Who stays poor? Who becomes poor?
Evidence from the British Household Panel Survey

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Abstract

We estimate a first-order Markov model of poverty persistence and entry rates for working-age Britons and demonstrate the importance of controlling for endogenous selection via initial poverty status and attrition. Predicted poverty transition rates reveal substantial heterogeneity in poverty transition rates, but there is also substantial genuine state dependence in poverty.

JEL codes: D31, I32, C23, C35

The philosophy of anti-poverty policy in Britain has shifted away from income supplementation of those currently poor and towards providing routes out of poverty and preventing falls into poverty. The motivation is that '[s]napshot data can lead people to focus on the symptoms of the problem rather than addressing the underlying processes which lead people to have or be denied opportunities.' (HM Treasury, 1999, p. 5.) If one takes the dynamic perspective, the salient research questions change from 'who is most likely to be poor at the moment?', to 'who is most likely to remain poor and who is most at risk of becoming poor?'. In this paper we provide new answers to these questions about poverty dynamics, using an econometric model of transition rates estimated with data from the British Household Panel Survey.

Our research has four distinctive features. First, multivariate models of poverty transitions in Britain are rare: ours is one of the first. Second, we take account of the fact that the set of individuals who are at risk of exiting poverty (or the set at risk of entering poverty) may be a non-random sample of the population, an example of an 'initial conditions' problem. Third, we model attrition. Household income data at two consecutive annual interviews (at waves t–1 and t) are not available for individuals who left the panel altogether between t–1 and t, or for those individuals at t living in a household with incomes missing for at least one household member. Our estimates of poverty transition rates allow for the potential non-random selection into the sub-sample of individuals with two consecutive incomes observed. Fourth, we provide estimates of the extent to which the experience of low income one year raises the risk of having low income in the following year ('state dependence'), while controlling for differences in observed and unobserved characteristics between individuals ('heterogeneity').

Related models have been applied to transitions into and out of low earnings. Stewart and Swaffield (1999), for example, modelled transitions controlling for the endogeneity of initial conditions and provided estimates of the degree of state dependence in low pay in Britain. Bingley *et al.* (1995) and Cappellari (1999) controlled for endogeneity in attrition as well as initial conditions in studies of Danish and Italian earnings mobility. Similar models for income do not exist, as far as we know. Most research has documented low income transition rates for different groups and used bivariate 'trigger event' methods to analyse causes (see for example Jarvis and Jenkins, 1997; Jenkins, 2000). Multivariate applications have either modelled poverty spell lengths using hazard regression techniques (see for example Devicienti, forthcoming, or Stevens, 1999, for the USA), or modelled transitions onto or off receipt of social assistance benefit rather than low income itself (see for example Böheim *et al*., 1999, or Noble *et al.*, 1998, for Britain, or Boskin and Nold, 1975, for the USA). State dependence in low income in Britain has not been studied before (Hill, 1981, is one US application).

1. An econometric model of poverty persistence rates and entry rates

We estimated a first-order Markov model of poverty persistence rates and entry rates allowing for the potential endogeneity of initial poverty status and for attrition. (See Cappellari and Jenkins, 2001, for full details.) Let P_{it} be a binary variable summarising individual *i*'s poverty status at interview *t*, equal to one if *i* is poor and zero otherwise. Similarly, let P_{it-1} summarise *i*'s poverty status at *t*–1. Define R_{it} to be a binary variable summarising retention in the sample, equal to one if *i*'s income was observed at both *t*–1 and *t*, and zero if observed only at *t*–1 (the attrition case). The model has equations for the probabilities of being poor at *t* conditional on poverty status at *t*–1 ('poverty transition'), of being poor at *t*–1, and of sample retention:

Poverty transition: $\mathrm{pr}(P_{it} = 1|P_{it-1},R_{it}) = \Phi[\{(P_{it-1})\lambda_1' + (1-P_{it-1})\lambda_2'\}Z_{it-1}]$ if $R_{it} = 1$. (1)

Initial poverty status: $\mathrm{pr}(P_{it-1} = 1) = \Phi(\beta' X_{it-1})$. (2)

Retention: $\mathrm{pr}(R_{it} = 1) = \Phi(\delta' Y_{it-1})$. (3)

Z_{it-1}, X_{it-1}, and Y_{it-1} are vectors of explanatory variables, and $\Phi(.)$ is the standard Normal cumulative distribution function. The model is parameterised so that each element of Z_{it-1} in poverty transition equation may have a different impact on poverty status at t depending on poverty status at t–1. Thus (1) provides estimates of the determinants of poverty persistence and poverty entry rates. The model also allows a simple test for genuine state dependence based on whether $\lambda_1 = \lambda_2$: if true, poverty status at t does not depend on poverty status at t–1. We allow an unrestricted correlation structure across equations. There are three correlations:

- ρ_1: the correlation between the unobservable factors affecting P_{it-1} and (P_{it} |P_{it-1}, R_{it}).
- ρ_2: the correlation between the unobservable factors affecting (P_{it} |P_{it-1}, R_{it}) and R_{it}.
- ρ_3: the correlation between the unobservable factors affecting R_{it} and P_{it-1}.

The estimate of ρ_1 provides a test of initial conditions exogeneity, and the estimate of ρ_2 a test of income retention exogeneity. The estimate of ρ_3 provides information about whether the poor are more or less likely than the non-poor to be retained in the sample, other things equal.

2. Data, definitions, and estimation

We use interview waves 1–9 (1991–9) of the British Household Panel Survey (see Taylor *et al.*, 2001, for details). Data from pairs of consecutive waves t–1 and t were pooled. The estimation sample was restricted to individuals aged 20 to 59 years who were not in full-time education; our focus is on poverty among adults of working age rather than child poverty or pensioner poverty.

Each individual's poverty status was measured using data about the income of the household to which he or she belonged. We use exactly the same definition of income as employed in the official low income statistics (Department of Social Security, 2000): post-tax post-transfer current household income, adjusted for differences in household needs using the McClements equivalence scale, in August 2000 prices (and without deducting housing costs). An individual was defined to be poor at t if he or she had an income below 60 per cent of median income at t, a poverty line that is widely used.

Because each individual's poverty status was measured using a household-level variable, all the covariates in our poverty transition equation (1) were also measured at the household level. (Standard errors were adjusted to account for the repeated observations within each household.) The covariates used refer either to the household head (age, age squared, sex, employment status, ethnic group) or to the household itself (several variables summarising household composition and work attachment). All variables were measured at the interview prior to a potential poverty transition (wave t–1).

The retention and initial poverty status equations included these same covariates, together with a number of additional variables: exclusion of these from the poverty transition equation identified the model. In particular we supposed that parental socio-economic status affected initial poverty status but not poverty transitions. (The set of indicators was similar to that employed by Stewart and Swaffield, 1999.) Three instruments were used for the retention equation: whether the respondent was an original sample member (rather than joined a panel household after wave 1), the proportion of respondents in the household who were assessed by the wave t–1 interviewer to have been 'very cooperative', and whether the interviewer changed between waves t–2 and t–1.

The first order Markov model summarised in equations (1)–(3) was estimated by simulated maximum likelihood, using a GHK-simulator to compute the trivariate cumulative Normal distribution function. The estimates of the poverty transition equations are given in the Appendix Table (available from the authors on request). See Jenkins and Cappellari (2001) for the complete model estimates, including the initial poverty status and retention equations. The instruments in the latter two equations were statistically significant.

3. Are initial poverty status and income retention exogenous to poverty transitions?

The answer is no. See Table 1 which reports the estimates of the cross-equation correlations of unobservables. Two of the three correlations were statistically significant at the one per cent level and the other statistically significant at the ten per cent level. Unobserved factors raising the chances of being poor in the first place also reduced the conditional probability of being poor. This is an analogous to a Galtonian 'regression to the mean' effect, and was also found by Stewart and Swaffield (1999) in their study of British low pay transitions. Had we ignored initial conditions endogeneity, we would have estimated poverty persistence on a sample with a conditional poverty propensity lower than the relevant population, thereby under-estimating persistence.

Those retained in the sample were more likely to make a poverty transition of either kind than those lost from the sample. If estimation had ignored attrition, poverty persistence and poverty entry rates would have been over-estimated – the retention endogeneity effect cited earlier. Judging by the yardstick of statistical significance, however, it is the differential retention effect that is more important. The estimated correlation implies that the poor are more likely to attrit from the sample. Thus using samples that exclude attritors would disproportionately exclude the poor, an exclusion that would lead to over-estimation of poverty persistence and poverty entry rates.

4. Rates of poverty persistence and poverty entry among working age adults

Who stayed poor and who became poor? According to model estimates (see Appendix Table), *poverty persistence rates* increased with household head's age and educational qualification, other things equal. And rates were significantly higher for individuals with household heads that were male, not in employment, of Pakistani or Bangladeshi ethnic origin. Rates were also higher for individuals living in a lone parent household or a multi-family household. The presence of elderly persons in the household was associated with lower poverty persistence rates, but the more children that were present (of whatever age), the greater the persistence rate. Interestingly, being in a lone parent household had no statistically significant impact (but observe that the household head's sex and work attachment were already controlled for). Nor did the number of workers in the household (though this did have a large association with initial poverty status).

The estimates suggest that there was substantial heterogeneity in poverty persistence rates. This can be seen explicitly from examination of predicted poverty persistence rates for different types of individual. See Table 2. The reference individual is a 40-year old married man with no A-levels, of European ethnic origin, who is working full-time (and the sole breadwinner in the household), with one child aged 5–11 years (type 1 in Table 2). His predicted poverty persistence rate is 0.64, slightly higher than the sample average rate, 0.58. Additional children in the family substantially raise the predicted poverty persistence rate: having two additional young children would boost the reference person's rate by ten percentage points to 0.74 (type 3). On the other hand, with no children the rate would fall to 0.57 (type 4). If there was an adult child living at home, the poverty persistence rate would fall dramatically, to only 0.42 (type 7). If the reference individual were of Pakistani or Bangladeshi rather than European ethnic origin, then the persistence rate would be 0.73 (type 9), i.e. of much the same magnitude as having two additional children.

The bottom half of the table considers instead a non-working female lone parent as the reference individual (type 10): the predicted poverty persistence rate in this case is 0.65. The effect on the poverty persistence rate of her getting a job, whether part-time or full-time, is negligible (type 11 and 12). Once one has controlled for other characteristics, work has little effect on whether she stays poor (but does have a large effect on her chances of being poor in the first place).

Consider now *poverty entry rates*. These were higher for individuals with a household head who was relatively young, had no educational qualifications, did not work full-time, or was of Bangladeshi, Pakistani or Chinese ethnic origin (see Appendix Table). Living in a lone parent household or in a household with many children was also associated with a higher poverty entry rate, but the presence of elderly persons was associated with a lower entry rate. Generally speaking, factors associated with higher poverty persistence rates were also associated with lower poverty entry rates.

It is also clear, however, that greater work attachment had a much larger impact on reducing entry rates than it had on reducing poverty persistence. For example, were the reference man to have two workers in his household rather than one, the predicted entry rate would fall by more than one third, from 0.08 to 0.05 (cf. types 1 and 6). There is a similarly large effect if the reference lone parent were to work full-time rather than not work: the predicted entry rate halves, from 0.23 to 0.11 (cf. types 10 and 12).

5. Is there genuine state dependence in low income?

Differences in personal characteristics and household circumstances are clearly associated with differences in transition probabilities. Does this leave a role for genuine state dependence in low income?

The degree of 'aggregate' state dependence is the difference between the conditional probability of being poor at *t* among those individuals who were poor at *t*–1 and the conditional probability of being poor at *t* among those who were non-poor at *t*–1. We estimate this to be 52 per cent (0.58 – 0.06: see Table 2). But this measure does not account for the effects of heterogeneity. In fact the null hypothesis of no genuine state dependence was easily rejected: the χ^2 test statistic derived from the difference between the coefficient vectors for poverty persistence and poverty entry was 590.9, with a *p*-value (d.f.=32) = 0.0000. To measure the degree of genuine state dependence we used the model to predict, for each non-attriting individual, the difference between the probability of being poor at *t* conditional on being poor at *t*–1 and the probability of being poor at *t* conditional on being non-poor at *t*–1. Our measure is the average of these predicted probability differences over all non-attriting individuals, and was estimated to equal 0.31, i.e. 58 per cent of aggregate state dependence.

While the degree of aggregate state dependence is similar to that found by Hill (1981) for the USA in the 1970s, our estimate of the extent of genuine state dependence is rather larger than hers (though she did not control for initial conditions or differential retention). Many of the sources of the scarring effect of low income are likely to lie in the labour market. See Stewart and Swaffield (1999) and Arulampalam *et al*. (2000) for further discussion.

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Table 1
Cross-equation correlations between unobserved effects

Estimated correlation	Estimate	| *t*-ratio |
corr(transition, initial poverty status)	-0.438	(6.64)
corr(initial poverty status, retention)	-0.052	(2.49)
corr(transition, retention)	0.174	(1.67)

Standard errors adjusted for repeated observations within households. Simulated maximum-likelihood estimates of the first-order Markov model described in (1)–(3). The correlations are defined in the text.

Table 2
Predicted poverty persistence and poverty entry rates

Type	Characteristics of household	Characteristics of household head	Persistence rate	Entry rate
0.	Sample average		0.58	0.06
1.	Married couple, one child aged 5–11 years, no other persons in household, 1 worker	40 years old, male, has no A-levels, in full-time work, European ethnic origin	0.64	0.08
2.	As (1), except also has child aged 3–4	As (1)	0.71	0.12
3.	As (1), except also has child aged 3–4 and child aged 0–2	As (1)	0.74	0.15
4.	As (1), except no children	As (1)	0.57	0.05
5.	As (1)	As (1), except has A-level(s)	0.59	0.05
6.	As (1), except 2 workers	As (1)	0.64	0.05
7.	As (1), except 2 workers and 2+ benefit units in household	As (1)	0.42	0.07
8.	As (1), except no workers	As (1), except not working	0.61	0.17
9.	As (1)	As (1), except head is of Pakistani or Bangladeshi ethnic origin	0.73	0.20
10.	Lone parent household, one child aged 5–11, no other persons in household, no workers	40 years old, female, has no A-levels, not working, European ethnic origin	0.65	0.23
11.	As (10), except 1 part-time worker	As (10), except works part-time	0.66	0.20
12.	As (10), except 1 full-time worker	As (10), except works full-time	0.69	0.11
13.	As (12), except also has child aged 3–4	As (10), except works full-time	0.75	0.17
14.	As (12), except also has children aged 3–4 and 0–2	As (10), except works full-time	0.78	0.21

Derived from simulated maximum likelihood estimates of the first-order Markov model described in (1)–(3).

Available from the authors on request:

Appendix Table
The correlates of poverty persistence and poverty entry rates

Covariate	Poverty persistence		Poverty entry	
	Estimate	| *t*–ratio |	Estimate	| *t*–ratio |
HoH[a] age	0.025	(1.86)	–0.055	(5.01)
HoH age squared	–0.0001	(0.93)	0.001	(4.79)
HoH is female	0.112	(1.99)	–0.066	(1.80)
HoH has A-levels	–0.116	(1.78)	–0.236	(6.86)
HoH in full-time work	0.072	(0.86)	–0.190	(3.43)
HoH in part-time work	–0.003	(0.03)	0.159	(2.45)
HoH ethnic group:				
Black Caribbean	0.255	(1.11)	–0.343	(1.31)
Black African	–0.071	(0.23)	0.040	(0.18)
Black other	0.077	(0.24)	0.163	(0.61)
Indian	0.093	(0.45)	0.073	(0.55)
Pakistani or Bangladeshi	0.253	(1.17)	0.554	(2.30)
Chinese	–0.471	(0.70)	1.127	(2.15)
Other group	0.378	(1.41)	–0.329	(1.71)
Lone parent HH[b]	0.005	(0.08)	0.258	(4.13)
'Other' HH type	0.152	(1.19)	–0.052	(0.68)
No. of workers in HH	0.008	(0.11)	–0.252	(7.76)
Elderly aged 60–75 in HH	–0.067	(0.49)	–0.305	(3.26)
Elderly aged 75+ in HH	–0.169	(0.69)	–0.937	(4.12)
Children aged 0–2 in HH	0.097	(1.44)	0.147	(2.91)
Children aged 3–4 in HH	0.171	(2.60)	0.223	(4.66)
Children aged 5–11 in HH	0.178	(3.08)	0.214	(5.77)
Children aged 12–15 in HH	0.019	(0.29)	0.190	(4.06)
Children aged 16–18 in HH	0.078	(0.57)	0.038	(0.43)
Multi-family HH	–0.518	(4.69)	0.202	(2.90)
Constant	–0.322	(1.13)	–0.031	(0.14)
ρ(transition, initial poverty)			–0.438	(6.64)
ρ(initial poverty, retention)			–0.052	(2.49)
ρ(transition, retention)			0.174	(1.67)
Number of observations		44,600		
Chi-squared (df)		6174.95 (141)		
prob>chi2		0.0000		

Standard error estimates account for repeated observations on income within households. Simulated maximum-likelihood estimates based on GHK-simulator for trivariate cumulative Normal distribution function (250 draws). [a]: HoH = head of household. [b]: HH = household. Reference categories for dummy variables: HoH male and European ethnic group, highest educational qualification less than A-level, single-family couple household, no elderly person or child present in household. Wave dummies were also included as covariates. For estimates of the equations for initial poverty status and retention, see Cappellari and Jenkins (2001).