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AN EMPIRICAL ANALYSIS OF LIFE SATISFACTION

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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 of 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 dealt with. In particular, our study revealed evidence of long memories among Albanian respondents about the collapse of that country's notorious pyramid scheme and the scarring effects of this episode continue to impact on life satisfaction even after 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, mostly, 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, identified by satisfaction with life, taking into account as many of the aforementioned components as possible. Moreover, it focuses on a few technical aspects which appear to have been neglected in most analyses 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 indicate national welfare in periods 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 the testing of 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 job satisfaction 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 wage levels, unemployment rates and the psychological damage of joblessness are similar across Europe. Alesina *et al.* (2004) compare the ideological attitude of Europeans and Americans to inequality as captured by income mobility in society. According to them, ideology is reflected in individuals’ behaviour and political orientation. They find that inequality concerns Europeans, particularly those politically on the left. It is argued that the lack of influence on Americans is due to equality being associated with a lack of stimuli and rewards for consistent hard work. Hayo and Seifert (2003) incorporate in their model the perception of the communist regime and the individual’s religion and church attendance. It might be argued that the latter two factors represent 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 expectations of the capitalist system, religion heavily influenced 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. Further, 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 stood up to 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 that 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 cohort 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 call socioemotional selectivity theory. In particular, they argue that young 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 greater positive affects and fewer negative ones are actually preferred.

[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.

The shift from relationships with high educational content to ones with high emotional assonance induces age-related behaviours. Although 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 to specify 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 that 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 his/her present circumstances so 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 people 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 indicate 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 being cohort ones. While Clark (2007) notes that cohort fixed effects do vary by gender and level of education, he does not investigate their origins and 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 their own (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 qualification 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 the divorced, compared either to married or single people, 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 a 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 on 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, indicates 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 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 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 investigated the relationship between relative rather than absolute measures of objective WB and individual SWB. The economics literature's preference for the dependence of SWB on relative measures, and of monetary adequacy in particular, is largely determined by interpreting SWB as a proxy for utility[5]. Relative rather than absolute money metric measures of WB express the monetary value of income or consumption enjoyed by the individual in relation to a reference group. A significant body of literature has been developed to investigate 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. However, 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 so we would like to keep it simple. Further research effort can improve on this.

[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 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 co-workers rather than its absolute value. Moreover, it indicates 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. Likewise, Clark (2008) points to corroborative evidence arising from the empirical investigation of several different datasets (Clark, 2006) and of the long run German panel data (Clark *et al.*, 2008b).

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

Clark and Oswald (1994) and Winkelmann and Winkelmann (1998) estimate the impact of unemployment over and above the mere disappointment arising from the

[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.

forgone income that this status entails. Clark and Oswald (1994) employ the '90s WB data for Britain and demonstrate that having lost a job causes worse psychological damage than getting divorced. Nonetheless, Winkelmann and Winkelmann (1998) suggest that the drop in social rewards associated with 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 differences between the results of either approach (Ferrer-i-Carbonell and Frijters, 2004), but 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 alone in justifying the choice of 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 in 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, only Hayo and Seifert (2003) 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 bridge this gap by 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 from 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 stage 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), situated in areas characterised by big/small towns and rural environment. Tirana is subject to oversampling using non-randomly selected EAs because of the large proportion of the population living there but we have ignored this additional piece of information and rely only on the 455 EAs randomly sampled. The Second Stage Sampling Units (SSUs) were the households. In each EA, 12 households were selected to be involved in the survey. Eight constituted the main interview sample while the other four were used as replacements in case of non response. In total the ALSMS provides 16,387 usable individual observations distributed in 3,638 households.

The household questionnaire is based on recommendations from Grosh and Glewwe (2000) and includes new 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 of the questionnaire.

We restrict individuals' age range between 15 and 65 and exclude students from the sample to model individuals' labour market status more easily. In 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 implications for the subjective measurement of WB. Clark (2008) submits that the finding that religious people report higher SWB than atheists is well established in 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 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 Spanish development during the Inquisition. McCleary and Barro (2006) report that Kuran (2004) supports the idea that countries with a Muslim majority adopt legal structures which generate a lot of red tape affecting contract enforcement, credit and insurance provision and corporate ownership. All of the above might discourage economic activity and is likely to diminish SWB. According to the aforementioned evidence we anticipate the OTHERREL dummy (accounting for the Protestants in our specification) to display a positive coefficient while the MUSLIM one could come up with a negative sign.

[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).

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 Catholic 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 minor 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 typical monthly outlays. Moreover, its inclusion in a multidimensional measure of expenditure can alter the welfare ranking previously obtained using an aggregate which excludes health expenditure. Possession of a health licence in Albania grants discounts on the cost of medicines or allows them to be obtained for free. The 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, mostly in Tirana. The significance and positive contribution to WB of a measure of per capita household expenditure might lead to the policy recommendation of raising resources available to individuals. If this happened through an expanded provision of free or subsidised healthcare, according to the above, the Albanian poor might experience a worsening in their conditions.

[10] We were suggested to include a measure of per capita floor space of the dwelling. We could not implement this suggestion because the survey's questionnaire does not provide us with the number of square metres.

[11] The visits to the public health provider are instead free of charge (World Bank, 2003).

Similarly, considerable and unexpected health costs under a tight budget constraint might force household members to sell off assets or borrow money to finance the increased outflow of money. Since it is difficult to distinguish between such an occurrence and the costs of a long term chronic illness, this type of expenditure has not been considered in order 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. The World Bank (2003) supports this choice indicating that different assumptions for economies of scales in consumption do not alter 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 literature concerned 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].

Recent Albanian history has been dominated by several negative events. Internal political conflict, the collapse of pyramid schemes, rampant corruption and the problems in Kosovo might be singled out as the most important (Jarvis, 1999). Among them, the schemes attracted a lot of research interest because of their nature, causes and consequences.

The Albanian pyramid schemes were fostered by depositors putting money into lending companies which, at the peak of their popularity and because 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 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 trafficking) but not from the schemes' real investments. They happened to be of negligible value and return. Besides the attractive rates of return, additional causes for the schemes' unprecedented popularity include:

[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 interested reader to it for a more complete picture.

- 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 an inability to proficiently mobilise 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 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%, GDP to fall by 7%, the inflation rate to skyrocket to 42%, the budget deficit to soar to 12% of GDP and unemployment to rise by 3% (ETF, 2006). On the social front, several riots occurred especially in the richer south of the country, the police and the army spiralled out of control, arms were stolen and around 2,000 people were killed in the unrest (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 wide social support provides better physical health while isolationism is detrimental to 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 common 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,.....,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]:

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

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

In general terms we can write $Prob\,[y_i = j] = \Phi\big(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}\big) - \Phi\big(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}\big)$ 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 $j = 0$, $\Phi\big(\theta_{j-1}\big) = \Phi(-\infty) = 0$ and for $j = 3, \Phi\big(\theta_j\big) = \Phi(+\infty) = 1$. 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:

$$Prob\,[y_i = j] = \Phi\big(\theta_j - \boldsymbol{x}_i'\boldsymbol{\beta}\big) - \Phi\big(\theta_{j-1} - \boldsymbol{x}_i'\boldsymbol{\beta}\big) \quad for \quad j = 0, \dots .J \tag{2}$$

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

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

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:

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

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 threshold 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 $\boldsymbol{x}_i$ except the constant. Once the results from the estimation of the likelihood function in (17) are 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) state 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 becomes 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 underestimation 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 that 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 in 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 his/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 average public sector earnings) of unemployment benefits (ETF, 2006) barely help to maintain sufficient confidence in the possibility of funding satisfactory levels of consumption. Interpreting this beta according to Winkelman and Winkelman (1998), the psychological damage associated with joblessness outweighs by far any other welfare reducing condition such as suffering because of the collapse of the pyramid schemes. We can refer to 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 that of Romania. Nonetheless, the Albanian result for unemployment is still about half the magnitude of that 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 those 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 that of having a university or higher degree. This result might highlight the inadequacy of the college and vocational education in providing 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 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 that of the US (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 countries like the US where a larger share of the workforce might have a university degree.

Muslims have a 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 been proved to be significantly associated with psychological and material strains 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 divorce. This, in turn, has been related to the impossibility for a woman to take the children with her when returning to her parents'

[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.

house and 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 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 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.

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[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 to 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 fewer 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 they are needed 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 those of women from higher social classes. A very controversial and peculiar feature of children as a 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 custom, the youngest child is expected to take care

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[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.

of the elderly parents (King and Mai, 2008). Our estimates seem to suggest that SWB data do not pick up on this relevant aspect correctly. It might be due to the need to require migrant children to give up consumption in the host country to send money home and to the parents' consciousness that their children will experience a regression in their standard of living when returning to Albania to care for them.

The negative beta might originate from the additional strain, from an 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 greater 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 of job insecurity, and becoming unemployed in particular, being significantly and negatively associated to SWB (Ravallion and Lokshin, 2000) while

– due to the overall size of the schemes and their popularity – might have exceeded even the benefits associated with the subsequent growth which, in 2004, had raised GDP to 36% 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 to paint 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 practice 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 assets and wealth, from the losers to the winners 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 which 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 size of the dwelling exerts a monotonically increasing effect on the probability of being fully satisfied with life. Living alone diminishes SWB but, since the

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windfall gains do not increase SWB (for lottery winners see Brickman *et al.*, 1978 in Clark *et al.*, 2008b).

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 with the increase in community population as a sign of the difficulties of local infrastructures and public services to cope with a rising number of settlers. Most of the population growth has happened in urban areas so 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 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 dispersed 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 with the logarithm of per capita real consumption reflects a fairly standard type of heteroscedastic relationship. It is possible to attribute this result to inequality in consumption. Sanfey and Teksoz (2005) and Senik (2004) explicitly controlled for 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 of 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 among the wealthy there are those who, because of the difference in floor space, might feel "super-rich" or "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 practice 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 that recorded in the Tirana area. This denotes fairly homogeneous economic and psychological conditions in these regions. This might be the result of the national and international migration of young individuals who have left behind adults and those unable to move (World Bank, 2006:50). The more homogeneous and “lonely” the local population, the more likely it is 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 emigration of young people. 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 situation of Albania allowed a number of additional matters to be treated. 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 enjoyed 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 in a

[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.

largely Muslim country. As anticipated, individuals who reside in communes with high crime levels enjoyed lower levels of satisfaction with the estimated effect 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 among individuals. Given that 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 allow controlling 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

[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.

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 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	β_{mean}	β_{mean}	$\beta_{variance}$
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.