploit the section of the 2005 ALSMS on the shocks which have affected the interviewed households between 1989 and 2005 to create a dummy for having suffered from the economic and social consequences of the collapse of the pyramid schemes. Comparing the historic accounts of the participation in the schemes and of the financial consequences of their collapse with the low mean of this variable, we expect it to capture the "memory" people have of the social, more than the economic, aftermath of the phenomenon as recorded in their levels of SWB.

The number of close friends of the respondent[14] will account for the individual or household belonging to informal insurance networks. It is well known, within the development literature, that they play an important part in limiting the extent of individual poverty. In fact, friends might be a good source of informal credit in the absence of a developed official market, of some help in running small businesses and in childminding. Easterlin (20003) reports that Berkman and Glass (2000) document that a large social support provides better physical health while isolationism is detrimental to the individuals[15]. Likewise, the number of friends seems a good variable to account for some unobservable personality traits which might allow outgoing individuals to be the happier the larger their social *entourage*.

We introduce a number of variables which characterize the community the individual inhabits. A measure of the perceived increase in population at the local level accounts for the additional strain put on underdeveloped local infrastructures and public services. For instance, UNEP (2000) reports that water supply and solid waste

[14] A couple of remarks we received stressed that it might be the quality, rather than the quantity, of friend which might increase SWB and this variable might be subject to several "extreme" occurrences. The section for the social capital in the 2005 ALSMS includes a question which aims to evaluate the "quality" of the interaction by asking the respondent to judge the likelihood that she would borrow money from her *entourage*. This question allows for five distinct qualitative answers, which once modelled using mutually exclusive dummies, might hinder the parsimoniousness of the current specification and increase the likelihood of multicollinearity (Piva, private conversation). A further unobservable element regarding the variable for the social entourage is the different emotional composition of the social networks which has been found to differ for people of different ages (Carstensen *et al.*, 1999).
The distribution of the number of close friends is indeed highly skewed toward zero while some non-trivial density is associated to having 10 friends. Nonetheless, the low mean for the variable appears consistent with the qualification that close friends are those "people [the respondent] feel[s] at ease with, can talk to about private matters, or call on for help" (INSTAT, 2004:76).

[15] In this respect we anticipate the variable for living in a single household dwelling to be associated with a decline in SWB.

treatment systems are utilized beyond capacity, especially in urban areas. On the other hand, a positive coefficient would highlight the localization incentive provided by a thriving local economy. The presence of a community organization denotes the degree of cohesion the local population experiences and its proactive attitude towards tackling and solving shared problems. In our model the latter are captured by the occurrence of thefts (Alesina *et al.,* 2004) and land disputes at the local level. The fight over land will detect the remaining instances of social unrest consequent to the transition from communism and "… its chaotic land privatisation …" (King and Mai, 2008:242).

After data cleaning across all the relevant variables, the data comprise 2,923 useable observations. The key dependent variable is categorical in nature and is based on the interviewee's response to the question *"How satisfied in general are you with your current life?"*[16]. The respondent or interviewee was either the head of household or someone deputed as the most knowledgeable person within the household to answer the questions on his/her behalf. The coding of the dependent variable uses decreasing numerical values which correspond to the following four mutually exclusive answers: *fully, rather, less than and not at all satisfied*. In the estimating sample, 2.16% of the respondents enjoy the highest satisfaction with life while 20.29% are rather satisfied. The largest percentage pertains to the *less than satisfied* group with 49.33% of the estimating sample and 28.22% of it answers that they are *not at all satisfied*.
The explanatory variables employed, above and beyond those already described, are detailed and summarized using their mean in the following table

TABLE 1 ABOUT HERE

[16] A referee remarked that recent work from Kahneman *et al*. (2004) and Kahneman and Krueger (2006) suggests scepticism regarding the respondent's ability to evaluate her overall SWB through a single question covering all life domains. They argue that multiple questions, each covering every relevant life domain, would be able to paint a more accurate picture of individual SWB. Yet, to minimise the survey and the researcher's bias, the ideal questionnaire should collect, not only the SWB answers for every domain but, also every individual's weighting scheme. While we acknowledge that the latter survey's setup provides more information to work with, the 2005 ALSMS does provide subjective evaluations only for the financial situation and, with a different categorisation, for food consumption and the level of expenditure for basic needs.

3. The Empirical Methodology

In comport with other studies undertaken by economists modelling individual-level survey responses on life satisfaction, an ordered probit model is used. Let y_i denote the observable ordinal variable, coded $1, 2, \ldots, J$ reflecting the J distinct SWB levels in the current application, and let y_i^* describe an underlying latent variable that captures the continuum in the satisfaction levels for the i^{th} individual. This can be expressed as a linear function of a vector $\boldsymbol{x}_i$ of the relevant explanatory variables and a constant as follows:

$$y_i^* = \boldsymbol{x}_i'\boldsymbol{\beta} + u_i \quad \text{where} \quad u_i \sim N(0, \sigma^2) \tag{1}$$

$\boldsymbol{\beta}$ is a vector of unknown parameters and σ^2 is the variance of the error term u_i.

Hayo and Seifert (2003) highlight that this specification assumes that the general utility function, which gives rise to the different SWB levels y_i, is separable in its arguments. As far as the econometric investigation is concerned, this means that the inclusion of additional explanatory variables should barely alter the estimates for the old ones (Blanchflower and Oswald, 1997). Blanchflower and Oswald (2007) suggest that the assumption of the independence of consumption from age might not be that innocuous due to the uncertainty which goes with it. Yet, Blanchflower and Oswald (1997) might pose as a surprising confirmation of the separability of the effect of the log family income, from that of the other covariates, on the happiness of the young and the adults in the US from 1970 to 1990.

Following Greene (2008) it is assumed that y_i^* is related to the observable ordinal variable y_i as follows[17]:

$$\begin{array}{llll}
y_i = 0 & [\text{"}\textit{not at all satisfied}\text{"}] & if & -\infty < y_i^* < \theta_0 \\
y_i = 1 & [\text{"}\textit{less than satisfied}\text{"}] & if & \theta_0 \leq y_i^* < \theta_1 \\
y_i = 2 & [\text{"}\textit{rather satisfied}\text{"}] & if & \theta_1 \leq y_i^* < \theta_2 \\
y_i = 3 & [\text{"}\textit{fully satisfied}] & if & \theta_2 \leq y_i^* < +\infty
\end{array}$$

[17] This is the formalisation of our interpretation of the Oswald (2005) reporting function.

In general terms we can write $Prob\,[y_i = j] = \Phi\left(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}\right) - \Phi\left(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}\right)$ for $j = 0, 1, 2, 3$ and where $\Phi(\cdot)$ denotes the cumulative distribution function operator for the standard normal. The first and the final intervals are open-ended, so for $j = 0$, $\Phi\left(\theta_{j-1}\right) = \Phi(-\infty) = 0$ and for $j = 3, \Phi\left(\theta_j\right) = \Phi(+\infty) = 1$. Therefore, the probability associated to the open ended intervals reads as $Prob\,[y_i = 0] = \Phi(\theta_0 - \boldsymbol{x}_i'\boldsymbol{\beta})$ for $j = 0$ and $Prob\,[y_i = 3] = 1 - \Phi(\theta_2 - \boldsymbol{x}_i'\boldsymbol{\beta})$ for $j = 3$.

If the $\boldsymbol{x}_i$ vector contains a constant term, the remaining set of threshold parameters $[\,\theta_0, \theta_1, \theta_2\,]$ is not identified. The exclusion of either the constant or one of the threshold parameters facilitates an arbitrary location for the scale of y_i^*. In our application we set $\theta_0 = 0$. Another identification restriction is also required as the parameters of the ordered probit are only identified up to some factor of proportionality. As with the standard probit, the convenient normalization that $\sigma^2 = 1$ is also imposed. In general terms, we can write:

$$Prob\,[y_i = j] = \Phi\left(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}\right) - \Phi\left(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}\right) \;\; for \;\; j = 0, \ldots .J \qquad (2)$$

The general expression for the log-likelihood function for this particular model is expressed as:

$$L = \sum_{i=1}^{n} \sum_{j=0}^{3} \delta_{ij} \; log_e \left[\Phi\left(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}\right) - \Phi\left(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}\right)\right] \qquad (3)$$

where $\delta_{ij} = 1$ if the i^{th} individual falls within the j^{th} category and 0 otherwise, and $log_e\,(\cdot)$ denotes the natural logarithmic operator. Conventional algorithms can be employed to provide maximum likelihood (ML) estimates for the $\boldsymbol{\beta}$ parameter vector and the remaining two threshold parameters $[\theta_1, \theta_2]$. The inverse of the regression model's information matrix provides the asymptotic variance-covariance matrix for the parameter vector and the threshold parameters.

The present paper explores the possibility that the model is mis-specified when $\boldsymbol{x}_i$ contains the explanatory variables in Table 1, σ^2 is not constant, the error term is not normal and the thresholds are heterogeneous for different values of the characteristics controlled for. To perform these tests we need to calculate the ordered probit (OP) pseudo-residuals using the ML estimates for $\boldsymbol{\beta}$, θ_j and the generalized formula:

$$\varepsilon_i = \frac{\phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}) - \phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta})}{\Phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}) - \Phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta})} \tag{4}$$

where $\phi(\cdot)$ denotes the probability density function operator for the standard normal and the other terms are as previously defined. To deploy a formal test for the omitted variables in our specification the higher order conditional moments for latent variables, as specified in Stewart (1983) cited in Machin and Stewart (1990), are calculated as:

$$M_{\tau i} = \frac{w_{(j-1)i}^{\tau}\phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}) - w_{ji}^{\tau}\phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta})}{\Phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}) - \Phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta})} \tag{5}$$

Hence the first four moments residuals according to Chesher and Irish (1987) are:

$$\begin{aligned} \varepsilon_i^{(1)} &= M_{0i} \\ \varepsilon_i^{(2)} &= M_{1i} \\ \varepsilon_i^{(3)} &= 2\varepsilon_i^{(1)} + M_{2i} \\ \varepsilon_i^{(2)} &= 3\varepsilon_i^{(2)} + M_{3i} \end{aligned} \tag{6}$$

The score tests which will constitute the core quantitative tool of the paper have the following generalised form:

$$\xi = 1'F(F'F)^{-1}F'1 \tag{7}$$

where 1 is an n-dimensional vector of ones and F is a matrix with row order n, each row of which contains the score contributions for all parameters of the model, both those estimated and those set to zero under the null.

We test for the correctness of the pseudo functional form developing a modified version of the RESET test formulated by Ramsey (1969) or Ramsey and Schmidt (1976) as follows:

$$y_i^* = \boldsymbol{x}_i'\boldsymbol{\beta} + \boldsymbol{q}_i'\boldsymbol{\psi} + u_i \tag{8}$$

where $\boldsymbol{q}_i$ is of length k and does not include a constant. A score test, of the form ξ given above, is constructed for the null hypothesis of $\boldsymbol{\psi} = 0$, with rows of F given by:

$$F_i = (\varepsilon_i^{(1)} x_i, \dots, \eta_{(j-1)i}, \varepsilon_i^{(1)} q_i) \tag{9}$$

Under the null hypothesis, ξ is distributed as $\chi^2(k)$. Peters (2000) provides some evidence on the power of the test for a number of different limited dependent variable models.

Similarly, the test for non-normality is a $\chi^2(2)$ for the significance of the skewness and/or kurtosis where the rows of F are:

$$F_i = (\varepsilon_i^{(1)} x_i, \eta_{2i} \dots, \eta_{(j-1)i}, \varepsilon_i^{(3)}, \varepsilon_i^{(4)}) \tag{10}$$

To evaluate the presence of heteroscedasticity in (1) we shall suppose that the variance of u_i is:

$$\sigma_i^2 = 1 + \boldsymbol{q}_i'\boldsymbol{\psi} \tag{11}$$

The null hypothesis is that $\boldsymbol{\psi} = 0$ and the rows of F are:

$$F_i = (\varepsilon_i^{(1)} x_i, \eta_{2i} \dots, \eta_{(j-1)i}, \varepsilon_i^{(2)} q_i) \tag{12}$$

Under the null hypothesis ξ is distributed as $\chi^2(k)$.

Finally, the homogeneity of the thresholds with respect to the variables included in $\boldsymbol{x}$, excluding the constant, is tested supposing that:

$$\mu_{ji} = \bar{\mu}_j + \boldsymbol{q}_i'\boldsymbol{\psi}_j \tag{13}$$

under the null hypothesis of $\boldsymbol{\psi}_j = 0$, $j=2, \ldots, J-1$ and $\boldsymbol{q}_i \equiv \boldsymbol{x}_i$. The rows of F are:

$$F_i = (\varepsilon_i^{(1)} x_i, \eta_{2i} \ldots, \eta_{(j-1)i}, \eta_{2i} q_i, \ldots, \eta_{(J-1)i} q_i) \tag{14}$$

where η_{ji} are the threshold contributions as defined here:

$$\eta_{ji} = \begin{cases} \frac{\phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta})}{\Phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}) - \Phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta})} & if\ y_i = j \\ \frac{-\phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta})}{\Phi(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}) - \Phi(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta})} & if\ y_i = j+1 \end{cases} \tag{15}$$

for $j=2, \ldots, J-1$. Under the null hypothesis ξ is distributed as $\chi^2(k(J-2))$.

If the null hypothesis of homoscedasticity is rejected, we propose to model σ according to the following:

$$\sigma_i = exp(\boldsymbol{z}_i\boldsymbol{\gamma}) \tag{16}$$

where $\boldsymbol{z}_i$ is a matrix of variables, likely to be a sub-set of the variables in $\boldsymbol{x}_i$, found to be the source of the residual dispersion and $\boldsymbol{\gamma}$ the associated vector of unknown parameters. Contrary to our approach, the use of a Huber (1967) type adjustment to correct for forms of potential and undiagnosed heteroscedasticity is fairly meaningless given the parameter inconsistency associated with a violation of homoscedasticity.

In turn this will imply re-specifying the log-likelihood function as follows:

$$L = \sum_{i=1}^{n} \sum_{j=0}^{3} \delta_{ij}\ log_e \left[\Phi\left(\frac{\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}}{exp(\boldsymbol{z}_i\boldsymbol{\gamma})}\right) - \Phi\left(\frac{\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}}{exp(\boldsymbol{z}_i\boldsymbol{\gamma})}\right)\right] \tag{17}$$

The ML procedure now involves the estimation of the parameter vectors $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$, and the thresholds θ_j .. The variance-covariance matrix for these vectors is provided by the inverse of the information matrix.

After the estimation of (17) we will repeat the tests in (9), (10), (12) and (14) to verify whether the specification of a separate variance function in (3) solves any of the misspecification, non-normality, heteroscedasticity and thresholds variability problems which might affect the model in (1).

The components of the vector of explanatory variables for the variance function (16) are to be identified employing the "general-to-specific" approach as follows. The initial specification of the vector $\boldsymbol{z}_i$ will include all the explanatory variables in $\boldsymbol{x}_i$ but the constant. Once the results from the estimation of the likelihood function in (17) will be retrieved, determinants of σ_i with p-values greater than 0.1 will be excluded from the specification of (16). The excluded restrictions will be tested for joint significance using a likelihood ratio (LR) test to identify a parsimonious version of (16).

The methodology outlined above, and summarised in (17), is, according to Reilly and Bellony (2009), the development by Caudill and Jackson (1993) of the work of Stewart (1983) cited in Machin and Stewart (1990).
Reilly and Bellony (2009) inform that Stewart (1983) adapts the likelihood function in McKelvey and Zavoina (1975) to allow for known threshold values (i.e., giving rise to the interval regression model) and the estimation of the parameters in a function for the variance of the error term. Nonetheless, Stewart (1983) maintains the crucial underlying hypothesis of homoscedasticity of the error term.
Caudill and Jackson (1993), concerned about the potentially biased and inconsistent estimated betas due to the violation of homoscedasticity, recast the Stewart (1983) procedure considering this mutated condition. Nonetheless, their contribution is limited by the objective of correcting the model for a problem (i.e., heteroscedasticity) which requires solution. We go beyond this approach and estimate the variance function to have a more complete understanding of the contribution of our explanatory variables to shaping SWB in Albania.

Finally, following Ravallion and Lokshin (2000) and their concern for income being endogenous to SWB, the present paper evaluates the need for an instrumental variable (IV) estimation procedure for the consumption measure. The investigation employs the two-stage conditional maximum likelihood (2SCML) procedure Rivers and Vuong

(1988) developed to estimate simultaneous probit models. They suggest estimating the reduced form, or first stage (1S), coefficients for $\boldsymbol{x}_i$ and the vector of instruments $\boldsymbol{\lambda}_i$ using OLS. In doing so we correct for heteroscedasticity between the natural logarithm of consumption (hereafter, $\ln(c)$) and the variables in $[\boldsymbol{x}_i; \boldsymbol{\lambda}_i]$ using the White (1980)/Huber (1967) adjustment. In the present application two different $\boldsymbol{\lambda}_i$ vectors are gathered. In the first specification the instruments are the dummies for owning a bike, a sewing machine, a tape or CD player, a refrigerator and a satellite dish in 1990. In the second, the $\ln(c)$ is regressed against the dummies for the individual speaking Italian, English, Greek and a variable for the age of the household head. Consistent tests for instruments' relevance and exogeneity as well as for the exogeneity of the $\ln(c)$ itself verify that $\ln(c)$ can be confidently treated as an exogenous determinant of SWB in Albania.

The specification of $\boldsymbol{\lambda}_i$ which relies on the dummies for possessing durables in 1990 is probably the more robust of the two due to the impossibility of reverse causality in the reduced form equation. In fact, it is impossible that the 2005 $\ln(c)$ could influence the purchase of durable goods in 1990.

Moreover, the usefulness of modelling the $\ln(c)$ using a non linear formulation based on piece-wise linear splines rather than its whole distribution was explored. The splines allow the $\ln(c)$ to exert different effects on SWB at different positions in the $\ln(c)$ distribution[18]. The difference between the betas for adjacent linear segments were confronted with the null hypothesis of being zero to test the superiority of five splines over the whole distribution of $\ln(c)$. None of the estimated betas was statistically different from the one for the adjacent spline. Moreover, a LR test on (1) being specified including the splines as opposed to including $\ln(c)$ proved that there is no gain in employing the piece-wise linear approximation of $\ln(c)$.

4. Results

The results from the estimation of (3) and of (17) appear in the following table

TABLE 2 ABOUT HERE

[18] See Gujarati (2003:317-319) for an introduction to the topic and Bale *et al.* (2007) for a successful application of the linear splines.

The first column of Table 2 reports the ML estimates for the ordered probit mean regression model (3), which has conventionally been used to model life satisfaction in the empirical literature. The estimated coefficients provide the average *ceteris paribus* effect of a characteristic on the standardised probit index measured in terms of standard deviations. The signs on the estimated coefficients provide the directional impact of the characteristics on an individual's degree of satisfaction with life. It is important to stress that a positive coefficient indicates a positive association between the variable of interest and the probability of being *fully satisfied*. The opposite of the estimated sign indicates the association between the variable and the probability of being *not at all satisfied*. It is difficult to use the signs of the ordered probit coefficients to infer anything about the probability of attachment to the two intermediate categories in the current application.

The mean regression estimates reported in column 1 are consistent with the findings in the existing literature.

However, the diagnostic test results reveal that the homoscedastic regression model violates the normality assumption, and the null hypothesis of homoscedasticity is decisively rejected by the data. In addition, the assumption of constant thresholds across the set of included variables is also fairly decisively rejected by the data. In order to respond to the model violations encountered, we re-estimated the model catering for the presence of heteroscedasticity using expression (17) to determine whether such an approach improves matters in regard to the set of other econometric assumptions.

The explicit modelling of heteroscedasticity generally inflates both the absolute magnitude of the mean estimates and their corresponding standard errors. In particular, self-employment, chronic illness, Orthodox religion, and living alone are the characteristics which were significant in (3) but are now insignificant in the mean equation estimated in (17). The presence of a community organization is the only variable which turns significant in the mean equation in (17) while the logarithm of per capita real consumption and the presence of an inside toilet maintain the significance level they recorded in (3).

While most of the results are qualitatively the same, the heteroscedastic model suggests that the more conventional, yet restrictive, approach leads to an underestimate in absolute terms of the importance of the determinants of life satisfaction. It should also be noted that the heteroscedastic model passes all the key diagnostic tests, which affords some degree of confidence in its estimated effects. In turn, the latter will be at the centre of our interpretational effort.

The only insignificant dummy for the labour market conditions is the one for being self-employed. This result is somewhat surprising since self-employment is regarded as an effective post transition strategy put in place by the most flexible and endowed people to cope with the persistently low job creation rate in the formal labour market (ETF, 2006; World Bank, 2006).

The coefficient for being employed on a household related business is associated with a drop in the likelihood of being fully satisfied with life compared to being employed in a non household related business[19]. This difference might originate from the absence of a financial stream rewarding the individual's effort due to the unpaid nature of her contribution to the household business.

The unemployed are by far the individuals who suffer the largest SWB shortfall compared to the wage employed. The short duration (12 months) and inadequate coverage (on average 65% of the legal minimum wage or 25% of the average public sector earnings) of the unemployment benefits (ETF, 2006) barely help maintaining sufficient confidence in the possibility of funding satisfactory levels of consumption. Interpreting this beta according to Winkelman and Winkelman (1998), the psychological damages associated to joblessness outweigh by far any other welfare reducing condition such as suffering from the collapse of the pyramid schemes. We would rely on an argument *a la* Andrén and Martinsson (2006), based on the limited adaptation to a novel post transition phenomenon, to explain this finding. In fact, the estimation of a coefficient far bigger than the one in Andrén and Martinsson (2006) might be due to the Albanian transition having been probably even more dramatic than the Romanian one. Nonetheless, the Albanian result for unemployment is still about half the magnitude of the one in Winkelman and Winkelman (1998). Furthermore, it is smaller than the values in Hayo and Seifert (2003) and Alesina *et al.* (2004); comparable with Eggers *et al.* (2006) and double the effect in Sanfey and Teksoz (2005).

[19] For expositional convenience we suppose that those who belong to this category receive a wage.

The only education dummy which increases WB is the one for having a university or higher degree. This result might highlight the inadequacy of the college and vocational education in providing the students with the skills and knowledge that the labour market values. According to ETF (2006) individuals who acquired tertiary education enjoyed, on average over the period 2002-04, 23% higher employment rates than those who achieved lower secondary and below. The successful completion of upper secondary is worth a meagre 3% more chances to be employed compared to finishing lower secondary and below. This disadvantaged condition has caused people with this level of education to be over-represented among the pool of international migrants (King and Mai, 2008). These findings might raise grave concerns that individuals with modest income could be excluded from the highest levels of education and in turn from the labour market. The coefficient for the university education in Albania complies with the evidence for Eastern Europe (Hayo and Seifert, 2003); is higher than in Sanfey and Teksoz (2005) and Eggers *et al.* (2006) and even double the US one (Alesina *et al.*, 2004). This marked difference may reflect the higher wage bargaining power Albanian graduates can exert in the local labour market compared to in countries like the US where a larger share of the workforce might have a university degree.

Muslims have higher probability of being fully satisfied with life compared to Catholics in Albania, *on average and ceteris paribus*. The sizable increase in SWB associated with belonging to the major Albanian religion seems to support the claim that Catholicism is providing its followers with feelings and attitudes which contrast with economic development and social cohesion (Putnam, 1993; La Porta *et al.*, 1997; Inglehart, 1999 and Landes, 1998 cited in Guiso *et al.*, 2003).

The low level of significance of the dummy for being Muslim is somewhat resonant of the insignificance of each of the dummies for religious affiliation in McCleary and Barro (2006)[20] and in La Porta *et al.* (1999), cited in McCleary and Barro (2006), when investigating the dependence of political and legal outcomes on, among others, adherence to a religion and per capita GDP.

The large and positive coefficient estimated for the Muslim religion might be due to the social capital/cultural component being fostered by the networking experience

[20] Despite the insignificance of each religion share used in McCleary and Barro (2006), the eight variables included in the actual specification of the model are jointly significant at the 1% level.

associated with attending formal religious services (McCleary and Barro, 2006). Interpreting the estimated coefficient according to the previous view requires supposing that the effect captured by the religion dummies is above and beyond the one controlled for by the NFRIENDS variable. Yet, according to Guiso *et al.* (2003), relying on the social capital/cultural component interpretation of the *role* of religion adherence in a within country analysis is dangerous since it suffers from the impossibility of separating the pure effect of creed from that of historical accidents with the latter shaping the former.

Lastly, interpreting the results for these variables according to the likelihood of religion being a source of discrimination, it is worth noting that only the followers of the most important religion experience higher SWB compared to Catholics. In this respect our result is consistent with the one in Clark and Lekles (2009). The relative segregation experienced by Catholics might have been induced by the isolationism a former colonial power imposed on the members of a minority religion[21].

Suffering from a chronic illness, perhaps surprisingly, is not found to impact on life satisfaction in the heteroscedastic model[22]. Likewise, being the household head does not entail distinctive responsibilities or rewards which induce shifts in satisfaction. On the other hand married individuals, compared to singles, enjoy a rise in the probability of being fully satisfied with life as in Sanfey and Teksoz (2005) and in the Peruvian model in Graham and Pettinato (2002). Despite having proved significantly associated to psychological and material strain which negatively affect SWB as a whole (e.g., Alesina *et al.*, 2004; Senik, 2004; Hayo and Seifert, 2003), the divorced condition (included in the variable OTHMARIT) does not impact significantly on the probability of being fully satisfied with life in Albania. The Albanian evidence might be due to the low incidence of divorces as a percentage of marriages. This, in turn, has been related to the impossibility for the woman to take the children with her when

[21] Clark and Lelkes (2005) and Hayo (2007) report a positive relationship between life satisfaction and engagement in religious activity. This is not something we can provide any insights on here though it may be the case that the degree of engagement is less if the religion is a minority one like Catholicism in Albania.

[22] Easterlin (2003) suggests that when individual health status is accounted for through a single characteristic or condition, concerns of reverse causality are reasonable. Conversely, when it is proxied by a number of variables, accounting for different aspects, it seems more likely that causality would run from the health status' variables to SWB. However, this reverse causality is not the subject of empirical investigation here given its failure to achieve statistical significance.

returning to her parents' house and to female reliance on husbands' income generating capacity (Young, 2006; King and Mai, 2008).

The relationship between SWB and age in Albania is estimated to be U-shaped. The minimum is located at 41.62 years of age[23]. This result is consistent with the empirical evidence from developed countries where satisfaction with life reaches its minimum around the mid-thirties (Hayo and Seifert, 2003) or forties (Sanfey and Teksoz, 2005) rather than around the early fifties (Sanfey, and Teksoz, 2005) or even mid-sixties (Namazie and Sanfey, 2001) as it happens in transition countries. Males are less likely to be fully satisfied with life than females *ceteris paribus*. This evidence is consistent with Alesina *et al.* (2004), Sanfey and Teksoz (2005) and Di Tella *et al.* (2001) and with the survey review of Dolan *et al.* (2008). It might be explained by male household members having to work to compensate for female members' low participation in the labour market.

Contrary to the literature reviewed here we estimate a large value for the impact of consumption[24] on SWB.

There is a significant gap between the estimates for Albania and for countries like Kyrgyzstan (Namazie and Sanfey, 2001) or Russia (Ravallion and Lokshin, 2000). It can be argued that the difference in the estimated and reference betas reflects the countries' development stage. In fact, despite periods of very high GDP growth, the transition from communism, the collapse of the pyramid schemes and the Kosovo war put Albania in such a distress that "by 2004 GDP was around 36% above its pre-transition level" (ETF, 2006:12). It is possible that part of the large positive impact of per capita consumption on SWB in Albania is due to the extra expenditure financed out of the income tax payments forgone by individuals who classify themselves as self employed to hide their condition of black labour market workers.

[23] Dr. Jeff Round was curious to know whether this result was related to the individual respondent's life expectancy. Accounting for cohort effects (i.e., including dummies for a number of classes of the years of birth) should take care of those unobservable phenomena which constitute common conditions for people born in close-by years. Besides the sources of cohort effects already reviewed in Section 1, King and Mai (2008) stress that watching Italian TV might have generated significant age-related characteristics typical of individuals under 40. It is difficult for us to include cohort effects *a la* Blanchflower and Oswald (2007) since we rely on a single survey. At best it could be implemented pooling the 2002 and 2005 ALSMS together allowing for differences in the wording of the questions with no predictable effect on the resulting estimates.

[24] In the current context, household expenditure could be taken to proxy for permanent income.

The beta we estimate might be reliable since the current per-capita GDP is set at US$ 3,400 and it falls short of the lowest value of the threshold above which the irrelevance of the impact of absolute money metrics of WB on SWB[25] is supported. Moreover, the recent detailed empirical work of Stevenson and Wolfers (2008) reaffirms the key importance of absolute income effects in determining life satisfaction. Thus, we use our estimated coefficient for the log of per capita household expenditure to provide some sense of the change in its value that would be required to move an average respondent from from category j to the above category $j+1$. In an OP framework, Holmes (2003) formalises it as $\frac{\theta_{j+1}-\theta_j}{\beta_k}$ where β_k is the OP estimate corresponding to the log of per capita household expenditure variable (x_k) and the θ_j are the estimated thresholds. Hence, the change in the log of per capita household expenditure required, in 2005, to move an average respondent from the *less than satisfied* to the *rather satisfied* category is $\frac{6.973-3.416}{2.027}=1.754$. Given the use of the log form for this covariate, the latter represents almost a six-fold increase in the per capita household expenditure level, *ceteris paribus*. A "back-of-the-envelope" calculation, using these ancillary findings and the per capita GDP data for 2007, suggests that to raise a person from the second from top to the top level of SWB, Albania should experience a rise in per capita GDP from its current level of about US $3,400 to close to US $20,000. In other words, this would signify transforming itself into Portugal, at least as far as current living standards are concerned. While our particular calculation for Albania should be treated as suggestive rather than compelling, it highlights the magnitude of the future challenges required in terms of economic growth to enhance life satisfaction levels in Albania. The speculative nature of this finding is due to the finding of long term economic policies devoted to the growth of per capita GDP being ineffective in yielding a stable rise in average satisfaction levels (Easterlin, 2003). In fact, in an evolving framework, adaptation neutralises these improvements while expectation of ever increasing levels of material WB causes a decline in satisfaction.

Once we control for household income using the per capita expenditure measure, the variables for the number of transfers the household receives display a negative and

[25] It seems that the value of per capita GDP is a good enough proxy for the per capita income used by Layard (2005).

monotonically increasing coefficient[26]. Interpreting the transfers as a purely monetary contribution to the household budget we would expect these variables to report positive and rising coefficients. Due to the negative signs we resort on an argument *a la* Winkelman and Winkelman (1998) to provide a useful interpretation. We argue that having controlled for the level of per capita real consumption, these mutually exclusive variables denote the awareness, of being disadvantaged or poor, an individual who lives in a household which receives welfare subsidies has. The comparison with individuals whose household receives less transfers causes the more in need to feel disadvantaged as the comparison effect (yielding a negative affect) prevails over the expectation (generating a positive affect through the Hirschman tunnel effect) of being able to autonomously finance household consumption.

The number of children in the household diminishes the probability of being fully satisfied. This finding is in accordance with Alesina *et al.* (2004) but is in clear conflict with the positive coefficient in Lelkes (2006).

A positive coefficient would have sprung the interpretation, common in the development literature, of children as a household safety net[27]. Accordingly poor households have more children than richer ones because pupils need to help with household chores or businesses. Besides the historical heritage of high birth rates (King and Mai, 2008), INSTAT (2006) provides evidence that, according to 2002 data, women from a poor background record fertility rates 12% higher than the one for women from higher social classes. A very controversial and peculiar feature of children as safety net in Albania is them being involved in trafficking for illegal purposes or in adoption for profit (see INSTAT, 2006 for supporting evidence and King and Mai, 2008 for a more conservative take on the problem). In Albania children might be the household's safety net through the remittances they are expected to send home once migrated. King and Mai (2008) summarise evidence of households with a large number of remitting migrants performing better than households with fewer. Moreover, according to Albanian customary rules, the younger son is expected to take care of the elderly parents (King and Mai, 2008). Our estimates seem to suggest that

[26] It is certainly the case that lower income households are less satisfied and also receive more state transfers. There is thus likely to be a high negative correlation between the household expenditure measure and state transfers but this just yields higher standard errors for the estimated effects for these inter-correlated variables. This is thus not seen as an important issue here given all relevant estimates are found to be statistically significant.

[27] We are grateful to Dr. Jeffery Round for suggesting we investigate this point.

SWB data do not pick up on this relevant aspect correctly. It might be due to the need for asking migrated children to give up consumption in the host country to remit home and to the parents' consciousness that their children will experience a regression in the standard of living when returning to Albania to care for them.
The negative beta might originate from the additional strain, by adults' perspective, put on a tight household budget by individuals younger than 15. Likewise, interpreting this coefficient altruistically, it would display the concerns that individuals older than nineteen[28] have for a younger generation which is facing an unsuitable education system, a stagnant labour market and is likely to migrate. Therefore, the worries about the future prospects of the youth outweigh the adults' emotional benefits of living in a "young" household.

Contrary to the effect for the number of children and the evidence in Senik (2004), an increase in household size is perceived by its active members as a source of higher satisfaction with life. A larger household can raise individual SWB by improving the chances of accumulating the monetary resources needed to finance consumption or by providing higher psychological contentment. According to the first interpretation, the results for Albania are consistent with the impact of household size on life domains in Winter *et al.* (1999). If a larger supportive family generates a higher probability of being fully satisfied with life, the coefficient for NFRIENDS is expected to and indeed does provide lower benefits compared to HHSIZE.

The Albanian WB data display a long memory of the collapse of the pyramid schemes. It is possible that the significance of this variable is driven by a cohort effect we failed to control for and similar to the one that Elder (1999) cited in Charles *et al.* (2001) ascribes to the Great Depression in inducing age related effects. In turn, this cohort effect might be indirectly confirming the loss aversion hypothesis for Albania (Tversky and Kahneman, 1991 cited in Clark *et al.*, 2008b). The latter states that the absolute negative effect of the unexpected loss of one dollar is larger than the positive effect of gaining a similar amount[29]. In particular, the negative impact of the collapse

[28] This is the minimum age of the respondents in the sample.
[29] This hypothesis has found extensive confirmation in job insecurity, and becoming unemployed in particular, being significantly and negatively associated to SWB (Ravallion and Lokshin, 2000) while windfall gains do not increase SWB (for lottery winners see Brickman *et al*., 1978 in Clark *et al*., 2008b).

– due to the overall size of the schemes and their popularity – might have exceeded even the benefits associated to the subsequent growth which, in 2004, had raised GDP to36% above its pre-transition levels (ETF, 2006).

Because this variable is very precisely determined and of non negligible size we would argue that its use as a SWB explanatory variable helps painting a better picture of the consequences of the collapse of the pyramid schemes than the one that emerges from the analysis of hard data on GDP, inflation, exchange rate, deposits and other monetary indicators. In fact, even if some uncertainty in the data due to political and social unrest is allowed, Jarvis (1999) using a wide range of quantitative indicators somewhat surprisingly concludes that "… the direct effects of [the] rise and fall [of the pyramid schemes] appear to have been limited" (Jarvis, 1999:21). Our findings contrast markedly with this statement and also with the claim that two years after the collapse the citizens' concerns had already shifted towards issues like the internal political conflict, the widespread practise of corruption and the problems in Kosovo (Jarvis, 1999). In particular we argue that the 2005 SWB data account for the psychological consequences of the social unrest which led to more than 2,000 civil victims and the reallocation of asset and wealth, from the losers to the gainers from the collapse. In doing this we succeed where Jarvis (1999) failed due to his use of hard economic data only and to the limited time intervened between his analysis and the actual event. Finally, Jarvis (1999) provides accounts of a decline in the personal confidence level of Albanians occurred when the pyramid schemes collapsed. It seems that the formal financial system is still nowadays regarded as not trustworthy (King and Mai, 2008). It is possible that this decline in confidence has spilled over to higher individual risk aversion, a different attitude to entrepreneurship and a diminished participation in the financial markets compared to before 1997. As a consequence, conspicuous present consumption is preferred to accumulating any savings for future investment (King and Mai, 2008).

All the household asset variables are significant and provide higher SWB. These variables complete the information arising from the per capita real consumption. The area of the dwelling exerts a monotonically increasing effect on the probability of being fully satisfied with life. Living alone would diminish SWB but, since the corresponding asymptotic absolute t-ratio is only 1.6, the beta is statistically insignificant at a conventional level using a two-tailed test – albeit only marginally so.

At the community level land disputes do not change the probability of being fully satisfied with life in a statistically significant way. On the contrary, thefts and the perceived increase in the local population induce a reduction in WB. The presence of a community organization is the only community variable which raises SWB. We interpret the negative effect on SWB associated to the increase in community population as the signal of the difficulties of local infrastructures and public services to cope with a rising number of settlers. The majority of the population growth has happened in urban areas such that "Tirana's population, for example, has more than doubled in the past ten years" (UNEP, 2000:14). This is reflected in our findings since the dummy for living in the cities impacts negatively on the probability of being fully satisfied with life while the same probability does not vary across the country's geographical regions.

The third column of Table 2 reports the variance model estimates. We experimented with the inclusion of the full set of regressors from the mean regression in the variance specification and tested down to a more parsimonious specification. The test for overall statistical significance of the variables excluded from the variance equation in Table 2 yielded an insignificant LR test of 20.46 $\sim \chi^2_{23}$ (prob-value = 0.61).

All the working conditions are associated with more disperse levels of SWB compared to working outside the household. In particular, the unemployed remain the individuals who experience the highest variability in WB conditions. This high dispersion in SWB levels can be induced for instance by the different skills, competences and duration of unemployment which can determine different probabilities of re-employment (i.e., the probability of restoring reasonable levels of SWB).

Similarly, being the head of the household is associated with a greater dispersion of SWB responses. This is likely to depend on whether, for instance, the household head is the only income provider or is alone in raising the children.

The increase in residual dispersion associated to the logarithm of per capita real consumption reflects a fairly standard type of heteroscedastic relationship. It is possible to attribute this result to the inequality in consumption. Sanfey and Teksoz (2005) and Senik (2004) explicitly controlled for the inequality in objective WB measures employing, for instance, the Gini coefficient. Inequality might either decrease SWB, due to the discouragement arising from pure comparison of WB conditions across individuals, or raise it due to the Hirschman tunnel effect. The positive coefficient estimated here supports the claim for the working of a modest, yet fairly well determined, "tunnel effect". The small influence on the overall variance in SWB might be due to the recent appearance of inequality in Albania. According to Kolberg *et al.* (2003) the Gini coefficient, between the '60s and 2002, climbed from virtually zero to a moderate 0.28.

The same variance increasing effect occurs for the largest dwelling area. The significance of the coefficient for the largest dwelling area only, which arguably is enjoyed by the wealthiest Albanians, might mean that even within the wealthy there is someone who, because of the difference in floor space, might feel "super-rich" and "relatively" poor. On the contrary, the dummy for an inside toilet reduces the variance in SWB responses. This dispersion reducing effect might be compatible with the practise of building houses with an inside toilet increasingly becoming a recognized national standard.

Living in a community which has an organisation responsible for dealing with local issues is associated with a reduction in the dispersion in SWB responses. This might be a reflection of the organisation's contribution towards higher compactness of the local population.

Living in an urban settlement is associated with higher residual dispersion in SWB responses probably due to the more varied housing conditions and public facilities respondents enjoy.

Finally, the variance in SWB responses in the central and mountainous region is smaller than the one recorded in the Tirana area. This result denotes fairly homogeneous economic and psychological conditions in these regions. This might be the result of the national and international migration of young individuals which has left behind the adult and those unsuitable to move (World Bank, 2006:50). The more homogeneous and "lonely" the local population, the more likely to report closer levels of SWB. King and Mai (2008) support our claim with personal direct evidence of the desolation, limited business thrive and focus on the day to day fulfilment of basic needs existing in villages and areas which experienced massive outward migration of the young. All these things considered, there is little scope for recording differentiated levels of SWB.

Summary and Conclusions

In common with much of the economics literature devoted to the empirical modelling of life satisfaction, this paper exploited an ordered probit model. However, in contrast with this literature, the current study placed an important emphasis on subjecting the estimated specifications to a battery of diagnostic tests. Not surprisingly for a cross-sectional application, the estimated mean regression model was found to be characterised by heteroscedasticity and, among other things, the assumption of threshold homogeneity was decisively rejected. The explicit incorporation of a heteroscedastic function into the modelling served to resolve all problems of mis-specification.

The tenor of our findings on the determinants of life satisfaction generally reflected that found for both advanced capitalist and transitional economies in the economics literature on this topic. In particular, we found an important and well determined positive role for the household welfare measure used.

The specific context of Albania allowed a number of additional themes to be interrogated. In particular, our study revealed evidence of very long memories among Albanian respondents with respect to the collapse of that country's notorious pyramid schemes. The estimated effect was sizeable and very well determined. It is clear that the psychological scars remain raw for those individuals who had direct experience of the collapse even with the passage of time[30]. Albanian Muslims registered higher life satisfaction levels compared to Catholics, *on average and ceteris paribus*, and this may reflect a sense of isolationism associated with following a minority religion within a largely Muslim country. As anticipated, individuals who reside in communes characterised by thefts enjoyed lower levels of satisfaction with the estimated effect

[30] Since life satisfaction is highly correlated with memories of the past through personality, this may represent the conflation of the pyramid effect and the role of personality. Given we cannot control for individual level heterogeneity in our analysis, the influence of personality is not something we can isolate in our cross-sectional dataset.

statistically indistinguishable from that associated with the pyramid scandal[31]. However, in communes where community organizations were active mean life satisfaction was higher and the variance in responses was found to be lower, *on average and ceteris paribus*.

The empirical approach adopted in this study was novel in its emphasis on econometric model evaluation and its use of an ordinal heteroscedastic model. The failure to explicitly control for the presence of heteroscedasticity was found to impact both the magnitude and the precision of the ordinal mean regression estimates. A useful additional feature of explicitly modelling heteroscedasticity was the additional empirical insights provided on how selected covariates impacted the variance in satisfaction responses across individuals. Given our preferred model successfully passed a battery of diagnostic tests, we argue that our approach has provided for some degree of confidence in the empirical evidence reported.

Nevertheless, we do acknowledge an important limitation associated with the work undertaken here. Our analysis used cross-sectional data, which obviously does not permit to control for individual unobservable effects. This type of neglected heterogeneity may be important in the analysis of life satisfaction. The work of Ferrer-i-Carbonell and Frijters (2004) highlights how controlling for such unobservables can influence the findings of what does and does not impact SWB. However, in order to introduce such individual-level heterogeneity effectively panel data with a reasonable temporal dimension are required. Such data are not currently available for Albania and are unlikely to be so even in the near future. Moreover, the use of panels with relatively short time dimensions may not prove entirely informative on this matter as time invariant factors (i.e., the individual fixed effects) may absorb most of the variance in the included covariates which, over a relatively short span of time, may possess minimal temporal variance. Similarly, even the inclusion of fixed effects will not cure the spurious correlation arising from the time-varying factors which simultaneously increase SWB and income (or any other explanatory variable) (Clark *et al.*, 2008b). Thus, in the absence of panel data, we argue that the use of

[31] The null hypothesis of no difference between the scarring effects of the pyramid collapse and that of the effect of crime on life satisfaction is upheld by the data using the heteroscedastic model with an absolute asymptotic t-value of 0.31.

cross-sectional data in conjunction with an adequately specified econometric model can provide important value-added to the study of the determinants of SWB.

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Table 1 Descriptive Statistics for the Estimating Sample

Individual level variables	**Description**	**Mean**
UNEMPL	=1 if the respondent is unemployed, =0 otherwise	0.0458
OLF	=1 if the respondent is out of the labour force, =0 otherwise.	0.2915
HEMPL	=1 if the respondent is employed on a household business, =0 otherwise	0.1632
SELF_EMP	=1 if the respondent is a self employed or an employer, =0 otherwise	0.1803
WEMPL	=1 if the respondent is employed in a non household related business, =0 otherwise	0.3192
PRIMARY_4	=1 if the respondent has no diploma or has a primary 4 diploma, =0 otherwise.	0.4122
PRIMARY_8	=1 if the respondent has a diploma from primary 8, =0 otherwise	0.1738
SEC_DIPLOMA	=1 if the respondent has a secondary diploma, =0 otherwise	0.2190
VOC_DIPLOMA	=1 if the respondent has a vocational diploma, =0 otherwise	0.1085
UNIV	=1 if the respondent has an university and more education title, =0 otherwise	0.0866
ILLNESS	=1 if the respondent has a medically ascertained chronic illness, =0 otherwise	0.2070
MUSLIM	=1 if the respondent is Muslim, =0 otherwise	0.7852
ORTHODOX	=1 if the respondent is Orthodox, =0 otherwise	0.0992
OTHERREL	=1 if the respondent is from other religions, =0 otherwise	0.0393
CATHOLIC	=1 if the respondent is Catholic, =0 otherwise.	0.0763
MARRIED	=1 if the respondent is married, =0 otherwise	0.8909
OTHMARIT	=1 if the respondent enjoys other marital statuses, =0 otherwise	0.0766
SINGLE	=1 if the respondent is single, =0 otherwise.	0.0325
HHHEAD	=1 if the respondent is the household head, =0 otherwise.	0.6189
AGE	Age of the respondent in years.	45.36
MALE	=1 if the respondent is male, =0 otherwise.	0.5696
Household level variables	**Description**	**Mean**
PRCONS	Real per capita household consumption (in 2002 Leks)	10,087.4
LNPRCONS	Natural logarithm of real per capita household consumption (in 2002 Leks)	9.0742
TRANSFER1	=1 if the household receives one state transfer, =0 otherwise	0.4034
TRANSFER2	=1 if the household receives two or more state transfers, =0 otherwise	0.0961
TRANSFER0	=1 if the household receives no state transfers, =0 otherwise.	0.5005
NCHILDR	Number of children in the household.	2.1444
NFRIENDS	Number of close friends of the respondent.	1.9319
HHSIZE	The number of individuals in the household.	4.6073
PYRAMID	=1 if the household of the respondent was affected by the collapse of the pyramid schemes between 1989 and 2005, =0 otherwise.	0.3134
LIVEALONE	=1 if the household of the respondent lives within a single household home, =0 otherwise.	0.6620
INSIDEWC	=1 if the dwelling the household of the respondent lives in has a toilet inside, =0 otherwise.	0.7465
AREA<70	=1 if the dwelling the household of the respondent lives in has an area of less than 70 square meters, =0 otherwise.	0.3900
AREA70_99	=1 if the dwelling the household of the respondent lives in has an area between 70 and 99 square meters, =0 otherwise.	0.3832

Table 1 Descriptive Statistics for the Estimating Sample *(continued)*

Household level variables	**Description**	**Mean**
AREA100_130	=1 if the dwelling the household of the respondent lives in has an area between 100 and 130 square meters, =0 otherwise.	0.1752
AREA130+	=1 if the dwelling the household of the respondent lives in has an area above 130 square meters, =0 otherwise	0.0517
Community level variables	**Description**	**Mean**
LAND_DISPUTE	=1 if the community the respondent belongs to is afflicted by disputes over land possession, =0 otherwise.	0.5631
THEFTS	=1 if the community the respondent belongs to is affected by thefts, =0 otherwise.	0.3579
POP_INCREASE	=1 if the community the respondent belongs to enjoys an increase in population, =0 otherwise.	0.5867
COMM_ORG	=1 if the community the respondent belongs to enjoys the presence of a community organization, =0 otherwise.	0.8980
URBAN	=1 if the respondent resides in an urban location, = 0 otherwise.	0.5631
COASTAL	=1 if the respondent resides in a coastal area, = 0 otherwise.	0.2887
CENTRAL	=1 if the respondent resides in the central area, = 0 otherwise.	0.2621
MOUNTAIN	=1 if the respondent resides in the mountain area, = 0 otherwise.	0.2713
TIRANA	=1 if the respondent is in the Tirana area, = 0 otherwise.	0.1779

Notes to Table 1:
(a) Albanian Living Standards Survey (2005) using 2,923 individual-level observations.
(b) The mean column reports the sample proportion for binary variables and conventional means for the continuous variables.
(c) In *Italics* excluded variables in estimation.

Table 2 Estimation Results for the Mean and Mean and Variance Equations

	Ordered Probit	Ordered Probit Corrected According the "General-to-Specific" Procedure	
Variables	β_{mean}	β_{mean}	$\beta_{variance}$
Constant	-7.9124*** (0.6565)	-16.7515*** (6.5927)	§
UNEMPL	-0.4540*** (0.1142)	-1.0058** (0.4643)	0.2652** (0.1145)
OLF	-0.2322*** (0.0657)	-0.4876** (0.2354)	0.1319*** (0.0540)
HEMPL	-0.2342*** (0.0771)	-0.4420* (0.2326)	0.1368** (0.0692)
SELF_EMP	0.1320** (0.0644)	0.2774 (0.1757)	0.0976* (0.0549)
PRIMARY_8	0.0026 (0.0881)	-0.0073 (0.1856)	§
SEC_DIPLOMA	0.0543 (0.1013)	0.0733 (0.2134)	§
VOC_DIPLOMA	0.0768 (0.0971)	0.1164 (0.2048)	§
UNIV	0.3525*** (0.1144)	0.7064** (0.3573)	§
ILLNESS	-0.0941* (0.0578)	-0.2021 (0.1459)	§
MUSLIM	0.2378*** (0.0876)	0.4676* (0.2568)	§
ORTHODOX	0.1849* (0.1078)	0.3901 (0.2654)	§
OTHERREL	0.0410 (0.1382)	0.0179 (0.3028)	§
MARRIED	0.4861*** (0.1288)	0.9571** (0.4472)	§
OTHMARIT	-0.1711 (0.1628)	-0.4267 (0.4021)	§
HHHEAD	0.0106 (0.0951)	-0.0474 (0.1945)	0.2018*** (0.0446)
AGE	-0.0495*** (0.0180)	-0.0999* (0.0523)	§
AGESQ	0.0006*** (0.0002)	0.0012** (0.0006)	§
MALE	-0.2070** (0.0955)	-0.4065* (0.2498)	§
LNPRCONS	0.9648*** (0.0549)	2.0270*** (0.7927)	0.0872** (0.0424)
TRANSFER1	-0.1084** (0.0504)	-0.2427* (0.1381)	§
TRANSFER2	-0.2151*** (0.0869)	-0.3976* (0.2261)	§
NCHILDR	-0.1234*** (0.0278)	-0.2435** (0.1095)	§
NFRIENDS	0.0278** (0.0117)	0.0574* (0.0337)	§
HHSIZE	0.1243*** (0.0225)	0.2442** (0.1054)	§
PYRAMID	-0.1672*** (0.0486)	-0.3627** (0.1769)	§
LIVEALONE	-0.1294** (0.0564)	-0.2595 (0.1627)	§

Table 2 Estimation Results for the Mean and Mean and Variance Equations *(continued)*

	Ordered Probit	Ordered Probit Corrected According the "General-to-Specific" Procedure	
Variables	$\boldsymbol{\beta_{mean}}$	$\boldsymbol{\beta_{mean}}$	$\boldsymbol{\beta_{variance}}$
INSIDEWC	0.3514*** (0.0611)	0.8170*** (0.3357)	-0.1207** (0.0575)
AREA70_99	0.1398*** (0.0513)	0.2907** (0.1508)	-0.0444 (0.0460)
AREA100_130	0.3704*** (0.0664)	0.7258** (0.3141)	0.0362 (0.0590)
AREA130+	0.5748*** (0.1066)	1.2657** (0.5576)	0.1716** (0.0842)
LAND_DISPUTE	-0.0765 (0.0461)	-0.1651 (0.1146)	§
THEFTS	-0.1526*** (0.0477)	-0.3181** (0.1634)	§
POP_INCREASE	-0.1676*** (0.0601)	-0.4138** (0.2015)	§
COMM_ORG	0.1103 (0.0773)	0.4151* (0.2331)	-0.2334*** (0.0697)
URBAN	-0.1638*** (0.0636)	-0.4153** (0.2079)	0.1467*** (0.0565)
COASTAL	-0.0025 (0.0726)	-0.0435 (0.1643)	0.0671 (0.0611)
CENTRAL	0.0980 (0.0745)	0.2576 (0.1858)	-0.1336** (0.0637)
MOUNTAIN	0.0335 (0.0878)	0.0648 (0.1830)	-0.1589*** (0.0632)
Thresholds			
θ_1	1.6478*** (0.0330)	3.4165*** (1.3282)	§
θ_2	3.2522*** (0.0643)	6.9726*** (2.7593)	§
Diagnostics			
Observations	2923	2923	
Log-Likelihood Value	-2742.8	-2702.8	
Pseudo-R^2	0.156	0.150	
RESET Test ~ χ^2_3	2.70	3.46	
Normality Test ~ χ^2_2	7.91**	2.80	
Threshold Homogeneity ~ χ^2_{76}	145.2***	63.3	
Homoscedasticity ~ χ^2_{38}	119.8***	§	

Notes to Table 2:
(a) ***, **, * denotes statistical significance at the 0.01, 0.05 and 0.10 level respectively using two-tailed tests.
(b) § denotes not applicable in estimation.
(c) See expressions (3) and (17) for the relevant log-likelihood functions.

AN EVALUATION OF WAGE INEQUALITY IN ALBANIA

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Abstract. *Inequality has developed in Albania after the collapse of Communism reaching moderate levels. This paper uses the nationally representative Albanian Living Standards Measurement Study surveys from 2002 and 2005 to investigate the significant correlates associated to individual log hourly wage in a sample of workers employed in their primary job. An augmented Mincer (1974) equation estimated for each year controls for the effects of demographic and job characteristics, human capital endowment, and professional and industry affiliation. Considerable care is put in developing a granular modelling of the industry effects in an attempt to detect potential national drivers of growth.*

The multivariate regression consistently estimates an inverted U shaped relationship between the log of real hourly wages, the existence of a wage premium accruing to males, international migrants and individuals with university and beyond education. Somewhat surprisingly the individuals employed in the public sector receive lower wages most likely because of their limited productivity. The most peculiar result regards individuals employed in agriculture, the reference category for industry affiliation, enjoying higher wages compared to any other industry controlled for. While this is an almost unique result it might be due to Albania still lagging behind several neighbouring South Eastern Europe and developing economies. Moreover, the implicit wage premium in agriculture could reflect the major relevance of this sector in the coping strategies of the poor and most vulnerable or a within-country comparative advantage with respect to industries which suffer the most from the backlog in infrastructural development.

A static comparison of the two sets of estimates reveals that while the wage premium accruing to male and more educated workers is higher in 2005 than in 2002, the one enjoyed by return migrants, individuals employed in the private sector and in agriculture is smaller in 2005 compared to 2002.

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Introduction

Wage receipts are the major component of personal income, therefore understanding their dynamics and within-country differences might provide a wealth of information capable to explain existing poverty levels, past and present migration as well as shifts in economic paradigms (González and Miles, 2001).

According to Fairris (2003), both developed and developing countries all over the world have experienced rising levels of wage inequality in recent years. On the other hand, the work of Zhu and Trefler (2005) on the Freeman and Oostendorp (2001) data quoted in Wang *et al.* (2009) reduces the relevance of the above bald statement in that "[…] just over half of 20 developing and newly industrialized countries […] experienced rising wage inequality throughout the 1990s" (Wang *et al.*, 2009:1208 fn 5).

Economics practitioners might interpret wage inequality as resulting from the interaction of wage gaps in gender (Mandel and Semyonov, 2005; Rosenfeld and Trappe, 2000; Nordman and Roubaud, 2009; Nielsen *et al.*, 2004; Deng and Li, 2009), race (Pettit and Ewert, 2009), human capital endowment (Lam and Levison, 1992; Sotomayor, 2004; Hartog *et al.*, 2001; González and Miles, 2001; Patrinos *et al.*, 2009), occupation (Deng and Li, 2009), industry affiliation (Deng and Li, 2009) and geographical location (Chamarbagwala, 2010; Deng and Li, 2009). Moreover, the specialised theoretical and empirical literature also considers the explanatory power of trade liberalisation (Arbache *et al.*, 2004; Hartog *et al.*, 2001), migration events (Reed, 2001), minimum wage (Pacheco, 2009; González and Miles, 2001) and the role of unions (Fairris, 2003).

The present paper aims to provide a preliminary picture of wage inequality in Albania using a human capital model as the explanatory framework of choice and the 2002 and 2005 Albanian Living Standards Measurement Study (ALSMS) surveys as the data sources. Since the collapse of the communist regime in 1992, Albania has been facing a non negligible, although still low, wage inequality. Using augmented Mincer equations (Mincer, 1974), comprising both the explanatory variables most commonly considered by the relevant empirical literature and some unique ones to our datasets, patterns of influence in the two years are described.

The remainder of the paper is organised as follows. The first section reviews some of the established literature on the possible sources of wage inequality. The second section draws a picture of the reforms Albania has undertaken since the transition and describes the country's current economic condition. The third section introduces the two ALSMS employed in the empirical analysis. Moreover it details the exact specification of the explanatory variables used to control for the variability in wage levels. The fourth section formalises the empirical tools while section five provides the estimated results and the related comments. Finally, the last section concludes.

1. Literature review

This section reviews some of the existing literature on the dependence of wage inequality from gender, race, returns from education, work experience, industry affiliation, the type of employer, regional localisation as well as structural variables.

Mandel and Semyonov (2005) argue that women are consistently penalised in their income chances on the global scene. Yet, the empirical evidence finds that the gender wage gap is lower in countries which enjoy more developed welfare states. A hierarchical linear model, accounting for the inequality-reducing effect of more institutionalised contract negotiations, increases the estimated size of the gender based wage discrimination in developed countries. This finding sustains the Mandel and Semyonov (2005) argument of a consistent level of gender discrimination around the world. Rosenfeld and Trappe (2000) use the structural differences between former East and West Germany to control for their unique family and work policies. In doing so it is verified that the statistical discrimination in the access to training and highly paid jobs that women suffer, due to their family commitments, is more sizeable than the one male would experience because of their military service (Rosenfeld and Trappe, 2000). The predominant role of discrimination in shaping gender gaps holds for Denmark, too (Nielsen *et al.*, 2004). Nordman and Roubaud (2009) note that, in developed countries, women's accumulated on-the-job-training provides significant returns after having taken a maternity leave. Neglecting the impact of the interruptions due to maternity on schooling or seniority might inflate the residual part of the gender gap which, in a decomposition exercise, is normally ascribed to pure discrimination (Nordman and Roubaud, 2009). Nonetheless, even accounting for these events, significant differences in the returns from education across gender remain in 1998 Madagascar. Deng and Li (2009) include a dummy for being a male in semi-log earning equation and find that the wage premium accruing to males has expanded between 1988 and 2002 in urban China.

Although the racial wage gap might be considered a component of "aggregate" wage inequality, Pettit and Ewert (2009) review several arguments which ascribe the former to different returns to individual characteristics (e.g., education, occupation and family structure), market forces (e.g., the diminished male wage rates which have increased female participation) and governmental policies (e.g., antidiscrimination laws in the access to public jobs). Pettit and Ewert (2009) verify that imputing the wage of women not participating in the labour market, in the US between 1979 and 2005, increases the estimate of a racial wage gap by 11 to 71%, compared to the one calculated on employed women only. Nonetheless, the rising – over time - racial gap can only tentatively be attributed to marriage, age and higher returns to education, rather than to gender differences in the endowments.

Lam and Levison (1992) remark that the relationship between schooling and income inequality has been thoroughly investigated in the economic literature. Frequently, researchers have used the human capital models (e.g., Mincer (1974)) as their preferred interpretational framework. The latter predict that at a single point in time individuals enjoying more education should receive higher wages. On the other hand, Lam and Levison (1992) report that the prediction, formalised by Becker and Chiswick (1966) and Chiswick (1971), that a decline in the inequality of education would reduce earnings inequality is only sometimes verified in cross country analyses. Patrinos *et al.* (2009) stress how the effect of education on the whole distribution of inequality is different between developed and developing countries. In the former education raises inequality benefitting more the individuals at the top of the income distribution while in the latter there is no unique effect. In fact while countries in Latin America display a picture of the education-wage inequality relationship resembling to the one typical of high-income countries, East Asia enjoys declining levels of inequality together with an increase in educational qualifications (Patrinos *et al.*, 2009). Kunze (2008) reports that studies on the US, employing the Oaxaca-Blinder decomposition, have conservatively estimated – due to a likely downward bias - that 25 to 50% of the national gender wage gap may depend on differences in human capital[1]. Hartog *et al.* (2001) affirm the centrality of the changes in the returns from schooling in explaining the substantial increase in Portuguese inequality during the 1980s and the early 1990s. In fact, a positive effect of education on the logarithm of monthly gross wages is estimated. Lam and Levison (1992) find that a lower dispersion in acquired education and declining rates of return from human capital would have progressive effects (i.e., reduce inequality) on wage differences in Brazil between 1976 and 1985 if only their impact were larger than the increase in inequality due to residual variance. Sotomayor (2004) confirms the tight and strong link existing between changes in the returns from education and the dynamics of wage inequality in Brazil but investigated it over the period 1976-2001. Deng and Li (2009) estimate returns from education which have augmented over time as market based wage-setting policies better remunerated skills in post-communist China. Increasing returns from education, accruing to the more educated workers, drive the increase in inequality in Uruguay, too (González and Miles, 2001).

Additional individual-level characteristics which might play a role in explaining wage levels and their inequality include work experience, industry and occupation affiliation, type of employer and regional localisation.
Deng and Li (2009) consider work experience as defined both by job seniority and on the job training and note that Knight and Song (1994) find that the former explains almost all of the wage inequality in planned economies[2]. Deng and Li (2009), including a second degree polynomial in work experience, estimate an inverted U shaped earning profile.
Melly (2005) evaluates how wages differ between the private and public sector in Germany and, using quantile regression decomposition, finds that the conditional wage distribution is less unequal in the latter. Moreover, the public sector wage gap favours females and penalises males while the gender gap in the public sector is smaller than in the private and competitive one. Patrinos *et al.* (2009) develop the estimation of the impact of education on wage inequality differentiating this effect for

[1] According to Lam and Levison (1992), Langoni (1973) attributed the impact of human capital on earnings to the occurrence of quasi-rents.
[2] Nonetheless according to Deng and Li (2009), Gustafsson *et al.* (2003) sustain that the effect of work experience in China surpassed the one in the former URSS. Although Albania has abandoned a planned economy since 1992, a few instances of dependence of wage levels from seniority might be maintained.

private and public employees. In particular, the reduction in wage inequality at the national level recorded in East Asia stems from the public sector. On the other hand, the private one is characterised by higher educational attainment yielding increasing levels of inequality, at the global level. According to empirical evidence, individuals working for foreign enterprises consistently enjoy a wage premium compared to employees in collective enterprises in China between 1988 and 2002 (Deng and Li, 2009).

The differences in wage levels due to industry affiliation are investigated in Deng and Li (2009) including 12 live dummies and using manufacturing as the reference category. A significant variability of wages across sectors is detected. In particular, in 1988 only the mining industry displayed a wage premium, compared to manufacturing, while all the other industries recorded wages which were either indistinguishable from the base category or significantly lower. By 2002 the picture of industry wage premia had changed with seven out of the nine significant dummies for industry affiliation suggesting the existence of a wage premium compared to manufacturing. Somewhat surprisingly, except for being employed in mining, all the positive wage gaps pertain to the service side of the economy. Deng and Li (2009) attribute this finding to certain sectors (i.e., telecommunications) exerting some monopolistic power. Similarly, higher responsibility and skilled individuals have a positive and widening, between 1988 and 2002, wage premium, compared to production workers in China.

Although Chamarbagwala (2010) analyses the spatial differences in consumption levels - rather than wages - in a liberalising India, the urban-rural gap in welfare conditions is deemed a crucial variable in the decision to move away from the countryside and into the cities. Deng and Li (2009) detect within-China differences in remunerations which significantly increase concerns for an expanding distance in the standards of living of the richest and poorest areas of the country.

The existence of an immigrant-natives wage gap, and its link with overall wage inequality, is mainly an empirical question which has returned different answers according to the country considered, the nature of the data (i.e., cross-section or panel data) and the estimation technique used. Nonetheless, a wage disadvantage penalising the immigrants, compared to the natives, and yet reducing with the permanence in the country seems to be a regular finding (Nielsen *et al.*, 2004). Reed (2001) suggests that while previous research posited the absence of any impact of immigration on the labour market's outcomes of the natives, inequality in the local labour market might rise due to the influx of low-skilled and low-paid workers. Reed (2001), using six counterfactual distributions of skills and wages, assesses that, between the late 1960s and the late 1990s in the US, immigration has determined around half of the regional variance in inequality in the last year of data (1997) and just above 25% percent of the increase over this period. Contrary to the established empirical evidence, Nielsen *et al.* (2004) report female migrants' wages being 10 to 15% higher compared to natives in Canada (Beach and Worswick, 1993; Shamsuddin, 1998). Nonetheless, accounting properly for selectivity bias the usual pattern of a negative, although reducing, gap afflicting the immigrants is restored (Field-Hendrey and Balkan, 1991). Nielsen *et al.* (2004) propose an elaborated decomposition of the wage gap of immigrants in Denmark, resting on panel data estimates corrected for self-selection, which highlights that differences in personal qualifications and the incomplete assimilation of the immigrants in the labour market are driving the size of the gap.

Hartog *et al.* (2001) allow for part of the increased earnings inequality in Portugal to depend upon the interaction of the increased Portuguese integration in the EU, an historical comparative advantage in labour intensive industries, an influx of European funds and FDIs devoted to technological upgrading the productive technology as well as a rise in the demand for skilled labour.

González and Miles (2001) argue that asymmetries, across developed and developing countries, might exist in the effect of, in particular, a fall in the real minimum wage on earnings inequality. While the expected rise in income inequality due to a decline in minimum wage is verified in many developed countries, it might not hold in some developing ones. A model for the inequality in hourly wages suggests that a rise in the real value of minimum wage limits, albeit in a marginally significant manner, only the inequality experienced by the New Zealand's youth in the lower tail of their wage distribution (Pacheco, 2009). Dickens and Manning (2004) find that the introduction of the UK national minimum wage in 1999 has affected the wage distribution with short lived effects due to the rising average market wage. González and Miles (2001) confirm that, at least for the lower tail of the income distribution in Uruguay over the period 1986 -97, the limited government capacity to enforce a statutory minimum pay has led to its decline being associated with a decline in income inequality. This is largely due to the legal minimum wage being irrelevant with respect to market outcomes. Using a sub-national Brazilian panel time-series for the ’80 and first half of the ’90, Bittencourt (2009) finds that minimum wage policies in Brazil have been progressive.

Lastly, Fairris (2003) estimates, using counterfactuals, a significant reduction in wage inequality because of the presence of workers’ unions in Mexico. In turn, their effect is found to depend on the number of workers represented and on their ability to remove wage differences, between and within plants, due to observable characteristics (i.e., gender, geographical, education, occupation and industry differences). Indeed, the decline in unions’ coverage, changes in bargaining procedures and the need for maintaining low wages and high occupational flexibility to attract US investments have led to a rise in national wage inequality by 5.6%.

2. The Albanian Reforming Experience and the Current Stage of Development

Besides experiencing a very unsettled transition from communism, Albania has been radically changed by the trade and agricultural reforms which have occurred since 1992. The effort and determination that Albania put into liberalizing trade is evident into its willingness to participate to the WTO, to several Regional Trade Agreements (RTAs) and to fulfil the requirements for the EU accession. It could be argued that, Albania was “forced” to liberalise, after having embraced a free and market-based economy, because at the time of its transition many of its European neighbours were already significantly integrated. At the time of the collapse of the communist regime the EU was the only sizeable market Albania could have a fruitful trade relationship with. The transition to a market based economy coincided with, according to World Bank (2002b) cited in Mancellari (2005), the country shifting from exporting - towards the EU - products highly reliant on natural resources to labour intensive ones (i.e., footwear and apparel). Moreover, World Bank (2004), building on the finding that light manufacturing is the major production exported to the EU, asserts that focusing on these goods could be a viable development strategy in the medium-run. On the other hand Gligorov *et al.* (2004) cited in Mancellari (2005) question the possibility that Albania has identified any comparative advantage able to drive its development.

The trade liberalisation reforms that Albania underwent mainly concerned the formation of prices, the abolition of state monopoly on foreign trade and the free conversion of the domestic currency. A consistent picture of the reforms appears in the following table.

TABLE 1 ABOUT HERE

The WTO accession is probably the most decisive move that brought Albania on the international scene while the participation to 8 RTAs is aimed at farther developing trade relationships with Albania's closest neighbours and the EU in particular (Mancellari, 2005). All the trade liberalisation initiatives attempt to increase the Albanian level of commercial integration to better identify any likely comparative advantage based on productivity differentials as well as on technological and knowledge transfers (World Bank, 2004). Yet, having implemented the RTAs in the form of bilateral FTAs, the nations involved have missed out on the long term gains arising from these very same "soft" benefits; economies of scale and from a better negotiating position on the global trading markets, in particular (Grupe and Siniša, 2005 in Mancellari, 2005)[3]. Industry wise, the RTAs that Albania has signed are more advantageous in manufacturing (almost completely liberalized) than in agriculture. In fact, World Bank (2004) notes that Albania has significantly increased, in the 2000s, its export of construction and mining manufactured machinery towards the South Eastern Europe (SEE-8[4]) economies while importing, from them, more raw materials.
In summary it seems that the picture of the Albanian export is quite fragmented, with respect to its destination: the SEE-8 countries seem to be a market suitable to absorb the rather capital intensive machinery for the construction and mining industries, the EU might continue to acquire the light manufacturing in leather and clothing while agricultural products may have a more international and global appeal (World Bank, 2004).

The Albanian agricultural land has historically been assigned to users without an adequate system of property rights (Deininger *et al.*, 2007). The two most radical agricultural reforms in Albania occurred according to the 1945 Agrarian Law and the

[3] Indeed, World Bank (2004) evaluates the Albanian experience in trade reforming suggesting that the most significant gains originated from the WTO accession and the preferential concessions from the EU.
[4] The SEE-8 group includes Bosnia and Herzegovina, Bulgaria, Croatia, Macedonia, Moldova, Romania and Serbia and Montenegro.

Law 7501. The former eliminated private ownership and instituted a collective regime through the confiscation of large possessions, redistribution of parcels and the consequent creation of agricultural production cooperatives and state farms. The latter initiated a post-communism redistribution of the agricultural land, on a per-capita basis, to those farm workers and farmers which used to work and own the land before its collectivisation. Therefore, the 550 formerly state and collective farms were split up into small parcels of about one hectare and distributed to 460,000 private owners[5] (World Bank, 2002a). Deininger *et al.* (2007) report that Kelm *et al.* (2001) note that the Albanian families which belonged to agricultural production cooperatives were granted inexpensive ownership rights while those working on state farms received use rights only but benefitted from the redistribution of livestock, orchards and fruit trees. Yet, the subsequent efforts at fine tuning the change in land structure were poorly managed such that the final result was a skewed allocation of agricultural machinery and farm equipment which sometimes were looted (Civici, 2001 in Deininger *et al.*, 2007). Moreover, the reform process which dismantled the collectivist land management caused major tensions on the location of village borders (Deininger *et al.*, 2007) and limited maintenance and innovation of several relevant agricultural infrastructures.

Unfortunately, despite the sweeping reforms undertaken so far, present Albania seems rather underdeveloped, compared to several neighbouring countries, according to a number of economic indicators (Mancellari, 2005). Kaltani (2007) provides a detailed assessment of the Albanian stage of development using the SEE-8 countries and the transition economies at large as a benchmark. The analysis concerns areas such as educational investment, financial depth, telecommunication infrastructures, governance, labour market flexibility and firm creation. Kaltani (2007) compares the country's current performance in each area with the one that would be predicted by its per capita GDP level. The resulting picture is summarised in the following table.

TABLE 2 ABOUT HERE

[5] Deininger *et al.* (2007) inform that, according to Mathijs and Swinnen (1998) the Albanian reform strategy was somewhat unique in the former communist world since it was replicated only in Romania.

Kaltani (2007) recognises that Albania is underperforming in some aspects of infrastructure provisioning, financial development, public expenditure for education and of the adequacy of the legal system. On the other hand, a few characteristics of the labour market seem fairly well developed.

Kaltani (2007) reports that, according to World Bank (2006b), Albania has recorded a significant rise in the growth rate of private credit from a very low initial level. Nonetheless, the whole financial sector still lacks the rules and regulations for an efficient functioning. Forms of credit rationing still exist such that it is claimed that formal institutions finance individuals already flushed with cash (King and Mai, 2008). Similarly, the copious remittances from international migrants, which according to IMF (2005) in Azzarri *et al.* (2008) have topped 14% percent GDP or 1.7 times the value of exports in 2004, cannot be considered a suitable source of development finance substituting for a formal, well-functioning credit system. In fact, only at times a significant share of these privately transferred funds has been devoted to productive investments (McCarthy *et al.*, 2006).

Despite the adequate level of acquired education in Albania, given the current level of per capita GDP[6], Kaltani (2007) echoes that UNICEF (2003) ascertains that in 2002 Albania spent only 2.6% of its GDP on education. Dropping out from school has increased largely due to the higher incidence of the cost of transportation to school and because of the government reneging on, among others, teacher training and school maintenance (Meurs *et al.*, 2008).

The legal system seems to have failed to generate trust because the courts are not impartial enough. This has caused a widespread violation of national laws (Broadman *et al.*, 2004 cited in Kaltani, 2007). For instance, Bozgo *et al.* (2002) report the resurgence of blood feuds to settle disputes over land ownership and tenure in northern Albania. Weak courts and legal institutions might hinder the finalisation of prospective investment agreements and disrupt business relations which rely on the straightforward and fair enforcement of contracts. Yet, the possibilities of starting up a business in Albania are greater than those predicted by the national overall level of development. In fact, entrepreneurs do not seem hassled by significant red tape.

On a rather positive note, Kaltani (2007) suggests that it is quite easy to hire workers in Albania. The use of term contracts and the remuneration at the minimum wage are

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[6] This might largely be due to the lead that Albania had built, under the communist regime, over several neighbouring, in particular non-communist, countries (Azzarri *et al.*, 2008).

practices which are not significantly discouraged. Likewise, it seems that firms are not always required by the law to notify, and get permission from, either the unions or the labour ministry to shed jobs. On the other hand, it seems that Albania has fallen behind in allowing more weekend and night shifts to be scheduled and in providing workers with a minimum number of vocation days.

3. The data

The present paper employs the 2002 and 2005 Albanian Living Standard Measurement Study (ALSMS) surveys to perform an analysis of the inequality in the distribution of wages in the two years. The two surveys include a household, community and a price questionnaire. A diary for recording household consumption provides valuable information on expenditure patterns. Both household questionnaires are based on recommendations from Grosh and Glewwe (2000). The 2005 wave incorporates novel sections and provides more comprehensive information, compared to the 2002 wave, for those topics surveyed in both years (i.e., in the labour section, the number of co-workers in the main job).

The 2002 ALSMS uses a two stages stratified sample. The Primary Sampling Units (PSUs) are the Enumeration Areas (EAs) used in the 2001 Population Census. The EAs were selected according to their geographic location (mountain, central, coastal and Tirana), being in areas characterised by big/small towns and rural or urban environment. The survey interviewed 8 households (out of the twelve candidates) in each of the 450 PSUs reaching in total 3,600 households and, in turn, 16,521 individuals.

The 2005 ALSMS builds on its previous iteration. It uses 455 randomly selected PSUs to collect nationally representative information and 20 non random PSUs to oversample the Tirana area and account for its astonishing development in recent years. As a whole the nationally representative 2005 ALSMS provides 16,387 usable individual observations distributed in 3,638 households. The present paper employs only the information arising from the nationally representative part of the ALSMS.

The estimating samples comprise individuals, aged 15 to 65, who have received a net last pay different from zero as a reward for the effort they have exerted in their main job. In turn, the main job is the one which has occupied the respondent the longest hours during the seven days before the interview. The main job is used as a proxy for the permanent one which Winters *et al.* (2008) associate to "[…] long term, stable and presumably high productivity work […]" (Winters *et al.*, 2008:6). The self-employed are dropped from the sample because it is well known that their earnings are frequently high because of consistently foregoing tax payments. Yet in doing this a significant part of the Albanian labour market is neglected (ETF, 2006; World Bank, 2006b).

Following Arbache *et al.* (2004) the net last monthly wage is turned in hourly figures by dividing for the number of hours worked in one week times 4.33[7]. The last paid wage is preferred to the usual one because the latter refers to a 12 month time span, and, as such, is more prone to measurement errors (Cragg and Epelbaum, 1996). Furthermore, the hourly wage is expressed in real terms by dividing it by the Paasche index calculated at the PSU level and standardized with respect to the capital. The hourly real wage is deflated by the national rate of inflation between 2002 and 2005 to enhance the comparability of the two years. The distribution of the real hourly wages, expressed in 2002 Leks, is trimmed at the first and 99th percentile to exclude those outliers which were recorded because of misreporting. Lastly, the departures from normality, which are typical of money metric measures of welfare (i.e., consumption or income), have been cured by applying a logarithmic transformation[8].

Since, according to Nordman and Roubaud (2009), there is very little consensus on the covariates which might fruitfully inform on the variation in wage levels, several productivity-related individual (e.g., education, labour market experience and marital status) and job characteristics (e.g., occupation and industry affiliation) are included as explanatory variables in a Mincer (1974) type equation. Moreover, the model is augmented considering the variables which are likely to play a significant role in the Albanian context[9].

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[7] The reported values of the last net pay can be referred to a monthly, 15 days, weekly and daily time span. Therefore, these figures are turned into monthly amounts by multiplying the two-weeks', weekly and daily salary by 2.165, 4.33 and 21.65, respectively. Once all the amounts were transformed into their monthly correspondent values, they were divided by the number of hours worked per week times 4.33 (Arbache *et al.*, 2004).

[8] A Box-Cox test rejected all the possible transformations that could be applied to the trimmed distribution of the real hourly wages (i.e., none, logarithmic and reciprocal). Yet, despite being strongly statistically rejected, the logarithmic transformation recorded the lowest value of the chi-square statistic.

[9] This section describes the right hand side of a Mincer (1974) non parsimonious equation.

Controlling for individual age and its squared value should account for the inverted U shaped relationship with the hourly wage due to the effect of physical decadence on personal productivity in blue-collar jobs and of skills' obsolescence in white collar ones. Ñopo (2004) suggests accounting for differences in workers' age to grasp the impact of early entrance in and exit from the labour market on wages. In particular, early entrance disrupts the final salary through lower levels of human capital accumulated while early exit through shorter job tenure (Ñopo, 2004).

A dummy variable for the worker being male should capture the gender wage premium. It is expected to be positive and significant especially due to differences in participation rates (74% for males and 49% for females, (ETF, 2006)) and to the finding that women are very likely to be engaged in the low-productivity farming of the family plot (Azzarri *et al.*, 2008).

Additional likely sources of variability include individual marital status and religious creed. The former may account for some of the productivity enhancing factors associated to the breadwinning condition of working household members, and by men in particular (Young, 2006). The latter could capture the expectation that Protestants are particularly active on the job[10] as well as might account for forms of wage discrimination rooted in ethnical, religious or casteless divisions.

The ten different diplomas surveyed in 2005 and the eight in 2002 were modelled using four distinct dummies for having no or primary diploma, secondary, vocational and university or higher degrees. Using these discrete variables, rather than the completed years of formal education allows for more flexibility in capturing the breadth of wage-increasing characteristics of the actual qualifications (Patrinos *et al.*, 2009). The dummies for the highest educational levels might embody the traditional return effect and a premium to compensate for taking up a job inland. It is plausible that a premium is paid to retain the brightest since the consequences of the international migration of the most skilled workers[11] on the survival of Albanian businesses have been very serious indeed. For instance, Kaltani (2007) reports that the

[10] Litchfield *et al.* (2009) review some of the evidence of the effects of religion on national and individual economic performance highlighting how the Protestants, among all other worshippers, might exert an higher working effort because of their religious belief. Yet, the effect in Albania might be small or even insignificant due to its status of a minority religion and this dummy catering for the atheists and others.

[11] Kaltani (2007) echoes the UNDP (2006) finding that in the years 1990-2005 "more than 50% of Albania's scientists and researchers left the country, and nearly 50% of those were under the age of 40" (Kaltani, 2007:248-249).

BEEPS (2002) data quantify the severity of this loss at 3.2 on a 1-5 scale (i.e., the highest score within the transition countries' pool and a level which grants Albania and Macedonia the joint lead of the SEE-8 group).

A dichotomous variable for having suffered for the last three months from a medically ascertained chronic illness is included as an explanatory variable to signal the impact of a poor health condition on the productivity of labour and, in turn, on the corresponding remuneration.

Since migration has been an important phenomenon in the Albanian history, two dummies are deployed to investigate the possibility of having moved nationally at least once and abroad for a minimum time span[12]. Outward migration has been an important feature of Albania to the extent that between 1989 and 2001 the national population fell by 3.5 percent (Azzarri *et al.*, 2008). The variable for being an international migrant controls for the possibility that the individual who migrated abroad found a job there and developed skills which are higher and/or newer compared to those required in the national labour market[13]. Azzarri *et al.* (2008) find that having accumulated migration experience by the year 2002 increased the probability, of young Albanians in particular, of working in an industry different from agriculture. Kilic *et al.* (2007), using the 2005 data, estimate that returning migration is associated with a higher probability of owning a non-farm business. Carletto and Kilic (2009), controlling for the endogeneity of the choice of migrating and returning in the 2005 ALSMS, find that the labour migrants' human and financial capital developed abroad has allowed them to access jobs requiring superior skills once returned. Hence, the on-the-job training accumulated abroad constitutes a competitive advantage that a return migrant may exploit to obtain an additional wage premium compared to those workers who did not move abroad and might even have been unemployed for quite some time (Çuka *et al.*, 2003). The significant savings that

[12] In the 2002 equation, workers who were born in a municipality different from the one they are currently living in are defined as having migrated nationally. Similarly, international migrants are those individuals who have lived abroad for at least three months since 1997. In the 2005 equation, having lived in a different municipality prior to the current one qualifies for being a national migrant while having lived abroad for at least one month since January, 1st 2004 qualified for being an international migrant.

[13] Both the 2002 and the 2005 ALSMS record migration episodes toward Greece, Italy, Germany, other countries in Europe, the USA, Canada and other countries in the world. Regarding the destination of migration, Carletto and Kilic (2009) find that moving to Italy and beyond (i.e., to any country other than Greece) is significantly associated with occupational uplifting while migrating to Greece is not.

migrants might have accumulated in foreign reserve could help financing independent sources of income across industries and the acquisition of more education (Azzarri *et al.*, 2008). Milanovic and Squire (2005) report that Warner (2002) argues that the ability of high skilled workers to relocate internationally grants them a wage set according to international standards. In turn, due to the inherent differences in the standard of living between rich and poor countries this increases the level of inequality within the boundaries of the latter.
The Albanian regions have been subject to different rates of economic and population growth. It is well known that the mountainous areas have experienced both national and international outward migration while "Tirana's population, for example, has more than doubled in the past ten years" (UNEP, 2000:14). A set of mutually exclusive dummies capturing the stratification in the data is deployed to have a better understanding of the differences in the wage levels. In particular these asymmetries could be due to the variability in cultural attitudes (i.e., gender roles (Meurs *et al.*, 2008)), the localisation of activities, the relative thrive of local communities and the state of infrastructure provision. Moreover, a dummy variable for the individual living in an urban area is necessary to control for the different organisation of rural and urban labour markets.

Professional affiliation is available in both surveys through the 3 digit ISCO 1988 code. Since some of the categories were marginally representative, a 2 digit classification was created and farther aggregated into six dummies for similar categories. Likewise, the industry an individual is working in is recorded according to the 3 digit NACE nomenclature. Due to underrepresentation, the classification was expressed at the 2 digit level and categories comprising less than five individuals were aggregated with the most pertinent neighbouring one. To farther analyse the source of wage inequality, the nature of the employer is modelled using five mutually exclusive dummies for being employed in the public sector, in a private or public firm, in a NGO and for being a collaborator in a private individual's business. Controlling for public employment allows to account for the declining importance of the sector in the Albanian economy, between 1991 and 2001 by three quarters (Kilic *et al.*, 2007) and continuing shedding, especially male, jobs till 2005 (Mendola and Carletto, 2009).

Controls for the worker receiving a bonus on the last net pay and additional payments are included to grasp the impact of the structure of the labour market on wage inequality that Attanasio *et al.* (2004) ascribe to minimum wage policies (see also González and Miles, 2001; Pacheco, 2009; Dickens and Manning, 2004), trade unions' power (see also Fairris, 2003) and to informality[14]. Nonetheless, due to the proxy nature of this control variable the estimated effects might be highly contaminated and caution is required when commenting.

The number of years spent in the reported main job, calculated as the difference between the survey's year and the year in which the responded first started that occupation, is included to allow for job tenure exerting an independent effect on labour income. Mincer and Jovanovic (1981) introduce this innovation to better account for the possible returns to specific on the job training (Nordman and Roubaud, 2009). Nielsen *et al.* (2004) argue that distinguishing between the two types of experience is crucial to correctly identify the effect of assimilation on the difference between immigrants and natives' wage. Moreover, its squared value is included to investigate whether a non linear relationship exists. Accounting for this additional source of wage variability should help in limiting the occurrence of those spurious industry wage premia Cragg and Epelbaum (1996) conceive.

Because firm size is likely to influence whether the wage negotiating power resides with the employer or the employee, seven dummies for the number of co-workers are added to the 2005 specification only. Similarly, the number of languages the worker speaks should help differentiating individuals according to skills which might not be accounted for completely by the diploma she received from formal schooling.

[14] Unfortunately, it is not possible to improve on such a broad proxy. In fact, despite ETF (2006) informs that "[…] according to the Social Security Institute, 59% of all declared [private firms'] employees receive an income equal to the national minimum wage of ALL 10,080 per month [in 2002]" (ETF, 2006:34), there is no question in the ALSMS asking the worker whether she is receiving the minimum wage. Likewise, there is only anecdotal evidence of a series of surges in the minimum wage between 2005 and 2007 (Gjokutaj, 2007). Furthermore, any detail on the relevance of the unions in the national labour market can be provided. Lastly, while it is well known that the Albanian informal economy is rather large (for instance, Muço *et al.* (2004) - applying different methodologies - estimate that it accounts for between 30 to more than 50% of the GDP) there is neither an established definition of informality in the literature nor a conclusive estimate of its relevance.

The full set of covariates employed in the estimation of the parsimonious[15] Mincer (1974) equations in 2002 and 2005 are summarised, using means, in the following tables.

TABLE 3 AND 4 ABOUT HERE

4. The empirical methodology

The empirical exercise in this paper evaluates the impact of the vector $(\boldsymbol{x}_i)$ of the characteristics in Table 3 and 4 on the log of the real hourly wages expressed in 2002 Leks (y_i), according to Mincer (1974), implementing the following model:

$$y_i = \boldsymbol{x}_i'\boldsymbol{\beta} + u_i \qquad (1)$$

where $\boldsymbol{\beta}$ is the corresponding vector of OLS estimates and $u_i \sim N(0,1)$ is a vector of residuals. The assumption of a well behaved u_i vector might be simplistic especially in the light of studies on wage discrimination which consider residuals as capturing the extent of discrimination (Kunze, 2008). Nonetheless, is deemed suitable to provide a preliminary static and cautious investigation of wage inequality in Albania.

The heteroscedastic nature of the vector u_i is tested using the standard Breusch-Pagan test and, in case, is corrected deploying the White (1980)/Huber (1967) sandwich estimator. Upon verifying the heteroscedastic nature of the disturbances, the discussion of the results will concern the corresponding robust estimates. Moreover, the extent of collinearity is evaluated implementing a correlation analysis of the explanatory variables and a variance inflation factor (VIF) investigation of the OLS's results.

Since the seminal work of Ashenfelter and Rouse (1998), which uses a sample of twins to correct the estimated rates of return from education for unobservable characteristics such as talent, the empirical research has dedicated a lot of effort to improve on their technique. The present study abstracts from evaluating this issue

[15] The parsimonious nature of the equations implied that for the year 2002 the variables for marital status, personal health condition, national migration, higher order polynomial in job tenure, informality in the labour market and urban settlement were excluded from estimation according to F (9, 1906) = 0.62, p-value = 0.78. Likewise, for the year 2005 the variables for marital status, religion, health condition, having nationally migrated, receiving a bonus, living in an urban settlement and the second degree term in job tenure were excluded in estimation according to F(10, 2219) = 0.78, p-value = 0.65.

mainly because it would limit the operative sample size even farther. Moreover, following Melly (2005), the results are interpreted conditional on the selected sample they originated from since the investigation and resolution of a possible sample selection bias is outside the scope of this work. Therefore, the components of the estimated $\boldsymbol{\beta}$ will be considered as statistically significant partial correlations. Similarly, the potential endogeneity of variables such as education, experience, industry choice and the pre-determinedness of others (Melly, 2005; Kunze, 2008) is not addressed here.

5. Results

Before reporting the estimates of the Mincer (1974) equations for the years 2002 and 2005, the level of the overall wage inequality in the two years is reported. The Gini coefficient is calculated at 0.34 for the trimmed distribution of the real hourly wage received by active people, who are not students in 2002. For the year 2005, the trimmed distribution of the real hourly wage converted in 2002 levels, displays a Gini coefficient of 0.28[16].

Since the specification of the Mincer (1974) equations for the years 2002 and 2005 are quite similar, the findings are commented comparing and contrasting the results for the two years which appear in the following tables.

TABLE 5 AND 6 ABOUT HERE

The Breusch-Pagan test for heteroscedasticity is significant and highlights the violation of the constant variance of the error term in (1). Hence the relevant column of results in the above tables is the one with robust standard errors. Moreover, collinearity seems neither widespread nor significant both according to the correlation analysis of the right hand side variables and to the VIF[17].

[16] The Gini for the year 2002 is calculated on 1,998 observations while the one for the year 2005 is calculated on 2,332 individuals. In both cases the Gini coefficients are statistically significant at the one percent level using a bootstrapped standard error arising from 100 replications.
[17] The results from this testing procedure are not presented here to conserve space but are available from the author upon request.

The log of the real hourly wage, expressed in 2002 Leks[18], displays in both years an inverted U shaped dependence with age consistent with the evolution of physical capabilities and skills during an individual's lifetime. Elementary algebra allows us to calculate the turning points at around 38 years of age (37.88) and, surprisingly, at around 69, for 2002 and 2005 respectively. The inverted U shaped trajectory is consistent with the one Lam and Levison (1992) estimate for the age profile of the variance in log earnings in Brazil. Both these and Lam and Levison (1992) results seem to conflict with the literature on the US which estimates U shaped age profiles of the earning inequality (Mincer, 1974; Schultz, 1975; Smith and Welch, 1979 cited in Lam and Levison, 1992).

Working males receive, *on average and ceteris paribus*[19], a higher real hourly wage, compared to females, by 18.13% in 2002 and by 21.73% in 2005[20]. This finding is consistent with a labour market and a culture biased, increasingly so over time, against women. According to World Bank (2006a) they suffered from a 5% higher probability of unemployment compared to males in 2004. Moreover, in some areas of the country, customary rules require them to move in their spouse's household limiting their employment opportunities even farther (Meurs *et al.*, 2008). Finally, Meurs *et al.* (2008) report that Silova and Magno (2004) have noted that the increase in early marriage, occurring in periods of slow economic growth, has diminished women's salary and returns from education.

The 2002 and 2005 results are in line with the findings of the original Mincer (1974) work in that the higher the education acquired, the higher the compensation the individual is entitled to for her work. A vocational diploma ensures, everything else being equal, a higher wage compared to having no or primary 8 education by 8.02% in 2002 while an individual with this qualification in 2005 has a wage which is indistinguishable from the one of someone with the base level of education. As expected, a university or higher degree is much more valuable in that it extends that premium to 29.27% and 33.87% for 2002 and 2005, respectively.

[18] Henceforth, the comments will omit the reference to the 2002 Old Lek unit of measurement.

[19] Despite the qualification *on average and ceteris paribus* is necessary every time the comment for a dummy variable is provided, this expression will be often omitted for expositional convenience.

[20] Due to the log linear specification, for dummy variables, the percentage change in the real hourly wage due to possessing that characteristic is calculated as $[e^{\beta_d} - 1] * 100$ where β_d is the estimated coefficient associated to the dummy included in estimation with respect to the omitted category.

The rate of return from obtaining a vocational and a university or higher diploma are derived supposing that they require the individual to stay in education for, on average, additional 4 and 11 years on top of the ones required for achieving a primary 8[21]. Studying to get the vocational diploma yields, according to the 2002 estimates, a return of 2.01% for every additional year of education while the university or higher is worth an annualized rate of return of 1.63%. The returns from university rise between 2002 and 2005 and reach 2.08%. The low rate of return to the vocational diploma seems at odds with the evidence in Azzarri *et al.* (2008) which suggests that more than ten years of education constitute the threshold above which acquiring farther human capital increases the probability of Albanians abandoning un(der)productive farming[22]. Nonetheless, according to the evidence in Meurs *et al.* (2008), the present result for vocational education is somewhat consistent, although on the lower bound, with the range of annualised returns (2.2 – 4.5%), from the vocational education, that Palomba and Vodopivec (2001) have previously obtained exploiting different estimation techniques and specifications[23]. The limited remuneration of the vocational diploma in the labour market might be connected to the government's decision of redirecting resources from the vocational technical institutes to more general-track ones (Berryman, 2000 in Meurs *et al.*, 2008). According to Berryman (2000), in Meurs *et al.* (2008), this "supply side" policy disrupted the educational options of the rural people and their ability to develop marketable skills. Lastly, it is possible that the government's failure in maintaining high quality curricula has affected, not only the access to secondary education (Meurs *et al.*, 2008), but also the returns to education.

The finding that the highest level of education provides a lower rate of return compared to a less qualifying one, for the year 2002[24], might reflect the national labour market being somewhat unprepared to absorb graduates who, on the other hand, have been widely appreciated abroad (Kaltani, 2007).

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[21] The calculation of these rates of return hinges upon these crucial hypotheses. Moreover, vocational training and secondary education are considered at the same level in the education hierarchy. Due to the nature of the hypotheses made, these results should be interpreted tentatively.

[22] To appreciate this comment note that the 10 and more years Azzarri *et al.* (2008) find to be a requirement to move out of farming roughly correspond to four more years of vocational diploma education on top of eight of primary.

[23] Yet, the returns from a vocational diploma in Albania are very low compared to international standards. In fact, Meurs *et al.* (2008) report that according to Psacharopoulos and Patrinos (2002) the global average rate of return associated to a vocational diploma is around 18%.

[24] The previously calculated rates of return might be deemed "theoretical" and upwardly biased also because they do not incorporate the probability of finding a job for a given education level obtained (Meurs *et al.*, 2008).

Every ISCO98 category included for the year 2002, except workers and individuals involved in elementary occupations, records a significant wage premium compared to the market oriented skilled workers. In particular, professionals enjoy a 60.83% premium while managers and clerks a 54.51%. Operators record the lowest premium at 42.25%. The picture is somewhat different in 2005 when high-skills occupations (e.g., professionals; managers and clerks) would enjoy higher wages compared to the market oriented skilled workers, if only the estimated coefficients were significant. Individuals performing elementary occupations are the only ones who receive a significantly lower wage, compared to the market oriented skilled workers, by 24.52% *on average and ceteris paribus*. Considering this body of evidence, the present results do not consistently sustain the Cragg and Epelbaum (1996) argument that the most progress-enhancing individuals (e.g., professionals and managers) drain a large part of the average wage increase associated with labour market reforms. In fact, while the 2002 results confirm that higher wages accrue to highly skilled individuals, this association disappears in the 2005 results.

Having migrated abroad is associated with a higher level of real hourly wage by 11.82%. in 2002 and by 9.52% in 2005. These results are supportive of the evidence presented above on return migrants faring better than non migrant individuals (King and Mai, 2008; Azzarri *et al.*, 2008; Kilic *et al.*, 2007; Carletto and Kilic, 2009). Furthermore, a reduction – over time - in the benefits from internationally sourced skills is apparent and might point to a declining “rate of return” from having migrated and returned.

It is rather surprising that working for any employer other than a private individual is associated, either significantly or insignificantly, with lower levels of real hourly wages both in 2002 and 2005. Employees in the public sector or firms are particularly disadvantaged since they record a drop in the real hourly wage received by 24.56% and 38.29% respectively in 2002. The public-private gap seems to have closed a great deal, for public firms’ employees in particular, since the shortfall diminished to 22.97% and 23.85% in 2005. This evidence is consistent with the account that Çuka *et al.* (2003) make of previous research that quantifies the wage premium existing between the private and public sector in Albania (e.g., Muço, 1996). The Albanian Institute of Statistics is reported to estimate that the average private sector’s wage in 1994 is twice the state’s one. Muço (1996), using a sample of large firms verifies that the gap stretches to three times as much[25]. This result might reflect genuine productivity differences between the private and the public sector with the latter being largely uncompetitive because it has been used to absorb excessive unemployment[26]. For instance, Bozgo *et al.* (2002) provide anecdotal evidence of the police being inefficient in preventing and repressing land disputes in the Mountain region of Albania. Moreover, individuals working outside the public sector might receive higher wages through foregoing, completely or in part, the contribution to the national pension and social assistance schemes they or their employers are supposed to make. In 2005 this argument is reinforced by the finding that receiving payments, other than the negotiated wage, impacts significantly on the final salary. In fact, other payments account for an additional 7.45% premium on the base real hourly wage. Likewise, the existence of such a large private-public premium could be a consequence of the significant reduction in the public departments' budgets mandated by the stabilisation programmes put in place to cure the macroeconomic instability which followed the transition from communism.

Due to the polarisation of the national economic activity in the Tirana area, it is not surprising that several dummies for regional localisation are significantly associated with a decline in the wage levels paid out compared to the capital. In particular, the largest shortfall is recorded in the mountainous area where wages are lower by 22.45% while in the other two regions the shortfall ranges between 13 and 14% in 2002. In 2005 only the real hourly wage in the coastal region is lower, compared to the Tirana area, by 4.7% *on average and ceteris paribus*.

The picture of wage variability due to industry affiliation provided here is somewhat limited since only a low percentage of the variables devoted to capturing this effect come up significant. In fact, 26 out of the forty dummies are significant in 2002 while only 14 out of the forty-one are statistically significant in 2005. All the significant dummies, but the one for being employed in the banking and financial sector, report that individuals employed in the associated industries receive lower wages, compared to those in agriculture, by an amount ranging between 25.57% (being employed in the public administration and defence sector) and 51.76% (mining of ores other than the

[25] Due to the extensive under-reporting of wages, Çuka *et al.* (2003) seem to favour the Muço (1996) estimation of the wage differential.
[26] This comment originates from a personal communication with Carlo Azzarri.

metal ones) in 2002. Being employed in the banking and financial sector grants a *ceteris paribus* higher wage, compared to being involved in agriculture, by almost 40%, instead. The negative wage differences, compared to being employed in agriculture in 2005, range from a minimum of 20.75% (individuals employed in recreational activities) to a maximum of 46.79% (individuals employed in the manufacture of intermediate goods for the textile industry). Only the refinement of petrol industry enjoys a premium in the hourly wages compared to working in agriculture, hunting and fishing by 40.28%.

The only consistent evidence arising from the present estimates does not confidently identify any sector paying wages above the ones in agriculture, both in 2002 and 2005.

Providing a suitable and detailed explanation of the estimated implicit wage premium for Albania in both years is not an easy task. In fact, while agriculture might provide high salaries, both in 2002 and 2005, because of the sector's historical relevance in the Albanian economy, the support for the existence of an international comparative advantage based on natural resources and cheap labour seems limited (Totev and Shahollari, 2001). On the one hand, World Bank (2002a) estimates that in 2002 agriculture contributed to 24.7% of the GDP, occupied 60% of the employed individuals, and accounted for 8.1% of the total export volume; INSTAT (2006) quantifies that at the end of 2006, the private firms in this sector were employing 58% of the individuals who had a job in that year[27] and CIA (2009) assesses that the agricultural contribution to GDP in 2008 was still 20.5%. On the other, Albania seems to have a comparative advantage only in somewhat minor productions: medical herbs, snails, olive products, honey and other bee products, cheese, some meat products (World Bank, 2002a), fruit and vegetables (Totev and Shahollari, 2001). While these productions have contributed to generate an Albanian export to import ratio in agriculture for the year 1999 accounting for half the national one (Totev and Shahollari, 2001), some of them seem to be rather occasional (i.e., medical herbs and olives being collected into the wilderness) (World Bank, 2002a).

The limited amount of marketed agricultural produce is another feature undermining the sector's ability to express an international comparative advantage. In fact, Azzarri

[27] According to these figures and to the evidence in World Bank (2008), Albania can be defined as an agriculture-based country.

et al. (2008) highlight that the majority of an otherwise rather ample variety of crops[28] and livestock is retained for household consumption. McCarthy *et al.* (2006) estimate that "[…] only between 4 and 8 percent of farm households […]" (McCarthy *et al.*, 2006:5) market the staple crops they produce while around 25% of the fruit and vegetables is sold. The most exchanged agricultural commodity is livestock which is put on the market by around 33% of the producers. The limited incentive to market farm produce is reinforced by a weak demand for raw material coming from the processing industry. In fact, the food and beverage manufacturing industry registers limited operating capacity, struggles with run down plants and cannot meet the high standards for food safety and quality required to have easier access to the European and global markets (World Bank, 2002a). For instance, the EU enforces sanitation rules on meat and dairy products which are incompatible with the family-level breeding of livestock practised in Albania (Totev and Shahollari, 2001).

Finally, the data uniquely available for 2005 suggest that every additional foreign language spoken grants a rise in the received hourly wage by 1.06 Leks. Similarly, there is a linear increasing trajectory linking the natural logarithm of the hourly wage and job tenure such that salaries rise by 1.08 Lek for every additional year spent in the same position. The level of the real hourly wage received in 2005 Albania seems to be positively, and increasingly, associated with the size of the operation and institution which employ the individual. In fact, individuals who are employed with additional ten to twenty-four co-workers earn, *on average and ceteris paribus*, 5.64% more than individuals who are part of micro-enterprises and institutions below ten workers. The wage premium in a fifty to ninety-nine co-workers environment rises to 7.99% while having 100 to 199 co-workers is associated with a wage premium of 13.21%, compared to the base category. Lastly, working with more than 200 people grants a 14.43% increase in the real hourly wage.

[28] McCarthy *et al.* (2006) note that farmers in the Central and Mountain regions are highly diversified. In the former the basket of produce includes wheat and potatoes; in the latter maize is the most popular crop. Finally, beans are typical of both areas while fruit and vegetables are more widespread throughout the country.

Conclusions

The empirical investigation developed here using two samples of data, extracted from the 2002 and 2005 ALSMS, suggests that the Gini coefficient on the real hourly wages in Albania is estimated at 0.34 for 2002 and 0.28 for 2005.

Despite the absence of panel data does not allow the estimation of a causal trajectory in the evolution of wage inequality between 2002 and 2005, deploying two augmented Mincer (1974) equations it is possible to compare and contrast the size of several gaps which might inform on the composition of the overall level of inequality in both years.

In particular, a persistent and increasing gender wage gap which favours men seems detectable. Similarly, higher returns to education accrue to individuals holding a university degree or higher. This inequality-increasing component has farther acquired relevance, since 2002, due to a likely stronger demand for graduates pervading the national labour market. According to these findings, the present work seems to comply with much of the literature on the persistence of lower wage opportunities for the women and on increasing returns from education.

The premium paid to return migrants for the higher quality of their internationally sourced on-the-job-training looks on the decline in more recent years. The increase in the public sector's productivity, ensuing from the significant reduction in the number of its employees, might be connected with the reduction in the wage shortfall associated to the nature of the employer.

In contrast with the abundant literature which describes the Albanian agriculture as being constrained in its efficiency, development and ultimately granting its employees and opportunity for survival rather than a viable way out of poverty (World Bank, 2002a; Azzarri *et al.*, 2008), the present work finds that almost no industry pays wages above the agricultural ones in 2002 and 2005. Attempting to disentangle the causes of these findings, the possibility that agriculture enjoys a within-country comparative advantage - with respect to industry and services - is quite an attractive proposition (Totev and Shahollari, 2001). The recent infrastructural development, the expansion of trade and the labour market liberalisation might have started improving the economic performance of sectors, other than agriculture, such that the negative wage gaps are smaller in 2005 compared to the ones in 2002. Yet, the availability of similar datasets in the future might provide the raw data instrumental in verifying the hypothesis that mostly positive wage gaps are associated with improved economic development.

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Table 1 The Albanian Roadmap to Trade Liberalisation

1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Price liberalisation of almost all commodities (almost all tradables). Excluded: electricity, water, telecommunications and public passengers transports, which were set close to cost recovery															
Significant expansion in private imports															
Cancellation of state monopoly on foreign trade															
		Convertibility of the national currency on most current account transactions, floating exchange rate regime													
Quite uniform tariff structure				Complex and more protectionist tariff structure			Simpler structure and lower rates tariff regime to meet the WTO accession criteria								
	Reduction and cancellation of the quantitative restriction on import and export to comply with the “tariffication” of the trade restriction measures														
										9/2000: WTO accession					
														Rescheduling of the tariff reduction for “sensitive” products	
															Involvement in 8 RFTAs[1]

Source: authors’ elaboration on Mancellari (2005)

[1] Albania’s Regional Free Trade Agreements (RFTAs) connect Albania with Bosnia and Herzegovina, Bulgaria, Croatia, Macedonia, Moldova, Romania, Serbia and Montenegro, and Kosovo/UNMIK (Mancellari, 2005).

Table 2 A picture of the current Albania vs. the "predicted" one in six reform fields

Policy Area	Proxy	Albania vs "predicted" Albania	Albania vs. SEE-8 and other transition countries (TCs)
Infrastructure	Per capita number of main telephone lines[1]	Adequate for its GDP per capita level	Far lower than the other SEE-8 countries
	Transmission and distribution losses of energy	Beyond what would be predicted by its per capita GDP	Only Moldova, among TCs, dissipates more
Financial Development	Private credit/GDP	Quite low by its per capita GDP	Lowest among SEE-8 countries
Educational Achievement	Secondary school enrolment	Broadly in line with the prediction of its per capita GDP	
	Quality of education (PISA results)	Consistent with the prediction of its per capita GDP	Worse than in other TCs
Governance	Composite index from Political Risk Services (2003).	Below what predicted by its per capita GDP	Lowest among the SEE-8 countries
Components:	prevalence of the rule of law	Below what predicted by its per capita GDP	Albania lagging behind compared to SEE-8 and other TCs
	democratic accountability of state actions		
	absence of corruption	Seems the most worrying problem	
	efficiency of bureaucracy		
Labour market	De jure index from World Bank (2005).	Satisfactory given its per capita GDP	Fares quite well compared to SEE-8
Components:	Difficulty of hiring workers	Satisfactory given its per capita GDP	Fares quite well compared to SEE-8
	Difficulty of dismissing workers	Satisfactory given its per capita GDP	Fares quite well compared to SEE-8
	Rigidity of working hours	Satisfactory given its per capita GDP	Fares quite well compared to SEE-8
Firm flexibility	Firm entry flexibility index from the Heritage Foundation (2003)	Satisfactory given its per capita GDP	Fares quite well compared to SEE-8
	Number of procedures for a start-up	Comparable to what would be predicted by its per capita GDP	Similar to the one for SEE-8 countries
	Costs associated with starting a business	Comparable to what would be predicted by its per capita GDP	Similar to the ones for SEE-8 countries
	Days necessary to complete a procedure	Comparable to what would be predicted by its per capita GDP	Similar to the ones for SEE-8 countries
	Overall regulation of the start-up process	Comparable to what would be predicted by its per capita GDP	Similar to the one for SEE-8 countries

Source: authors' compilation on Kaltani (2007).

[1] The appropriateness of this measure to proxy national infrastructure development with respect to transportation possibilities, energy, roads might be questioned when evaluating the impact of improved infrastructures on the development of trade. Kaltani (2007) anticipates this criticism and justifies this choice underscoring that this variable preserved sample size the most in the cross-country regression framework she deals with. It is believed that the result in Kaltani (2007) is highly sensitive to the variable employed.

Table 3 Summary Statistics for the 2002 Estimating Sample

Variable	Description	Mean
ltrhwag2002	continuous variable for the log of the real hourly wages distribution trimmed at the 1st and 99th percentile, expressed in 2002 New Leks	4.5335
trhwag2002	continuous variable for real hourly wages distribution trimmed at the 1st and 99th percentile, expressed in 2002 New Leks	111.6622
agey	Age of the respondent in years	39.5433
agesq	Squared variable for the age of the respondent in years	1663.2580
male	=1 if the respondent is male, =0 otherwise	0.6555
muslim	=1 if the respondent is Muslim, =0 otherwise	0.7734
orthodox	=1 if the respondent is Orthodox, =0 otherwise	0.1315
otherrel	=1 if the respondent is from other religions, =0 otherwise	0.0334
catholic	=1 if the respondent is Catholic, =0 otherwise	0.0617
noorpridip	=1 if the respondent has no or a primary diploma, =0 otherwise	0.3030
secdiploma	=1 if the respondent has a secondary diploma, =0 otherwise	0.2226
vocatdipl	=1 if the respondent has a vocational diploma, =0 otherwise	0.2352
univandmo	=1 if the respondent has an university or higher education title, =0 otherwise	0.2393
profess	=1 if the respondent is a professional (ISCO2 =21+22+23+24+31+32+33+34), =0 otherwise	0.3273
mancler	=1 if the respondent is employed as either a manager or a clerk (ISCO2 =11+12+13+41+42), =0 otherwise	0.0830
workers	=1 if the respondent is employed as a worker (ISCO2 =51+52+71+72+73+74), =0 otherwise	0.3515
operators	=1 if the respondent is employed as an operator (ISCO2 =81+82+83), =0 otherwise	0.1209
elemoccup	=1 if the respondent is employed in an elementary occupation (ISCO2 =91+92+93), =0 otherwise	0.1007
mktorskwor	=1 if the respondent is a market oriented skilled worker (ISCO2 =61), =0 otherwise	0.0167
migrabr	=1 if the respondent lived abroad for at least three months since 1997, =0 otherwise	0.0703
publsec	=1 if the respondent's employer is the public sector, =0 otherwise	0.4471
privfir	=1 if the respondent's employer is a private firm, =0 otherwise	0.2276
publfir	=1 if the respondent's employer is a public firm, =0 otherwise	0.0941
NGOs	=1 if the respondent's employer is a NGO, =0 otherwise	0.0167
privind	=1 if the respondent's employer is a private individual, =0 otherwise	0.2145
coastal	=1 if the respondent resides in the coastal region, =0 otherwise	0.2645
central	=1 if the respondent resides in the central region, =0 otherwise	0.2403
mountain	=1 if the respondent resides in the mountain region, =0 otherwise	0.2215
tirana	=1 if the respondent resides in the Tirana region, =0 otherwise	0.2736
agrhunfish	=1 if the respondent is employed in agriculture, hunting and fishing (NACE2 =01 + 05), =0 otherwise	0.0162
forestry	=1 if the respondent is employed in forestry (NACE2 =02), =0 otherwise	0.0051
mincrude	=1 if the respondent is employed in extraction of crude petroleum (NACE2 =10+11), =0 otherwise	0.0162
minion	=1 if the respondent is employed in mining of iron ore (NACE2 =13), =0 otherwise	0.0152
minotore	=1 if the respondent is employed in mining of other metal ore (NACE2 =14), =0 otherwise	0.0040
fdbevtobman	=1 if the respondent is employed in food, beverage and tobacco manufacture (NACE2 =15+16), =0 otherwise	0.0177

Table 3 Summary Statistics for the 2002 Estimating Sample *(continued)*

Variable	Description	Mean
textman	=1 if the respondent is employed in textile manufacture (NACE2 =17+18), =0 otherwise	0.0298
footwman	=1 if the respondent is employed in footwear manufacture (NACE2 =19), =0 otherwise	0.0147
woodman	=1 if the respondent is employed in wood manufacture (NACE2= 20), =0 otherwise	0.0076
papermanpub	=1 if the respondent is employed in paper manufacture and publications (NACE2 =21+22), =0 otherwise	0.0040
petrchemru	=1 if the respondent is employed in petrol refinement, chemical, rubber and plastic manufacture (NACE2 =23+24+25), =0 otherwise	0.0056
glasbuilmat	=1 if the respondent is employed in manufacture of glass, ceramics and other building materials (NACE2 =26+27), =0 otherwise	0.0106
othmetman	=1 if the respondent is employed in other metal manufacture (NACE2 =28), =0 otherwise	0.0071
machiman	=1 if the respondent is employed in metal, structures and engine manufactures (NACE2 =29), =0 otherwise	0.0040
manelecmed	=1 if the respondent is employed in manufacture of electricity and medical goods (NACE2 =31+32+33), =0 otherwise	0.0025
furnman	=1 if the respondent is employed in furniture manufacturing (NACE2 =36), =0 otherwise	0.0061
elecgaspro	=1 if the respondent is employed in electricity and gas supply (NACE2 =40), =0 otherwise	0.0324
watcolpurdis	=1 if the respondent is employed in water distribution and treatment (NACE2 =41), =0 otherwise	0.0172
construction	=1 if the respondent is employed in construction (NACE2 =45), =0 otherwise	0.1452
automsales	=1 if the respondent is employed in sale, assistance of motor vehicles (NACE2 =50), =0 otherwise	0.0218
wholeagent	=1 if the respondent is employed in wholesale trade (NACE2 =51), =0 otherwise	0.0116
retailsale	=1 if the respondent is employed in retail trade (NACE2 =52), =0 otherwise	0.0496
hotrest	=1 if the respondent is employed in hotels and restaurants (NACE2 =55), =0 otherwise	0.0395
landtrans	=1 if the respondent is employed in land transportation (NACE2 =60), =0 otherwise	0.0450
watairtrn	=1 if the respondent is employed in water transportation (NACE2 =61+62), =0 otherwise	0.0025
trnssupp	=1 if the respondent is employed in support services to the transport sector (NACE2 =63), =0 otherwise	0.0056
posttel	=1 if the respondent is employed in post and telecommunication (NACE2 =64), =0 otherwise	0.0157
bnkfin	=1 if the respondent is employed in bank and financial intermediation (NACE2 =65), =0 otherwise	0.0056
insurance	=1 if the respondent is employed in insurance sector (NACE2 =66), =0 otherwise	0.0020
reestfmkts	=1 if the respondent is employed in real estate and financial markets' support services (NACE2 =67+70), =0 otherwise	0.0025
itconsrandd	=1 if the respondent is employed in IT consultancy and research and development (NACE2 =72+73), =0 otherwise	0.0111
otrbussupac	=1 if the respondent is employed in other business activities (NACE2 =74), =0 otherwise	0.0091
pubadmndef	=1 if the respondent is employed in public administration (NACE2 =75), =0 otherwise	0.1340
education	=1 if the respondent is employed in education (NACE2 =80), =0 otherwise	0.1406
hlthsocwk	=1 if the respondent is employed in health and social workers (NACE2 =85), =0 otherwise	0.0840

Table 3 Summary Statistics for the 2002 Estimating Sample *(continued)*

Variable	Description	Mean
sewrefdisp	=1 if the respondent is employed in sewage and garbage disposal (NACE2 =90), =0 otherwise	0.0116
actvorgan	=1 if the respondent is employed in membership of organization (NACE2 =91), =0 otherwise	0.0142
recrperserv	=1 if the respondent is employed in recreational activities and personal private services (NACE2 =92+93), =0 otherwise	0.0248
privhousempl	=1 if the respondent is employed in private household employee (NACE2 =95), =0 otherwise	0.0025
extrterorg	=1 if the respondent is employed in extra territorial organizations (NACE2 =99), =0 otherwise	0.0056

Note to Table 3:

(a): means based on 1,977 observations.

(b): in *Italics* variables excluded in estimation.

Table 4 Summary Statistics for the 2005 Estimating Sample

Variable	Description	Mean
ltrhwag2002	continuous variable for the log of the real hourly wages distribution trimmed at the 1^{st} and 99^{th} percentile, expressed in 2002 New Leks	4.6159
trhwag2002	continuous variable for real hourly wages distribution trimmed at the 1^{st} and 99^{th} percentile, expressed in 2002 New Leks	114.4635
agey	Age of the respondent in years	40.1601
agesq	Squared variable for the age of the respondent in years	1725.7100
male	=1 if the respondent is male, =0 otherwise	0.6577
numlang	number of languages spoken by the respondent	0.5620
noorpridip	=1 if the respondent has no or a primary diploma, =0 otherwise	0.3062
secdiploma	=1 if the respondent has a secondary diploma, =0 otherwise	0.2166
vocatdipl	=1 if the respondent has a vocational diploma, =0 otherwise	0.2597
univandmo	=1 if the respondent has an university or higher education title, =0 otherwise	0.2175
profess	=1 if the respondent is a professional (ISCO2 =21+22+23+24+31+32+33+34), =0 otherwise	0.3071
mancler	=1 if the respondent is employed as either a manager or a clerk (ISCO2 =11+12+13+41+42), =0 otherwise	0.0770
workers	=1 if the respondent is employed as a worker (ISCO2 =51+52+71+72+73+74), =0 otherwise	0.3545
operators	=1 if the respondent is employed as an operator (ISCO2 =81+82+83), =0 otherwise	0.1122
elemoccup	=1 if the respondent is employed in an elementary occupation (ISCO2 =91+92+93), =0 otherwise	0.0974
mktorskwor	=1 if the respondent is a market oriented skilled worker (ISCO2 =61), =0 otherwise	0.0518
migrabr	=1 if the respondent lived abroad for at least one month since January, 1^{st} 2004, =0 otherwise	0.2092
publsec	=1 if the respondent's employer is the public sector, =0 otherwise	0.4263
privfir	=1 if the respondent's employer is a private firm, =0 otherwise	0.2980
publfir	=1 if the respondent's employer is a public firm, =0 otherwise	0.0370
NGOs	=1 if the respondent's employer is a NGO, =0 otherwise	0.0122
privind	=1 if the respondent's employer is a private individual, =0 otherwise	0.2266
jobten	number of years in the current position	8.1757
othpaym	=1 if the respondent received payments other than the wage and other bonuses, =0 otherwise	0.0583

Table 4 Summary Statistics for the 2005 Estimating Sample (*continued*)

Variable	Description	Mean
n1to9	=1 if the respondent has between 1 and 9 co-workers in the present position, =0 otherwise	0.3536
n10to24	=1 if the respondent has between 10 and 25 co-workers in the present position, =0 otherwise	0.2127
n25to49	=1 if the respondent has between 24 and 49 co-workers in the present position, =0 otherwise	0.1557
n50to99	=1 if the respondent has between 50 and 99 co-workers in the present position, =0 otherwise	0.1005
n100to199	=1 if the respondent has between 100 and 199 co-workers in the present position, =0 otherwise	0.0813
nmore200	=1 if the respondent has more than or 200 co-workers in the present position, =0 otherwise	0.0692
ndontknw	=1 if the respondent did not know the number of co-workers in the present positon, =0 otherwise	0.0270
coastal	=1 if the respondent resides in the coastal region, =0 otherwise	0.2527
central	=1 if the respondent resides in the central region, =0 otherwise	0.2501
mountain	=1 if the respondent resides in the mountain region, =0 otherwise	0.2244
tirana	=1 if the respondent resides in the Tirana region, =0 otherwise	0.2727
agrhun	=1 if the respondent is employed in agriculture and hunting, =0 otherwise	0.0526
forestry	=1 if the respondent is employed in forestry, =0 otherwise	0.0057
mincrude	=1 if the respondent is employed in extraction of crude petroleum, =0 otherwise	0.0130
minion	=1 if the respondent is employed in mining of iron ore, =0 otherwise	0.0083
minotore	=1 if the respondent is employed in mining of other metal ore, =0 otherwise	0.0026
fdbevtobman	=1 if the respondent is employed in food, beverage and tobacco manufacture, =0 otherwise	0.0187
textintpr	=1 if the respondent is employed in manufacture of intermediate products for the textile sector, =0 otherwise	0.0022
textappman	=1 if the respondent is employed in textile and apparel manufacture, =0 otherwise	0.0261
footwman	=1 if the respondent is employed in footwear manufacture, =0 otherwise	0.0148
woodman	=1 if the respondent is employed in wood manufacture, =0 otherwise	0.0052
papermanpub	=1 if the respondent is employed in paper manufacture and publications, =0 otherwise	0.0039
petrref	=1 if the respondent is employed in petrol refinement, =0 otherwise	0.0057
chemrubpla	=1 if the respondent is employed in chemical, rubber and plastic manufacture, =0 otherwise	0.0043
glasbuilmat	=1 if the respondent is employed in manufacture of glass, ceramics and other building materials, =0 otherwise	0.0126
othmetman	=1 if the respondent is employed in other metal manufacture, =0 otherwise	0.0043
metstrengman	=1 if the respondent is employed in metal structures and engine manufactures, =0 otherwise	0.0083
manwoodgds	=1 if the respondent is employed in manufacture of wooden goods, =0 otherwise	0.0109
elecgassup	=1 if the respondent is employed in electricity and gas supply, =0 otherwise	0.0209
watdistrea	=1 if the respondent is employed in water distribution and treatment, =0 otherwise	0.0148
construct	=1 if the respondent is employed in construction, =0 otherwise	0.1831

Table 4 Summary Statistics for the 2005 Estimating Sample (*continued*)

Variable	Description	Mean
motveisale	=1 if the respondent is employed in sale, assistance of motor vehicles, =0 otherwise	0.0174
agentwholes	=1 if the respondent is employed in wholesale trade, =0 otherwise	0.0096
retailsales	=1 if the respondent is employed in retail trade, =0 otherwise	0.0413
hotelrest	=1 if the respondent is employed in hotels and restaurants, =0 otherwise	0.0492
landtrans	=1 if the respondent is employed in land transportation, =0 otherwise	0.0304
watertran	=1 if the respondent is employed in water transportation, =0 otherwise	0.0022
airtransp	=1 if the respondent is employed in air transportation, =0 otherwise	0.0022
trsecsuac	=1 if the respondent is employed in support services to the transport sector, =0 otherwise	0.0057
posttel	=1 if the respondent is employed in post and telecommunication, =0 otherwise	0.0104
bankfinandre	=1 if the respondent is employed in bank, financial intermediation and real estate, =0 otherwise	0.0096
penlifins	=1 if the respondent is employed in life insurance and pension, =0 otherwise	0.0043
itconrandd	=1 if the respondent is employed in IT consultancy and research and development, =0 otherwise	0.0113
othbussupact	=1 if the respondent is employed in other business activities, =0 otherwise	0.0096
othpubadmdef	=1 if the respondent is employed in public administration, =0 otherwise	0.1144
education	=1 if the respondent is employed in education, =0 otherwise	0.1296
healsocwk	=1 if the respondent is employed in health and social workers, =0 otherwise	0.0809
sanitact	=1 if the respondent is employed in sewage and garbage disposal, =0 otherwise	0.0100
activorg	=1 if the respondent is employed in membership of organization, =0 otherwise	0.0070
recractiv	=1 if the respondent is employed in recreational activities, =0 otherwise	0.0270
persserv	=1 if the respondent is employed in personal private services, =0 otherwise	0.0030
privhempl	=1 if the respondent is employed in private household employee, =0 otherwise	0.0030
extrterorg	=1 if the respondent is employed in extra territorial organizations, =0 otherwise	0.0039

Note to Table 4:

(a): means based on 2,299 observations.

(b): in *Italics* variables excluded in estimation.

Table 5 Estimates for the Mincer Equation for the Year 2002

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
agey	0.0303*** (0.0076)	0.0303*** (0.0078)
agesq	-0.0004*** (0.0001)	-0.0004*** (0.0001)
male	0.1666*** (0.0288)	0.1666*** (0.0291)
muslim	0.0418 (0.0457)	0.0418 (0.0447)
orthodox	0.0208 (0.0529)	0.0208 (0.0509)
otherrel	0.1397* (0.0752)	0.1397* (0.0788)
secdiploma	0.0018 (0.0329)	0.0018 (0.0356)
vocatdipl	0.0771** (0.0340)	0.0771** (0.0355)
univandmo	0.2567*** (0.0425)	0.2567*** (0.0454)
profess	0.4752*** (0.1589)	0.4752*** (0.1674)
mancler	0.4351*** (0.1623)	0.4351*** (0.1692)
workers	0.2363 (0.1589)	0.2363 (0.1661)
operators	0.3524** (0.1589)	0.3524** (0.1663)
elemoccup	0.0596 (0.1596)	0.0596 (0.1669)
migrabr	0.1117*** (0.0426)	0.1117** (0.0485)
publsec	-0.2818*** (0.0466)	-0.2818*** (0.0555)
privfir	-0.0529 (0.0356)	-0.0529 (0.0408)
publfir	-0.4827*** (0.0546)	-0.4827*** (0.0590)
NGOs	-0.1636 (0.1257)	-0.1636 (0.1510)
coastal	-0.1490*** (0.0314)	-0.1490*** (0.0321)
central	-0.1441*** (0.0317)	-0.1441*** (0.0326)
mountain	-0.2542*** (0.0342)	-0.2542*** (0.0338)
forestry	-0.4902*** (0.1934)	-0.4902** (0.2074)
mincrude	-0.0417 (0.1812)	-0.0417 (0.1785)
minion	0.0441 (0.1868)	0.0441 (0.1729)
minotore	-0.7290*** (0.2307)	-0.7290*** (0.1927)

Table 5 Estimates for the Mincer Equation for the Year 2002 (*continued*)

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
fdbevtobman	-0.6112*** (0.1765)	-0.6112*** (0.1861)
textman	-0.5804*** (0.1723)	-0.5804*** (0.1743)
footwman	-0.5636*** (0.1825)	-0.5636*** (0.1918)
woodman	-0.5458*** (0.1993)	-0.5458*** (0.2123)
papermanpub	-0.6277*** (0.2317)	-0.6277*** (0.2384)
petrchemru	-0.4652** (0.2134)	-0.4652** (0.2002)
glasbuilmat	-0.2903 (0.1892)	-0.2903 (0.1871)
othmetman	-0.4807** (0.2028)	-0.4807*** (0.1947)
machiman	-0.5347** (0.2310)	-0.5347*** (0.2047)
manelecmed	-0.2671 (0.2655)	-0.2671 (0.2376)
furnman	-0.5099** (0.2103)	-0.5099*** (0.1920)
elecgaspro	-0.1020 (0.1719)	-0.1020 (0.1758)
watcolpurdis	-0.3177* (0.1790)	-0.3177* (0.1833)
construction	-0.2313 (0.1606)	-0.2313 (0.1648)
automsales	-0.4682*** (0.1742)	-0.4682*** (0.1796)
wholeagent	-0.4887*** (0.1862)	-0.4887*** (0.1951)
retailsale	-0.5273*** (0.1651)	-0.5273*** (0.1738)
hotrest	-0.6942*** (0.1671)	-0.6942*** (0.1773)
landtrans	-0.2172 (0.1649)	-0.2172 (0.1714)
watairtrn	-0.0642 (0.2655)	-0.0642 (0.1984)
trnssupp	-0.3692* (0.2132)	-0.3692 (0.2465)
posttel	-0.1927 (0.1811)	-0.1927 (0.1789)
bnkfin	0.3363 (0.2139)	0.3363* (0.2057)
insurance	-0.3118 (0.2851)	-0.3118* (0.1811)
reestfmkts	-0.3942 (0.2645)	-0.3942* (0.2247)
itconsrandd	-0.4208** (0.1846)	-0.4208** (0.1836)

Table 5 Estimates for the Mincer Equation for the Year 2002 (*continued*)

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
otrbussupac	-0.5324*** (0.1933)	-0.5324*** (0.1918)
pubadmndef	-0.2953* (0.1628)	-0.2953* (0.1698)
education	-0.2190 (0.1626)	-0.2190 (0.1692)
hlthsocwk	-0.5809*** (0.1641)	-0.5809*** (0.1692)
sewrefdisp	-0.5051*** (0.1826)	-0.5051*** (0.1824)
actvorgan	-0.0393 (0.2020)	-0.0393 (0.2161)
recrperserv	-0.4151** (0.1714)	-0.4151** (0.1819)
privhousempl	-0.5481** (0.2660)	-0.5481*** (0.2042)
extrterorg	0.2805 (0.2143)	0.2805 (0.2222)
constant	4.0307*** (0.1738)	4.0307*** (0.1891)
Observations	1977	1977
R^2	0.33	0.33
Adjusted R^2	0.31	§
Breusch-Pagan / Cook-Weisberg test for heteroskedasticity	8.76***	§

Notes to Table 5:

(a) *** denotes statistical significance at the 1% level, ** denotes statistical significance at the 5% level, * denotes statistical significance at the 10% level.

(b) § denotes not applicable in estimation.

(c) homoscedastic standard errors arising from the robust option in STATA v10.

Table 6 Estimates for the Mincer Equation for the Year 2005

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
agey	0.0138*** (0.0054)	0.0138** (0.0060)
agesq	-0.0001** (0.0001)	-0.0001** (0.0001)
male	0.1966*** (0.0230)	0.1966*** (0.0230)
numlang	0.0577*** (0.0121)	0.0577*** (0.0127)
secdiploma	-0.0077 (0.0259)	-0.0077 (0.0267)
vocatdipl	0.0306 (0.0260)	0.0306 (0.0259)
univandmo	0.2917*** (0.0356)	0.2917*** (0.0360)
profess	0.0589 (0.0993)	0.0589 (0.1287)
mancler	0.0301 (0.1016)	0.0301 (0.1313)
workers	-0.1261 (0.0969)	-0.1261 (0.1284)
operators	-0.0900 (0.0971)	-0.0900 (0.1263)
elemoccup	-0.2813*** (0.0984)	-0.2813** (0.1280)
migrabr	0.0909*** (0.0242)	0.0909*** (0.0270)
publsec	-0.2610*** (0.0405)	-0.2610*** (0.0474)
privfir	-0.0218 (0.0285)	-0.0218 (0.0333)
publfir	-0.2724*** (0.0597)	-0.2724*** (0.0583)
NGOs	0.1365 (0.1005)	0.1365 (0.0982)
jobten	0.0026** (0.0013)	0.0026** (0.0011)
othpaym	0.0719** (0.0373)	0.0719** (0.0337)
n10to24	0.0549** (0.0269)	0.0549* (0.0298)
n25to49	0.0101 (0.0305)	0.0101 (0.0303)
n50to99	0.0769** (0.0356)	0.0769** (0.0344)
n100to199	0.1241*** (0.0381)	0.1241*** (0.0348)
nmore200	0.1348*** (0.0430)	0.1348*** (0.0383)
ndontknw	0.0535 (0.0543)	0.0535 (0.0619)

Table 6 Estimates for the Mincer Equation for the Year 2005 (*continued*)

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
coastal	-0.0481** (0.0246)	-0.0481** (0.0248)
central	-0.0152 (0.0243)	-0.0152 (0.0257)
mountain	-0.0053 (0.0264)	-0.0053 (0.0259)
forestry	-0.2614** (0.1297)	-0.2614** (0.1313)
mincrude	0.1310 (0.1213)	0.1310 (0.1431)
minion	0.1535 (0.1356)	0.1535 (0.1466)
minotore	-0.1278 (0.1860)	-0.1278 (0.1938)
fdbevtobman	-0.1864* (0.1111)	-0.1864 (0.1429)
textintpr	-0.6309*** (0.1998)	-0.6309*** (0.1583)
textappman	-0.3816*** (0.1110)	-0.3816*** (0.1405)
footwman	-0.4107*** (0.1180)	-0.4107*** (0.1374)
woodman	-0.2499* (0.1475)	-0.2499 (0.1622)
papermanpub	-0.4037*** (0.1627)	-0.4037** (0.1892)
petrref	0.3385** (0.1526)	0.3385** (0.1410)
chemrubpla	-0.2259 (0.1558)	-0.2259 (0.2047)
glasbuilmat	-0.1740 (0.1194)	-0.1740 (0.1476)
othmetman	-0.2105 (0.1576)	-0.2105 (0.1827)
metstrengman	-0.0810 (0.1302)	-0.0810 (0.1648)
manwoodgds	-0.2867** (0.1233)	-0.2867* (0.1597)
elecgassup	0.0468 (0.1127)	0.0468 (0.1381)
watdistrea	-0.2443** (0.1178)	-0.2443* (0.1458)
construct	0.0000 (0.0965)	0.0000 (0.1317)
motveisale	-0.1889* (0.1129)	-0.1889 (0.1537)
agentwholes	-0.1986 (0.1261)	-0.1986 (0.1572)
retailsales	-0.2672*** (0.1024)	-0.2672** (0.1382)
hotelrest	-0.4327*** (0.1012)	-0.4327*** (0.1386)

Table 6 Estimates for the Mincer Equation for the Year 2005 (*continued*)

	OLS Estimates	
Variables	**β**	**β and Robust S.E.**
landtrans	-0.1786* (0.1047)	-0.1786 (0.1361)
watertran	0.0235 (0.2025)	0.0235 (0.1748)
airtransp	0.3374* (0.2007)	0.3374 (0.2468)
trsecsuac	-0.0426 (0.1441)	-0.0426 (0.2025)
posttel	-0.1007 (0.1249)	-0.1007 (0.1468)
bankfinandre	0.0840 (0.1270)	0.0840 (0.1490)
penlifins	-0.2815* (0.1565)	-0.2815 (0.2201)
itconrandd	-0.1314 (0.1212)	-0.1314 (0.1468)
othbussupact	-0.4035*** (0.1261)	-0.4035*** (0.1639)
othpubadmdef	-0.0039 (0.0998)	-0.0039 (0.1333)
education	-0.1226 (0.0991)	-0.1226 (0.1324)
healsocwk	-0.2608*** (0.1012)	-0.2608** (0.1340)
sanitact	-0.1041 (0.1223)	-0.1041 (0.1599)
activorg	-0.0999 (0.1571)	-0.0999 (0.1794)
recractiv	-0.2326** (0.1067)	-0.2326* (0.1426)
persserv	-0.5304*** (0.1763)	-0.5304*** (0.1918)
privhempl	0.1856 (0.1778)	0.1856 (0.3082)
extrterorg	0.0068 (0.1655)	0.0068 (0.3332)
_cons	4.3353*** (0.1109)	4.3353*** (0.1250)
Observations	2299	2299
R^2	0.37	0.37
Adjusted R^2	0.35	§
Breusch-Pagan / Cook-Weisberg test for heteroskedasticity	25.90***	§

Notes to Table 6:
(a) *** denotes statistical significance at the 1% level, ** denotes statistical significance at the 5% level, * denotes statistical significance at the 10% level.
(b) § denotes not applicable in estimation.
(c) homoscedastic standard errors arising from the robust option in STATA v10.