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Persistance du bien-être subjectif : bonheur permanent, malheurs passagers ?

Résumé
Ce travail s'attache à séparer les rôles respectifs de la dépendance d'état et de l'hétérogénéité inobservée dans l'évaluation du bien-être subjectif réalisée par les individus eux-mêmes. Il estime un modèle dynamique ordonné sur des données de panel, principalement à partir de sources françaises, mais en mobilisant également des sources allemandes, anglaises et australiennes. La satisfaction dans la vie renseignée par les individus est persistante au cours du temps, persistance omise, par définition, par des modèles statiques. Cette persistance est surtout hétérogène puisqu'elle concerne principalement le haut de la distribution de bien-être subjectif ; les insatisfactions déclarées semblent, au contraire, plus transitoires. Enfin, les conditions initiales jouent un rôle prépondérant en comparaison avec les déterminants usuels du bien-être subjectif, dépendance d'état incluse.

Mots-clés : Bonheur ; bien-être subjectif ; satisfaction dans la vie ; modèle dynamique ; dépendance d'état ; effets aléatoires corrélés ; conditions initiales.

The persistence of subjective well-being: permanent happiness, transitory misery?

Abstract:
This paper disentangles the roles played by state dependence and unobserved heterogeneity in self-assessed happiness. It estimates a dynamic nonlinear model of subjective well-being on longitudinal data, primarily from France, but also from Australia, Germany, and the UK. Life satisfaction is persistent over time, which static models ignore. This persistence is heterogeneous across individuals: it concerns mostly those already happy with their lives while, in contrast, unhappiness seems more transitory. The impact of initial conditions is large in comparison with usual determinants of happiness, or with state dependence.

Keywords: Happiness; subjective well-being; life satisfaction; dynamic model; state dependence; correlated random effects; initial conditions.

Classification JEL : I31.

1 Introduction

Longitudinal surveys on subjective well-being have enabled researchers to isolate the role played by different time-varying factors (e.g., income, unemployment, marriage, widowhood) on self-assessed happiness. On top of observed characteristics, unobserved heterogeneity accounts for a large part of the dispersion in individual subjective well-being: omitted variables, optimism or pessimism bias, scale effects or pure heterogeneity in preferences can nevertheless be controlled for thanks to panel data combined with appropriate econometrics, e.g., fixed effects. Surprisingly, the dynamics of self-assessed happiness has not been the object of much attention: though subjective well-being likely exhibits a strong inertia, the quantification of such an autocorrelation is rare in the literature. Yet missing the dynamic part of the picture, including initial conditions, is misleading.

In this paper, I document the persistence of subjective well-being as well as its heterogeneity. In particular, I disentangle the respective roles played by state dependence and by initial conditions. Subjective well-being is often proxied by life satisfaction and self-assessed by individuals on a Cantril scale, i.e., on a discrete scale that ranges from 0 to 10. I estimate therefore a dynamic, ordered Logit model with correlated random effects; this nonlinear model allows me to separate state dependence from unobserved heterogeneity, hence to emphasize the role of persistence in reported life satisfaction. It also enables me to deal with self-assessed life satisfaction as a polytomous variable, to model the initial condition, and thus to approximate unobserved heterogeneity as parsimoniously as possible. The current analysis is primarily based on a French panel dataset, SRCV, from 2013 to 2017, but I show that the main results also hold in Australia, in Germany and in the UK, these countries disposing of longitudinal surveys that can be easily accessed to by researchers and which include some measure of subjective well-being. The main findings can be summarized as follows: (i) state dependence is significant at usual levels; (ii) its magnitude is strong when compared to usual determinants of happiness; (iii) it is asymmetric: happier people tend to remain happy more than less happy people; put differently, happiness looks rather persistent while unhappiness is rather transitory, and (iv) even more important in the dynamics of self-assessed happiness is the role played by initial conditions, which relates to fixed unobserved heterogeneity. Usual determinants of subjective well-being remain significant once state dependence and unobserved heterogeneity have been controlled for: having a partner improves life satisfaction while unemployment has a depressing effect -and so do deprivation indicators measuring quality of life, or negative capabilities (Sen, 1979). Current (transitory) income tends to matter less than average (permanent) income. Education, gender and occupation turn out to be mostly non-significant at usual levels. However, the empirical evidence at stake suggests that misspecification of static models is substantial, given the magnitude of the autocorrelation at stake: for instance, top past happiness would cushion the impact of unemployment, poor health and weak social ties altogether. The corresponding impact of initial conditions is even higher. More generally, controlling for persistence and unobserved heterogeneity achieves both a better fit and a more accurate description of the whole story of individual happiness.

These results are robust to parametric assumptions, to endogenous attrition concerns as well as to balancing issues. Moreover, in order to be sure that these findings are specific neither to my dataset, nor to France, I resort to three other panels: (i) the Household, Income and Labour Dynamics in Australia (HILDA) survey, (ii) the UK Understanding Society (UKUS) survey that took over the British Household Panel Survey (BHPS), with 8 waves from 2009 to 2018, and (iii) the German SOcio Economic Panel (GSOEP) available from 1984 to 2017. Both descriptive transition matrices of subjective well-being at the individual level and estimations from the same econometric model as before concur to similar findings. Once again, happiness is significantly persistent, especially at the top of the distribution. These results are obtained after neutralizing the role played by the institutional setting, including survey design; they strengthen previous empirical evidence and give some credit to the idea of unveiling some empirical regularity of individual behavior in the data.

Where does state dependence in self-assessed life satisfaction come from? Autocorrelation might arise from people not changing their mind every year about their subjective well-being. A possible explanation could lie in individuals evaluating once-and-for-all their average, permanent satisfaction with life, from which they would rarely deviate across different waves of longitudinal surveys, but depart from it when they experience good or bad shocks. Such a behavior would resemble to anchoring effects according to which agents would stick to initial or past self-evaluation. Interestingly, the revealed heterogeneity in happiness persistence, namely the asymmetry between upward and downward mobility in reported subjective well-being, suggests that unhappiness is more transitory. This might indicate that it is more difficult to revise downwards one's self-evaluation of life satisfaction, due to psychological costs for instance.

The rest of the paper is organized as follows. The next section is devoted to a literature review. Section 3 describes the French SRCV data and the econometric model is presented in Section 4. Results are discussed in Section 5 and Section 6 investigates some robustness checks, including the estimation on data from other countries. Section 7 concludes.

2 Literature

Key determinants of subjective well-being, often proxied by life satisfaction, have been widely documented by the literature: see, for instance, the excellent survey by Layard et al. (2015). The main types of determinants of life satisfaction are: (i) individual determinants (income: Easterlin (1974), Clark et al. (1996), Clark et al. (2005), Senik (2005), Clark and Senik (2010), age: Clark et al. (1996), labor force status[1] family status, objective indicators of the quality of life: Godefroy and Lollivier (2014); education, gender, occupation are seldom significant, but can be invoked depending on the country); (ii) macroeconomic determinants (GDP growth, unemployment rate, inflation, inequality, environmental issues, government and unions, not pretending to be exhaustive; their identification requires variation across countries, hence they are often absent in a one-country econometric analysis): see, e.g., Tella et al. (2003) on that topic; (iii) spatial determinants (rural *versus* urban areas: Easterlin et al. (2011), regional effects: Oswald and Wu (2010), the price of gasoline: Boyd-Swan and Herbst (2012), among others). However, the dynamics of subjective well-being, often proxied by life satisfaction, has been often ignored.

From a methodological perspective, subjective well-being is specific in the sense that individuals are asked to report their life satisfaction on an ordered, discrete scale. To deal with ordinality, researchers have estimated ordered polytomous models that rely on a latent, unobserved but cardinal propensity to happiness. In

[1]Being unemployed causes significant, persistent losses (Clark and Oswald, 1994). Interestingly, Clark et al. (2008) show that the recovery after such unfortunate career shocks is much slower than it is after personal events like widowhood.

practice though, estimating linear models does not affect the sign of the covariates,[2] and yields qualitatively similar results.

Moreover, many studies on subjective well-being relied first on cross-sectional data, which limits the ability of the researcher to control for unobserved heterogeneity. As soon as panel data have been available on that topic, econometric specifications have included individual effects, which do a better job at controlling for unobserved heterogeneity, and hence limit omitted variable biases (see Ferrer-i Carbonell and Frijters (2004) on that topic). For instance, the problem of optimism (or pessimism) arises as soon as one seeks to explain the level of subjective well-being from a cross-sectional analysis: two individuals may well report very different answers as regards their life satisfaction, though they look close in the sense that their observed characteristics are similar. To take unobserved heterogeneity into account, but also to avoid the problem of incidental parameters, the literature has estimated conditional ordered Logit models with fixed effects, see, e.g., Frijters et al. (2004). Though such models are powerful tools to capture unobserved heterogeneity, their identification relies on a subset of movers; empirically, this restriction induces likely a dramatic selection.

According to Hsiao (2003), state dependence and individual heterogeneity offer "diametrically opposite" explanations of persistence in life satisfaction outcome. Few dynamic models have been considered so far: the role of state dependence has rarely been explored *per se*, though recommended by Clark (2018). Exceptions include Frijters et al. (2011) who considered the dynamics of covariates and assessed how current life satisfaction depends on past events like getting married or divorced, becoming unemployed, widowhood, etc. Bottan and Truglia (2011) tries to disentangle a "specific habituation" channel, namely the causal effect of lagged covariates, from a "general habituation" channel, which has trait to the persistent nature of subjective well-being *per se*. According to Wunder (2012) who exhibits an "adaptation" channel, individuals update their expectations as a response to changing circumstances. However, none of these papers address the ordinal nature of the data, which I propose to do here by estimating a nonlinear, ordered polytomous model that fits better the data. On top of that, I investigate the heterogeneity of persistence across reported levels of life satisfaction: to the best of my knowledge, this has not been the object of any attention up to now.

[2]That sign is identified non-parametrically.

3 Data

Many papers in happiness economics have used data from Germany, from the UK or from Australia since all these countries dispose of longitudinal surveys (resp. GSOEP, BHPS and HILDA) that enable researchers to follow individuals over time and to learn about changes in their subjective well-being. Following the recommendation of the Stiglitz et al. (2009) commission, France has also started to ask individuals directly how they felt about their lives. The Insee produces the SRCV survey (*enquête Statistique sur les Ressources et Conditions de Vie*) targeting about 10,000 households every year. From 2010 onwards, it has included several questions related to individual life satisfaction, job satisfaction, and satisfaction with family and friends. On top of these measures of subjective well-being, it offers usual information at the individual level: gender, age, education, occupation, family status and labor force status. Income is measured at the household level; in what follows, I consider the logarithm of the CPI-deflated annual household income, i.e., the sum of real incomes from all members in the household divided by the number of units of consumption as defined by the OECD scale.[3]

The unit in charge of SRCV at Insee indicates that, though the survey has started in 2010, its reliability casts doubts before 2013. The questionnaire has been modified in 2013: questions relative to life satisfaction have been placed after those relative to income. *De facto*, a break in the time series of life satisfaction can be observed from that date. Hence I assume that the first reliable wave, common to all individuals, is 2013.

Table 1 shows some descriptive statistics relative to the working sample, an unbalanced panel of 19,253 individuals followed from 2013 to 2017 with at least two consecutive observations, which will matter for a proper identification of the role played by state dependence and whose annual income exceeds €1. Subjective well-being is measured on a Cantril scale ranging from 0 to 10. It has an average score of 7.17, it is rather concentrated around levels 7 and 8 (see Figure 1), it nevertheless uses the whole support of the distribution and it has a cross-sectional coefficient of variation as small as .23. Women and elders are slightly over-represented (62% of the sample aged 55 on average), which is usual in French household surveys.

[3]According to this scale, the first adult in the household has weight 1, the other adults or children aged at least 14 have a weight equal to .5, and children aged less than 14 have a weight equal to .3.

The average income amounts to nearly €26,000 per year. Income exhibits sizeable dispersion: its coefficient of variation is roughly 1.4 and the top 1% earns more than €86,000 a year. As regards education, 32% of individuals in the sample have a vocational degree, 29% graduated from high-school, 16% from college while 23% don't have any degree. As far as labor force status is concerned, about 47% of the sample is employed while 39% is retired. The remaining part of the sample is either unemployed or inactive. One half of the sample is made up of current or former clerks (28%) or individuals with an intermediate occupation (22%); the others are (or were) mainly blue collars (17%) or white collars (13%). Regarding family status, singles account for one quarter of the sample, while nearly 40% of individuals are living with a partner; then come parents of two children (11%), one child (9%), three children at least (5%) and single parents (4%). Finally, indicators measuring the objective quality of life are available in the survey: the exposition to psycho-social hazard, health problems, environment troubles, poor living conditions, (economic or general) insecurity as well as the weakness of social ties. Between 2% and 14% of individuals are exposed to at least one of such problems.

Turning to the persistence of subjective well-being, the main insight from Figure 1 which depicts the evolution of *aggregate* subjective well-being from 2013 to 2017 in France is that the distribution of answers is rather stable from year to year.

Focusing now at the *individual* level, about 1/5 of surveyed individuals report systematically the same level of life satisfaction over the five waves (Table 2). For slightly more than 1/3 of them, the difference between the highest and the lowest self-assessed level of life satisfaction is equal to one, while for about 1/4 of them, this difference is equal to two. That difference exceeds three for a last 1/4 of individuals, which suggests that subjective well-being is persistent over time at the individual level. It is therefore consistent with the hypothesis that individuals have some anchor in mind, from which they depart in case of favorable or less favorable shocks.

Table 3 provides with the transition matrix of individual levels of life satisfaction, and confirms that persistence of subjective well-being over time is strong. Its diagonal is heavy, which means that the probability of reporting the same level of life satisfaction as the year before is high. The most plausible past level of life satisfaction, given any current, self-assessed level of life satisfaction, is that very same level, few exceptions aside (namely, levels 2, 4 and 9). Yet most models of subjective well-being are static and ignore state dependence, i.e., they do not include any lagged variable as an explanatory variable. Hence they assume that each destination state has the same probability, regardless of the initial state. The main lesson from this descriptive analysis is that this assumption is rejected on the data: this transition matrix suggests that happiness is almost an absorbing state.

Interestingly, this persistence looks asymmetric: happiest individuals tend to stay happy, while unhappiness tends to be more transitory. As one gets higher in the distribution of reported levels of life satisfaction, the annual probabilities of reporting the same level of subjective well-being as before increase: the coefficients on the diagonal may be as high as 52.2% (level 8), and do not fall below 35.4% (level 9). By contrast, they are comprised between 11.4% (level 2) and 24.5% (level 0) at the bottom of that distribution. Even though upward mobility is mechanically more frequent at the bottom, it is striking to see that individuals reporting a level less than 4 have at least nearly 3/4 chance to see their life satisfaction increase next year. Intermediate levels 5 and 6 exhibit a slightly upward-biased trend, too: their annual persistence ranges from 26% to 31% with a probability of having a higher subjective well-being higher than 1/2. To sum up, from levels 0 to 6, upward mobility is more likely while inertia is more plausible from level 7 onwards. Put differently, defining happiness as reporting a level 8 or higher, the transition matrix of the latter would be symmetric with 3/4 on its diagonal and equally distributed states among the population; doing the same with unhappiness defined as reporting a level 5 or less would yield a 40% chance of departing from unhappiness for the 6% concerned individuals.

To confirm previous eyeball impressions and to check that they are not the mere consequence of both observed and unobserved heterogeneity, an econometric model that disentangles carefully state dependence from heterogeneity is however needed.

4 Model

To take both state dependence and unobserved heterogeneity into account, I consider an econometric specification that relies on a dynamic, ordered Logit with correlated random effects and unknown thresholds. Let y_{it} be the dependent variable, i.e., subjective well-being, ranging from $j = 0$ to $j = J \equiv 10$.[4] To deal with ordinal preferences, the ordered polytomous model assumes the existence of an explicit relationship between the observed variable y_{it} and some unobserved, latent variable y_{it}^* such that $\forall j \in [\![0, J]\!]$,

$$y_{it} = j \iff y_{it}^* \in [s_j, s_{j+1}[$$

or equivalently,

$$y_{it} = \sum_{j=0}^{J} j \; \mathbb{1}\{s_j \leq y_{it}^* < s_{j+1}\}. \tag{1}$$

The main advantage of this approach consists in retrieving both linearity and cardinality for the latent variable y_{it}^*. $\{s_j\}_{j=1}^J$ are the unknown thresholds with $s_0 = -\infty$ and $s_{J+1} = +\infty$.

I consider a dynamic model on the latent variable of the form:

$$y_{it}^* = \sum_j \rho_j \mathbb{1}[y_{i,t-1} = j] + x_{it}\beta + \alpha_i + \varepsilon_{it},$$

where idiosyncratic shocks ε_{it} follow the logistic distribution with mean 0 and variance $\frac{\pi^2}{3}$. As in Wooldridge (2005), state dependence is allowed to be nonlinear, too – namely specific to every value of past subjective well-being: the ρ_j coefficient is related to lagged j-value of life satisfaction.

At this stage, a first option could be to posit individual fixed effects, i.e., in making no parametric assumption on the distribution of α_i. This solution requires however to overcome the incidental parameter problem (Neyman et al., 1948; Lancaster, 2000). When the model is linear, differencing enables the econometrician to get rid of individual fixed effects. By contrast, in nonlinear models, the maximum likelihood estimator (MLE) is generally not consistent and asymptotically normal (CAN) due to the presence of numerous incidental parameters. In the Logit case, a well-known trick consists in conditioning the likelihood of an observed sequence $(y_{i1}, \ldots, y_{iT})$ by a sufficient statistics in order to make the fixed effects disappear of the likelihood. This so-called conditional likelihood estima-

[4] It is possible to consider a model with a smaller number of groups, say 7, to mimic the UK case after grouping 0-2, 3-4 and 5-6 levels for instance. The corresponding estimates are available upon request: the results are robust to such an aggregation choice.

tion (CLE) has been used by Rasch (1960); Andersen (1973); Chamberlain (1980); Honoré and Kyriazidou (2000); Magnac (2000); Frijters et al. (2004). In the case of a dynamic Logit model with fixed effects, a sufficient statistics corresponds to the number of occurrences of each state in the observed sequence of outcomes, initial and terminal conditions aside; in the binary case, Bartolucci and Nigro (2010, 2012) refer to total scores. This method is the analog, in spirit, to first-differencing in linear models. However, its cost is rather high since it requires to compute the denominator of the conditional likelihood which is composed of numerous terms. Moreover, the identification of the model relies on a subset of individuals only, the "movers", i.e., individuals whose sequence is not constant over time; these supplementary exclusion restrictions may be problematic in practice since they often constrain the estimation to rely on small sub-samples.

Another option consists in assuming some parametric form for the individual effects α_i, typically a normal distribution: this solution corresponds namely to the random effect approach. To enrich the latter, I consider rather the correlated random effects (CRE) solution *à la* Chamberlain (1982). Its main advantages are (i) to approximate fixed effects as much as possible by allowing for an explicit relationship between the individual effect and the covariates; (ii) to solve the initial condition problem that arises in dynamic models. Once again, two options are possible. As put by Arulampalam and Stewart (2009), "the Heckman (1981) estimator approximates the joint probability of the full observed y sequence $(y_0, y_1, \dots, y_T)$. Wooldridge (2005) on the other hand, has proposed an alternative conditional maximum likelihood estimator that considers the distribution of $(y_1, y_2, \dots, y_T)$ conditional on the initial period value y_0 (and exogenous variables)". I follow the latter approach and assume that

$$\alpha_i | y_{i0}, \boldsymbol{x_i} \sim \mathcal{N}\left(\sum_{j=0}^{J} \rho_j^0 \mathbb{1}[y_{i0} = j] + \boldsymbol{x'_{i0}}\boldsymbol{\gamma^0} + \overline{\boldsymbol{x'_i}}\gamma, \sigma_u^2\right), \tag{2}$$

where $\boldsymbol{x_i} = (x_{i0}, \dots, x_{i,T_i-1})$ if the researcher disposes of T_i observations for individual i. This parametric restriction enables me (i) to get rid of the incidental parameter problem; (ii) to model the initial condition; (iii) to avoid programming the maximization of the conditional likelihood. Lastly, I follow Rabe-Hesketh and Skrondal (2013) who propose a more parsimonious specification of the individual effect. They show that including initial $\boldsymbol{x_{i0}}$ and mean values of covariates $\overline{\boldsymbol{x_i}}$

is sufficient,[5] as opposed to including the whole set of covariates at all dates (Wooldridge, 2005). This approximation of unobserved heterogeneity is reminiscent of Mundlak (1978).

In the end, the estimating equation is:

$$y_{it}^* = \sum_{j=0}^{J} \left(\rho_j \mathbb{1}[y_{i,t-1} = j] + \rho_j^0 \mathbb{1}[y_{i0} = j]\right) + \boldsymbol{x'_{it}\beta} + \boldsymbol{x'_{i0}}\gamma^{\boldsymbol{0}} + \overline{\boldsymbol{x'_i}}\boldsymbol{\gamma} + u_i + \epsilon_{it} \quad (3)$$

with $u_i \sim \mathcal{N}(0, \sigma_u^2)$.[6]

As usual in dynamic models, strict exogeneity can't be assumed because of the presence of lagged variables in (3), which is a source of endogeneity, i.e., of correlation between current shocks and past outcomes. The identification of the model requires predetermination, but also strict exogeneity of the covariates $\boldsymbol{x_i}$ conditional on the individual effects α_i.

Two normalizations are required for the joint identification of agents preferences and of unknown thresholds viewed as parameters to be estimated: (i) location: $\beta_0 = 0$, for shifting the constant and the thresholds simultaneously by some constant yields an observationally equivalent model; (ii) scale $\sigma_\varepsilon^2 = \pi^2/3$ (Logit) or $\sigma_\varepsilon^2 = 1$ (Probit), for multiplying the latent and all its parameters yields the same likelihood. Under these normalizations, the vector of parameters $\boldsymbol{\theta} = (\boldsymbol{\beta}, \boldsymbol{\gamma}, \boldsymbol{\gamma^0}, \boldsymbol{\rho}, \boldsymbol{\rho^0}, \boldsymbol{s}, \sigma_u)$ is identified.

As regards estimation, Wooldridge (2005) shows that the MLE is CAN as N grows large even for small, fixed T. This holds as soon as $T \geq 3$, which is required in order to disentangle the role of initial from that of past subjective well-being.

Two ways may still be ahead as far as the idiosyncratic shocks ε are concerned, a standard normal distribution (Probit) or a logistic distribution (Logit). Empirically, the latter produces a better fit, i.e., yields a higher likelihood.[7] Robustness checks are nevertheless provided with respect to that choice in section 6.1. Besides, average partial effects are close in both specifications. Moreover, the Logit permits an interpretation in terms of odds ratios, which the Probit does not allow.

[5]Time-constant variables and time dummies are dropped from the list of initial and mean covariates.

[6]Lee (2016) considered an equation of this form when focusing on health status in Korea.

[7]Due to the fatter tails of the logistic distribution, the Logit model puts more weight on extreme events.

A last concern is the selection of covariates, i.e., of explanatory variables $\boldsymbol{x_{it}}$ in the estimating equation. First, the literature devoted to the individual determinants of subjective well-being provides with some guidance. Second, statistical methods based either on the BIC, on the (rigorous) Lasso, or on a stepwise algorithm provide with useful tools to select the most relevant variables. In practice, I find that both the literature and statistical criteria (estimated on the static model) are globally consistent: the list of relevant covariates include objective quality of life, labor force status, family status, income and age. To be exhaustive though, education, occupation, gender and year dummies are also included.

5 Results

Tables 4 to 6 display the results from the main specification, namely equation (3), estimated on the unbalanced panel.[8] Carefully comparing columns allows me to disentangle the role played by unobserved heterogeneity from that played by state dependence. While column (1) omits individual effects in a pooled, cross-sectional regression fashion, column (2) includes a pure, random effect that is uncorrelated with covariates. Column (3) consists of the same correlated random effect approach as in column (4) but does not include the lagged dependent variable as an explanatory covariate, which column (4) does. Put differently, column (3) imposes the constraint $\boldsymbol{\rho} = 0$ with respect to the dynamic model of column (4), the preferred specification; column (2) assumes further that $\boldsymbol{\gamma} = \boldsymbol{\gamma^0} = \boldsymbol{\rho^0} = 0$ and column (1) adds up $u_i = u, \forall i$. On the one hand, state dependence is encompassed by $\boldsymbol{\rho}$: hence its role can be isolated by a direct comparison between columns (3) and (4), given that the sample is voluntarily identical (hence the need of restricting our attention to individuals with at least two consecutive observations), so that observations contributing to the identification of the parameters of the model are the same in all columns. On the other hand, $\boldsymbol{\gamma}, \boldsymbol{\gamma^0}, \boldsymbol{\rho^0}$ as well as the residual variance account for unobserved heterogeneity, the role of initial conditions being encompassed by $\boldsymbol{\rho^0}$. As a *caveat*, an eyeball, quantitative comparison across columns would be misleading since the coefficients do not have a common scale (see, e.g., Contoyannis et al., 2004); however, this warning concerns neither rel-

[8]For the sake of readability, the same Table of results has been cut into three parts. Though available upon request, $\boldsymbol{\gamma}$ and $\boldsymbol{\gamma_0}$ coefficients are not reported, at the exception of the estimate corresponding to mean income.

ative, nor qualitative comparisons (namely, significance). Average partial effects permit nevertheless a quantitative comparison across columns.

I find empirical evidence of state dependence, which confirms the eyeball impression given by Table 3. The estimated autocorrelation vector $\boldsymbol{\rho}$ is statistically significant at usual levels: $H_0 : \rho_1 = \ldots = \rho_{10} = 0$ is rejected at 5%, the $\chi^2(10)$ statistic being 196.99. Moreover, state dependence turns to have a nonlinear impact on the latent propensity to happiness: $H_0 : \rho_j - \rho_{j+1} = \rho_{j+1} - \rho_{j+2}, \forall j = 0, \ldots, K-2$, is also rejected at 5% with a $\chi^2(9)$ statistic of 57.57, which justifies the specification of state dependence adopted here with respect to a more parsimonious, linear one. It is striking to notice that only the highest levels of life satisfaction exhibit state dependence, while the lowest levels of subjective well-being have less inertia: this evidence confirms the asymmetry observed in the descriptive analysis. Formally, $H_0 : \rho_1 = \ldots = \rho_6 = 0$ is rejected at 5%, the $\chi^2(6)$ statistics being 22.29 with a p-value of .001, while $H_0 : \rho_1 = \ldots = \rho_5 = 0$ cannot be rejected at 5%, the $\chi^2(5)$ statistics being 10.55 with a p-value of .06. In that sense, happiness, defined as a reported level of life satisfaction higher than 7, is persistent, while unhappiness, defined as a reported level of life satisfaction lower than 5, is not.

Moreover, the role played by the initial condition, measured by $\boldsymbol{\rho}^0$, is strong, as in Wooldridge (2005). In particular, it is a first-order issue to explain the persistence of subjective well-being even when compared to state dependence: the magnitude of the impact of initial conditions $\boldsymbol{\rho}^0$ exceeds the one related to the lagged dependent variable $\boldsymbol{\rho}$.

Average partial effects (APEs)[9] enable me to quantify previous statements, to compare the relative effects of state dependence and of other explanatory variables, and to answer the question: by how much is the impact of the main determinants of subjective well-being attenuated when taking initial conditions and state dependence into account? Though APEs are computed for all covariates and for each of the eleven levels of self-assessed life satisfaction, Table 7 summarizes the APEs of selected variables only (current and mean income, unemployed, some kinds of marital status, poor health, weak social ties and lagged life satisfaction) on the probability to report the highest level, i.e., level 10, of life satisfaction – for the

[9]For continuous regressors such as income, these APEs are obtained by taking the derivative of the ordered Logit probabilities with respect to the variable in question. For discrete regressors, such as lagged life satisfaction, they are obtained by taking differences.

sake of readability.[10] We already knew that unemployment diminished that probability, we now learn that controlling neither for unobserved heterogeneity, nor for state dependence leads to overestimate its impact by 1/3. The same holds as regards weak social ties and poor health, for which the overestimation bias is even more pronounced, the attenuation factor being strictly higher than 2. It is also confirmed that parents of more than 2 children do not report more frequently level 10 than other couples, contrary to what a naive analysis would conclude. Having a partner raises by 2pp the propensity of being most satisfied with one's life. A 1% increase in permanent income raises that probability by 1.6pp. Interestingly, the effect of pure state dependence, i.e., of reporting level 10 the year before, increases the probability of reporting that very level by 7.1 points: for the sake of comparison, unemployment, poor health and weak social ties altogether make this probability decrease by nearly the same amount.

The current analysis encompasses main stylized facts in happiness economics. It must yet be acknowledged that the identifying source of variation for the effects of covariates differs substantially when moving from a RE specification (column (2)) to a CRE specification (column (3)). The former relies on inter-individual differences while the latter stems mostly from intra-individual, or temporal, differences. On top of that, the CRE specification considered here allows to disentangle the role played by average covariates (say, the *level*) from the one played by initial covariates (say, the *growth*). First, objective deprivation indicators measuring the quality of life are the strongest predictors of a lower life satisfaction, especially the weakness of social ties, poor living conditions, and poor health, which is consistent with the previous evidence found by Godefroy and Lollivier (2014). The $\boldsymbol{\gamma}^{\mathbf{0}}$ coefficients[11] for these indicators are significantly positive while corresponding $\boldsymbol{\beta}$ and $\boldsymbol{\gamma}$ coefficients[12] are significantly negative, which suggests that the effect stems from both between and within dimensions. Put differently, people with a poor objective quality of life tend to report lower levels of subjective well-being, and an individual who becomes deprived declares herself more unhappy, too -as individuals with a trend towards deprivation do. Second, unemployment is a major cause of misery, and stems mostly from the fact of losing one's job. Third, the role of family status can be precised: partners and parents do not differ much from singles in the sense

[10] All other APEs are available upon request.
[11] not reported in the Tables for the sake of readability, available upon request.
[12] Same as footnote 11.

that their $\boldsymbol{\gamma}$ coefficients are not significantly different from zero at usual levels. However, finding a partner and having children increases self-assessed well-being: their $\boldsymbol{\beta}$ (and $\boldsymbol{\gamma}^{\mathbf{0}}$) coefficients are positive and significant. In other words, these people are not happier *per se* but unions and children make them (temporarily) happier. However, this interpretation is complicated by composition effects, the reference category here being the "always singles", and not the "once singles" as in the RE specification. Fourth, the CRE approach makes it possible to identify the channel through which income matters, i.e., average income. In other words, *transitory income*, i.e., current shocks of income viewed as deviations from average or *permanent income*, matter less than the latter, which is reminiscent of the result obtained by Frijters et al. (2004), among many others. Moreover, the corresponding $\boldsymbol{\gamma}^{\mathbf{0}}$ being negative suggests that income growth is correlated with a higher life satisfaction. Fifth, the U-shape with age is retrieved. Sixth, gender does not matter. Seventh, neither does education. Eighth, neither does occupation – farmers aside, who are significantly far less satisfied with their lives.

Finally, from a purely statistical perspective, the fit of the model, as measured by the average individual log-likelihood, reaches -1.51 in the restricted model. Put differently, the average likelihood of an observation is .22, which is quite satisfying. By construction, the fit improves as one moves from most constrained specification (column 1) to least constrained specification (column 4). A likelihood-ratio test confirms that this improvement is significant at usual levels for (i) the random effect, i.e., moving from column 1 to column 2: the test statistic is 8,380 while the .95 quantile of the $\chi^2(1)$ is 3.84, (ii) the CRE specification, i.e., moving from column 2 to column 3: the test statistic is 7,661 while the .95 quantile of the $\chi^2(66)$ is 85.97, and (iii) the lagged dependent variable, i.e., moving from column 3 to column 4: the test statistic is 316 while the .95 quantile of the $\chi^2(10)$ is 18.31. Moreover, the residual variance in subjective well-being, i.e., the dispersion of subjective well-being that remains unexplained by the full model including covariates, correlated random effects and state dependence, shrinks also mechanically as one moves from column (2) to column (4): it has been divided by a factor 2 (resp. 3 in the dynamic CRE specification) in comparison with a pure RE specification.

6 Robustness checks

I perform several robustness checks to assess the sensitivity of previous results with respect to (i) functional form; (ii) attrition; (iii) data. The last point deserves particular attention since it guarantees that the point made in this paper unveils some empirical regularity that is not specific to the French database; on the contrary, I find that it is common to several countries.

6.1 Parametric assumptions

First, I estimate an alternative parametric specification, namely a Probit model. I replicate the entire analysis by assuming that the idiosyncratic error terms ε_{it} follow a normal distribution instead of a logistic distribution, although both the fit, as measured by the log-likelihood, and the parsimony, proxied by the BIC, would be worsened. From a qualitative point of view, they turn out to be very close to the previous ones. From a quantitative point of view, the same holds as regards average partial effects.[13]

6.2 Attrition

Second, I address statistical concerns related to endogenous attrition. Tables 8 to 10 display the estimates obtained on the balanced panel. From a qualitative point of view, they yield similar results as those obtained on the unbalanced panel. The main qualitative difference is related to the age effect, which disappears after controlling for state dependence on top of unobserved heterogeneity. However, this loss of significance stems from a lack of statistical power that is due to low sample size, which is less worrying.

I explore next the role played by sample attrition. I resort to a statistical test for possible attrition bias as recommended by Verbeek and Nijman (1992). Their test consists in introducing a dummy for being part of the balanced panel and the number of times an individual is present in the unbalanced sample as further explanatory variables in the previous model. In practice, both covariates turn out to be non-significant, which indicates that endogenous attrition is not too much of a problem here.[14]

[13] All these estimates are available upon request.

[14] In case it were, a method to deal with it could have been to use the inverse probability weighting solution proposed by Wooldridge (2002).

6.3 Replication: Australia, Germany and the UK

Third, I replicate the current analysis on other databases issued from three other countries (Australia, Germany and the UK), which permits to neutralize the role played by the institutional setting (including survey design). In all these countries, stylized facts of subjective well-being are retrieved.

First, I use the Household, Income and Labour Dynamics in Australia (HILDA) Survey for which I resort to 17 waves from 2001 to 2017.[15] More than 20,000 individuals report their self-assessed overall satisfaction with life on a similar Cantril scale. Table 11 depicts annual transitions of individual answers to that question; it suggests once again that state dependence can't be ignored, in particular for individuals who are already satisfied with their life. An econometric analysis confirms this impression. The same model as previously is estimated, except that quality of life indicators are missing in the survey. Controls include age, age squared, gender, income (measured at the household level), education, labor force status and family status. Table 12 suggests that the very same results as those found in France hold in Australia. To be precise, state dependence is all the more pronounced that one gets higher in the distribution of subjective well-being, and the respective effects of initial conditions and of state dependence look pretty similar.

Second, I resort to the GSOEP in Germany. This exceptional longitudinal survey has been available from 1984 to 2017[16] and has no less than 34 waves, which permits to follow accurately the evolution of life satisfaction for more than 50,000 individuals. As in Australia and in France, the latter is self-reported on a Cantril scale. Tables 13 and 14 display very similar results as before, from a qualitative point of view. State dependence is increasing all over the distribution of subjective well-being. The GSOEP is the unique case where the initial condition (namely, the vector of coefficients $\boldsymbol{\rho}^{\mathbf{0}}$) matters less than state dependence encompassed by $\boldsymbol{\rho}$, which is conform to the rationale: the initial condition in the GSOEP may date back up to 33 years ago.

Third, I use the UK Understanding Society (UKUS) panel. This survey takes over the British Household Panel Survey (BHPS), starting from 2009. Eight waves are available for about 50,000 individuals. These people are asked about their overall satisfaction with life and their answer is available on a discrete, ordered scale

[15] My code is adapted from the one provided by PanelWhiz on http://www.panelwhiz.eu.

[16] In West Germany only; East Germany has been surveyed from the reunification onwards.

ranging from 1 to 7. Table 15 suggests as previously that state dependence can't be ignored, in particular for individuals who are already satisfied with their life, which is confirmed by a *ceteris paribus* analysis. Following the recommendations of Clark and Georgellis (2013) on how to estimate subjective well-being equations on the BHPS, I control for age, age squared, gender, income (measured at the household level), the number of children, dummies for married, unemployed and self-employed individuals, as well as dummies for having a high or a medium degree, on top of year dummies. Table 16 exhibits very similar results to those that prevail in the other countries. Yet state dependence exhibits some non-monotonicity, namely a U-shape here: it is more pronounced for level 2 than it is for level 3, which suggests that bad states of life satisfaction are somehow more persistent in the UK. Interestingly, the asymmetry observed in Australia, in France and in Germany as regards persistence at the top and at the bottom of the distribution of happiness was thus neither a mechanical effect, nor a statistical artifact. Finally, initial conditions matter more than state dependence, as was the case in Australia and in France.

7 Conclusion

This paper has disentangled the roles played by state dependence and unobserved heterogeneity in self-assessed life satisfaction, a much ignored issue in happiness economics. Not only is subjective well-being persistent over time at both aggregate and individual levels, but this persistence turns out to be more pronounced at the top of the distribution of self-reported happiness. Thanks to the estimation of a dynamic, ordered Logit model with correlated random effects which permits to take unobserved heterogeneity into account, I have shown that this empirical evidence held in at least three other countries (Australia, Germany and the UK), neutralizing hence the role played by the institutional setting (including survey design). In the UK, the asymmetry is slightly less pronounced: bad states are persistent, too. I have also quantified the impact of initial conditions, which is rather large compared with the roles played by other usual determinants of happiness.

From an econometric perspective, there are at least two limits of the current approach. First, the dynamic, nonlinear model estimated here does not include fixed effects. A natural extension would thus consist in considering a dynamic, ordered model with fixed effects in the vein of Frijters et al. (2004),Bartolucci and

Nigro (2010) or Carro and Traferri (2014). Second, state dependence could be modelled by higher-order Markov processes than the first-order process used here: more lags could be included in the estimating equation.

From a social science perspective, further research should try to understand which mechanisms explain such an inertia. Are cognitive biases at stake? Do anchoring effects matter? Why is it revealed as more costly for individuals to revise downwards their self-assessed evaluation of life satisfaction? Why does unhappiness look more transitory? Determining the profound causes that lie behind persistence of happiness sounds like a challenging and exciting task.

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A Figures

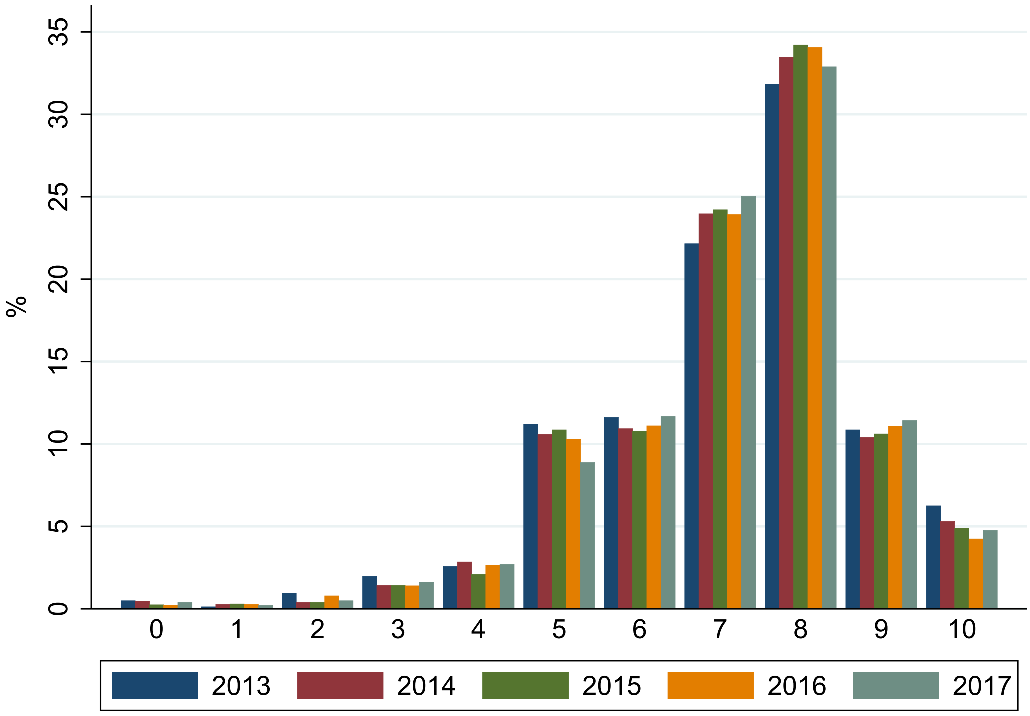

Figure 1: Evolution of life satisfaction in France

B Tables

Table 1: Summary statistics

	mean	sd	min	max
Life satisfaction	7.17	1.67	0	10
Female	0.62	0.49	0	1
Age	55.0	16.8	17	100
Income	25,799	36,408	20	4,468,733
Education				
No degree	0.23	0.42	0	1
High-school	0.29	0.45	0	1
Vocational	0.32	0.47	0	1
College	0.16	0.37	0	1
Other degree	0.00	0.06	0	1
Labor force status				
Employed	0.47	0.50	0	1
Unemployed	0.06	0.23	0	1
Student	0.02	0.12	0	1
Inactive	0.05	0.23	0	1
Retired	0.39	0.49	0	1
Undetermined	0.02	0.13	0	1
Occupation				
Clerk	0.28	0.45	0	1
Farmer	0.03	0.17	0	1
White collar	0.13	0.34	0	1
Self-employed	0.06	0.23	0	1
Intermediate	0.22	0.42	0	1
Blue collar	0.17	0.38	0	1
Other	0.10	0.30	0	1
Family status				
Single	0.24	0.42	0	1
Two adults, w/o child	0.39	0.49	0	1
Two adults, 1 child	0.09	0.29	0	1
Two adults, 2 children	0.11	0.32	0	1
Two adults, 3+ children	0.05	0.22	0	1
Single parent	0.04	0.20	0	1
Others w/o child	0.05	0.22	0	1
Others with children	0.03	0.16	0	1
Undetermined	0.01	0.08	0	1
Quality of life				
Poor living conditions	0.10	0.30	0	1
Environmental troubles	0.03	0.18	0	1
Psycho-social hazard	0.12	0.33	0	1
Economic insecurity	0.02	0.16	0	1
Poor health	0.09	0.29	0	1
Insecurity	0.14	0.35	0	1
Weak social ties	0.14	0.34	0	1
Observations		44,085		

Source. French SRCV survey, 2013-2017.

Sample. Unbalanced panel of 19,253 individuals with at least two consecutive observations.

Table 2: Within-individual heterogeneity in reported life satisfaction

Absolute maximal difference in reported life satisfaction	Frequency (%)
0	19.4
1	36.0
2	23.7
3	12.0
4	4.6
5	2.8
6-10	1.6

Source. French SRCV survey, 2013-2017.
Sample. Unbalanced panel of 19,253 individuals with at least two consecutive observations.
Lecture. Highest minus lowest level of life satisfaction over the period.

Table 3: Life satisfaction in France: annual transitions

Destination → Initial ↓	0	1	2	3	4	5	6	7	8	9	10
0	24.5	5.1	10.2	9.2	9.2	22.5	6.1	7.1	4.1	0.0	2.0
1	3.5	15.5	13.8	10.3	3.5	22.4	5.2	5.2	17.2	0.0	3.5
2	4.4	7.0	11.4	10.8	13.3	23.4	12.7	12.0	2.5	2.5	0.0
3	2.7	2.2	6.2	15.4	19.2	26.8	12.2	9.2	4.9	1.1	0.3
4	1.7	0.3	4.3	9.8	15.1	32.3	16.1	11.7	6.3	1.9	0.6
5	0.6	0.3	1.8	4.2	7.5	32.4	20.3	18.4	11.3	1.9	1.4
6	0.2	0.3	0.5	2.0	4.0	16.6	25.8	31.6	15.7	2.4	1.1
7	0.1	0.1	0.2	0.7	1.4	7.7	13.8	40.0	30.3	4.5	1.3
8	0.0	0.1	0.1	0.3	0.6	4.0	5.5	20.9	52.2	12.9	3.5
9	0.1	0.0	0.1	0.2	0.5	1.1	2.3	10.1	38.9	35.4	11.3
10	0.1	0.0	0.1	0.4	0.2	2.2	2.2	6.5	24.0	24.9	39.3
Total	0.4	0.3	0.7	1.6	2.6	10.2	11.1	23.8	33.1	11.1	5.2

Source. French SRCV survey, 2013-2017.
Sample. Unbalanced panel of 19,253 individuals with at least two consecutive observations.

Table 4: Dynamic ordered Logit model - unbalanced panel (1)

Dependent	Life satisfaction (LS_t)			
	(1)	(2)	(3)	(4)
Current log income (β)	0.408***	0.510***	0.073	0.046
	(0.028)	(0.038)	(0.055)	(0.053)
Mean log income (γ)			0.430***	0.400***
			(0.114)	(0.105)
Initial log income (γ^0)			-0.154**	-0.141**
			(0.076)	(0.067)
Age	-0.039***	-0.060***	-0.025***	-0.022***
	(0.005)	(0.007)	(0.006)	(0.006)
$Age^2/100$	0.026***	0.039***	0.013**	0.011**
	(0.004)	(0.007)	(0.006)	(0.005)
Female	-0.005	-0.003	-0.023	-0.021
	(0.026)	(0.039)	(0.033)	(0.029)
Education (ref=no degree)				
High-school	-0.051	-0.028	-0.124	-0.114
	(0.041)	(0.060)	(0.258)	(0.246)
Vocational	-0.092**	-0.096*	0.071	0.065
	(0.036)	(0.053)	(0.229)	(0.219)
College	0.018	0.104	0.038	0.046
	(0.050)	(0.074)	(0.304)	(0.289)
Other	-0.051	0.206	0.411	0.434
	(0.199)	(0.300)	(0.345)	(0.313)
Labor force status (ref=employed)				
Unemployed	-0.817***	-1.164***	-0.771***	-0.716***
	(0.054)	(0.073)	(0.108)	(0.105)
Student	-0.012	-0.006	-0.338	-0.381
	(0.101)	(0.144)	(0.270)	(0.262)
Inactive	-0.168***	-0.374***	-0.439**	-0.426**
	(0.064)	(0.091)	(0.174)	(0.167)
Retired	-0.150***	-0.194***	-0.256**	-0.258**
	(0.043)	(0.060)	(0.113)	(0.108)
Undetermined	-0.508***	-0.690***	-0.392**	-0.394**
	(0.098)	(0.126)	(0.189)	(0.181)
Occupation (ref=clerk)				
Farmer	-0.503***	-0.715***	-1.761**	-1.715**
	(0.077)	(0.113)	(0.696)	(0.671)
Blue collar	0.006	-0.011	-0.227	-0.204
	(0.039)	(0.056)	(0.152)	(0.145)
Intermediate	0.033	0.125**	-0.146	-0.160
	(0.034)	(0.050)	(0.133)	(0.128)
White collar	0.154***	0.329***	0.142	0.089
	(0.045)	(0.065)	(0.169)	(0.162)
Self-employed	-0.126**	-0.142*	-0.114	-0.175
	(0.058)	(0.084)	(0.235)	(0.226)
Other	0.034	0.017	-0.311**	-0.287*
	(0.052)	(0.076)	(0.154)	(0.147)
Undetermined	0.010	-0.183	-0.760**	-0.739**
	(0.183)	(0.232)	(0.320)	(0.312)
Family status (ref=single)				
Two adults (no child)	0.554***	0.881***	0.732***	0.637***
	(0.031)	(0.046)	(0.117)	(0.112)
Other (no child)	0.495***	0.741***	0.479***	0.380**
	(0.058)	(0.078)	(0.155)	(0.148)
Single parent	-0.049	-0.026	0.246	0.184
	(0.060)	(0.086)	(0.169)	(0.161)
Two adults (1 child)	0.557***	0.862***	0.692***	0.575***
	(0.044)	(0.064)	(0.144)	(0.137)
Two adults (2 children)	0.672***	0.978***	0.592***	0.481***
	(0.043)	(0.064)	(0.165)	(0.157)
Two adults (3+ children)	0.763***	1.112***	0.695***	0.568***
	(0.058)	(0.085)	(0.205)	(0.196)
Other (children)	0.690***	0.938***	0.653***	0.563***
	(0.071)	(0.093)	(0.176)	(0.168)
Undetermined	0.657***	1.088***	0.910***	0.811***
	(0.126)	(0.163)	(0.221)	(0.215)

Table 5: Ordered Logit model - unbalanced panel (2)

Quality of life				
Poor living conditions	-1.258***	-1.494***	-0.718***	-0.693***
	(0.041)	(0.055)	(0.070)	(0.068)
Environmental troubles	-0.367***	-0.391***	-0.076	-0.071
	(0.055)	(0.069)	(0.083)	(0.082)
Insecurity	-0.233***	-0.229***	-0.053	-0.056
	(0.029)	(0.037)	(0.043)	(0.043)
Weak social ties	-1.157***	-1.269***	-0.687***	-0.670***
	(0.032)	(0.041)	(0.049)	(0.048)
Psycho-social hazard	-0.614***	-0.703***	-0.393***	-0.384***
	(0.031)	(0.041)	(0.050)	(0.049)
Economic insecurity	-0.344***	-0.358***	-0.049	-0.023
	(0.063)	(0.087)	(0.108)	(0.104)
Poor health	-1.087***	-1.252***	-0.713***	-0.675***
	(0.042)	(0.053)	(0.065)	(0.063)
Initial life satisfaction (ρ^0) - (ref=0)				
$LS_{2013} = 1$			-0.127	-0.385
			(0.441)	(0.440)
$LS_{2013} = 2$			0.388	0.168
			(0.327)	(0.333)
$LS_{2013} = 3$			1.154***	0.878***
			(0.296)	(0.297)
$LS_{2013} = 4$			1.240***	0.922***
			(0.281)	(0.290)
$LS_{2013} = 5$			1.963***	1.504***
			(0.273)	(0.283)
$LS_{2013} = 6$			2.399***	1.839***
			(0.274)	(0.284)
$LS_{2013} = 7$			3.103***	2.363***
			(0.274)	(0.286)
$LS_{2013} = 8$			4.228***	3.234***
			(0.275)	(0.291)
$LS_{2013} = 9$			5.469***	4.157***
			(0.280)	(0.300)
$LS_{2013} = 10$			6.479***	4.817***
			(0.286)	(0.312)
Past LS (ρ) - (ref=0)				
$LS_{t-1} = 1$				0.558
				(0.383)
$LS_{t-1} = 2$				0.367
				(0.274)
$LS_{t-1} = 3$				0.348
				(0.246)
$LS_{t-1} = 4$				0.465*
				(0.247)
$LS_{t-1} = 5$				0.580**
				(0.242)
$LS_{t-1} = 6$				0.712***
				(0.244)
$LS_{t-1} = 7$				0.936***
				(0.246)
$LS_{t-1} = 8$				1.200***
				(0.251)
$LS_{t-1} = 9$				1.555***
				(0.258)
$LS_{t-1} = 10$				2.023***
				(0.270)

Table 6: Ordered Logit model - unbalanced panel (3)

Cut-offs				
s_1	-3.551***	-5.406***	-2.610***	-2.324***
	(0.301)	(0.415)	(0.721)	(0.665)
s_2	-3.082***	-4.846***	-2.046***	-1.781***
	(0.297)	(0.411)	(0.719)	(0.663)
s_3	-2.305***	-3.894***	-1.087	-0.861
	(0.293)	(0.408)	(0.717)	(0.661)
s_4	-1.462***	-2.819***	-0.011	0.171
	(0.291)	(0.406)	(0.717)	(0.661)
s_5	-0.709**	-1.824***	0.984	1.121*
	(0.291)	(0.405)	(0.717)	(0.661)
s_6	0.680**	0.110	2.924***	2.967***
	(0.291)	(0.404)	(0.718)	(0.662)
s_7	1.504***	1.329***	4.153***	4.134***
	(0.291)	(0.404)	(0.719)	(0.663)
s_8	2.723***	3.217***	6.067***	5.948***
	(0.292)	(0.404)	(0.719)	(0.664)
s_9	4.491***	5.978***	8.873***	8.610***
	(0.293)	(0.406)	(0.722)	(0.667)
s_{10}	5.724***	7.763***	10.682***	10.332***
	(0.294)	(0.408)	(0.724)	(0.670)
σ_u^2		3.952***	2.196***	1.379***
		(0.105)	(0.069)	(0.083)
Year dummies	Yes	Yes	Yes	Yes
Individual effects	No	RE	CRE	CRE
# of individuals	19,253	19,253	19,253	19,253
# of observations	44,085	44,085	44,085	44,085
$\log(L)/N$	-1.699	-1.604	-1.517	-1.514

Source. French SRCV survey, 2013-2017.

Sample. 19,253 individuals with at least two consecutive observations.

Robust standard errors clustered at the individual level.

Estimates for initial and average covariates not reported (income inside).

Table 7: Average partial effects on probability of reporting level 10 of life satisfaction (selected variables only)

Dependent	Life satisfaction (LS_t)			
	(1)	(2)	(3)	(4)
Current log income - (β)	0.022***	0.019***	0.003	0.002
	(0.002)	(0.001)	(0.002)	(0.002)
Mean log income - (γ)			0.017***	0.016***
			(0.004)	(0.004)
Unemployed	-0.034***	-0.034***	-0.027***	-0.025***
	(0.002)	(0.002)	(0.003)	(0.003)
Two adults (no child)	0.025***	0.028***	0.026***	0.023***
	(0.001)	(0.002)	(0.004)	(0.004)
Two adults (1 child)	0.025***	0.028***	0.024***	0.021***
	(0.002)	(0.002)	(0.005)	(0.005)
Two adults (2 children)	0.032***	0.033***	0.020***	0.017***
	(0.002)	(0.002)	(0.006)	(0.006)
Two adults (3+ children)	0.038***	0.039***	0.025***	0.020***
	(0.004)	(0.004)	(0.008)	(0.008)
Weak social ties	-0.061***	-0.048***	-0.027***	-0.027***
	(0.002)	(0.002)	(0.003)	(0.002)
Poor health	-0.057***	-0.047***	-0.028***	-0.027***
	(0.003)	(0.002)	(0.003)	(0.003)
Past life satisfaction (ρ) - (ref=0)				
$LS_{t-1} = 1$				0.011
				(0.008)
$LS_{t-1} = 2$				0.006
				(0.005)
$LS_{t-1} = 3$				0.006
				(0.004)
$LS_{t-1} = 4$				0.009**
				(0.004)
$LS_{t-1} = 5$				0.011***
				(0.004)
$LS_{t-1} = 6$				0.015***
				(0.004)
$LS_{t-1} = 7$				0.021***
				(0.004)
$LS_{t-1} = 8$				0.030***
				(0.004)
$LS_{t-1} = 9$				0.045***
				(0.005)
$LS_{t-1} = 10$				0.071***
				(0.006)
Individual effects	No	RE	CRE	CRE
# of individuals	19,253	19,253	19,253	19,253
# of observations	44,085	44,085	44,085	44,085

Source. French SRCV survey, 2013-2017.
Sample. Unbalanced panel of 19,253 individuals with at least two consecutive observations.
Robust standard errors clustered at the individual level.

Table 8: Dynamic ordered Logit model - balanced panel (1)

Dependent	Life satisfaction (LS_t)			
	(1)	(2)	(3)	(4)
Current log income (β)	0.412***	0.438***	0.139*	0.094
	(0.050)	(0.062)	(0.082)	(0.081)
Mean log income (γ)			0.337*	0.322*
			(0.183)	(0.169)
Initial log income (γ^0)			-0.052	-0.048
			(0.127)	(0.113)
Age	-0.036***	-0.054***	-0.017	-0.015
	(0.009)	(0.014)	(0.013)	(0.011)
$Age^2/100$	0.026***	0.038***	0.008	0.007
	(0.009)	(0.013)	(0.012)	(0.010)
Female	-0.028	-0.067	-0.039	-0.033
	(0.049)	(0.076)	(0.062)	(0.054)
Education (ref=no degree)				
High-school	0.012	0.140	-0.349	-0.329
	(0.077)	(0.116)	(0.467)	(0.443)
Vocational	-0.023	0.045	-0.315	-0.306
	(0.068)	(0.101)	(0.375)	(0.356)
College	0.075	0.220	-0.470	-0.451
	(0.094)	(0.141)	(0.507)	(0.484)
Other	0.092	1.000*	1.355**	1.424***
	(0.446)	(0.519)	(0.572)	(0.519)
Labor force status (ref=employed)				
Unemployed	-0.681***	-0.976***	-0.694***	-0.641***
	(0.108)	(0.133)	(0.165)	(0.163)
Student	-0.138	-0.268	-0.158	-0.243
	(0.289)	(0.311)	(0.387)	(0.390)
Inactive	-0.079	-0.168	0.160	0.131
	(0.110)	(0.148)	(0.242)	(0.234)
Retired	-0.133*	-0.121	-0.070	-0.089
	(0.076)	(0.102)	(0.155)	(0.150)
Undetermined	-0.471***	-0.492**	0.026	-0.018
	(0.181)	(0.212)	(0.280)	(0.271)
Occupation (ref=clerk)				
Farmer	-0.424***	-0.609***	-1.458	-1.322
	(0.142)	(0.221)	(1.034)	(0.989)
Blue collar	0.010	-0.114	-0.547**	-0.505**
	(0.073)	(0.105)	(0.221)	(0.211)
Intermediate	-0.001	0.028	-0.222	-0.223
	(0.063)	(0.092)	(0.199)	(0.196)
White collar	0.142*	0.291**	0.211	0.189
	(0.084)	(0.121)	(0.251)	(0.244)
Self-employed	-0.041	0.020	0.141	0.095
	(0.108)	(0.157)	(0.330)	(0.325)
Other	-0.032	-0.196	-0.669***	-0.616***
	(0.085)	(0.124)	(0.232)	(0.227)
Undetermined	0.048	-0.161	-0.626	-0.599
	(0.441)	(0.425)	(0.489)	(0.483)
Family status (ref=single)				
Two adults (no child)	0.533***	0.840***	0.694***	0.609***
	(0.055)	(0.079)	(0.162)	(0.156)
Other (no child)	0.495***	0.663***	0.334	0.254
	(0.112)	(0.146)	(0.231)	(0.224)
Single parent	0.003	0.140	0.269	0.218
	(0.108)	(0.149)	(0.232)	(0.224)
Two adults (1 child)	0.631***	0.927***	0.617***	0.498**
	(0.081)	(0.117)	(0.210)	(0.203)
Two adults (2 children)	0.713***	1.042***	0.580**	0.463**
	(0.083)	(0.120)	(0.239)	(0.229)
Two adults (3+ children)	0.928***	1.200***	0.405	0.270
	(0.114)	(0.163)	(0.291)	(0.283)
Other (children)	1.039***	1.142***	0.464*	0.366
	(0.139)	(0.156)	(0.253)	(0.245)
Undetermined	0.169	0.727***	0.595**	0.527*
	(0.171)	(0.228)	(0.292)	(0.290)

Table 9: Ordered Logit model - balanced panel (2)

Quality of life				
Poor living conditions	-1.237***	-1.249***	-0.776***	-0.755***
	(0.074)	(0.095)	(0.110)	(0.108)
Environmental troubles	-0.480***	-0.343***	-0.092	-0.098
	(0.099)	(0.113)	(0.118)	(0.116)
Insecurity	-0.236***	-0.128**	0.058	0.050
	(0.050)	(0.057)	(0.062)	(0.061)
Weak social ties	-1.207***	-1.155***	-0.771***	-0.751***
	(0.058)	(0.068)	(0.074)	(0.073)
Psycho-social hazard	-0.564***	-0.624***	-0.475***	-0.473***
	(0.057)	(0.067)	(0.074)	(0.074)
Economic insecurity	-0.237**	-0.206	-0.019	0.022
	(0.120)	(0.147)	(0.157)	(0.154)
Poor health	-1.105***	-1.083***	-0.724***	-0.694***
	(0.073)	(0.085)	(0.093)	(0.092)
Initial LS (ρ^0) - (ref=0)				
$\text{LS}_{2013} = 1$			-1.084	-1.443*
			(0.675)	(0.745)
$\text{LS}_{2013} = 2$			-0.112	-0.248
			(0.673)	(0.672)
$\text{LS}_{2013} = 3$			0.485	0.355
			(0.619)	(0.615)
$\text{LS}_{2013} = 4$			0.335	0.227
			(0.606)	(0.610)
$\text{LS}_{2013} = 5$			1.210**	0.999*
			(0.590)	(0.598)
$\text{LS}_{2013} = 6$			1.454**	1.197**
			(0.590)	(0.598)
$\text{LS}_{2013} = 7$			2.158***	1.758***
			(0.589)	(0.598)
$\text{LS}_{2013} = 8$			3.126***	2.543***
			(0.590)	(0.601)
$\text{LS}_{2013} = 9$			4.348***	3.533***
			(0.596)	(0.611)
$\text{LS}_{2013} = 10$			5.530***	4.432***
			(0.611)	(0.631)
Past LS (ρ) - (ref=0)				
$\text{LS}_{t-1} = 1$				1.049
				(0.750)
$\text{LS}_{t-1} = 2$				0.490
				(0.492)
$\text{LS}_{t-1} = 3$				0.405
				(0.414)
$\text{LS}_{t-1} = 4$				0.380
				(0.414)
$\text{LS}_{t-1} = 5$				0.432
				(0.410)
$\text{LS}_{t-1} = 6$				0.597
				(0.413)
$\text{LS}_{t-1} = 7$				0.821**
				(0.413)
$\text{LS}_{t-1} = 8$				1.095***
				(0.417)
$\text{LS}_{t-1} = 9$				1.418***
				(0.425)
$\text{LS}_{t-1} = 10$				1.889***
				(0.442)

Table 10: Ordered Logit model - balanced panel (3)

Cut-offs				
s_1	-3.576***	-5.734***	-2.379*	-2.284*
	(0.545)	(0.714)	(1.413)	(1.284)
s_2	-2.980***	-5.053***	-1.698	-1.616
	(0.534)	(0.705)	(1.406)	(1.277)
s_3	-2.316***	-4.283***	-0.925	-0.858
	(0.526)	(0.697)	(1.401)	(1.272)
s_4	-1.389***	-3.168***	0.197	0.241
	(0.524)	(0.695)	(1.401)	(1.272)
s_5	-0.581	-2.159***	1.208	1.227
	(0.523)	(0.693)	(1.402)	(1.273)
s_6	0.846	-0.248	3.121**	3.084**
	(0.524)	(0.693)	(1.404)	(1.276)
s_7	1.689***	0.971	4.343***	4.269***
	(0.525)	(0.693)	(1.405)	(1.277)
s_8	2.950***	2.890***	6.272***	6.135***
	(0.526)	(0.693)	(1.406)	(1.279)
s_9	4.794***	5.729***	9.138***	8.912***
	(0.528)	(0.696)	(1.409)	(1.283)
s_{10}	6.144***	7.648***	11.082***	10.799***
	(0.530)	(0.701)	(1.414)	(1.288)
σ_u^2		3.729***	2.044***	1.379***
		(0.154)	(0.094)	(0.100)
Year dummies	Yes	Yes	Yes	Yes
Individual effects	No	RE	CRE	CRE
# of individuals	4,081	4,081	4,081	4,081
# of observations	16,324	16,324	16,324	16,324
$\log(L)/N$	-1.667	-1.534	-1.473	-1.468

Source. French SRCV survey, 2013-2017.
Sample. Balanced panel of 4,081 individuals.
Robust standard errors clustered at the individual level.
Estimates for initial and average covariates not reported (income inside).

C Appendix

C.1 Australia

Table 11: Life satisfaction in Australia: annual transitions

Destination → Initial ↓	0	1	2	3	4	5	6	7	8	9	10
0	14.1	11.1	13.1	9.1	4.0	21.2	7.1	3.5	7.6	2.5	6.6
1	4.3	12.7	12.4	16.2	8.5	17.8	7.3	8.1	7.0	4.6	1.2
2	3.1	6.2	12.2	13.7	8.8	20.7	9.4	11.3	10.8	2.3	1.5
3	1.3	2.9	6.6	13.0	11.5	22.2	14.6	14.0	9.3	2.9	1.7
4	0.6	0.9	3.2	7.5	12.8	23.0	18.2	17.9	10.2	3.7	2.0
5	0.6	0.7	1.7	3.2	6.2	23.9	17.9	23.0	15.7	4.1	3.0
6	0.1	0.1	0.6	1.7	3.9	11.4	21.9	34.2	19.3	4.7	2.0
7	0.1	0.1	0.3	0.5	1.1	4.4	10.4	40.0	34.2	7.2	1.9
8	0.0	0.0	0.1	0.2	0.4	1.9	3.3	19.4	49.7	20.4	4.5
9	0.0	0.1	0.1	0.1	0.2	0.8	1.2	6.5	32.1	46.2	12.8
10	0.1	0.0	0.1	0.1	0.2	1.1	1.1	3.5	14.9	26.2	52.8
Total	0.1	0.2	0.4	0.7	1.1	3.9	5.9	19.6	34.7	22.3	11.2

Source. The Household, Income and Labour Dynamics in Australia (HILDA), 2001-2017.
Sample. Unbalanced panel of 24,305 individuals with at least two consecutive observations.

Table 12: Dynamic ordered Logit model of life satisfaction in Australia

Dependent	Life satisfaction (LS_t)
Initial life satisfaction (ref=0)	
$LS_0 = 1$	-0.384
	(0.498)
$LS_0 = 2$	-0.030
	(0.359)
$LS_0 = 3$	-0.212
	(0.357)
$LS_0 = 4$	0.187
	(0.343)
$LS_0 = 5$	0.501
	(0.332)
$LS_0 = 6$	0.669**
	(0.330)
$LS_0 = 7$	1.194***
	(0.329)
$LS_0 = 8$	1.857***
	(0.329)
$LS_0 = 9$	2.589***
	(0.330)
$LS_0 = 10$	3.507***
	(0.331)
Past life satisfaction - (ref=0)	
$LS_{t-1} = 1$	0.313
	(0.266)
$LS_{t-1} = 2$	0.644**
	(0.252)
$LS_{t-1} = 3$	1.005***
	(0.243)
$LS_{t-1} = 4$	1.229***
	(0.242)
$LS_{t-1} = 5$	1.534***
	(0.240)
$LS_{t-1} = 6$	1.811***
	(0.239)
$LS_{t-1} = 7$	2.176***
	(0.239)
$LS_{t-1} = 8$	2.631***
	(0.240)
$LS_{t-1} = 9$	3.108***
	(0.241)
$LS_{t-1} = 10$	3.765***
	(0.243)
Controls	Yes
Year dummies	Yes
Individual effects	CRE
# of individuals	24,305
# of observations	208,003
$\log(L)/N$	-1.374

Source. HILDA, 2001-2017.
Sample. Unbalanced panel of 24,305 individuals with at least two consecutive observations.
Model. Dynamic ordered Logit with correlated random effects (CRE): see Tables 4 to 6, column 3.
age, age squared, gender, income, education, labor market status, family status.
Robust standard errors clustered at the individual level.

C.2 Germany

Table 13: Life satisfaction in Germany: annual transitions

Destination → Initial ↓	0	1	2	3	4	5	6	7	8	9	10
0	21.9	8.7	11.4	11.6	6.7	19.2	5.3	5.5	6.0	1.7	2.1
1	8.7	12.1	15.0	14.9	9.4	17.0	6.2	6.9	5.0	3.3	1.7
2	3.6	5.3	15.2	17.5	11.3	18.9	8.8	8.3	8.0	2.3	1.0
3	1.6	2.4	8.2	16.9	15.0	22.9	11.9	10.9	7.8	1.7	0.8
4	0.8	1.2	4.2	10.8	16.2	26.2	15.8	14.2	8.5	1.7	0.5
5	0.7	0.6	2.0	5.0	7.9	33.3	17.5	17.4	12.2	2.2	1.3
6	0.2	0.3	1.0	2.7	5.0	18.2	23.6	28.2	16.9	2.7	1.2
7	0.1	0.1	0.5	1.5	2.3	9.3	14.4	35.7	29.8	4.9	1.4
8	0.1	0.1	0.3	0.7	1.0	4.8	6.5	22.0	47.5	13.7	3.3
9	0.1	0.1	0.2	0.4	0.6	2.4	3.0	9.8	37.7	36.7	9.1
10	0.2	0.1	0.3	0.5	0.6	3.7	2.6	6.8	22.5	23.9	38.8
Total	0.4	0.4	1.2	2.6	3.5	11.7	11.2	22.1	30.6	11.5	4.7

Source. The German Socio-Economic Panel (GSOEP), 1984-2017.
Sample. Unbalanced panel of 57,637 individuals with at least two consecutive observations.

Table 14: Dynamic ordered Logit model of life satisfaction in Germany

Dependent	Life satisfaction (LS_t)
Initial life satisfaction (ref=0)	
$LS_0 = 1$	-0.090
	(0.173)
$LS_0 = 2$	-0.047
	(0.137)
$LS_0 = 3$	0.108
	(0.124)
$LS_0 = 4$	0.316***
	(0.119)
$LS_0 = 5$	0.514***
	(0.114)
$LS_0 = 6$	0.744***
	(0.115)
$LS_0 = 7$	1.063***
	(0.114)
$LS_0 = 8$	1.526***
	(0.114)
$LS_0 = 9$	2.071***
	(0.115)
$LS_0 = 10$	2.472***
	(0.117)
Past life satisfaction (ref=0)	
$LS_{t-1} = 1$	0.198**
	(0.096)
$LS_{t-1} = 2$	0.531***
	(0.081)
$LS_{t-1} = 3$	0.798***
	(0.079)
$LS_{t-1} = 4$	1.049***
	(0.078)
$LS_{t-1} = 5$	1.377***
	(0.077)
$LS_{t-1} = 6$	1.664***
	(0.078)
$LS_{t-1} = 7$	2.026***
	(0.078)
$LS_{t-1} = 8$	2.454***
	(0.078)
$LS_{t-1} = 9$	2.944***
	(0.080)
$LS_{t-1} = 10$	3.477***
	(0.082)
Controls	Yes
Year dummies	Yes
Individual effects	CRE
# of individuals	57,637
# of observations	469,408
$\log(L)/N$	-1.615

Source. GSOEP, 1984-2017.
Sample. Unbalanced panel of 57,637 individuals with at least two consecutive observations.
Model. Dynamic ordered Logit with correlated random effects (CRE): see Tables 4 to 6, column 3.
age, age squared, gender, income, education, labor market status, family status.
Robust standard errors clustered at the individual level.

C.3 The UK

Table 15: Life satisfaction in the UK: annual transitions

Destination → Initial ↓	1	2	3	4	5	6	7
1	24.1	16.0	8.7	7.6	6.7	19.9	17.0
2	6.5	20.3	14.8	10.0	12.7	31.3	4.3
3	2.6	11.1	28.9	18.2	20.4	17.2	1.7
4	1.7	5.8	15.4	28.2	23.8	22.3	2.7
5	0.8	4.3	9.8	12.8	31.3	37.7	3.4
6	1.0	3.7	3.0	4.5	14.1	63.9	9.8
7	3.4	2.3	1.3	2.5	5.4	40.5	44.7
Total	2.3	5.6	7.9	9.3	17.2	46.5	11.2

Source. The United Kingdom Understanding Society (UKUS) survey, waves 1 to 8.

Sample. Unbalanced panel of 54,593 individuals with at least two consecutive observations.

Table 16: Dynamic ordered Logit model of life satisfaction in the UK

Dependent	Life satisfaction (LS_t)
Initial life satisfaction (ref=1)	
$LS_0 = 2$	0.234***
	(0.083)
$LS_0 = 3$	0.156**
	(0.078)
$LS_0 = 4$	0.523***
	(0.077)
$LS_0 = 5$	0.931***
	(0.077)
$LS_0 = 6$	1.691***
	(0.077)
$LS_0 = 7$	2.745***
	(0.082)
Past life satisfaction (ref=1)	
$LS_{t-1} = 2$	0.120**
	(0.054)
$LS_{t-1} = 3$	-0.061
	(0.053)
$LS_{t-1} = 4$	0.098*
	(0.053)
$LS_{t-1} = 5$	0.309***
	(0.053)
$LS_{t-1} = 6$	0.688***
	(0.055)
$LS_{t-1} = 7$	1.297***
	(0.062)
Controls	Yes
Year dummies	Yes
Individual effects	CRE
# of individuals	54,593
# of observations	213,256
$\log(L)/N$	-1.403

Source. UKUS, waves 1 to 8.
Sample. Unbalanced panel of 54,593 individuals with at least two consecutive observations.
Model. Dynamic ordered Logit with correlated random effects (CRE): see Tables 4 to 6, column 3.
Controls: age, age squared, gender, income, education, labor market status, family status.
Robust standard errors clustered at the individual level.

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