The Fade-Away Effect of initial nonresponse in panel surveys: Empirical results for EU-SILC

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Abstract

Nonresponse in surveys may result in a distortion of the distribution of interest. In a panel survey the participation behavior in later waves is different from the participation behavior at the start. With register data that cover also the information for non-respondents one can observe a fade away of the distributional differences between the distribution of the full sample, including nonresponders, and the respondent sample, without the nonrespondents.

The mechanics of this effect may be explained by a Markov chain model. Under suitable regularity conditions the distribution on the state space converges to the steady state distribution of the chain, which is independent from the starting distribution of the chain. Therefore the fade-away effect is considered here as the swing-in into the steady state distribution.

A sufficient condition for the fade-away effect assumes the same transition law for the responders and the nonresponders. Such a hypothesis is investigated here for the Finnish subsample of EU-SILC for the equivalized household net-income. The income is grouped into income brackets which divides the starting sample into quintiles. This analysis is based on register information. For this analysis null-hypothesis of equal transition behavior between income quintiles for responders and nonresponders cannot be rejected. This finding restates a result for Finland for the ECHP (European Community Household Panel).

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The velocity of the swing-in into the steady state distribution depends on the stability to stay in the same income state. The stability may vary among the European countries. Therefore we investigated the transition matrices for 25 EU-SILC countries. We simulated 6 different pattern of nonresponse bias and investigated the fade-away effect across the waves 2006 to 2009. We found remarkable differences between these 25 countries. Expressed by the relative bias, i.e. bias in 2009 divided by bias at start in 2006, we found a reduction down to 26 percent for Bulgaria (foremost reduction) up to 61 percent for Finland (least reduction).

Keywords: Panel surveys, nonresponse, panel attrition, Markov chains, income mobility.

1 Introduction

Nonresponse may distort the results of survey analysis. It was a surprising result of Sisto (2003) that these distortions diminish in later waves of a panel survey. Such an observation can only be made by the use of some information on the nonrespondents. This information was available in the case of Finnish subsample of the European Community Household Panel (ECHP), where it was possible to use information for income and household composition also for the first wave nonrespondents.

Rendtel (2003) used a Markov chain approach to give a statistical explanation for this phenomenon. He coined this effect by the term "Fade-Away Effect". The core of this approach is the steady state distribution of a Markov chain. If the transition law of the Markov chain is stable over time, then (under regularity conditions) the distribution on the state space of the Markov chain converges to a stable distribution, the steady state distribution. The convergence takes place irrespective from the starting distribution. However, the transition law between successive states must be the same for responders and nonresponders.

In this paper we repeat the analysis for the Finnish subsample of EU-SILC, the European Statistics of Income and Living Conditions. The motivation is to show that the fade-away effect is not a singular result for a special panel. Furthermore, the constraints for the convergence of the Markov chain under panel attrition are investigated in detail.

The speed of the swing-in into the steady state distribution depends on the off-diagonal elements of the Markov chain. The larger these elements are the higher is the dynamic between different states and the faster is is the convergence into the steady state distribution. Thus one may expect that the swing-in process is different for the different countries that participate in EU-SILC. For that reason we estimate the transition matrices for 25 EU-SILC subsamples. However, in this case we have no information about the non-respondents and the size of a potential nonresponse bias. Therefore we use 6 different scenarios of a potential nonresponse bias and observe how fast this bias fades-away until the 4th wave of EU-SILC. Due to the rotation panel scheme of EU-SILC this is the longest duration that can be observed in this survey.

Finally we discuss the practical implications of the fade-away effect for the maintenance of a panel.

2 An analysis of responders and non-responders of the Finish subsample of EU-SILC

The purpose of this section is to investigate the stability of the findings from Rendtel (2005) and Sisto (2003) for the Finish subsample of ECHP, where the fade-away effect was investigated for the disposable equivalized household income and other measures of income disparity, like the quintile ratio, the percent poor and the Gini-coefficient.

For that purpose Junes (2012) investigated the 2006 rotation quarter of the Finish subsample of EU-SILC. This rotation group remained in SILC until wave 2009[1]. Of these 2353 persons 584 persons, about 25 percent, did refuse to participate in SILC. It is possible to compute the disposable household income for the entire sample. In order to make comparisons compatible across time and across persons with a different household composition the disposable household income is divided by the OECD weights, which are 1 for the head of the household, 0.5 for every additional adult and 0.3 for every child under 14 years. This results in the equivalized disposable income.

As we are interested in a discrete Markov chain framework we use income brackets to create 5 income states. The income brackets are chosen such that the full sample[2] with 2353 persons are separated into 5 income quintiles[3] Finally, we we had to correct the bracket limits for inflation to avoid a trend in the distribution on the income states. Here we used the increase of the median, which was 1.16 from wave 2006 to wave 2009. All income brackets boundaries are multiplied by the reciprocal of the corresponding inflation ratio[4].

Table 1 compares the distribution on the income brackets for the full sample and on the respondent part of the sample[5] that participated in wave one of the ECHP (1996) or SILC (2006). By construction the distribution for the full sample is identical for both surveys. While we have a virulent under-representation of high incomes in the ECHP, there is virtually no bias in the SILC survey. The reason for this different behavior will be probably due to a different organisation of the field work: while the ECHP questionnaire was run as a separate survey meaning some extra respondent burden, the SILC questionnaire was completely integrated into the general Finnish income survey, which is a well established survey. This raises the question whether the superiority of SILC with respect to ECHP holds also for other countries. With the data at hand, we cannot answer this question for SILC. With respect to the ECHP Ehling/Rendtel (2003) have compared newly started panels with ongoing panels. They found lower attrition

[1] As the register income is based on taxation records they refer to the previous year. For this reason Junes (2012) refers to the income years 2005 to 2008.

[2] We use the term full sample in accordance with our notation for the ECHP. Alternatively one may use the term gross-sample of wave 1.

[3] There were two alternatives here: One is to chose the income brackets from a design-based analysis. In this case the intervals refer to the population quintiles. Or, we may refer to the level of the full sample without using the sample weights. In this case we refer to the pure biasing effect of the nonresponse at the start of the panel. As the survey weights use calibration techniques that try to compensate for a nonresponse bias, the use of the survey weights might obscure the biasing effect of nonresponse, which is the issue here. For that reason we did use the unweighted results in our analysis of the Finnish subsample.

[4] There are apparent other methods of deflation, for example, one might use a consumer price index. However, deflation is not the main topic of this paper and we simply want to avoid trends in the distribution of the income states.

[5] This part of the sample is often referred as the net-sample.

Table 1: Comparison of the initial bias for income quintiles in the ECHP and SILC

Sample	FULL	ECHP	SILC
Quintile	Sample	RESPONDENT Sample	RESPONDENT Sample
1	20.0	21.8	19.3
2	20.0	20.7	20.1
3	20.0	21.8	20.0
4	20.0	20.1	20.5
5	20.0	15.6	20.1
Results from Junes (2012) and Rendtel (2005)			

rates for the ongoing panels.

In the Finnish SILC sample it is not worth to think about a fade-away effect. Therefore in a later analysis we simulated a substantial nonresponse bias and studied the velocity of the swing-in into the steady state distribution.

3 Regularity conditions for the Markov chain and the attrition process

We now have to be precise about the conditions under which first wave nonresponse R_1 and attrition in later panel waves, say R_2, R_3, R_4, where $R_t = 1$ indicates response and $R_t = 0$ indicates nonresponse at wave t, can be ignored. The distribution on the income states at wave 4 in the observed sample at wave 4 is $P(Y_4|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1)$. The fade-away hypothesis assumes:

$$P(Y_4|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1 \approx P(Y_4) \tag{1}$$

Now we have:

$$\begin{aligned}
& P(Y_4 = i|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \\
= & \sum_{j_3} P(Y_4 = i|Y_3 = j_3, R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \\
\times & P(Y_3 = j_3|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \qquad (2) \\
= & \sum_{j_3} P(Y_4 = i|Y_3 = j_3, R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \\
\times & \frac{P(R_4 = 1|Y_3 = j_3, R_1 = 1, R_2 = 1, R_3 = 1)}{P(R_4 = 1|R_1 = 1, R_2 = 1, R_3 = 1)} \\
\times & P(Y_3 = j_3|R_1 = 1, R_2 = 1, R_3 = 1) \qquad (3)
\end{aligned}$$

In order to proceed have to assume that the transition behavior does not depend on the participation behavior (**Assumption A**):

$$P(Y_4 = i|Y_3 = j_3, R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) = P(Y_4 = i|Y_3 = j_3) \tag{4}$$

Assumption A is equivalent to the missing at random (MAR)assumption, which states that the probability of response must not depend on the unobserved value.

Furthermore we need **Assumption B** stating that the previous income state does not have an additional effect on the participation in the present wave:

$$P(R_4 = 1|Y_3 = j_3, R_1 = 1, R_2 = 1, R_3 = 1) = P(R_4 = 1|R_1 = 1, R_2 = 1, R_3 = 1) \tag{5}$$

By using assumptions **A** and **B** one gets:

$$\begin{aligned} & P(Y_4 = i|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \\ = & \sum_{j_3} P(Y_4 = i|Y_3 = j_3) P(Y_3 = j_3|R_1 = 1, R_2 = 1, R_3 = 1) \end{aligned} \tag{6}$$

Using the same kind of analysis for $P(Y_3 = j_3|R_1 = 1, R_2 = 1, R_3 = 1)$ and inserting into eq. 6 one obtains:

$$\begin{aligned} & P(Y_4 = i|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \qquad (7) \\ = & \sum_{j_3, j_2} P(Y_4 = i|Y_3 = j_3) P(Y_3 = j_3|Y_2 = j_2) P(Y_2 = j_2|R_1 = 1, R_2 = 1) \end{aligned}$$

Finally we arrive at:

$$\begin{aligned} & P(Y_4 = i|R_1 = 1, R_2 = 1, R_3 = 1, R_4 = 1) \\ = & \sum_{j_3, j_2, j_1} P(Y_4 = i|Y_3 = j_3) P(Y_3 = j_3|Y_2 = j_2) P(Y_2 = j_2|Y_1 = j_1) \\ \times & P(Y_1 = j_1|R_1 = 1) \end{aligned} \tag{8}$$

where the last term $P(Y_1 = j_1|R_1 = 1)$ is the starting distribution for the respondents of wave 1 and the summation is done over 3 cycles of the Markov chain between the income states.

Assumptions **A** and **B** and the corresponding expressions for wave 3 and 2 may be regarded as restrictive, as they state that attrition must not be linked to transition between income states nor to the income state of the previous period. However, these statements have to be fulfilled by only those persons who have participated already 3, 2 or at least one year in the panel. Previous experience from other panels states a very different participation behavior in the first wave and in the subsequent waves. Behr et al. (2005) found for the ECHP that the most important variables for panel attrition are related to field-work, like change of the interviewer in an interviewer-based panel, or item nonresponse to sensitive questions. Also residential mobility, which is clearly linked to field-work, is an important variable. These results hold for all European countries in the ECHP. In many instances these field-work related variables will be independent from the income state. These findings suggest similar results also for the other European participants of SILC. Furthermore the attrition rates between waves tend to be small when compared to the initial nonresponse. Junes (2012) reports attrition rates of 8 % (wave 2), 7 % (wave 3) and 5 % (wave 4) which compare to 30 % in wave 1 in the case of the Finnish SILC.

Table 2: Transition rates in percent between income states. Upper panel: transitions for wave 1 respondents, lower panel: transitions for wave 1 non-respondents

Quintile	Responders				
1	76.5	16.2	4.4	2.1	0.7
2	15.7	57.6	19.1	5.7	1.8
3	4.6	17.2	51.4	22.9	3.9
4	3.0	5.9	16.1	58.9	16.1
5	2.8	1.2	3.3	14.0	78.6
Quintile	Non-responders				
1	73.9	17.9	5.0	2.1	1.0
2	16.8	58.4	17.1	5.8	1.7
3	4.2	16.7	55.9	18.5	4.6
4	1.2	5.5	15.7	63.9	13.7
5	3.7	2.0	3.9	10.1	79.4

Usually the hypothesis of the independence of attrition and changes between income states cannot be tested, because the income state at the current wave is not known for the attriters. However with register data, we can check such a hypothesis directly. For that purpose we pooled the transitions of the wave 1 non-responders over the panel waves. The estimated transition matrix is given in the lower panel of Table 2.

A likelihood ratio test[6] on differences of the transition matrices between the two groups resulted in $2*(-12189.03+12197.07) = 16.06$ with $5*4 = 20$ degrees of freedom. This results in a p-value of 0.72. Hence the null-hypothesis of equal transition matrices cannot be rejected[7].

A comparison of the income distribution of those who participated in wave 1 (sample RESP) and those who participated in the last wave (sample OBS) may reveal a possible attrition effect. Table 3 compares the two distributions for the ECHP (last wave =wave 5) and EU SILC (last wave= wave 4). While for the ECHP we see only minor discrepancies between the two distributions, the findings for SILC might indicate an attrition effect with an over-representation of the above median incomes and under-representation of low incomes. As the the OBS sample is a subset of the RESP sample standard test-routines for systematic differences don't apply here. So the conclusion of Junes (2012) of a manifest attrition effect still needs some statistical underpinning.

A substantial part in our model is the time homogeneity of the Markov chain to achieve a steady state distribution after a sufficient number of transitions. The deflation was an essential tool in this context. From a subject matter point of view there is no reason for a sudden change in the transition law between income states in Finland. A formal test, which checks the equality of the three transition matrices, gives a likelihood ratio of $2*(-12169.7+12197.1) = 54.8$ with $2*20 = 40$ degrees of freedom. The corresponding p-value is 0.06. The separate estimates of the transition probabilities, however, don't exhibit a meaningful trend over time, see the Appendix. Therefore we will not reject the hy-

[6]The computations were done with the *l*em package of Vermunt (1997)

[7]An analysis on the impact of pre-wave income position on attrition is missing here. The format of the analysis might be: all participants of wave t-1. Dependent variable R_t. Independent variable Y_{t-1}. One may pool the analysis over the waves t=2,3,4.

Table 3: Comparison of the distribution on income states for the three samples FULL (All selected persons wave 1), RESP (All respondents wave 1) and OBS (All observed persons in last wave)

	ECHP Sample			EU-SILC Sample		
Quintile	FULL 14616	RESP 7809	OBS 5192	FULL 2353	RESP 1769	OBS 1448
1	23.9	22.2	22.4	20.4	20.5	18.9
2	16.9	16.6	17.4	19.8	19.3	18.7
3	18.3	17.9	17.6	18.7	18.2	18.1
4	20.6	21.4	21.8	21.1	21.7	22.2
5	20.4	22.0	20.9	20.1	20.4	22.1
Results from Junes (2012) and Rendtel (2003)						

pothesis of time homogeneous transition matrices.

Table 4 compares the estimates of the transition matrices between the income states for the ECHP and SILC. Furthermore the last column of Table 4 displays the steady state distribution of the the corresponding Markov chains. A steady state distribution exists in both cases as the chains are irreducible and aperiodic. A sufficient condition is that all states can be reached within one step from all other states, i.e. all entries of the transition matrix P are strictly positive. The steady state distribution π can be characterized by an eigenvector equation for the eigenvalue 1:

$$P'\pi = 1 * \pi \qquad (9)$$

Equation 9 states that the distribution π is not changed by one transition of the Markov chain.

Although there are some differences between the transition tables of the ECHP (period 1996 to 2000) and SILC (period 2006 to 2009) the steady state distributions in Table 4 are quite similar. The largest difference is a shift from the lower quintile position, which has become less frequent, to the highest quintile position, which has become more frequent. Table 4 exhibits an interesting structural pattern, that is displayed for the ECHP and SILC: the above diagonal elements are larger than the corresponding entries of the below diagonal elements[8]. This means that one step increases are more frequent than one step decreases. This pattern is well reflected for both transition tables that refer to different decades. Once again this points to some stability in the transition behaviour over time.

4 The design of the simulation study

The distribution of income states is one of the core variables of EU-SILC. Here we use the equivalized household income which establishes comparability over households with different composition. In order to establish comparability over time we have also to use some deflation of the income brackets that define the income classes. The results are taken from Dietz (2012), who used the

[8]The only exception is the transition from the lowest to the second quintile.

Table 4: Comparison of the transition probabilities in the ECHP and in SILC

quintiles	P for ECHP					steady state distribution
	1	2	3	4	5	
1	72.2	18.3	5.4	2.5	1.6	23.9
2	20.6	49.9	21.4	6.3	1.9	17.9
3	6.9	16.7	49.1	23.2	4.1	18.2
4	4.5	5.1	16.3	57.1	17.0	20.9
5	4.0	2.6	4.0	16.0	73.4	18.9
quintiles	P for SILC					steady state distribution
	1	2	3	4	5	
1	75.8	16.6	4.5	2.1	0.8	20.9
2	16.0	57.8	18.6	5.8	1.8	19.1
3	4.5	17.1	52.6	21.8	4.0	18.1
4	2.6	5.8	16.0	60.1	15.5	20.8
5	3.0	1.4	3.5	13.3	78.9	21.2

consumer price index for deflation[9] In order to facilitate the comparison across countries we used here the cross-sectional design weights to establish national quintile intervals. For the longitudinal analysis we used the longitudinal SILC weights[10].

The following analysis uses 25 sub-samples based on the EU-SILC User Data Base, taking into account 23 EU Member States[11] plus Norway and Iceland.

Figure 1 displays some characteristics of these 25 states included in this analysis. The case numbers refer to the longitudinal cohorts that participate over 4 waves (2006 – 2009). There is an apparent variation of the median income and the Gini-coefficient.

Figures 2 and 3 display the estimated transition matrices between the successive income quintiles. The time period covers the years 2006 to 2009. There are remarkable differences between the transitions matrices within the EU. For example, the probability to stay in the lowest quintile ranges from 0.39 for Latvia (LV) to 0.84 for Cyprus (CY). With respect to the highest income quintile Slovenia (SI) is the most stable country with a probability of 0.90 while Island (IS) has become the most risky country for high incomes with a probability of 0.57. Island is also a country where transitions to the next lower quintile are more frequent than transitions to the next higher quintile position. Just the opposite pattern can be found in Norway (NO). Here the risk to reach the next higher quintile is always higher than the risk to fall down one position.

[9]This differs from the approach in the previous section where we used the ratio of the medians of the equivalence income. To our knowledge the effect of different deflations very small.

[10]These methodological differences may explain the different values for Finnland, where the results of the previous section will differ from the transition matrix displayed in Figure 2 below.

[11]Some EU Member states, for example Germany, are missing as they do not provide longitudinal data for this data base.

Cntry	N	Median	Min	Max	Gini
SE	4 173	19 857	1	268 782	21.49
DK	3 714	25 347	614	284 042	22.49
SI	9 410	9 237	128	43 413	23.89
NL	5 572	17 958	166	155 200	24.62
CZ	8 646	4 743	269	93 068	25.46
AT	4 964	17 273	11	114 750	25.60
FI	4 825	18 008	154	422 234	26.43
NO	9 124	27 488	1	3 679 263	26.87
BE	4 916	16 867	67	308 000	27.00
FR	18 175	16 250	20	193 690	27.39
MT	3 374	9 239	155	113 006	27.40
IS	1 772	28 458	1 323	529 424	27.75
LU	10 403	29 433	7	1 118 114	27.77

Cntry	N	Median	Min	Max	Gini
ES	10 047	11 101	3	116 807	29.86
BG	3 038	1 391	160	16 478	30.18
CY	2 853	14 575	4 929	297 977	30.42
HU	6 522	3 794	106	63 370	30.99
IT	15 304	14 727	50	218 733	31.80
UK	7 591	19 735	59	325 666	32.27
LT	4 173	2 574	25	23 376	33.44
PL	12 305	3 154	44	30 640	33.86
EE	4 615	3 655	137	35 285	34.12
GR	4 702	9 174	367	133 800	35.36
PT	3 035	7 153	193	123 899	36.76
LV	3 598	2 637	37	51 025	38.04

Figure 1: No. of Observations, Median, Min and Max in Euros and Gini coefficient of net-equivalence income (2006), (Source: EU-SILC, Calculations taken from Dietz (2012))

AT				
58.84	23.97	8.84	4.59	3.76
17.52	45.47	21.91	11.20	3.90
4.58	19.10	38.98	26.90	10.44
3.00	8.29	18.11	41.31	29.29
2.62	3.02	5.59	20.21	68.57

BE				
66.89	21.73	6.62	2.65	2.12
13.92	53.27	23.87	5.78	3.16
4.24	15.68	51.24	22.22	6.61
1.62	4.71	18.15	55.90	19.62
1.49	3.12	3.62	15.70	76.08

BG				
47.27	25.88	11.88	8.25	6.72
15.86	36.74	25.69	13.80	7.92
5.84	12.33	30.34	25.08	26.41
3.10	7.29	14.13	33.03	42.45
2.44	2.44	3.33	12.37	79.42

CY				
84.23	11.72	3.37	0.53	0.15
32.96	37.95	25.99	2.80	0.30
23.9	20.73	35.27	17.65	2.45
3.06	20.16	19.35	45.11	12.32
1.34	2.66	9.32	18.4	68.27

CZ				
46.00	28.29	15.35	6.87	3.49
7.81	28.37	39.34	17.98	6.50
2.07	7.32	32.58	43.28	14.75
0.59	1.75	7.82	40.10	49.75
0.14	1.06	2.04	7.42	89.34

DK				
68.16	23.51	4.73	2.14	1.45
10.10	60.36	23.47	4.35	1.72
2.51	14.96	42.37	33.26	6.90
2.71	2.39	18.93	43.25	32.72
1.22	0.81	3.77	13.25	80.95

EE				
53.21	29.41	12.20	3.74	1.44
9.71	43.71	31.25	12.71	2.62
3.17	13.03	39.35	38.30	6.16
1.22	4.83	12.24	53.79	27.92
0.95	0.61	2.05	19.51	76.88

ES				
54.01	25.18	13.56	4.44	2.82
20.53	35.66	27.11	12.41	4.29
8.69	18.93	36.26	25.57	10.55
3.30	7.31	18.01	41.61	29.77
1.48	2.22	4.46	15.05	76.80

FI				
66.37	25.66	4.11	3.03	0.84
10.14	53.69	25.04	9.71	1.42
1.010	13.08	53.37	28.98	3.56
1.65	2.13	11.74	56.43	28.05
0.91	0.70	0.91	7.37	90.11

FR				
54.76	29.75	7.77	3.85	3.86
13.00	47.88	25.81	9.87	3.45
3.67	12.28	46.03	30.03	7.99
1.98	3.37	12.58	54.34	27.72
1.34	1.09	3.07	11.40	83.10

GR				
57.72	25.81	10.85	4.43	1.19
15.49	53.36	25.56	5.10	0.49
5.08	13.19	54.99	23.85	2.89
1.72	2.65	15.98	61.52	18.14
1.23	1.26	3.93	16.79	76.79

HU				
55.74	27.77	10.00	5.53	0.96
12.32	41.75	33.36	9.8	2.77
4.16	18.01	40.73	30.02	7.08
2.12	5.00	17.45	48.93	26.50
1.65	3.51	6.87	16.07	71.90

Figure 2: Transition matrix between income quintiles for 12 EU-Member states (Source: EU-SILC, values taken from Dietz (2012))

IS				
79.04	10.60	6.93	1.38	2.05
40.87	33.79	16.72	6.05	2.57
25.73	22.58	29.85	17.15	4.69
13.05	17.16	22.25	29.75	17.78
5.52	5.29	11.31	20.53	57.35

IT				
68.35	21.98	6.72	1.96	0.99
13.03	55.91	23.40	4.79	2.87
5.06	14.10	51.93	22.83	6.08
1.59	4.05	13.85	53.27	27.24
0.85	1.92	5.37	16.37	75.48

LT				
43.76	32.77	14.23	6.89	2.35
6.82	38.73	36.96	14.69	2.80
3.61	9.26	42.08	37.74	7.31
0.69	2.52	13.02	47.82	35.94
0.09	0.86	3.02	9.22	86.80

LU				
77.4	17.49	3.04	0.99	1.07
20.07	53.26	21.07	4.14	1.45
4.43	15.94	53.47	20.00	6.16
1.41	4.02	19.65	53.59	21.33
1.07	1.84	3.78	16.44	76.87

LV				
39.11	32.00	15.37	12.05	1.46
3.50	37.91	34.01	17.67	6.91
1.53	7.52	34.51	39.66	16.79
1.01	4.94	7.30	31.13	55.62
0.37	1.25	4.26	7.15	86.98

MT				
61.18	25.85	9.44	2.41	1.11
21.34	44.98	25.67	5.27	2.74
6.54	18.83	49.77	19.45	5.41
3.19	8.34	14.83	52.85	20.79
0.76	2.23	6.92	20.9	69.2

NL				
66.26	24.17	5.62	2.68	1.27
9.09	57.83	25.03	4.36	3.70
3.35	6.70	58.88	27.96	3.11
1.35	1.22	10.84	60.14	26.44
0.55	0.91	2.28	8.63	87.63

NO				
63.21	23.08	7.62	4.26	1.83
9.67	47.99	29.56	8.10	4.68
3.18	9.19	47.21	34.87	5.54
2.48	3.03	9.97	53.32	31.20
1.98	1.28	2.75	8.83	85.16

PL				
50.54	33.66	9.97	3.95	1.89
9.62	46.11	30.77	10.55	2.95
1.99	10.42	40.53	36.17	10.89
1.02	2.77	11.72	49.67	34.81
0.69	0.79	2.70	9.54	86.28

PT				
65.38	24.54	6.54	2.78	0.76
13.07	45.92	29.72	9.91	1.39
5.34	16.56	46.85	27.77	3.48
1.37	4.53	11.41	66.4	16.29
0.18	0.97	1.42	9.94	87.49

SE				
70.37	15.26	7.58	4.52	2.28
13.82	56.3	22.87	4.31	2.71
2.67	13.6	48.77	30.14	4.83
0.97	5.34	13.15	54.5	26.04
1.08	0.69	3.2	10.81	84.22

SI				
64.47	25.98	6.83	2.23	0.50
9.55	45.55	33.6	9.53	1.78
1.13	9.61	46.08	35.49	7.69
1.22	1.86	8.96	53.96	34.00
0.35	0.59	0.82	7.87	90.37

UK				
59.1	27.72	7.40	3.33	2.45
25.56	47.11	19.04	5.83	2.45
9.03	27.19	44.06	16.11	3.61
3.58	9.21	28.19	45.25	13.77
3.38	3.47	10.48	20.14	62.53

Figure 3: Transition matrix between income quintiles for 11 EU-Member states plus NO and IS (Source: EU-SILC, values taken from Dietz (2012))

With the SILC user data base (UDB), the initial non-response bias can not be observed, therefore we decided to run a simulation experiment to demonstrate how fast the distribution on the quintiles swings back into the steady state distribution. For the transitions between the quintiles we assume the assumptions A and B.

The nonresponse bias is simulated by using six different distorted initial distributions, which are displayed in Figure 4. Scenario 1 (first row of Table in Figure 4) is exactly the situation of the first wave of the Finnish subsample of the ECHP described in Table 1. In Scenario 2 we assume the starting distribution to be even more skew. Scenario 3 is even more pronounced. Here the two lowest income quintiles are over-represented while the two highest quintiles are under-represented. Scenario 4 just reverses the situation: here the high income quintiles are over-represented while the two lowest quintiles are under-represented. Scenario 5 displays a non-monotonic response behaviour. Here we have an over-representation of the middle quintiles. Finally, Scenario 6 refers to a situation where the extreme quintiles are over-represented.

Scenario	Income brackets				
	1	2	3	4	5
1	0.218	0.207	0.218	0.201	0.156
2	0.235	0.200	0.225	0.210	0.130
3	0.320	0.250	0.190	0.150	0.090
4	0.135	0.165	0.215	0.225	0.260
5	0.150	0.225	0.240	0.225	0.160
6	0.300	0.160	0.100	0.150	0.290

Figure 4: Simulation of six different distorted initial distributions in 2006

5 Results of the simulation study

Of particular interest is the speed of the fade-away effect. We are interested whether the swing-in process from the initial distributions is fast enough to give reliable results within the four year period of the SILC longitudinal files. The development of the distribution on the income quintiles is demonstrated here for the French sub-samples under Scenario 1, see Figure 5. According to this scenario an initial bias of 4.4 percentage points in the up-most income quintile has reduced to 2 percentage points three waves later.

In the following tables we will use a global measure over all income states. For this purpose we use the absolute bias B, which is $B = \sqrt{\sum_q b_q^2}$ where b_q is the bias of the quintile q. The absolute bias has therefore a direct interpretation on the scale of percentage points. Therefore, the absolute bias is equal for all countries is in the base year 2006 in each of the 6 scenarios. The value differs between the scenarios. It's value is respectively 0.0513, 0.0828, 0.1778, 0.0995 0.0834, 0.1794. Figure 6 displays the decrease of absolute bias with each panel wave. Until wave 4 in 2009 we observe for all subsamples substantial decreases of this absolute bias. However, the speed of the fade-away process

2006 (t=0)		2007 (t=1)		2008 (t=2)		2009 (t=3)	
$\Pi_{0,F}(i)$	$\Pi_{0,R}(i)$	$\Pi_{1,F}(i)$	$\Pi_{1,R}(i)$	$\Pi_{2,F}(i)$	$\Pi_{2,R}(i)$	$\Pi_{3,F}(i)$	$\Pi_{3,R}(i)$
0.2	0.218	0.150	0.160	0.121	0.128	0.103	0.109
0.2	0.207	0.189	0.199	0.168	0.178	0.150	0.157
0.2	0.218	0.191	0.201	0.183	0.191	0.175	0.181
0.2	0.201	0.219	0.221	0.229	0.231	0.235	0.236
0.2	0.156	0.250	0.218	0.298	0.272	0.336	0.316

Figure 5: Development of the initial non-response bias on income quintiles for the French sub-sample. Left column: true distribution, right column: distribution based on biased starting distribution based on Scenario 1.(Source: EU-SILC, Calculations taken from Dietz (2012))

varies substantially between the different scenarios. For the Scenarios 3 and 6 we also see substantial differences across the different SILC sub-samples. Scenario 3 refers to a monotonic over-representation of the low incomes, while Scenario 6 refers to the under-representation of the middle income classes.

The speed of the convergence can be best displayed by the ratio of the absolute bias in two subsequent waves. This is demonstrated again for French sub-sample of SILC in Figure 7. There we see that the absolute bias decreases like a geometric sequence. The speed factor depends on the scenario: for the monotonic pattern (Scenario 1 to 4) the decrease is slower than is the two-sided pattern (Scenario 5 and 6).

Finally we compare the relative bias B_{2009}/B_{2006}. Figure 8 orders the countries according to their relative bias in Scenario 1. Here, Bulgaria is the country with the largest reduction of the absolute bias (factor 0.26), while Finland turned out to be the country with the least bias reduction (factor 0.61). The ranking of the countries with respect to this reduction factor is quite stable across the different scenarios. Figure 8 displays also for Bulgaria and Finland the absolute bias in 2009 B_{2009}. Countries where this value is smaller than 2.24 % are displayed with a grey field.

Finally Figure 9 compares the relative bias of column 1 (Scenario 1) of Figure 8 according to their geographical distribution. While there seems to be a pattern of countries with a slow fade-away effect in Northern Europe there is no clear geographical pattern with respect to medium or fast fade-away effect. However, the speed of the swing-in into the steady state distribution depends on the diagonal of the transition matrix, i.e. the probability to stay in the same income quintile. Figure 10 classifies the SILC countries according to their average stability figures which are computed by the mean of the diagonal of the transition matrix. This classification resembles the fade-away effect of Figure 9 quite well. Thus we have arrived at economic interpretation of the fade-away effect: If the regularity assumption for fade-away apply, the effect will be largest in countries with the lowest stability in income positions.

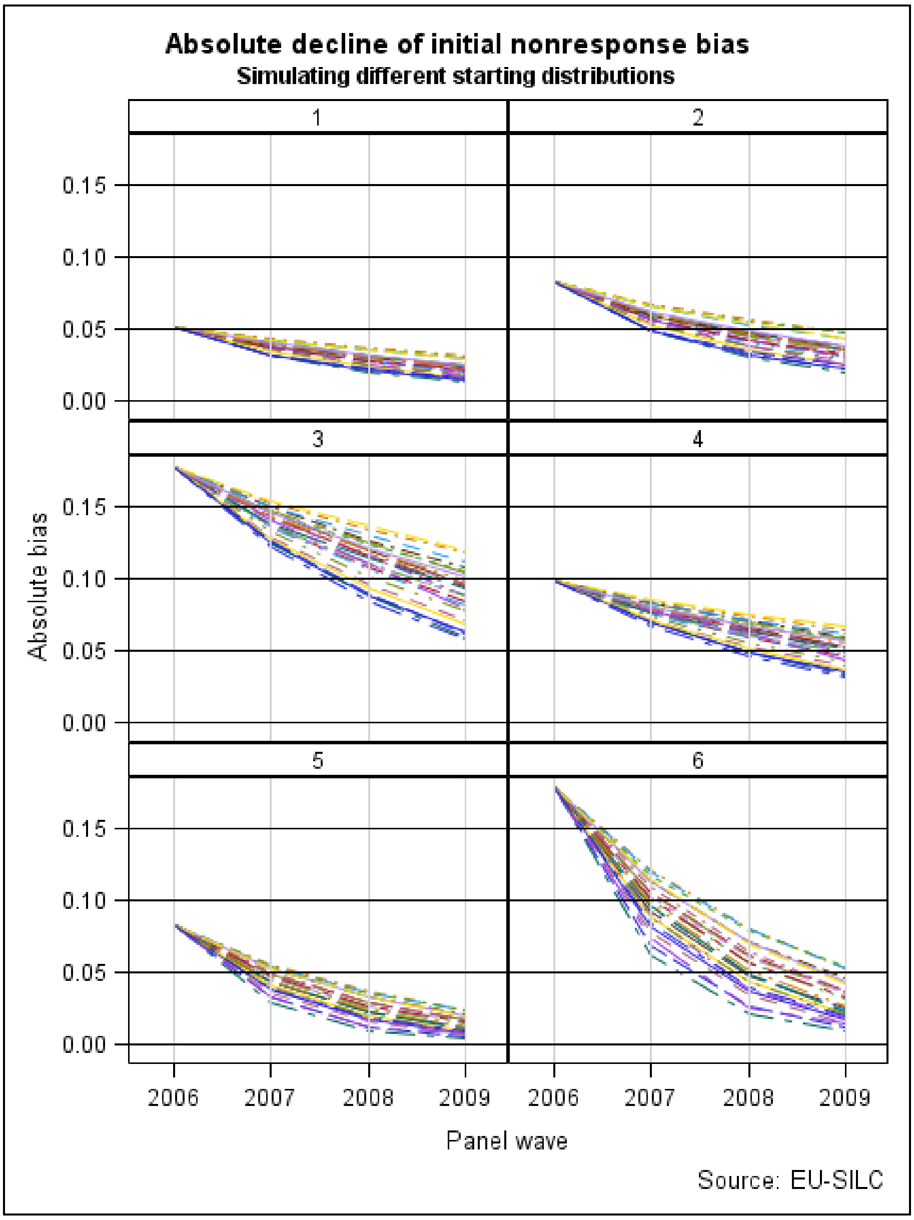

Figure 6: Decline of the absolute global bias in the national sub-samples under six different scenarios for initial distribution at wave one (2006). Each country is displayed by a different color (Source:EU-SILC, Calculations taken from Dietz (2012))

Year	Scenario					
	1	2	3	4	5	6
2007	0.75	0.71	0.77	0.77	0.55	0.56
2008	0.76	0.76	0.80	0.80	0.56	0.57
2009	0.77	0.76	0.80	0.80	0.55	0.55

Figure 7: Speed of convergence to the steady state distribution across the 6 Scenarios for the French sub-sample. Each row displays the ratio B_t/B_{t-1} where B_t is the absolute bias in wave t. (Source:EU-SILC, Calculations taken from Dietz (2012))

No.	Country	Bias in 2009 in percent of bias in 2006 by country and scenario					
		1	2	3	4	5	6
	B(2006)	*(0.0513)*	*(0.0828)*	*(0.1778)*	*(0.0995)*	*(0.0834)*	*(0.1794)*
1	BG	0.26	0.24	0.34	0.34	0.05	0.05
	B(2009)	*(0.0132)*	*(0.0201)*	*(0.0600)*	*(0.0337)*	*(0.0044)*	*(0.0090)*
2	AT	0.28	0.27	0.35	0.35	0.10	0.10
3	UK	0.29	0.28	0.33	0.32	0.11	0.11
4	IS	0.31	0.30	0.35	0.34	0.06	0.06
5	HU	0.32	0.30	0.38	0.38	0.12	0.12
6	LV	0.33	0.31	0.44	0.44	0.09	0.08
7	ES	0.36	0.34	0.40	0.40	0.08	0.08
8	MT	0.37	0.35	0.44	0.43	0.12	0.12
9	CZ	0.38	0.36	0.50	0.50	0.11	0.10
10	EE	0.39	0.37	0.46	0.46	0.14	0.14
11	BE	0.40	0.38	0.47	0.47	0.21	0.21
12	IT	0.43	0.40	0.53	0.53	0.18	0.18
13	FR	0.44	0.41	0.50	0.49	0.17	0.17
14	GR	0.45	0.42	0.50	0.49	0.19	0.20
15	LT	0.46	0.43	0.52	0.52	0.12	0.13
16	PL	0.46	0.43	0.56	0.55	0.15	0.15
17	DK	0.47	0.44	0.59	0.59	0.20	0.20
18	NO	0.47	0.44	0.54	0.53	0.21	0.21
19	LU	0.48	0.45	0.61	0.60	0.26	0.26
20	CY	0.49	0.47	0.54	0.53	0.14	0.14
21	SE	0.50	0.47	0.57	0.57	0.24	0.25
22	NL	0.57	0.53	0.63	0.63	0.28	0.29
23	SI	0.57	0.53	0.68	0.67	0.24	0.23
24	PT	0.59	0.57	0.60	0.59	0.29	0.30
25	FI	0.61	0.57	0.67	0.66	0.29	0.29
	B(2009)	*(0.0315)*	*(0.0475)*	*(0.1187)*	*(0.0653)*	*(0.0238)*	*(0.0525)*

Figure 8: Bias in 2009 as ratio of the bias in 2006 by country and scenario (Source:EU-SILC, Calculations taken from Dietz 2012)

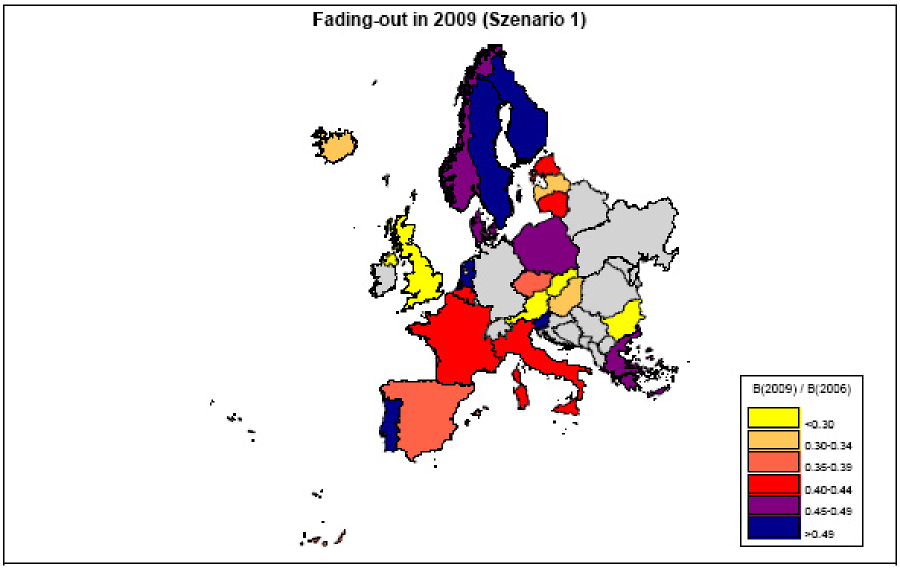

Figure 9: The stability of the relative of the relative initial bias (2006) in 2009 (Source:EU-SILC, Calculations taken from Dietz(2012))

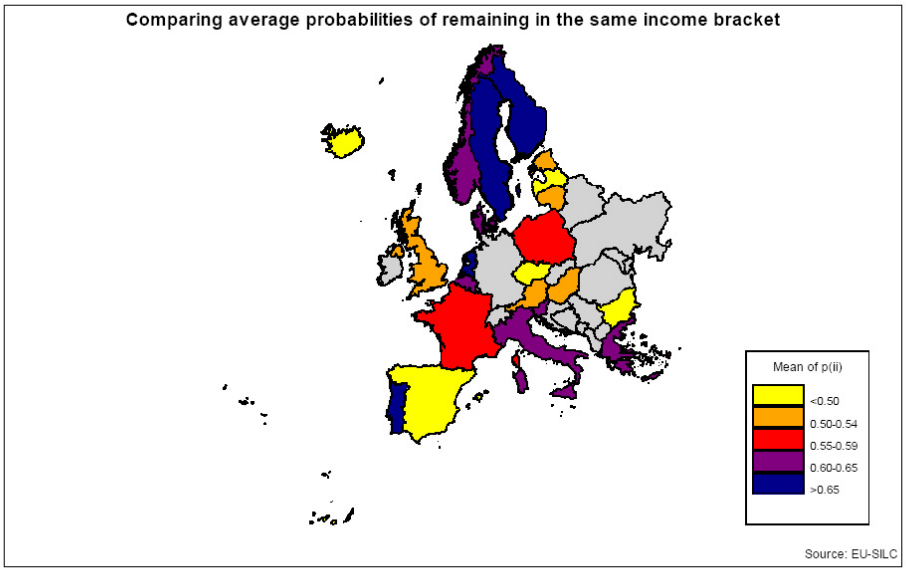

Figure 10: Classification of SILC countries according to their average stability to stay in the same income quintile. (Source:EU-SILC, Calculations taken from Dietz(2012))

6 Conclusions

The results on the fade-away effect base on an assumption that cannot be checked with the net-sample alone. Only in countries where there is a link to information from registers for the nonrespondents one can check the assumptions that are necessary to establish a fade-away effect for panels. Up to now we could verify these assumptions only for two surveys in Finland. Thus a similar analysis should be replicated for other register countries. As the size of the fade-away effect depends on the stability of the income position, it would be useful to include one register country with low income stability. Good candidates under this perspective are Island and Latvia.

Furthermore our results indicate that panel attrition is sensitive to details of the field-work, for example, whether EU-SILC is integrated in a well accepted standard survey. These details should be better documented.

Besides direct information about non-respondents via register information one could also use a comparison with census or microcensus counts that are thought to be of high reliability. This strategy was chosen by Fitzgerald et al (1998) who compared population estimates from the Panel Study of Income Dynamics (PSID), which started in 1968, with population counts in later years. The population estimates were not corrected of selective attrition. However, the χ^2-distance to the population counts decreased with increased duration of the PSID. Such an effect is to be expected under the fade-away hypothesis. Admittedly the evaluation of the PSID used longer time intervals like periods of 5 years, which are not relevant for EU-SILC.

One essential of this approach is the stability of the transition matrices with respect to time. Otherwise the existence of the steady state distribution is no longer guaranteed. The results for Finland suggest that transitions laws between income quintiles are quite stable over time. However, in the case of a sudden economic crisis, for example, like in Island in 2008, this assumption will probably not hold. But the stability of the transition law can be checked from the observed data.

Even in the case where the sufficient conditions for a fade-away effect do not hold we may observe a decrease of the initial bias in later panel waves. This may simply result from the fact, that the case numbers of attriters are relatively small with respect to initial nonresponse or the differentials of participation are small. On the other hand, the swing-in towards the steady state distribution holds for the entire sample. Thus one may end up with the observation of a fade-away effect despite some of the regularity conditions do not hold.

What are the consequences for the design of a survey like EU-SILC, which incorporates cross-sectional as well as longitudinal components? The ECHP was designed as a panel of unlimited length. It was stopped for various reasons. One argument was the cumulative size on nonresponses after wave one. It was argued that potential biases might increase as a result of cumulative nonresponse. However, this argument is wrong under the fade-away hypothesis. On the contrary, a "fresh" sample that is taken as a starting point of a new panel rotation group always incurs a "fresh" initial wave nonresponse bias. As long as this initial nonresponse bias is not successfully treated by calibration the results of the first panel waves will be probably more unreliable than the results from later panel waves.

Even if there is no apparent nonresponse bias at the start of the panel, as in the case of income of the Finnish SILC sample, one might be interested in longer observational periods. The rotational scheme of EU-SILC overs only a period of 4 years at best. This may be regarded as sufficient for the computation of the so-called Laeken indicators. However, the analytical power of longitudinal analysis increases with the duration of the panel. The technical feasibility of long running panel surveys has been well established. What matters is the interest in longitudinal results.

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7 Appendix

Separate estimation of transition probabilities between income positions:

State at		State at End				
Start	Transition	1	2	3	4	5
1	$Y_2\|Y_1$	75.1	18.1	3.8	1.9	1.0
1	$Y_3\|Y_2$	77.8	14.1	4.4	2.7	1.0
1	$Y_4\|Y_3$	74.7	17.7	5.4	1.8	0.4
2	$Y_2\|Y_1$	16.7	56.7	18.8	5.7	1.9
2	$Y_3\|Y_2$	17.6	57.2	16.7	7.3	1.1
2	$Y_4\|Y_3$	13.6	59.5	20.4	4.1	2.5
3	$Y_2\|Y_1$	4.5	15.5	50.6	24.7	4.7
3	$Y_3\|Y_2$	4.4	19.9	53.6	18.3	3.7
3	$Y_4\|Y_3$	4.6	16.2	53.6	22.2	3.7
4	$Y_2\|Y_1$	2.7	5.7	14.0	59.5	18.0
4	$Y_3\|Y_2$	2.1	3.7	19.8	59.8	14.6
4	$Y_4\|Y_3$	2.9	7.9	14.1	61.1	13.9
5	$Y_2\|Y_1$	3.4	1.5	3.4	10.2	81.5
5	$Y_3\|Y_2$	3.4	1.6	3.0	13.7	76.4
5	$Y_4\|Y_3$	2.2	1.0	4.0	15.9	76.8