The dynamics of social assistance benefit receipt in Britain

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Abstract

We analyze the dynamics of social assistance benefit (SA) receipt among working-age adults in Britain between 1991 and 2005. The decline in the annual SA receipt rate was driven by a decline in the SA entry rate rather than by the SA exit rate (which actually declined too). We examine the determinants of these trends using a multivariate dynamic random effects probit model of SA entry and exit probabilities applied to British Household Panel Survey data. The model estimates and accompanying counterfactual simulations highlight the importance of two factors – the decline in the unemployment rate over the period, and other changes in the socioeconomic environment including two reforms to the income maintenance system in the 1990s. The results also reveal substantial heterogeneity in SA transition rates.

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1. Introduction

For at least two decades, 'welfare to work' ideas have strongly influenced social policy thinking on both sides of the Atlantic and in English-speaking OECD nations more generally. There has been substantial interest in reorienting the design of systems of cash transfers for poor working-age families away from schemes involving arguably passive receipt of benefits ('welfare') towards schemes in which individuals are more actively involved in meeting minimum income requirements, by increasing their labour market participation ('work'). In the USA, this reorientation is illustrated by the abolition of the Aid for Families with Dependent Children program by the 1996 Personal Responsibility and Work Opportunity Reconciliation Act and introduction of the Temporary Assistance for Needy Families program with time limited benefit payments. The Earned Income Tax Credit program supporting low income working families was expanded during the 1980s and 1990s, and is a major anti-poverty policy.

Britain's Labour Government, elected in 1997, was strongly influenced by the US reforms. It was responsible for substantial extensions to the provision and generosity of in-work benefits through the Working Families Tax Credit (WFTC) program introduced in 1999 and modified and extended in 2003. Major social assistance benefit programs for working age families remain in place, however, though with some modifications in the mid-1990s that tightened eligibility requirements for unemployed people of working age.

This background raises questions such as: How much did dependence on social assistance benefits decline over the last two decades in Britain? To what extent were policy reforms responsible for observed trends and what was the role played by other factors such as changes in the availability of jobs? This paper documents what happened to social assistance benefit receipt dynamics in Britain over the period 1991–2005, and investigates the determinants of trends using an econometric model fitted to household panel data together with counterfactual simulations based on the model estimates.

To set the scene, look at Figure 1 which shows trends in receipt of social assistance benefit (SA) receipt in Britain between 2001 and 2005. (Definitions and data are explained in more detail later.) Apart from the rise in receipt at the beginning of the 1990s when Britain went into recession, the percentage of working age adults in receipt of SA halved, falling from a peak of around 12 per cent in 1993 to around 6 per cent in 2005. If the definition of SA is widened to include housing benefits, the proportion in receipt each year is consistently 2–3 per cent higher, but follows a similar downward trend. Two leading explanations for

these trends are the reforms to the benefit system intending to 'make work pay' and changes in the availability of jobs.

Figure 1 shows the substantial increase over the period in receipt of in-work cash assistance ('tax credits'). The proportion of working-age adults in receipt was consistently about 2–3 per cent during the 1990s, until the introduction of WFTC in October 1999 after which the proportion in receipt rose dramatically to almost 7 per cent in 2002. The receipt rate then rose again significantly with the extension of eligibility in 2003. Observe that the turning points in the SA receipt rate series do not correspond closely with the turning points in the series for tax credit receipt, suggesting that in-work benefit reforms were not a major driver of the former. By contrast, note the relatively close correspondence between the trends in the unemployment rate and in the SA receipt rate. These and other explanations are examined in greater detail later in the paper. We show that the story about what drove the trends is more complicated than Figure 1 suggests.

<Figure 1 near here>

Since changes in SA receipt rate from year to year reflect the combination of changes in annual rates of entry to or exit from receipt (which are processes with different determinants), we analyze entry and exit rates. Figure 2 shows the trends in these SA transition rates over the period 1991–2005. The entry rate fell from above 4 per cent in 1993 to below 2 per cent in 2005. The exit rate fell from around 40 per cent to nearly 25 per cent (the greater variability in the rate at the end of the period may simply reflect small sample sizes).

<Figure 2 near here>

We conclude that the secular decline in annual (cross-sectional) SA receipt rates was driven primarily by a decline in entry rates: the fall in the entry rate was sufficiently large that it offset the decline in the exit rate over the same period (which would increase receipt rates, other things being equal). This conclusion follows directly from the stock-flow identity that links the proportion receiving SA in year t to entry and exit rates.[1] The importance for trends in cross-sectional receipt of changes in the entry rate rather than the exit rate echo findings reported for the USA by Grogger (2004) and Haider and Klerman (2005).

To investigate the determinants of these trends, we propose a multivariate dynamic random effects probit model of individuals' annual SA entry and exit probabilities and fit it to data from waves 1–15 of the British Household Panel Survey (BHPS) covering 1991–2005.

[1] The proportion receiving SA in year t, p_t, is given by: $p_t = (1-x_t)\, p_{t-1} + e_t\,(1-p_{t-1})$, where x_t is the exit rate and e_t is the entry rate at t.

(As explained later, suitable data for later years are unavailable.) Dynamic random effects probit models have been used to examine benefit (and unemployment) dynamics in the past but we employ them in a manner that we have not seen before, exploiting the Markov properties of the model to the full. We use an extended empirical specification that allows the impact on the SA entry and exit probability of each covariate to differ, and use the model parameter estimates to derive year-by-year predictions of SA entry and exit rates, thereby assessing model fit and also, under various counterfactual scenarios, the relative importance of the several potential determinants of observed trends in transition rates. The results highlight the importance of both the decline in the unemployment rate and also other secular changes in the socioeconomic environment including two reforms to the social security benefit system. The results also point to substantial individual heterogeneity in SA transition rates.

In Section 2, we explain the benefit system in Britain over the period 1991–2005, referring to both SA and in-work benefits. We introduce the British Household Panel Survey (BHPS) data used in the analysis in Section 3, and present the statistical model in Section 4. Estimates and simulations are discussed in Section 5.

The focus throughout is on individuals of working age. More specifically, we consider only individuals below the age of 60. (The state retirement pension age in Britain is currently 60 for woman and 65 for men.) To avoid complications associated with education and training, we also exclude individuals aged less than 25, or individuals in families in which there are any adults of working age who are full-time students.

2. Benefits and tax credits in Britain, 1991–2005

Social assistance benefits are income-tested 'safety net' cash benefits, sometimes called 'welfare benefits'. They are paid to bring incomes up to some minimum income level – they refer to income maintenance. By contrast, social insurance benefits refer to income replacement – payments made in response to the occurrence of particular risky events such as sickness or unemployment and for which an appropriate record of social insurance contributions exists.[2]

[2] There are also benefits for individuals who are ill or injured that are not discussed here: Statutory Sick Pay for employees, Employment and Support Allowance (replacing Incapacity Benefit since October 2008) for those unable to work because of illness or disability and with a suitable national insurance contributions record, Industrial Injuries Disablement Benefit for those ill or disabled because of an accident or event that happened at work or in connection with work. Berthoud (2011) argues from analysis of General Household Survey data that

2.1 The system of benefits and tax credits, 2005

The principal social assistance benefits in Britain in 2005 for people of working age are those shown in Table 1. Income Support (IS) and income-based Job Seekers Allowance (JSA) differ from Housing Benefit (HB) and Council Tax Benefit (CTB) because receipt depends on employment status. Put differently, receipt of HB and CTB depends on income (and some other conditions), but not on employment or job-search status. As Figure 1 shows, the SA receipt rates according to both the narrower and broader definitions move in tandem over time. The populations receiving IS and JSA on the one hand, and HB and CTB on the other hand, overlap substantially, and so the choice of whether to include housing benefits in the definition of social assistance benefits is of little practical importance in the current context. In the analysis presented below, we do not include housing benefits (HB and CTB) in our definition of SA.

<Table 1 near here>

Alongside these benefits for non-working families, there is extensive cash support available for low income working families, currently through the Working Tax Credit program. (It plays a similar role to the Earned Income Tax Credit in the USA.) The eligibility conditions relate to having an income below a specified minimum level, and at least one family member in 'full-time work' defined to mean working at least 16 hours per week.

2.2 Changes in benefits and tax credits, 1991–2005

Between 1991 and 2005, the system of benefits and tax credits in Britain changed substantially. The main reforms are summarized in Table 2 by year of introduction. Prior to 1996, a non-working family could be receiving Income Support (social assistance), Unemployment Benefit (UB, social insurance), or both. In October 1996, cash benefits for unemployed jobseekers were unified under the Job Seekers Allowance (JSA) program, with a distinction made between 'income-based' JSA corresponding to the former Income Support and 'contribution-based' JSA corresponding to the former UB, which was a flat-rate non-means-tested social insurance benefit paid to unemployed individuals with a satisfactory National Insurance contribution record. JSA also incorporated more stringent job search

changes in disabled peoples' employment rates or benefit payments have not coincided with major changes in the social security system's rules and procedures.

requirements for those assessed as available for work. Unemployed individuals with an incomplete national insurance contribution record and a sufficiently low income were also eligible to claim contribution-based JSA on a means-tested basis. Because UB payments were relatively low, most recipients' families were also eligible for IS, and it remains the case today that most JSA recipients receive income-based benefits.[3] Official statistics on JSA numbers no longer distinguish between contribution-based and income-based JSA, and it is also difficult to identify them separately in household surveys. For this reason, we include both types of JSA in our definition of SA: see below.

<Table 2 near here>

The other main changes were introduced by the Labour government that was elected in May 1997. The most significant reform was the replacement of the existing in-work benefits program, Family Credit, by the Working Families Tax Credit (WFTC) program modeled more closely on the US EITC. Aiming to 'make work pay', WFTC had more generous payments and extended eligibility notably by lowering the number of work hours required for qualification. Take-up was substantial, as Figure 1 illustrates. In 2003, the child allowance elements of WFTC were spun out into the Child Tax Credit program, which aimed to unify child support payments across the income maintenance system more generally. The WFTC component supplementing earnings became Working Tax Credit, and eligibility was extended to single people and to families without children. This gave another fillip to the proportion of working-age adults receiving tax credits: see Figure 1.

Among other policy reforms introduced by the Labour government to make work pay was a national minimum wage rate per hour. And, as part of its aim to reduce child poverty, there was an increase in support for families with children through increases in Child Benefit (a universal non-income-tested benefit paid per child) and in the child allowance element of other benefits.[4]

Brewer and Shephard's (2004) summary assessment, focusing on families with dependent children, is that '[e]xamining outcomes of Labour's ultimate objectives would lead one to conclude that the make work pay policies have been a success. … Academic studies

[3] According to the most recent administrative statistics, referring to February 2005, there were 687,400 JSA recipients in total, of whom 80.4 per cent were eligible to received income-based JSA and 19.6 per cent were eligible to receive contribution-based JSA. (The latter group includes some claimants also in receipt of income-based JSA.) See Table JSA 3.1 at http://statistics.dwp.gov.uk/asd/asd1/jsa/index.php?page=jsa_quarterly_feb05.

[4] There were also a number of active labour market programs for specific groups, the New Deals for unemployed young people and for lone parents, providing individualized help to improve job readiness and job search. Because of their targeted focus, and since they do not directly affect incomes, they are less relevant in the current context.

agree that government policies were partially responsible for these changes, at least among lone parents.' (2004, p. vii.)[5]

To the extent that policies making work pay are successful, we would expect them to be accompanied by corresponding reductions in SA receipt, and for the turning points and inflections in SA receipt trend lines to correspond with the major changes in tax credits (1999 and 2003). Similarly, we would expect the tightening of job search requirements for unemployed people accompanying the introduction of JSA in 1996 to lead to a decline in SA receipt, other things equal. Analysis of administrative record data by Petrongolo (2009) of men aged 16–64 suggests that tighter job search requirements were successful in moving individuals off unemployment benefits. Manning (2009) derives the same conclusion using Labour Force Survey data.[6]

3. Data: the British Household Panel Survey (BHPS)

We track SA receipt among working age adults using BHPS data from survey years 1991 to 2005 (waves 1–15).[7] The first wave of the BHPS was a nationally representative probability sample of the private household population, with interviews in Autumn 1991. The achieved sample consists of more than 10,000 individuals in some 5000 households, who have been re-interviewed annually. Individuals in split-off families have been followed, as in other household panels such as the US Panel Study of Income Dynamics or the German Socio-Economic Panel.[8] Our analysis is restricted to 1991–2005 because a local unemployment rate series that we rely on in the econometric analysis is not available beyond 2005 (see below). Note also that the BHPS ended in its current form with wave 18 (survey year 2008), with the subsample being subsumed into the new and much larger Understanding Society household panel survey (but with interviews not beginning for the BHPS sample until some 18 months later).

[5] For an overview of the impact of WFTC on labour supply and other outcomes, see the *Economic Journal* Features issue on 'In-work benefit reform in a cross-national perspective' (Brewer et al. 2009). Earlier research focusing on labour supply effects includes Blundell (2001), Blundell and Hoynes (2004), Brewer et al. (2006), Francesconi and van der Klaauw (2007), and Gregg et al. (2009).

[6] The reform did not lead to an unambiguous increase in the job-finding rate because some recipients moved instead to other benefits such as incapacity benefits (Manning 2009; Petrongolo 2009).

[7] We use respondents to the original ('Essex') sample only. Respondents from the extension samples for Scotland, Wales and Northern Ireland incorporated in the BHPS at the end of the 1990s are not used. Taking account of the differential sample inclusion probabilities would be a large task, beyond the scope of this project, and the number of observations in the original 1991 sample is relatively large in any case (see below).

[8] For full documentation of the BHPS, see http://www.iser.essex.ac.uk/survey/bhps.

We define an individual to be in receipt of SA if any individual in his or her family is receiving social assistance benefits at the time of the BHPS interview. In Britain, assessment of benefit eligibility is based on the income of the nuclear family, the so-called 'benefit unit', which is a single person or a couple living together with or without dependent children. (A dependent child is aged less than 16 years, or more than 16 years but under 19 years and unmarried, in full-time non-advanced education and living with his/her parents.) It is not legal marital status that distinguishes a 'couple' from two single adults; it is living arrangements (cohabiting unions are treated like legal marriages). In the sample of SA recipients we analyze, only one quarter are lone parents, 54 per cent live with a partner (37 per cent have a partner and children), and 21 per cent are childless single adults. We track individuals over time, not families, since families and households cannot be followed over time as a unit in any consistent manner. Families and households change their composition over time as individuals arrive (e.g. via partnership formation) or depart (e.g. via partnership dissolution). These types of change are common (Jenkins 2000).

We define SA to include IS and either UB or JSA (of either type). This is a matter of practical necessity: it is the only definition of SA that can be measured consistently over time using the BHPS (and other British surveys). As mentioned earlier, it is difficult to reliably distinguish between receipt of contribution-based JSA and income-based JSA. Indeed, since JSA's introduction in 1996, the BHPS interview has not asked respondents receiving JSA to distinguish between the two types for precisely this reason. See Cappellari and Jenkins (2008*a*) for further discussion.

Our definition of SA receipt for a given year *t* refers to receipt at the time of the BHPS interview in survey year *t* (typically September or October) – this is the definition used for Figure 1. The entry rate refers to the proportion of individuals not receiving SA at the year *t*–1 interview that are receiving SA at the year *t* interview and the exit rate refers to the proportion of individuals receiving SA at the year *t*–1 interview that are not receiving SA at the year *t* interview. Thus, the dynamics of SA receipt analyzed in this paper refer to transitions to and from receipt between successive annual interviews.

An alternative approach to receipt dynamics would be to take a spell-based approach, where spells are defined in terms of either consecutive 'benefit years' (meaning receipt at least once during the relevant year) or consecutive sub-annual periods such as 'months' where data are available. For the USA, the benefit year approach has been used to define spells in studies using PSID data from the pioneering analysis of AFDC benefit receipt by Bane and Ellwood (1983, 1994) onwards, while spell-based approaches based on sub-annual

data have been applied in studies based on the Survey of Income and Program Participation (e.g. Chay and Hyslop 2000, Grogger 2004).

By focusing on transitions between annual interviews, we play to the BHPS's strengths. Wishing to minimize measurement error and respondent burden, most BHPS effort is devoted to collecting information about the various income sources received at the time of the interview and the corresponding amounts. To be sure, at each interview the survey also asks about receipt of each of a large number of cash benefits for each month between the interview month and the September of the year prior to the current survey year using the respondent's retrospective recall. There are, however, substantial complications arising in the creation of consistent monthly histories of SA receipt. Not only are there 'seam' problems to deal with (an implausible number of transitions at the seam where successive histories overlap and have to be spliced together), but there are also issues arising from the family-based measure of receipt since histories are required for all the individuals who were in each person's family month by month. See Cappellari and Jenkins (2008*a*) for further discussion. Addressing these issues is an important and major task, and beyond the scope of this paper.[9] Our approach focusing on annual SA transitions rather than spells of receipt is effectively the same as the one taken by Andrén and Andrén (2013) and Hansen and Lofstrum (2003, 2006) using Swedish administrative record data, by Hansen, Lofstrum, and Zhang (2006) using Canadian household panel data, and by Königs (2013) using German household panel data.

With fifteen waves of BHPS data, our analysis data set contains a maximum of 15 observations per individual on SA receipt and other variables. At least two consecutive waves of data are required to estimate transition rates and any multivariate model of dynamics. We track individuals from when they are first observed as BHPS respondents until the first wave at which they drop out of the panel, either completely non-responding or with item non-response of sufficient degree that the individual's data cannot be used for estimation. If individuals rejoin the panel at some later wave, leading to gaps in benefit receipt sequences, we exclude them because taking account of intermittent participation complicates modeling substantially. Thus we focus on what is known as the 'absorbing attrition' case. The sample used for the empirical analysis is restricted to individuals of working age and not in full-time education (see earlier), and without missing data for some important explanatory variables. The basic estimation sample is an unbalanced panel of 75,988 person-wave observations for 9,036 adults. The majority of the sequences start at wave 1 (56 per cent), but there are

[9] See Bhuller and Königs (2013) for a study of social assistance dynamics in Norway, comparing estimates derived from monthly and benefit year data.

sequences that begin at each subsequent wave as well (roughly 200–300 adults each year). See Cappellari and Jenkins (2008*a*) for further details concerning sample selection, and demonstration that conclusions are robust to the use of a balanced rather than an unbalanced panel.

4. A dynamic random effects probit model of transition probabilities

Our data consist of a temporally ordered sequence of ones (representing SA receipt at a particular interview) or zeros (representing non-receipt), for every adult included in the analysis sample. To analyze these data, we propose use of panel data methods for binary sequences, specifically a version of the dynamic random effects probit model popularized by Heckman (1981*a*) which also accounts for unobserved individual heterogeneity. Dynamic random effects probit models have been used to analyze social assistance dynamics by Andrén and Andrén (2013), Bhuller and Königs (2013), Hansen and Lofstrom (2006), Hansen et al. (2006), and Königs (2013). One US application is by Chay and Hyslop (2000). Our model specification differs from those cited because, first, we allow each covariate to affect both exit and entry probabilities but with potentially different impacts in each case.[10] Most previous studies constrain the effects to be same. Second, we know of no previous study that has used the dynamic random effects probit model estimates to simulate entry and exit rates over time in the way that we do.

4.1 The statistical model

Let p^*_{it} represent the latent propensity of SA receipt in each year of the sequence of T_i years for which an individual is observed in the analysis panel data, excluding the first year ($t = 1$), where

$$p^*_{it} = (\gamma + \lambda y_{it-1})'Z_{it-1} + \tau_i + \zeta_{it}; \quad t = 2, \ldots, T_i. \tag{1}$$

Each individual, $i = 1, \ldots, N$, is observed to receive SA ($y_{it} = 1$) in year t if $p^*_{it} > 0$, and not to receive it ($y_{it} = 0$) otherwise. Observed individual heterogeneity is measured by the vector of variables represented by Z_{it-1} (which also includes an intercept term). These variables are

[10] Ribar's (2005) endogenous switching model of AFDC transitions shares this feature but he models unobserved heterogeneity differently. On endogenous switching models of transitions, see also Cappellari and Jenkins (2004, 2008*b*).

measured in the year at which the individual is at risk of making a transition (year t–1). As we show shortly, the interactions between the lagged dependent variable (y_{it-1}) and each element of Z_{it-1} allow characteristics to affect SA entry and exit rates differently.

Unobserved heterogeneity is characterized by a fixed individual-specific component (τ_i) and a white noise error component (ζ_{it}), where the error terms are assumed to be uncorrelated with each other and with each element of Z_{it}. The errors are each assumed to have a mean of zero and be normally distributed, with the variance of ζ_{it} normalized to be one, and variance of τ_i estimated from the data. In order to relax the assumption that the unobserved individual-specific components are uncorrelated with the observed explanatory variables, we follow Mundlak (1978) and Chamberlain (1984), and many researchers since, in allowing for correlations between u_i and Z_{it} by supposing that

$$\tau_i = \xi' \overline{Z}_i + u_i \tag{2}$$

where u_i is distributed $N(0, \sigma_u^2)$ and is assumed independent of Z_{it} and ζ_{it} for all persons and time periods. The $\overline{Z}_i$ may be defined in several ways – we follow the common practice of defining them as the longitudinal average for each individual of each characteristic within the vector Z_{it} (with the exception of intrinsically time-varying variables like age). Intuitively, differences in longitudinally-averaged characteristics are informative about underlying individual-specific characteristics, so that the unobserved individual differences that are left (u_i) may be more plausibly supposed to be independent of observed characteristics. For brevity in notation, we subsume the longitudinally-averaged variables into the vector Z_{it} henceforth.

There is an issue for estimation concerning the 'initial conditions' of the sequence of observations for each individual – whether y_{i1} is independent of u_i. If receipt in the initial year is correlated with the time-invariant individual-specific effect, a correlation is induced between the error term and the lagged dependent variable in (1), leading to bias in parameter estimates.

We handle initial conditions using the conditional maximum likelihood estimator proposed by Wooldridge (2005). Rather than modeling the joint distribution of the sequence of binary receipt indicators from the initial one to the final one conditioning on the set of explanatory variables, Wooldridge showed that one may model the distribution of the binary receipt indicators from $t_i = 2, \ldots, T_i$, conditioning on the set of explanatory variables and the binary receipt indicator for the initial year. Wooldridge proposed modeling the distribution of τ_i conditional on y_{i1} and either $Z_i = (Z_{i1}, Z_{i2}, \ldots, Z_{iT})$, or $\overline{Z}_i$. His model for the individual-

specific component (abstracting from $\overline{Z}_i$ already incorporated using the Chamberlain-Mundlak specification) can be written as:

$$\tau_i = a_0 + a_1 y_{i1} + u_i + \zeta_{it}, \tag{3}$$

so that equation (1) becomes

$$p^*_{it} = (\gamma + \lambda y_{it-1})' Z_{it} + \xi' \overline{Z}_i + a_0 + a_1 y_{i1} + u_i + \zeta_{it}; \quad t = 2, \ldots, T_i. \tag{4}$$

The Wooldridge estimator has the advantages that initial conditions do not have to be modeled and estimation can done using standard random-effects probit software.[11]

The model outlined incorporates a relatively simple dynamic structure. It characterizes a first order Markov process: transition behavior does not depend on receipt history beyond the year before the current one. Higher order Markov models can be fitted, as Chay and Hyslop (2001) and Stewart (2007) demonstrate, but we find that our model characterizes aggregate trends in transition rates relatively well (see below). More flexible approaches to duration dependence can also be modeled using survival analysis methods applied to spell data. We eschew those methods because of the difficulties of deriving consistent monthly histories (see earlier) and because using the annual interview data to define spells would underestimate the prevalence of short spells. For spell length predictions from our first order Markov model ignoring the latter issue and employing a 'steady state' assumption, see Cappellari and Jenkins (2008*a*).

4.2 Model implications for entry and exit probabilities

The dynamic random effects probit model characterizes transition probabilities for individuals of different types given appropriate conditioning on receipt status at t–1. The implied SA entry rate for non-recipients at t–1, e_{it}, is:

$$e_{it} \equiv \Pr(y_{it} = 1 | y_{it-1} = 0, Z_{it-1}) = \Phi((\gamma' Z_{it-1}) (1-\rho)^{0.5}). \tag{5}$$

The implied SA persistence rate for recipients at t–1, s_{it}, is

$$s_{it} \equiv \Pr(y_{it} = 1 | y_{it-1} = 1, Z_{it-1}) = \Phi([(\gamma + \lambda)' Z_{it-1}] (1-\rho)^{0.5}) \tag{6}$$

[11] In our earlier work (Cappellari and Jenkins 2008*a*), we showed that the Wooldridge estimator provided almost identical parameter estimates to the estimators of Heckman (1981) and Orme (2001) that also account for the initial conditions issue, for both balanced and unbalanced panels. We attribute this robustness to the long length of our panel. See also Arkay (2009) and Arulampalam and Stewart (2009). We focused on estimates based on the Heckman estimator in our earlier work. Compared to the model specification employed by Cappellari and Jenkins (2008*a*), the main differences in the current paper are that observed characteristics are now measured at t–1 rather than t, and a more extensive set of variables is included in the interaction between Z_{it-1} and past receipt.

and the SA exit rate for recipients at t–1, x_{it}, is

$$x_{it} \equiv \Pr(y_{it} = 0 | y_{it-1} = 1, Z_{it-1}) = 1 - s_{it} \quad (7)$$

where $\rho = \sigma_u^2/(1+\sigma_u^2)$. The cross-sectional SA receipt rate at t, p_{it}, is:

$$p_{it} \equiv \Pr(y_{it} = 1 | Z_{it-1}) = (1 - y_{it-1})e_{it} + y_{it-1}s_{it} \quad (8)$$

The impact on the SA entry rate of a factor included in Z_{it-1} depends on the coefficient in γ corresponding to the factor. For example, we would expect larger unemployment rates to be associated with a larger SA entry rate and a smaller exit rate (a larger persistence rate), other things being equal. In this case, the coefficient on the unemployment rate in γ (call it γ_{unemp}) would be positive. (As the model is non-linear, the precise effect of changing unemployment rates also depends on the values of other parameters and the other characteristics in Z_{it-1}.) To ascertain impacts on exit rates is not quite as straightforward because they depend on sums of corresponding elements in γ and λ. For example, a negative association between the exit rate and the unemployment rate requires $\gamma_{unemp} + \lambda_{unemp}$ to be negative. Observe that many previous studies using dynamic random effects probit models constrain the impact of covariates on entry and exit rates to be the same: the λ vector contains an intercept term and zeros otherwise. (In the example just discussed, the assumption would be that $\lambda_{unemp} = 0$.)

For each year within our sample period, we simulate aggregate SA entry and exit rates, as well as cross-sectional receipt rates. We substitute our estimates of γ, λ and σ_u^2 into equations (5)–(8) to derive predicted transition and receipt probabilities for each individual in the estimation sample, year by year, and then we average the individual-specific probabilities across those at risk to derive the aggregate rates.

The model accounts for trends in aggregate entry and exit rates over the period in two ways. One is through time-varying parameters: we allow the intercepts in γ and λ to be year-specific. Second, there are changes over time in the characteristics (Z) among those at risk of SA exit or entry (such as changes in the local unemployment rates that individuals face), and these affect our predictions of aggregate transition probabilities because we derive them by averaging individual-specific probabilities across the individuals at risk in each year.

We assess the model's within-sample goodness of fit by comparing predicted and observed aggregate SA transition and receipt probabilities, year by year. Having shown that the model fits relatively well, we use it to undertake some counterfactual simulations in order to investigate the relative impacts of changes such as benefit reforms and unavailability of

jobs. In the first case, we compare observed aggregate SA entry, exit, and receipt rates with the corresponding rates predicted by holding the year-specific intercepts in γ and λ fixed at their values at the start of the period, with all other factors (parameters and characteristics in Z_{t-1}) being held fixed. (The reason for this specification is discussed in the next Section.) In the second case, we compare observed rates with the rates predicted with local unemployment rates held at their levels at the start of the period. In both cases, the larger is the difference between the counter-factual and within-sample predictions, the larger is the effect that we attribute to the factor examined. We would emphasise that this is a descriptive assessment of 'impact' rather than a causal one but we believe that it is valuable nonetheless, because it helps one judge the economic significance of the estimated model parameters in a more helpful way than simply looking at each parameter in isolation. For more details of the simulations, see Section 5; the characteristics included in *Z* are discussed next.

4.3 Explanatory variables

Observed characteristics are summarized by sex and age, highest educational qualification, and health status. Age refers to whether the individual is aged 50 or more years: we expect older workers to have less work attachment, other things being equal, and hence higher SA entry and lower SA exit rates. We distinguish four categories of educational qualification: none, low, high, and missing. 'Low' refers to having passes in examinations taken at age 16 (CSE(s) and/or O-levels); 'high' refers to having passes in examinations taken at age 18 (A-level(s)) or higher qualifications such as a degree. Around one tenth of respondents have missing data on educational qualifications: these are mostly respondents for whom only a proxy interview was gained but sufficient information was derived about other characteristics from the proxy respondent so that the individual could be included in the sample. The missing qualifications indicator is better interpreted as a control for response propensity than as a measure of educational qualifications. Health status refers to whether the respondent stated that s/he had one of more of 13 health problems.[12]

[12] The problems refer to: (1) Problems or disability connected with: arms, legs, hands, feet, back, or neck (including arthritis and rheumatism); (2) Difficulty in seeing (other than needing glasses to read normal size print); (3) Difficulty in hearing; (4) Skin conditions/allergies; (5) Chest/breathing problems, asthma, bronchitis; (6) Heart/blood pressure or blood circulation problems; (7) Stomach/liver/kidneys; (8) Diabetes; (9) Anxiety, depression or bad nerves; (10) Alcohol or drug related problems; (11) Epilepsy; (12) Migraine or frequent headaches; (13) Other health problems.

The characteristics of each respondent's family are summarized by the number of dependent children, whether the age of the youngest child is less than five years, and by family type (single adult, couple, or lone parent). We also control whether the respondent lives in the London area (as the labour market is very different from elsewhere in the country), and housing tenure (whether the respondent lived in owner-occupied housing rather than other tenures such as social housing or renting privately). We do not suggest that tenure itself necessarily has an impact; rather it is a marker for other factors including differences in wealth and assets and local area disadvantage.

The remainder of the explanatory variables relate to the factors cited earlier in the descriptions of trends in aggregate receipt and transition rates (Figures 1 and 2), specifically changes in the unavailability of jobs and in tax and benefit policies.

Our measure of unemployment is the unemployment rate in the individual's local area, by which we mean travel-to-work area (TTWA). A TTWA is defined with reference to commuting patterns, and corresponds broadly to a city and surrounding area. More precisely, the local area unemployment rate is the ratio of the number of unemployed people to the number in the labour force in the individual's TTWA at the time of the annual interview, derived from the Joint Unemployment & Vacancies Operating System Cohort (a 5 per cent sample of all computerized claims for unemployment-related benefits selected by reference to a claimant's National Insurance Number).[13]

The impacts of benefit policy changes are accounted for by allowing the intercept terms in the SA entry and exit equations to differ by survey year. We follow this strategy rather than including measures of (say) programme generosity over time for two reasons. First, benefit rates in the UK are set nationally; there is no spatial variation across regions or states as in the USA or some European countries. Second, there have many program changes within the 1991–2005 period (Tables 1 and 2), and so it is difficult to identify the impact of any specific policy reform. Our approach is therefore more descriptive than causal and we cannot identify the impact of specific policies. Nonetheless, we expect that if the policy changes cited in Table 2 had had an effect on SA entry and exit rates, they would be reflected in changes in the time-varying intercepts, especially around 1996, 1999 and 2003 (see the earlier discussion).[14]

[13] See Cappellari and Jenkins (2008*a*) for more details. It is not possible to construct a consistent series that both covers the period 1991–2005 and extends further into the future.

[14] We also considered a quasi-differences-in-differences approach in which we also included interactions between the survey year indicators in Z_{it-1} and the presence of children: the reasoning is that the Labour government's reforms were directly primarily at families with children. As it happened, very few coefficients on

Derivation of unbiased estimates relies on the assumption that all explanatory variables are strictly exogenous. This assumption is relatively innocuous for factors such as local unemployment rates or changes to the benefit system (changes to which are what we consider in our counterfactual simulation exercises). More debatable is our supposition that the impacts of a shock in an individual's SA eligibility on the individual's future housing tenure or household type are sufficiently small that they may be ignored. To account for potential endogeneity is particularly difficult and hampered, for example, by a lack of plausible instruments. For valuable discussion of the issues, see Biewen (2009).

4.4. Characteristics of at-risk groups, and trends over time

Changes in the aggregate SA transition rates depend in part on changes in the composition of the populations at risk (see above). Consider first those at risk of entry (SA non-recipients). Averaging over the period as a whole, the mean age was 41 years, half the sample were women, and just over one half reported at least one health problem. About thirteen per cent had no educational qualifications and a quarter had only low educational qualifications. Almost a half of non-recipients had at least one dependent child and just under a fifth had a child aged less than five years. Four-fifths were part of a couple, some 17 per cent were single adults and 3 per cent were lone parents. Just over 80 per cent lived in owner-occupied accommodation, one in ten lived in the London area, and the average local area rate was just over 5 per cent. These whole-period averages disguise some marked trends over time. In particular, there was a rise in the proportion reporting a health problem (up from around 50 per cent in 1991 to around 60 per cent in 2005), and even larger changes in the distribution of educational qualifications. The fraction of non-recipients with no or low educational qualifications declined from around 53 per cent in 1991 to 26 per cent in 2005. The other main variation over time was in local area unemployment rates, which declined from almost 10 per cent on average at the start of the 1990s to just over 2 per cent in 2005. In contrast, the means of virtually all the demographic variables (age, number and age of children, family type) changed very little.

These changes suggest that one reason for the decline in SA entry rates may be the improvement in educational qualifications (improving employability) and the decline in local area unemployment rates reflecting improved availability of jobs. The rate of decline in the

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interaction variables were statistically significant, and the temporal pattern of the estimates did not correspond with expectations regarding the timing of policy changes.

average local area unemployment rate leveled off around 1997, which is consistent with the leveling off in the decline in the entry rate around that year (Figure 2).

Consider now those at risk of SA exit. Averaging over the period as a whole, we find that compared to non-recipients, the proportion of female recipients is larger (60 per cent rather than 50 per cent), and the proportion with relatively low educational qualifications is higher (40 per cent of recipients have no qualifications compared to 13 per cent of non-recipients). A quarter of recipients belong to families with three or more children (compared to 8 per cent of recipients) and the proportion with a child aged less than five years is 25 per cent rather than 18 per cent. About one quarter of recipients are lone parents but only 3 per cent of non-recipients. The proportion of SA recipients living in owned accommodation is only one third compared with 80 per cent for non-recipients. In addition, recipients tend to live in areas with slightly higher unemployment rates than non-recipients. In sum, SA receipt is concentrated among individuals with characteristics commonly associated with labour market disadvantage.

As far as trends are concerned, there are both similarities and differences between SA recipients and non-recipients. For both groups, the prevalence of health problems rose between 1991 and 2005, but the increase is larger for recipients (from 53 per cent to 76 per cent, compared to from 50 per cent to 59 per cent). Local area unemployment rates fell for both groups; so too did the proportion with no educational qualifications but the decline in the latter was smaller for SA recipients than non-recipients. There are some distinctive demographic trends for recipients: the proportions with children and with young children in particular declined over time. There was also a decline in number of lone parents among recipients (from 35 per cent in 1991 to 18 per cent in 2005). In principle this might be explained by a shift in low income families with children (and lone parents in particular) from SA receipt to receipt of in-work benefits such as WFTC, but the proportions of couples with children and of lone parents was largely unchanged over the period. Another marked trend among recipients is the decline in the proportion living in owned accommodation, from 41 per cent in 1991 to 22 per cent in 2005. Put another way, the association between SA receipt and living in rented accommodation (much of which is subsidized social housing) has increased.

These patterns suggest that two trends in particular may help account for the decline in the aggregate SA exit rate over time: the rise in proportion of recipients with health problems, and the large rise in the proportion living in non-owned accommodation. Both trends are consistent with a ‘creaming’ hypothesis – the most skilled and work-ready

individuals left SA for a job, whereas the group left on SA increasingly consisted of individuals who were less well equipped for work, and for whom the probability of SA exit is relatively low.

5. Model estimates and counterfactual simulations

5.1 Model parameter estimates

Estimates of the dynamic random effects probit model specified in (4) are set out in Table 3. In the top half of the table, the first column of numbers refers to estimates of γ and the second column to estimates of λ. In the bottom half of the table are the estimates of the effects of the longitudinally-averaged variables on SA receipt propensities, the impact of being in SA receipt when initially observed, and the variance of the unobserved heterogeneity distribution.

Table 3 shows that SA entry probabilities are lower for women, individuals living with a partner, and without a child aged less than 5 years. There is a gradient in entry rates with educational qualifications – highest for those with no qualification and lower the higher the qualifications attained (and higher still for those with missing qualifications data). There is no statistically significant association between entry rates and age, the presence of health problems, number of children, home ownership, and residential location.

<Table 3 near here>

The greater the local area unemployment rate, the larger is the SA entry rate – which is consistent with the aggregate trend data shown in Figures 1 and 2.

The pattern of estimates of the survey year intercepts also corresponds with the trends in aggregate data. (The year labels shown in Table 3 refer to year *t* for transitions between year *t*–1 and year *t*.) Between the mid-1990s and the mid-2000s, the coefficients become increasingly negative implying a lower SA entry rate, other things being equal. These appear to be large changes: the estimated intercept doubled in magnitude between 1995 and 2005, from –0.14 to –0.28. But are there variations in intercepts corresponding to years of major benefit reform (survey years 1997, 2000, and 2004 according to the labeling convention used in Table 3)?

The estimates suggest that the introduction of JSA, accompanied by tightening of eligibility conditions, was associated with a decline in SA entry rates. For those at risk of entry in 1996 (year $t = 1997$), the estimated coefficient is –0.17, but it is –0.37 for those at risk of entry in 1997 (year $t = 1998$). There is also some evidence consistent with a WFTC introduction effect as the coefficient for those at risk of entry in 1999 (year $t = 2000$) is –0.24, but –0.43 for those at risk of entry a year later. In contrast, there is no similar change in the coefficients for years round the change from WFTC to WTC. We therefore find a smoking gun for some effects on SA entry rates of benefit policy reforms. We refrain from drawing stronger conclusions about causality for the reasons described earlier.

What about the determinants of exit rates? Observe, first of all, that relatively few of the coefficients on interactions between characteristics and lagged SA receipt status are statistically significant. The exceptions concern the number of children in the family and whether the respondent has a partner. Having more children is associated with a smaller exit rate, whereas being a member of a couple is associated with larger exit rate. Otherwise, we find, for example and as expected, that having better educational qualifications is associated with a higher exit rate, and a larger unemployment rate is associated with a lower exit rate.

By contrast with the results for entry rates, there appears to be little evidence of benefit policy reform effects on exit rates. For the relevant survey years, observe that the sum of corresponding γ and λ intercepts is close to zero: the estimates are of approximately the same magnitude and of the opposite sign. Hence the change between successive years is also negligible.

The estimates of the parameters associated with unobserved heterogeneity are at the bottom of Table 3. The table shows that individuals with a disposition to health problems or many children are more likely to receive SA. Family type and housing tenure also matter. Individuals who are more likely to be lone parents are more likely to receive SA, whereas individuals with a partner are less likely to. Home-owners are less likely to receive SA than renters. Experiencing persistently high local area unemployment rates does not appear to be associated with a high probability of SA receipt, other things being equal. (Thus it is the year-to-year variations in unemployment that drive changes in SA receipt, by changing exit and entry propensities.) There is statistically significant unobserved heterogeneity in addition to the heterogeneity captured by the longitudinally-averaged variables: the estimate of σ_u^2 is 0.42. Finally, observe that initial conditions matter. Individuals who are receiving SA when

initially observed are much more likely than non-recipients to be receiving SA in some subsequent year.

5.2 Counterfactual simulations

We now show that the fitted model tracks the aggregate trends in aggregate SA transition and receipt rates and then assess the main drivers of the trends using counterfactual simulations.

Figure 3 compares observed and within-sample predictions of SA entry, exit and receipt rates year by year. (The series for the observed rates differ slightly from those shown in Figures 1 and 2 because they are based on the estimation sample. This is smaller than the samples for Figures 1 and 2 because of missing values on elements in Z_{it-1}.) Panel (a) compares entry rates; panel (b) compares exit rates; and panel (c) compares receipt rates.

<Figure 3 near here>

The within-sample predictions are good.[15] Entry rates are tracked remarkably well by the model. Exit rates are tracked less well, with the gap between observed and predicted rates increasing between 1992 and 1999 to around 10 percentage points and remaining at that level thereafter. On the other hand, the confidence intervals around the estimates of observed and especially predicted exit rates are relatively large (sample sizes are relatively small) and overlap.[16] Moreover, the turning points in the two exit rate series correspond. And, as emphasized earlier, it is entry rate trends rather than exit rate trends which are the main driver of receipt rate trends. For these various reasons, we are relatively sanguine about the within-sample prediction of exit rates. The importance of entry rates is highlighted by the comparisons of predicted and observed receipt rates shown in panel (c). Within-sample predictions under-estimate the aggregate receipt rate in each year, but the difference is small (at most around half a percentage point).

Given the good within-sample fit, we proceed to our counterfactual simulations. The first exercise concerns the impact of the decline in unemployment rates after the early-1990s recession. To assess this, we consider what would have happened to SA entry and exit rates were unemployment to have remained at its 1993 peak and everything else were to stay the same (including the values of the longitudinally-averaged variables and initial receipt status). More precisely, individual-specific local area unemployment rates in each year are set equal

[15] The fit is much better than reported in the working paper versions of this paper because we have corrected a programming error.
[16] Confidence intervals are not shown in the figures for legibility's sake.

to their 1993 values. The simulation generates the predictions of aggregate rates shown by the series labeled '1993 unemployment rates' in Figure 4.

<Figure 4 near here>

Figure 4 suggests that the fall in unemployment rates over the period had a large impact on SA entry rates. If unemployment rates had remained at their peak level, the entry rate would have been almost 1 percentage point higher by 2005. This is a large impact given that the aggregate (within-sample predicted) entry rate was just over 1 per cent in 2005. Observe that most of impact of falling unemployment rates occurred before 2000, as suggested by the aggregate unemployment rate trends reported in Figure 1. The difference between the counter-factual and within-sample simulated rates increases sharply between 1993 and 1997 (to 0.7 percentage points) and then increases at a slower rate over the rest of the period, reaching a rate some 0.9 percentage points higher in 2005 than in 1993.

For SA exit rates, we also see the impact expected. If unemployment rates had remained at their early-1990s high, the exit rate would have been lower but, again, with the gap increasing fastest before 2000 and less thereafter. However, the magnitude of the simulated effect on exit rates is relatively small compared with the impact on entry rates. The difference between the series in panel (b) is never more than 4.7 percentage points (reached in 2004), which is a small amount compared to the predicted within-sample aggregate exit rate (which fluctuates around 40 per cent).

The simulated impact of persistently high unemployment rates on aggregate SA receipt rates is shown in panel (c) of Figure 3. The gap between predicted series increases over the 1990s, and from around 1998 through to 2005, the receipt rate is predicted to be around 1 percentage point higher. This is a relatively large magnitude given the size of the receipt rate in the 2000s.

Our second counterfactual exercise investigates factors associated with the passage of time, including benefit policy reform effects. We compare the within-sample prediction series with aggregate rates predicted by fixing the survey year intercepts at their estimated values for 1993. (All other factors remain unchanged.) The resulting series are shown in Figure 5, panels (a)–(c).

For SA entry rates, the effects associated with the passage of time are large, and larger in magnitude than the effect associated with persistently high unemployment rates. By 1998, the gap between series had increased to around 1.1 percentage points and fluctuated around that size until 2005. The simulations for exit rates (panel b) confirm the impressions derived from discussion of the regression table estimates: there is little clear evidence of policy

reform or other time effects in this case. The counterfactual and within-sample series overlap substantially. Hence, unsurprisingly, the results for the receipt rate are similar to those for the entry rate. If there had been no year effects, the proportions receiving SA would have been higher, notably from around 1996 onwards. The predicted increase fluctuates over following decade between around 0.7 and 1.4 percentage points, and thus about the same size as the effect predicted were 1993 local unemployment rates to have persisted. Clearly, this exercise does not identify the size of the impact of any specific benefit policy reform, rather it indicates changes associated with the passage of time and so, at most, it provides an upper bound on the cumulative impact of policy changes.

<Figure 5 near here>

In Section 4, we noted how changes in distributions of characteristics over time could account for changes in SA transition and receipt rates and, for example, drew attention to a marked upgrading in educational qualifications among the working-age population and a halving between 1991 and 2005 in the proportion of owner-occupier SA recipients from around 40 per cent to around 20 per cent. We investigated the impact of these two factors using counterfactual simulations as well, fixing at 1991 levels first the distribution of educational qualifications and then, second, the proportion of individuals who were owner-occupiers. In the first exercise, the impacts were small compared to the effects of falling unemployment rates or of year-specific effects that were discussed earlier. In the second exercise, there was no perceptible impact at all. For brevity's sake, we do not show the predictions here. (They are reported in the working paper versions of this paper.)

5.3 Heterogeneity in transition rates

So far, the simulations have been concerned with generating predictions of aggregate transition rates and their trends over time. This involved averaging of individual-level predictions and hence information about the degree of heterogeneity in transition rates across individuals with different characteristics was hidden. We now illustrate this heterogeneity by predicting SA entry and exit rates for an individual with a specific 'base' set of characteristics, and compare these predictions with those derived by changing each of a number of characteristics one at a time. The predictions refer to a particular year. The results are shown in Table 4. The 'base' set of characteristics refers to a woman aged 40 years who is an owner-occupier living outside London, married with one child under 5, has no health

problems and no educational qualifications, and the local area unemployment rate is 3 per cent, and the year is 2005. She did not receive SA in the year she was initially observed.

The predicted SA entry probability for this person is 1.2 per cent, and the predicted exit probability is 65.9 per cent. Men and women have much the same entry and exit probabilities (row 2). More important than gender is family type. If the reference woman is a lone mother rather than married, then her entry probability more than trebles to 3.9 per cent, and her exit probability falls by about a quarter to 43.7 per cent (row 6). If, in addition, the woman is not an owner-occupier, the entry probability is a massive 14.6 per cent and the exit probability is less than one third that of the base case, 19.6 per cent (row 7). Differences in housing tenure alone are associated with large differences in transition probabilities (row 5). Not having children almost halves the entry probability, and slightly increases the exit probability, relative to the base case (row 8). Finally, observe that if the year is changed from 2005 to 1993 and the local area unemployment rate from 3 per cent to 9 per cent, then the entry probability more than doubles (from 1.2 per cent to 3.3 per cent) and the exit probability hardly changes. This underlines our earlier remarks regarding the importance of changes in entry rates rather than changes in exit rates for explaining trends in SA receipt rates.

<Table 4 near here>

6. Summary and conclusions

We have analyzed the dynamics of SA receipt among working-age adults in Britain between 1991 and 2005, arguing that the near-halving in the aggregate annual receipt rate over this period was driven by a decline in the annual entry rate rather than changes in the annual exit rate (which actually declined). To examine the determinants of trends in aggregate SA transition rates, we have developed a dynamic random effects probit model of SA receipt entry and exit probabilities, fitted it to BHPS data, and used the estimates to provide counterfactual simulations of trends in SA transition rates.

Our analysis indicates that two sets of factors had important effects on transition rate trends over the period 1991–2005. First, there was the increasing availability of jobs, summarised by the fall in unemployment rates between 1993 and the mid-2000s. Given the substantial increase in unemployment associated with the recession that began in 2007/8, it is

unsurprising that the number of individuals in receipt of SA has also increased sharply as well (Department for Work and Pensions 2013).

Second, there were a number of other changes between 1991 and 2005 in the socioeconomic environment including various reforms to the benefit system. Our results suggest that two changes had an effect on SA entry rates: the introduction of JSA in 1996 (by making it harder to claim), and the introduction of WFTC in 1999 (making work pay and hence SA less attractive). Investigating the extent to which these effects on SA receipt are genuinely causal is a task for future research; so too is analysis of the nature of other changes in the socioeconomic environment that were associated with the fall in the SA entry rate.

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Table 1. The principal social assistance benefits in the UK for working age adults, 2005

Benefit	Eligibility conditions (main)
Income Support	Income less than a specified minimum level, and unavailable for full-time work (e.g. lone parent, registered sick or disabled, caring for someone who's sick or elderly).
Job Seekers Allowance (income based)	Income less than a specified minimum level, and unemployed but able to work and available to work (which has to be regularly declared).
Housing Benefit	Income less than a specified minimum level, and needing financial help to pay all or part of one's housing costs.
Council Tax Benefit	Income less than a specified minimum level, and needing financial help to pay all or part of one's Council Tax bill.

Notes: Income Support was introduced in 1988 (its predecessor was called Supplementary Benefit). Housing Benefit was introduced in 1983 and Council Tax Benefit in 1993. Job Seekers Allowance was introduced in 1996. See Table 2 below regarding changes between 1991 and 2005.

Table 2. Principal changes to the UK system of cash benefits and tax credits, 1991–2005

Year of introduction	Change
1996	Job Seekers Allowance (JSA) introduced in October 1996 with 'income based' and 'contribution based' components. JSA replaced Income Support (IS) and Unemployment Benefit (UB) for unemployed jobseekers. Accompanied by more stringent job search requirements for those assessed as available for work. IS became available only to those not available for work.
1999	Working Families Tax Credit (WFTC) introduced in October 1999, and fully phased in by April 2000. This in-work benefit program for low income families was more generous and widened eligibility relative to its predecessor, the Family Credit program (FC). FC, introduced in 1988, replaced Family Income Supplement (FIS) which began in 1971. Administered by the income tax authorities (HM Revenue and Customs) rather then the benefits authorities (Department for Work and Pensions, and Benefits Agency).
1999	Increased support for families with children, including increases in Child Benefit (flat-rate payment per child, paid regardless of parental work status or income), and increases in the child allowances in other benefits.
1999	National Minimum Wage introduced.
2003	WFTC replaced by the Working Tax Credit (WTC) and Child Tax Credit (CTC) programs from April 2003. WTC extended eligibility to single people and to families without children. CTC unified child allowances across benefits.

Note. See Brewer and Shephard (2004) for a concise overview of the Labour government's welfare to work policies and associated changes in the benefit system.

Table 3. The probability of SA receipt (dynamic random effect probit model estimates)

Explanatory variables (measured at *t*–1)	γ		λ	
Female	–0.0895	***	0.0896	
	(0.034)		(0.060)	
Aged 50 years or more	0.0529		0.0463	
	(0.044)		(0.081)	
Has health problem(s)	0.0278		–0.0143	
	(0.038)		(0.057)	
Educational qualifications				
Low: O-level(s), CSE, etc	–0.2073	***	–0.0534	
	(0.046)		(0.071)	
High: A-level(s) or higher	–0.3623	***	–0.0680	
	(0.045)		(0.073)	
Missing data	–0.4215	***	0.0318	
	(0.063)		(0.150)	
Number of children in family	–0.0264		0.0751	***
	(0.024)		(0.027)	
Age of youngest child < 5 years	0.1540	***	–0.1084	
	(0.049)		(0.073)	
Family type: lone parent	–0.0030		0.0724	
	(0.095)		(0.110)	
Family type: couple	–0.1214	*	–0.2342	***
	(0.069)		(0.079)	
House tenure: owned	–0.0491		0.0145	
	(0.056)		(0.059)	
Lives in London (inner or outer)	0.2402		0.0472	
	(0.163)		(0.091)	
Unemployment rate in local area (%)	0.0323	***	–0.0107	
	(0.010)		(0.015)	
Survey year (year *t*)				
1993	–0.0321		0.0049	
	(0.058)		(0.118)	
1994	–0.0964		–0.0659	
	(0.060)		(0.121)	
1995	–0.1372	**	0.0462	
	(0.060)		(0.122)	
1996	–0.1559	**	0.0013	
	(0.061)		(0.124)	
1997	–0.1734	***	0.1401	
	(0.067)		(0.135)	
1998	–0.2914	***	0.2572	*
	(0.077)		(0.147)	
1999	–0.3655	***	0.3811	**

	(0.082)		(0.154)	
2000	–0.2357	***	0.2633	*
	(0.080)		(0.157)	
2001	–0.4262	***	0.4703	***
	(0.089)		(0.165)	
2002	–0.3834	***	0.3103	*
	(0.089)		(0.167)	
2003	–0.3152	***	0.5179	***
	(0.088)		(0.176)	
2004	–0.3991	***	0.2557	
	(0.092)		(0.173)	
2005	–0.2847	***	0.4109	**
	(0.090)		(0.181)	
Intercept	–1.5693	***	1.2370	***
	(0.102)		(0.180)	
Longitudinally-averaged variables				
Has health problems	0.2495	***		
	(0.055)			
Family type: couple	–0.1473	*		
	(0.081)			
Family type: lone parent	0.6074	***		
	(0.121)			
Number of children in family	0.0964	***		
	(0.031)			
Age of youngest child < 5 years	–0.0054			
	(0.085)			
House tenure: owned	–0.7970	***		
	(0.068)			
Lives in London (inner or outer)	–0.2498			
	(0.171)			
Unemployment rate in local area (%)	0.0118			
	(0.009)			
Received SA at $t = 1$	0.7658	***		
	(0.051)			
σ_u^2	0.4201	***		
	(0.036)			

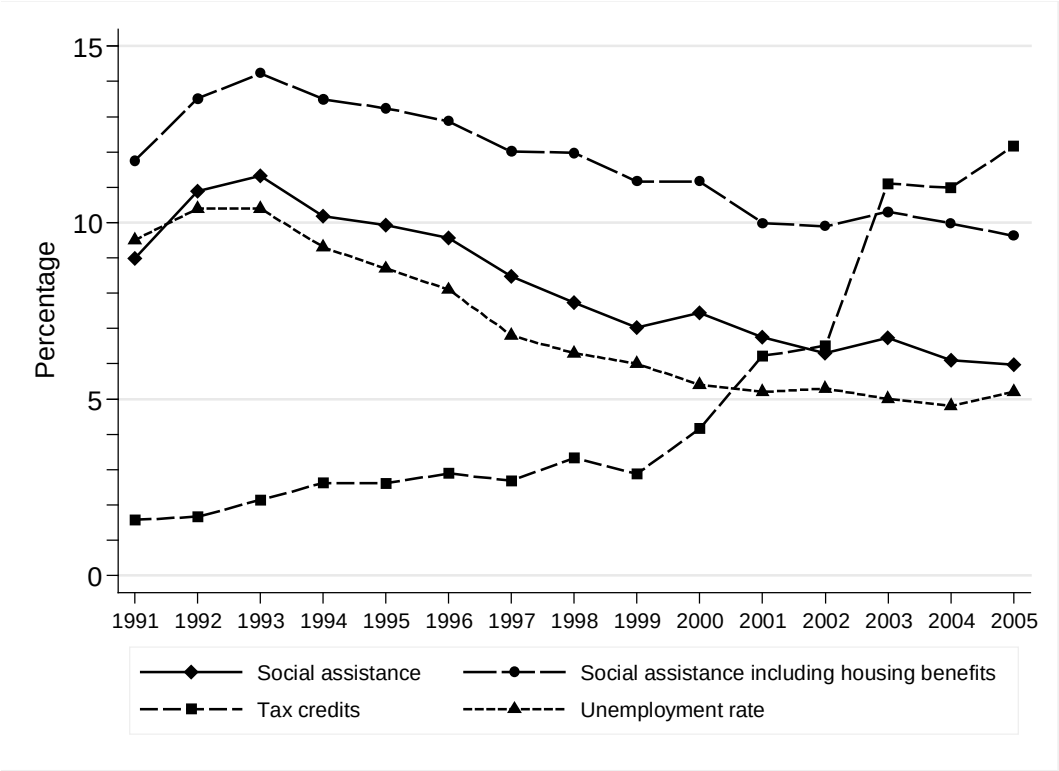
Notes. The table shows estimates of equation (4) model fitted to data from waves 1–15 of the BHPS using the Wooldridge (2005) estimator. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Log-likelihood = –8792.293. Number of person-year observations = 66,952. Number of persons = 9,036. Reference categories: male, aged 25–50, has no health problems, has no educational qualifications, family type is single, lives in non-owned accommodation outside the London area, and the survey year is 1992. The outcome is measured in year *t*, and explanatory variables in year *t*–1, with the exception of the survey year indicators for which the indicator label refers to year *t*.

Table 4. Heterogeneity in SA annual entry and exit rates: predicted transition probabilities for individuals with different sets of characteristics

Characteristics	Entry probability (%)	Exit probability (%)
1. Base set of characteristics*	1.2	65.9
As Base, except:		
2. Man	1.4	65.9
3. Has health problems	2.1	57.5
4. Has educational qualifications to A-level or higher	0.5	76.3
5. Non-owner	5.9	38.7
6. Lone mother	3.9	43.7
7. Lone mother and non-owner	14.6	19.6
8. No children	0.7	71.4
9. Year is 1993, local area unemployment rate = 9%	3.3	64.4

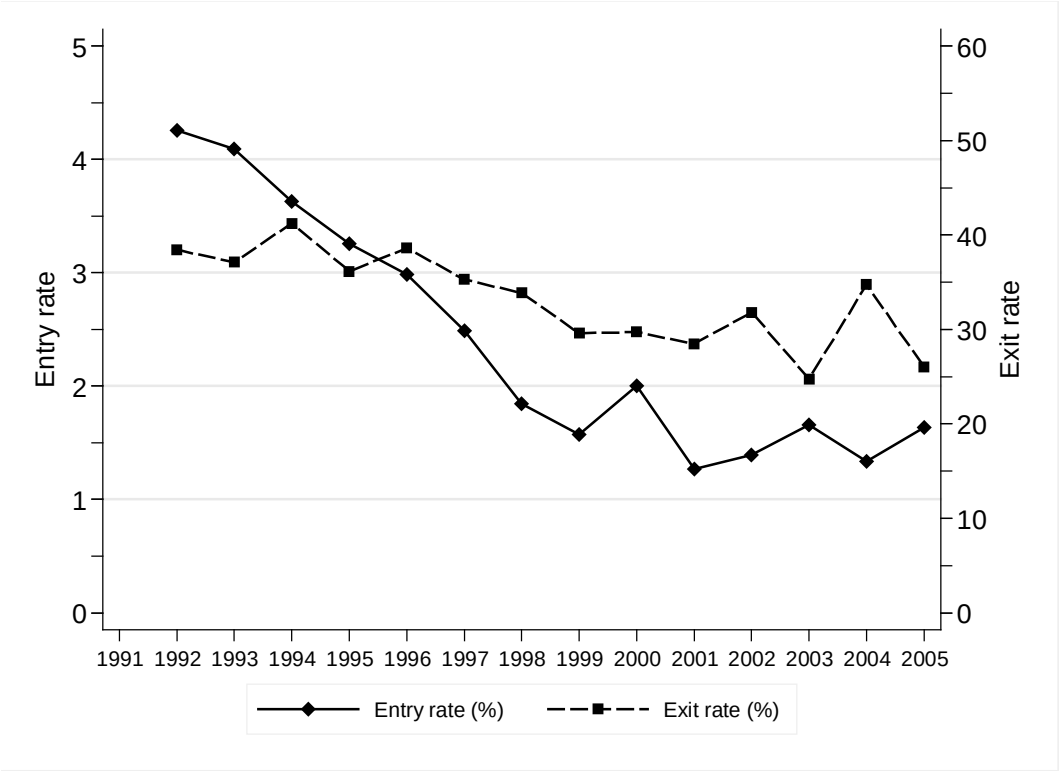
Notes. Predictions derived using equations (5) and (7). * Base characteristics: woman, 40 years old, living outside London, one child under 5, married , no health problems, no educational qualifications, local area unemployment rate = 3%, owner-occupier, year = 2005, not receiving SA in year initially observed.

Figure 1. Proportion of working-age adults receiving social assistance benefits and tax credits, and the unemployment rate, by year



Notes: Authors' calculations using data from waves 1–15 of the British Household Panel Survey (BHPS), except for the unemployment rate which is series YBTI from the UK Office for National Statistics, http://www.ons.gov.uk. The unemployment rate is the ILO unemployment rate for all adults (men aged 16–64, women aged 16–59), derived from the Labour Force Survey, and is a three-month average centered on the October of the year in question. The definitions of social assistance, housing benefits, and tax credits are explained later.

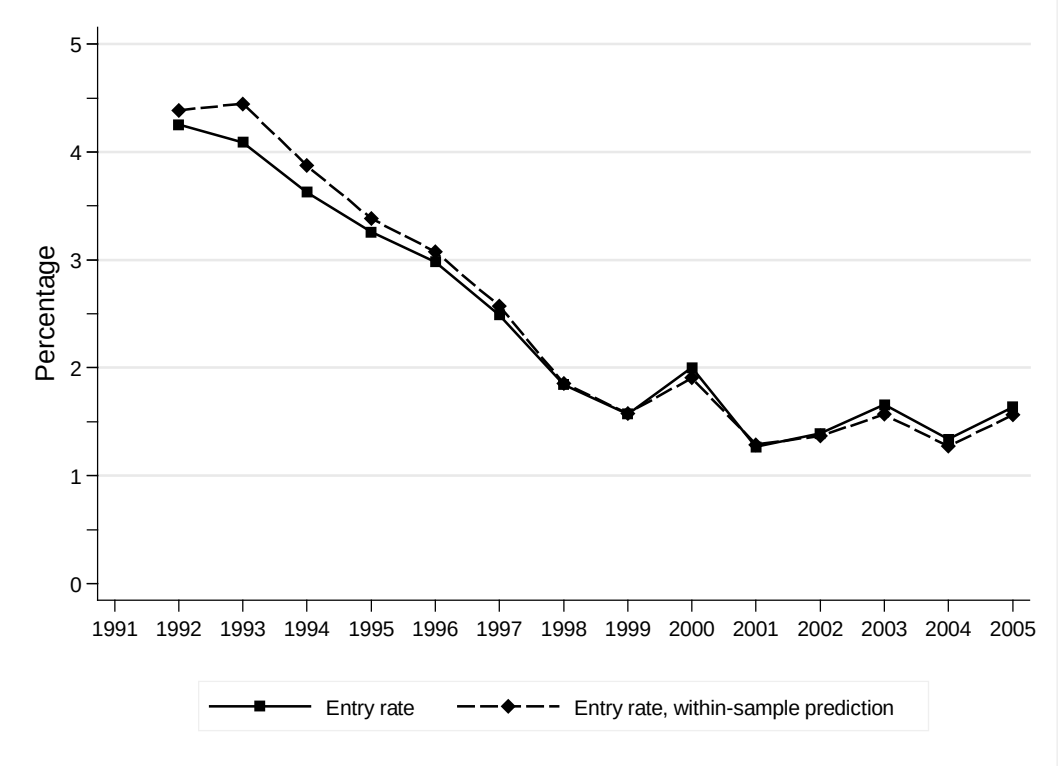
Figure 2. Annual rates of entry to and exit from social assistance benefit receipt: working-age adults, by year



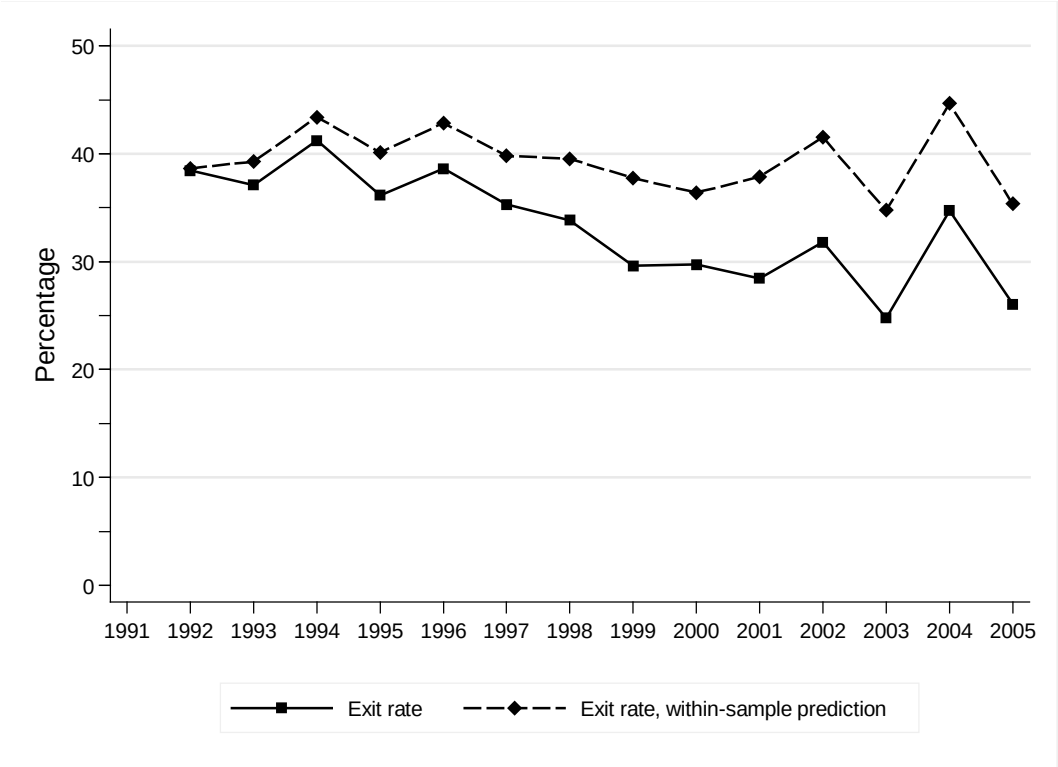
Notes: Authors' calculations using BHPS data. The definition of social assistance benefit receipt is explained later.

Figure 3. Within-sample predictions of SA transition rates compared to aggregate transition rates

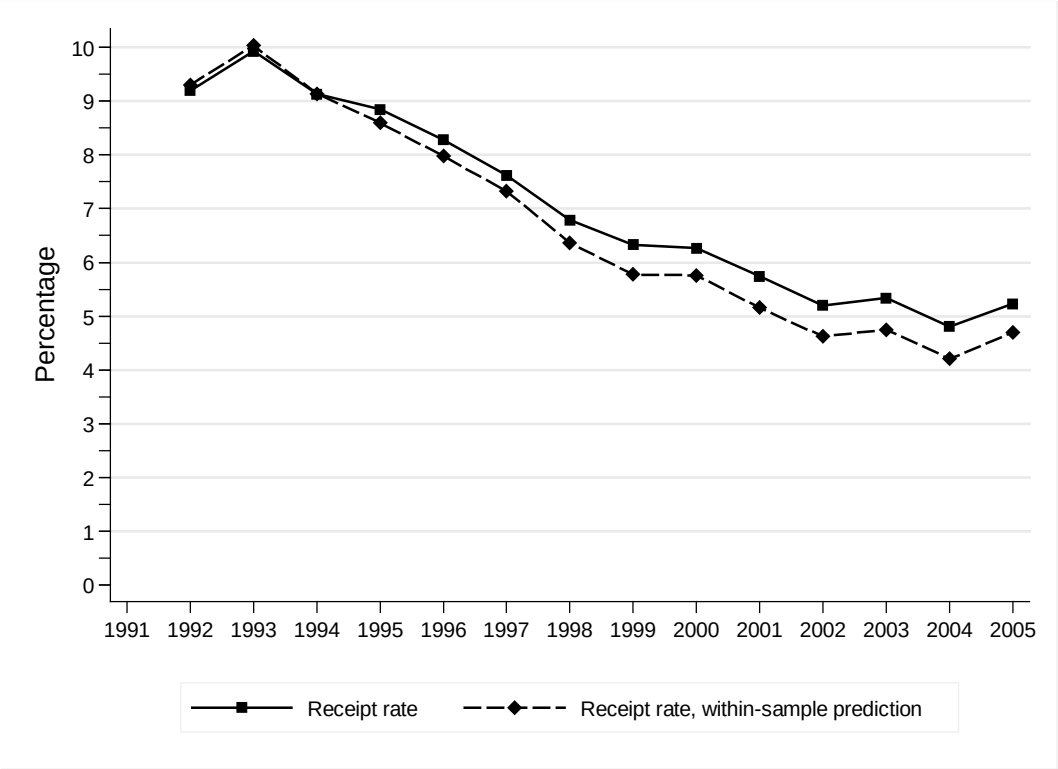
(a) Entry rates



(b) Exit rates



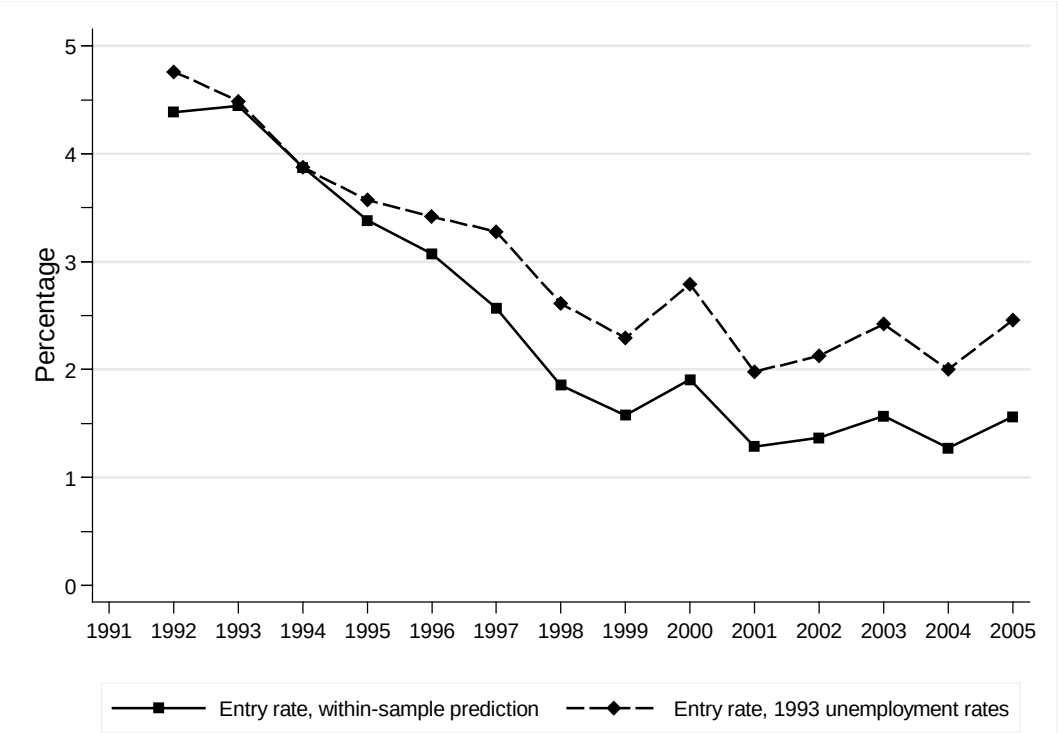
(c) Receipt rates



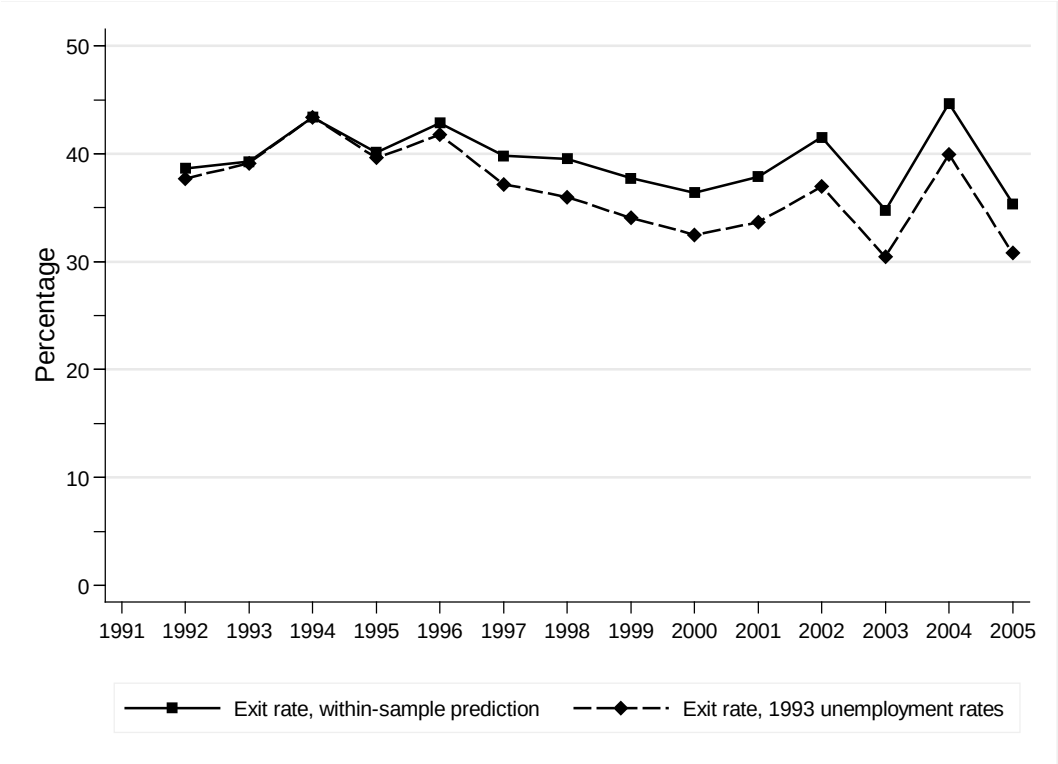
Note. Authors' calculations from BHPS data using the parameter estimates shown in Table 3. The construction of the series is explained in the text.

Figure 4. Counterfactual simulations of SA transition rates: what if local area unemployment rates were fixed at their 1993 values?

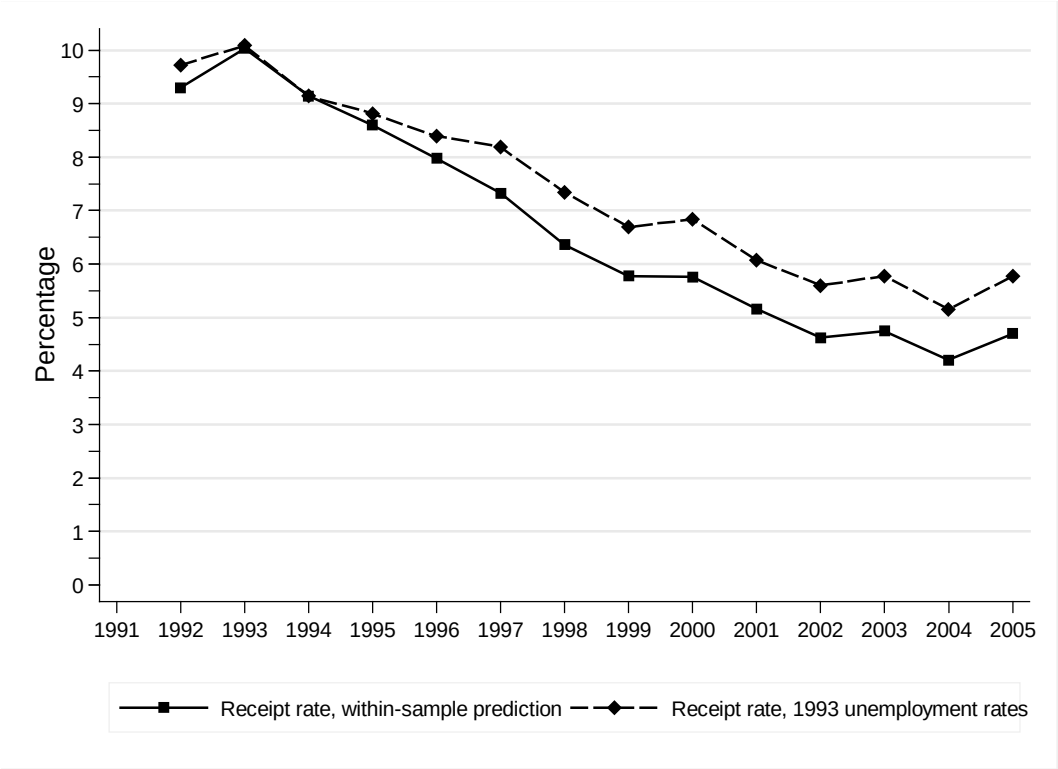
(a) Entry rates



(b) Exit rates



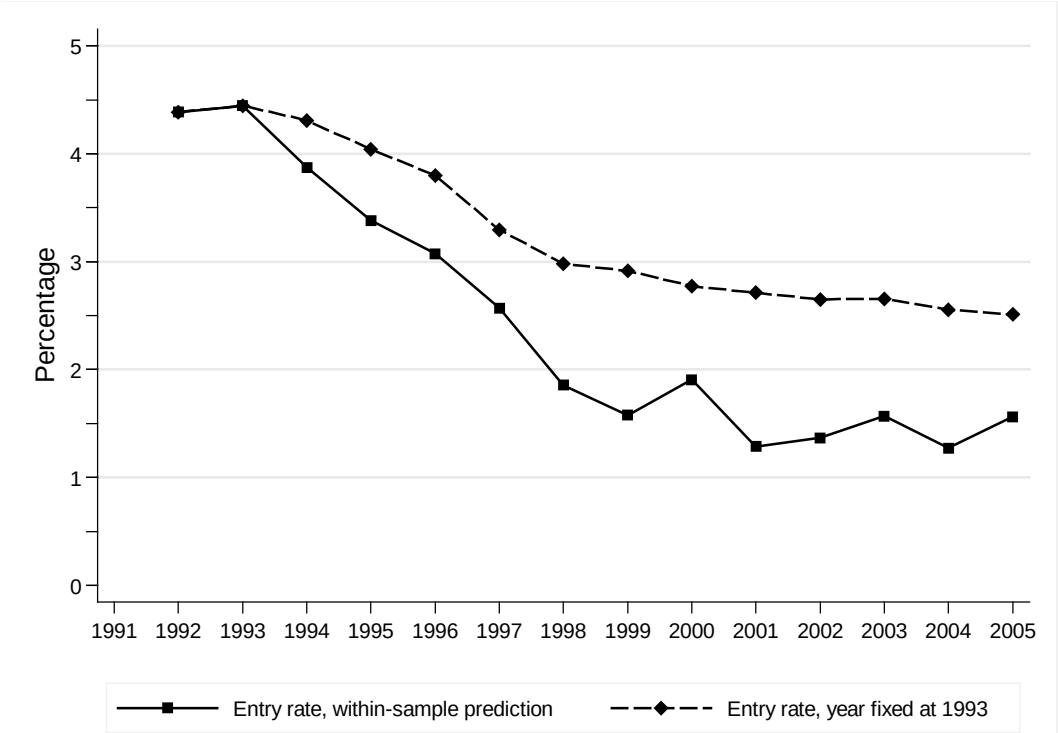
(c) Receipt rates



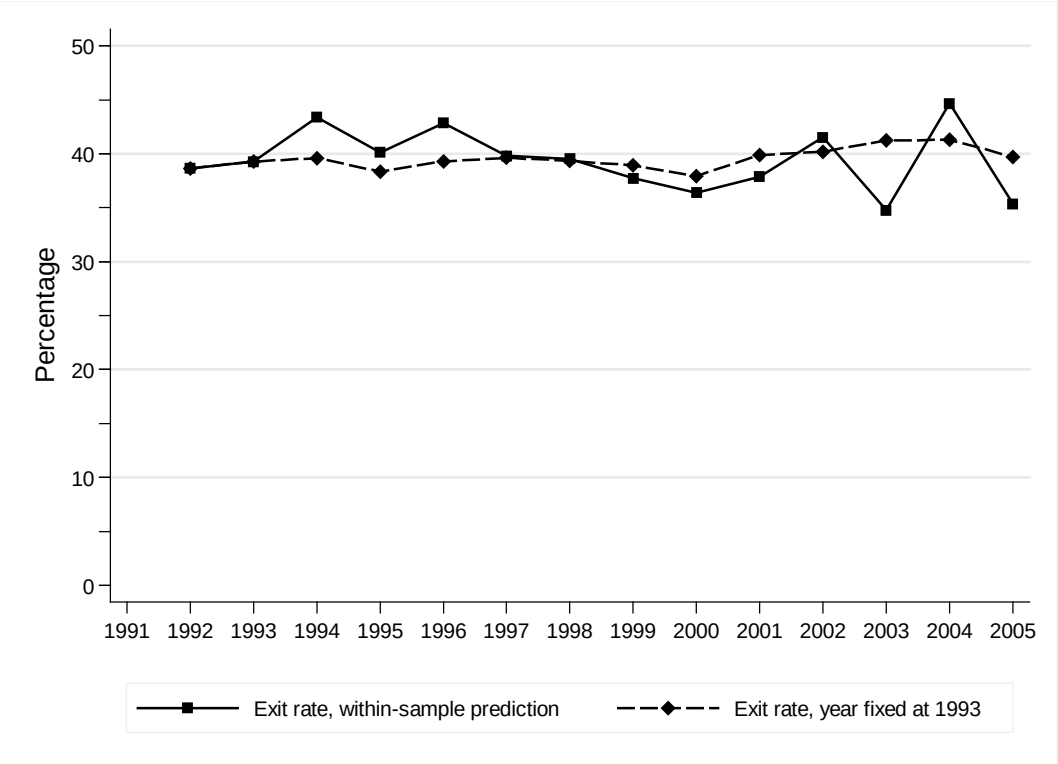
Note. Calculations from BHPS data using the parameter estimates shown in Table 3. The construction of the series is explained in the text.

Figure 5. Counterfactual simulations of SA transition rates: what if the survey year intercepts were all fixed at the value for 1993?

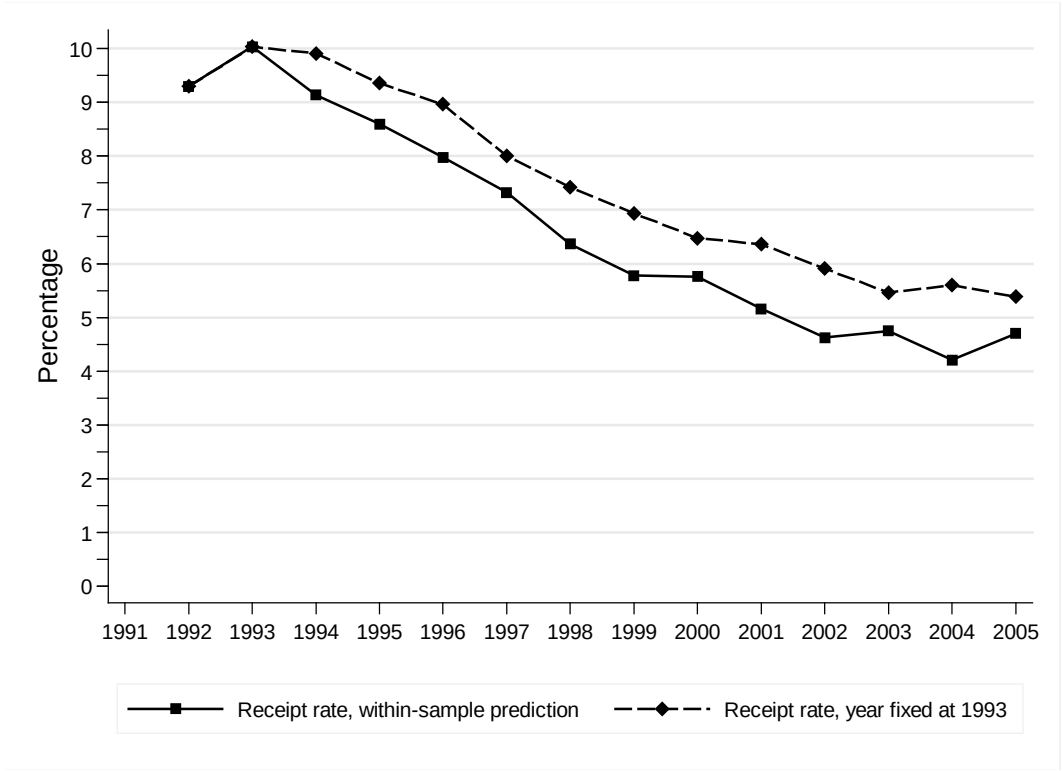
(a) Entry rates



(b) Exit rates



(c) Receipt rates



Note. Calculations from BHPS data using the parameter estimates shown in Table 3. The construction of the series is explained in the text.