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Zhiyang Jia and Trine E. Vattø

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Abstract:
The standard static labor supply model ignores that it takes time for individuals to adjust to a tax-benefit reform. A labor supply decision model is developed that allows for lagged responses in terms of state dependence, stemming from preferences, labor market constraints and adjustment costs. The parameters of the model are estimated using panel data on working hours for Norwegian females. We find evidence of all three sources of state dependence, with adjustment costs as the most dominant component. When using the model to simulate the path of adjustment to a general tax cut, we find that state dependence brings down responses to only one third of the estimated full effect in the first year. The females reach the proximity of the full effect after about seven years.

Keywords: labor supply, path of adjustment, state dependence, adjustment costs, discrete choice model, tax-benefit microsimulation

JEL classification: C35, C51, H24, H31, J22

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Address: Zhiyang Jia, Statistics Norway, Research Department. E-mail: jia@ssb.no

Trine Engh Vattø, Statistics Norway, Research Department. E-mail: tev@ssb.no

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Sammendrag

Det følger av resultatene fra arbeidstilbudsmodellen LOTTE-Arbeid i SSB at aktørene antas å endre tilpasning på arbeidsmarkedet umiddelbart etter skatteendringen. Det er en forenkling, og vi har derfor ønsket å se nærmere på hvordan det faktiske tidsforløpet kan se ut. Formålet med denne studien er derfor å utvikle en strukturell arbeidstilbudsmodell, som tillater et tidsetterslep av tilpasninger etter reformer i skatte- og overføringssystemet. Vi definer tre former for tilstandsavhengighet i arbeidstilbudet som gir en treghet i tilpasningene. Disse kan tilskrives preferanser, restriksjoner på arbeidsmarkedet og faste tilpasningskostnader. Mens tilstandsavhengighet i preferansene er definert som et skift i preferanser for fritid, er tilpasningskostnadene definert som en fast engangskostnad ved å skifte arbeidstid (som ofte innebærer skifte av jobb). Dette betyr at tilstandsavhengighet i preferansene vil påvirke arbeidstakere til å reagere gradvis på endringer i skatte- og overføringssystemet, ved for eksempel å bevege seg fra ikke-deltakelse, til deltid, og videre til fulltid, etter hvert som preferansene for arbeid langsomt endres. Tilpasningskostnader, derimot, vil fange opp at et skift i arbeidstid er kostbart, uavhengig av hvor mye du beveger deg bort fra det opprinnelige arbeidstilbudsvalget. Tilstandsavhengighet i restriksjoner på arbeidsmarkedet kan identifiseres som en ytterligere treghet i tilpasningene når en arbeidstaker ønsker å komme i arbeid, etter å ha stått utenfor arbeidsmarkedet en periode. Modellparameterne er estimert ved å benytte et panel med observert arbeidstid over perioden 2003-2007, for norske, samboende kvinner. Vi finner at alle tre former for tilstandsavhengighet har betydning, selv om vi kontrollerer for observerbar og uobserverbar heterogenitet mellom arbeidstakerne. Det ser ut til at faste tilpasningskostnader dominerer. Simuleringsresultatene viser at tilstandsavhengighet reduserer første års arbeidstilbudsresponser etter et skattekutt til bare en tredjedel av responsene på lang sikt. Først om lag 7 år etter en reform i skatte- og overføringssystemet har aktørene etablert seg i nye tilpassede arbeidstilbudsvalg.

I Introduction

Structural labor supply models are typically silent on whether it takes time for individuals to adjust to a tax-benefit reform. A few studies provide evidence on gradual rather than immediate adjustment in labor earnings using quasi-experimental identification (Holmlund and Söderström, 2011; Gelber et al., 2013; Schroeder, 2016 and Vattø, 2016) or by using simulation techniques (Haan, 2010). In this study, we aim to improve the understanding of the path of labor supply adjustment to tax-benefit reforms, by developing a decision model which explicitly accounts for possible sources of lagged responses.

When workers' prior labor market attachments have a causal effect on their current labor market decision, it is labelled as 'true state dependence' (Heckman, 1981). This should not be confused with so-called spurious state dependence, caused by individual heterogeneity, which does not imply lagged labor supply responses. Previous studies on state dependence and labor supply dynamics, such as Hyslop (1999) and Prowse (2012), typically apply dynamic probit or logit models to obtain alternative-specific coefficients for individual heterogeneity and state dependence. Whereas most of these studies only consider participation in the labor market, Prowse (2012) provides evidence of state dependence and unobserved heterogeneity in a multinomial choice framework, with three employment states (non-participation, part-time and full-time).

In order to develop a behavior model for policy evaluation, it is appropriate to model individual choices as a function of the characteristics of the choice alternative, rather than as a function of individual characteristics, as pointed out by Hoffman and Duncan (1988). Along these lines, Haan (2010) applies a dynamic conditional logit model, in which individuals' choice is determined by characteristics of the choice alternative, such as working time and level of disposable income. He thereby extends a standard labor supply model for policy evaluation to allow for state dependence. After controlling for unobserved heterogeneity in preferences for leisure and consumption, he derives positive parameter estimates of state dependence using German survey data. Subsequently, he uses the estimated model to simulate a gradual path of adjustment to a new prospective tax schedule. State dependence is, however, modelled in a reduced form manner, by allowing for interaction dummies for each combination of transitions. Haan et al. (2015) estimate a similar model by Bayesian techniques, whereas Haan and Uhlendorff (2013) extend the model to account for involuntary unemployment (for males).

We follow a similar approach as Haan (2010), but go further in modelling state dependence. We define three possible sources of state dependence in labor supply decisions: first, prior labor market attachment affects individuals' current evaluation of leisure (state dependence in preferences); second, prior participation affects labor market constraints such as the individual's job opportunities (state dependence in labor market constraints); and third, prior labor market decisions define status quo for the individual, and consequently a move away from it comes with a cost (adjustment costs). Our study is, to our knowledge, the first decision model to interpret labor market constraints and adjustment costs, as discussed by Blundell et al. (2008), Chetty et al. (2011) and Chetty (2012), as components of state dependence in a random utility framework. The benefit of modeling is twofold: Firstly, we provide new insights into sources of lagged tax responses. Secondly, we provide a structural decision model which is better equipped to simulate the path of labor supply responses under prospective tax-benefit reforms.

Our decision model is estimated on panel data obtained from Norwegian administrative registers over the period 2003-2007. We focus on the behavior of married and cohabiting women, as this group is characterized by relatively high variation in working time. Moreover, married women are often the secondary earner in the household and typically assumed to be more responsive to financial incentives. The data show strong individual persistence in labor market attachment over time, both in terms of the participation decision (extensive margin) and the number of hours worked (intensive margin).

We find evidence for all three sources of state dependence in labor supply decisions, although controlling for persistence caused by observed and unobserved individual heterogeneity. According to our estimated model, adjustment costs seem to be the most important contributor to state dependence. We simulate (out-of-sample) the effect of a tax cut on the path of labor supply adjustment. The simulation results suggest that state dependence brings down the first year effect to only one third of the long run effect. We find that the females reach the proximity of the full effect after about seven years. According to our results, state dependence, including adjustment costs, does not prevent individuals from adjusting in the long run, as the predictions from our model eventually reach predictions from a static model without state dependence.

This paper is organized as follows. In Section II, we discuss the sources of state dependence and lagged tax responses in more detail. The model and its empirical specification are described in Section III, whereas the data used for estimation is presented in Section IV. In Section V we discuss the estimated model and the simulation results. In particular, we describe how each source of state dependence contributes to persistence, and present simulated paths of adjustment. Section VI concludes the paper.

II Sources of State Dependence and Lagged Responses to Tax-Benefit Reforms

Before turning to a more formal outline of our model, we discuss in more detail the possible sources of state dependence in labor supply decisions. When we use the term 'state dependence' in the following, we refer to the causal effect of previous labor supply decisions, or so-called true state dependence as defined by Heckman (1981). Such effects imply lagged responses to tax-benefit reforms.

Although there is strong evidence of positive state dependence in labor supply decisions (Eckstein and Wolpin, 1989; Hyslop, 1999; Okamura and Islam, 2009; Haan, 2010; Prowse, 2012), there is surprisingly little empirical evidence on what causes state dependence.

Recent influential studies by Chetty et al. (2011) and Chetty (2012) emphasize the role of frictions, such as adjustment costs, as explanations for the modest microeconometric estimates of labor supply elasticities. The focus is on equilibrium effects, in a static manner, without discussing whether the effect of adjustment costs attenuate over time. Our study represents a new perspective on this issue by interpreting adjustment costs in a random utility framework.

Remember that we define three sources of state dependence, each of which can be separately sorted out in our model framework, under suitable assumptions. First, state dependence in preferences arises if leisure, or more precisely, the marginal rate of substitution between consumption and leisure, is altered by past experience. For instance, individuals who start to work part-time may find that the extra free time is more desirable than they had previously thought, thus placing higher value to the alternatives with part-time working hours in the future. The standard assumption in economic theory that preferences are stable can be questioned, and empirical evidence seems to suggest that past experience does indeed affect preferences (Neuman et al., 2010; Adams et al., 2015).

Second, state dependence in labor market constraints appears if job opportunities are altered by past labor market experience. For instance, an individual without job experience in the previous period may have less job opportunities in the current period than an otherwise identical worker. This may result from employers considering previous non-participation as means of depreciation of human capital or as a signal of low productivity. Reduced job opportunities, in turn, limit the possibilities for a non-participant to re-enter the labor market, and thus lead to persistence in non-participation, and lagged responses to policy reforms that encourage participation.[1]

Third, in our framework, adjustment costs capture fixed costs of deviating from status quo, defined as the previous labor supply decision. These costs might include the cost of negotiating a new contract with an employer, time and financial cost of job search (Gelber et al., 2013), informational costs (Chetty et al., 2013; Matejka and McKay, 2014), and mental costs of even considering and recognizing the choice situation (Samuelson and Zeckhauser, 1988).

Although we are not able to distingush between the detailed explanations within each category of state dependence, we believe that our framework represents an enhanced understanding of the main sources of lagged responses to tax-benefit reforms.[2]

III The Model

In order to develop a decision model that accounts for lagged responses to tax-benefit reforms, we extend a static model framework described in Dagsvik and Jia (2016). We provide a brief review of the static model, before extending the model to allow for sources of state dependence.

[1] Note that labor market constraints arise from the demand side of the labor market, under the implicit assumption of inelastic demand for productive workers.

[2] An additional source of state dependence, which is not covered in our baseline model, is state dependence in the budget constraint, arising from state dependence in wages. In a robustness check (reported in Section V), we find that accounting for this additional source of state dependence does not substantially influence our results.

The static model

Discrete choice models of labor supply derived from a stochastic utility representation and a discretized set of feasible hours of work, see e.g., van Soest (1995) and Creedy and Kalb (2005), have gained popularity because they are more practical than the traditional approaches based on marginal calculus. We use a particular type of discrete choice models, denoted the job choice model, see e.g. Dagsvik and Strøm (2006), Dagsvik et al. (2014) and Dagsvik and Jia (2016). According to this framework, labor supply decisions are viewed as an outcome of individuals choosing among jobs, with additional constraints on the set of available jobs.

The job choice model is specified as follows: Each individual is assumed to have preferences over a set of 'jobs', where each job (indexed by $z = 1, 2, ...$) is characterized by disposable income $C(z)$, hours of work $h(z)$, and other non-pecuniary job attributes, such as e.g. the nature of the job-specific tasks to be performed and the location of the workplace. Disposable income for a given job is defined as $C(z) = f(h(z)w(z), I)$, where $w(z)$ is the offered wage rate for the given job z, I is the non-labor income and $f(.)$ is the net-of-tax function. For simplicity the offered wage rate $w(z)$ is assumed to be constant across jobs for a given individual.[3] The individual's utility of choosing job z is represented as $U(C,h,z)$, where the utility function is assumed additively separable, namely $U(C,h,z) = v(C,h) + \varepsilon(z)$.

The sets of available jobs, from which the individuals choose, are individual-specific. Dagsvik and Jia (2016) show that it is sufficient to identify the model by introducing a measure of job opportunities, which represents the number of available jobs for a given working time alternative h, $m(h)$, where the number of non-working opportunities is normalized to one, i.e., $m(0) = 1$.

Hours of work for each job takes value in a given set H. It can be shown that under the assumption of extreme value distributed error terms, $\varepsilon(z)$, the probability that a worker chooses one of the jobs with working time $h \in H$, can be written as,

$$\varphi(h) = \frac{m(h)exp(v(C(h),h))}{exp(v(C(0),0)) + \sum_{x \in H'} m(x)exp(v(C(x),x))} \tag{1}$$

where H' represents the subset of positive working hours. This expression is analogous to a

[3] It turns out that when allowing for job-specific wage rates, it is more difficult to account for individual differences in the wage rates, see Aaberge et al. (1995) for an alternative approach.

multinomial logit model, with representative utility, $v(C,h)$, weighted by the number of available jobs, $m(h)$. Since $m(h)$ is not observable, it is estimated simultaneously with the systematic part of the utility function, $v(C,h)$.

Extending the static model to allow for state dependence

Now, we extend the static model framework to introduce dependence in decisions over time $(t = 1,...,T)$. Individuals are still assumed to be myopic, and to make their decisions based on a period to period optimization. But as discussed in Section II, we allow for the path of decisions being influenced by state dependence. More specifically, we let prior labor market attachment influence the current labor supply decision in terms of preferences, job opportunities and adjustment costs. Following e.g. Haan (2010) and Hyslop (1999), we assume, for simplicity, that state dependence forms a Markov chain over time, which essentially means that only the last period, and not the whole work history, matters for current decisions.

Following Dagsvik and Strøm (2006) and Dagsvik and Jia (2016), we assume that the deterministic part of the utility function can be represented by a Box-Cox function,

$$v(C_t,h_t\,|\,h_{t-1}) = \alpha_0\frac{(C_t-C_0)^{\alpha_1}-1}{\alpha_1} + \beta_0(h_{t-1})\frac{(\bar{h}-h_t)^{\beta_1}-1}{\beta_1} \tag{2}$$

where C_t is disposable income given by $C_t = f_t(h_t w_t, I_t)$. C_0 represents the minimum or subsistence household-adjusted consumption level, here set to $C_0 = 50,000$ NOK (about $6,200$ EUR). $\bar{h}$ is defined as 80 hours per week, such that $(\bar{h}-h_t)$ measures leisure time.

The extension to allow for state dependence in preferences is obtained by allowing β_0 to depend on the previous working time decision, h_{t-1}. We let $\beta_0(h_{t-1})$ be specified as follows,

$$\beta_0(h_{t-1}) = b_0 + \mathbf{b_1'x_t} + \mathbf{b_2'I}(h_{t-1}) + \mathbf{b_3'I}(h_0) \tag{3}$$

where $\mathbf{x_t}$ is a vector of observed individual characteristics (age and number of children), $\mathbf{I}(h_{t-1})$, is a unit vector of the individual's previous working time alternative, and $\mathbf{I}(h_0)$, is a unit vector of the initial working time.[4] It follows that if $\mathbf{b}_2' \neq \mathbf{0}$, then the previous working time decision affect current preferences for leisure, and thus represents what we refer to as state dependence in preferences.

A novelty of the discrete choice model applied here is the job opportunity measure, $m(h_t \,|\, h_{t-1})$, which allows us to explicitly model state dependence in labor market constraints. According to the static framework, we assume that the job opportunity measure can be decomposed as follows,

$$m(h_t) = \theta_t g(h_t) \tag{4}$$

where θ_t describes the total number of jobs available (relatively to non-participation), while $g(h_t)$ is interpreted as the fraction of jobs available to the agent with offered hours of work equal to h.

We let $g(h_t)$ be uniformly distributed among working time alternatives, except for a possible peak (estimated within the model) for full-time jobs. This essentially means that there are equally many jobs with short and long part-time to choose from, and a larger number of full-time jobs available.

An important identifying assumption for this demand side restriction is that the relative number of jobs with various working time is not individual- and time-specific. This can be justified as $g(h_t)$ is considered to stem from institutional restrictions as a result of centralized negotiations between labor unions and employers' associations. Institutional settings are relatively stable over time, and it is not likely that an individual's past experience could have a strong effect on this.

State dependence and individual heterogeneity in job opportunities is accounted for in the total number of jobs available (relatively to non-participation). We specify $\theta_t = \theta_t(h_{t-1})$ according to how it enters the log-likelihood function as,

[4] Wooldridge (2005) suggests estimating unobserved heterogeneity (see next subsection) conditionally on the initial period observations. A similar strategy for the structure of state dependence and initial conditions can be found in Haan et al. (2015).

$$ln(\frac{1}{\theta_t(h_{t-1})}) = \gamma_0 + \gamma_{11}(V_r/U_r)_t + \gamma_{12}edu_t + \gamma_2\mathbf{I}(h_{t-1}) + \gamma_3\mathbf{I}(h_0) \tag{5}$$

where V_r/U_r refers to the vacancy to unemployed ratio which is a measure of regional labor market tightness, and edu_t symbolizes the individual's education level.

Note that $\theta_t(h_{t-1})$ does not differ across current working time alternatives, only between the working and non-working alternatives, as the number of non-working opportunities is normalized to one, i.e., $\theta_t(h_{t-1})g(h_t = 0) = 1$. This means that state dependence in opportunities has an impact on the current labor market decision only in terms of participation, leading to possibly lagged responses in the transition from non-participation to participation.

Lastly, we introduce adjustment costs, $AC(h_t|h_{t-1})$, which represents a fixed disutility caused by a change in working time from period $t-1$ to t,

$$AC(h_t|h_{t-1}) = \begin{cases} ac \text{ if } h_t \neq h_{t-1} \\ 0 \text{ otherwise} \end{cases} \tag{6}$$

where ac is a constant.[5]

We let adjustment costs enter the decision model as an additive part of the deterministic utility function.[6]

$$\tilde{v}(C_t, h_t \,|\, h_{t-1}) = v(C_t, h_t \,|\, h_{t-1}) - AC(h_t \,|\, h_{t-1}) \tag{7}$$

Remember that we define state dependence in preferences as an upward or downward shift in the preferences for leisure, whereas adjustment costs are characterized as one-time fixed costs of altering working time from last period's decision. Note that although adjustment costs might as well originate from preferences, it represents another explanation for lagged responses. State dependence in preferences induces individuals to slowly react to incentive changes in a stepwise fashion, going from non-participation to part-time and further to full-

[5] Our definition of adjustment costs relates to so-called switching cost in market science, where consumers have costs of switching between competing firms' product (see e.g., Klemperer, 1995 and Osborne, 2011).

[6] Chetty (2012) defines adjustment costs similarly, as the utility loss agents can tolerate to deviate from the frictionless optimum. Note that his definition might also cover what we define as state dependence in preferences.

time, as the preferences for leisure are influenced by previous working time. Adjustment costs on the other hand, simply captures that a change is costly, independently on in what direction the adjustment takes place or how radical the change is. Also adjustment costs lead to gradual responses in our model, although not at the individual level, because each individual has a positive probability of overcoming the negative effect of the adjustment costs each period.

Now, $\tilde{v}(C_t, h_t \mid h_{t-1})$ and $m(h_t \mid h_{t-1}) = \theta_t(h_{t-1})g(h_t)$ enter into eq. (1), to define the probability (conditional on h_{t-1}) of choosing a job with working hours h in period t.

$$\varphi(h_t \mid h_{t-1}) = \frac{g(h_t)exp(\tilde{v}(C_t(h_t), h_t \mid h_{t-1}))}{\theta_t(h_{t-1})^{-1}exp(\tilde{v}(C_t(0), 0 \mid h_{t-1})) + \sum_{x \in H'} g(x_t)exp(\tilde{v}(C_t(x_t), x_t \mid h_{t-1}))} \tag{8}$$

Recall that $\theta_t(h_{t-1})^{-1}$ appears only at the extensive margin, as reflected by eq. (8). It follows that state dependence in job opportunities is identified as an additional lag in responses, when deciding to participate in the labor market after a period of non-participation.

Unobserved heterogeneity and related sources of persistence

Persistence in unobserved individual heterogeneity leads to spurious state dependence, which might bias our estimates of true state dependence (Heckman, 1981). In order to avoid this problem, we model persistent unobserved heterogeneity in a latent class framework. We assume that individuals can be classified into K different (unobserved) types, for which some key parameters differ. The type of an individual is assumed to be constant over time. The fraction of individuals of type $k = (1, .., K)$ is estimated within the model by $p_k \in (0,1)$. This leads essentially to a finite mixture model and has the advantage that unobserved heterogeneity can be handled with flexibility, without imposing a parametric structure.

Persistent individual heterogeneity creates a pattern of serial correlation in the error terms of the utility functions. A more general structure of serial correlation or the related concept of habit persistence, as defined by Heckman (1981), cannot be separately identified in our model. Remember that state dependence implies that it is the *actual choice* made in the past, h_{t-1}, which has a direct effect on the valuation of the current choice. In contrast, habit persistence implies that it is the *valuation* of different alternatives in past periods, $U_{t-1}(C,h,z)$, which influences the valuations in this period. This might be related to serial correlation in the error terms, $\varepsilon_t(z)$, which arises as a result of random shocks that last longer than one period, or if unobserved individual characteristics change slowly over time. As these sources of persistence are not separately identified in our conditional logit framework, we cannot rule out that our estimates of state dependence are influenced by this. However, a number of empirical studies seem to find that serial correlation and habit persistence is of less importance than state dependence (Hyslop, 1999; Seetharaman, 2004; Prowse, 2012). Moreover, as we have defined state dependence quite restrictively, we do not think the exclusion of these remaining aspects of persistence is crucial for our interpretation of the results.

In our baseline model, we let preferences, opportunity measure and adjustment costs differ between two unobserved types of individuals ($K = 2$) in terms of heterogeneity by the following four parameters: $\{\alpha_0, b_0, ac, \gamma_0\}$.[7]

Now, the joint likelihood for an individual to choose the sequence $(h_1, h_2 \cdots h_T)$ can be written as

$$\varphi(h_1, h_2 \cdots h_T) = \sum_k p_k \prod_t \varphi_k(h_t | h_{t-1}) \tag{9}$$

where the conditional probabilities for each type are given by eq. (8).

IV Data

Data and summary statistics

We use data from different connected administrative registers of Statistics Norway, over the period from 2003 to 2007, in order to estimate the model parameters. The focus is on married and cohabiting Norwegian women aged 25−62 years. Unemployed, self-employed, disabled persons and students are excluded from the sample. We further limit the analysis to women

[7] We provide robustness checks with K=3 and K=4 to demonstrate that our results are not sensitive to the number of unobserved types. For simplicity, and in the interest of an apparent and clear-cut presentation of the results, we use K=2 in our baseline model.

whose partners' total pre-tax income level exceed NOK 150,000 (about EUR 19,000) per year, in order to abstract from effects of various welfare transfers. The remaining data set consists of about 300,000 observations each year.

We have information on working time for about 70 percent of the observations each year, which is based on employers' questionnaire reports and administrative registers (Statistics Norway, 2006). In order to avoid attrition and selection effects, we impute the remaining missing information (see Appendix).

For all individuals, we discretize the information on working time by dividing into five categories based on weekly hours of work: $h \in \{0, 1-19, 20-34, 35-40, 40+\}$. Each individual's disposable income ('consumption'), for each possible working time alternative for each year, is constructed by adding simulated wage income (based on the hourly wage rate and median working time in each alternative) to capital income, transfers and the partner's income, and deduct taxes.

For practical reasons, we select a random ten percent balanced sample for use in the estimations. Table 1 provides an overview of the main characteristics of this sample.

Figure 1 describes the observed fraction of individuals in each working time category over the period 2004-2007. On average about 6 percent are not participating in the labor market, 19 percent work short part-time, 31 percent work long part-time, 36 percent work full-time and 8 percent work overtime. It seems to be a trend towards increased participation and longer working hours, which is partly a result of the characteristics of a balanced sample. The pattern can also be explained within our model framework as a result of tax reductions and increased labor market tightness.[8]

[8] The model parameters are estimated over a reform period in order to have additional identifying variation in the tax rates.

TABLE 1

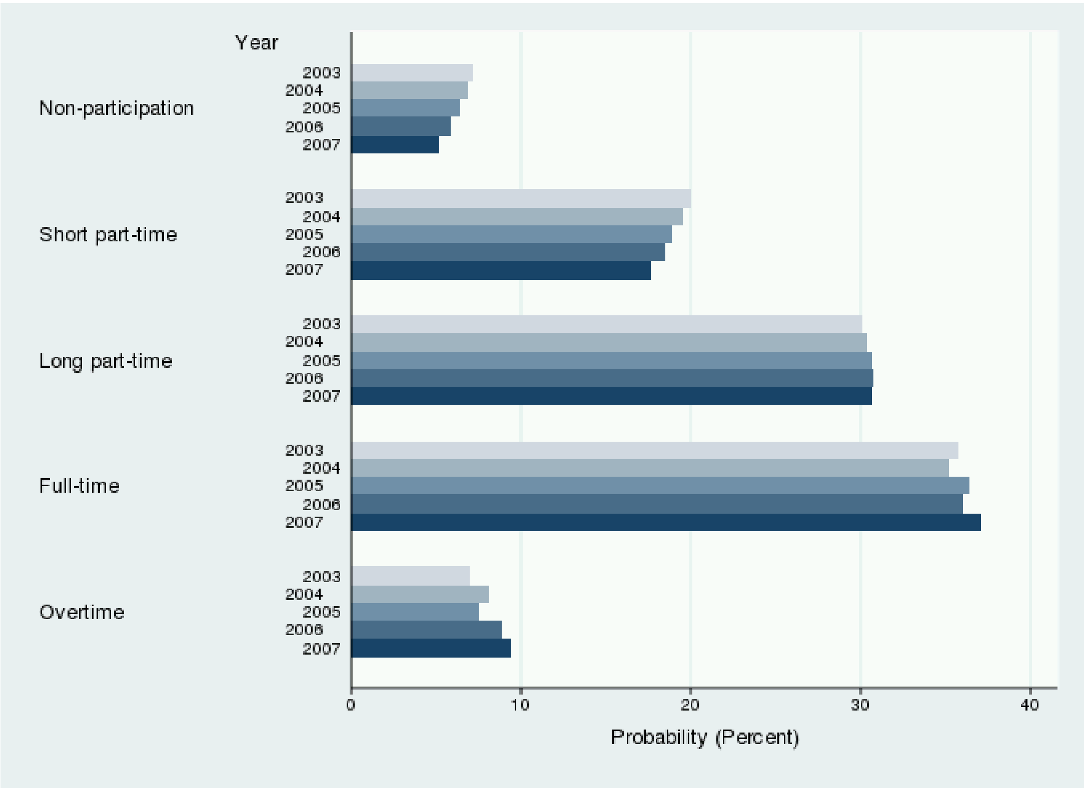
Characteristics of the pooled sample 2003-2007

Variable	*Mean*	*Std. Dev.*	*Min.*	*Max.*
Imputed wage rate (NOK)	165.1	35.5	58.4	256.4
Non-labor income (10^3 NOK)	32.7	566	-528.6	1.76e+05
Partner's gross income (10^3 NOK)	592.7	1664	150.0	2.97e+05
Age	44.7	8.9	26	62
Child(ren) under age 6	0.29	0.45	0	1
Child(ren) under age 12	0.48	0.50	0	1
Low education ($\leq$10 years)	0.40	0.49	0	1
Regional vacancy-to-unemployed ratio	0.28	0.20	0.07	1.08
Number of individuals		29,757		

Notes: Income and wage rates are adjusted to 2007-NOK (1 EUR≈8 NOK).

FIGURE 1

Observed choice probabilities



Observed persistence in working time

Transitions can be presented in a Markov transition matrix, which provides information about the fraction of individuals starting out in a certain working time alternative, who stay there the next period, and how many switch to other working time alternatives. The diagonal elements of the transition matrix can be seen as a measure of persistence, as it expresses the number of individuals staying (in percent), relative to those initially present in that working time alternative.

Persistence can also be measured in terms of stability as the number of individuals who are in a specific working time alternative both periods divided by the number of individuals who have been present in the working time alternative at least one of the two periods.[9] The persistence rate takes a value between 0 and 1, where higher value implies higher persistence and less mobility across alternatives. If there is completely mobility, where no one chooses the same alternative two periods in a row, the persistence rate equals 0, whereas if there are no mobility, where everyone chooses a given alternative over and over again, the persistence rate equals 1.

Table 2 presents the observed transition probabilities from 2006 to 2007 and the persistence rates for each working time alternative, as defined above. Similar patterns are observed for the other year-by-year transitions. We see that there is a strong persistence in observed labor supply decisions. The persistence rate is highest for non-participation (0.76) and lowest for overtime work (0.43). When comparing the extensive margin (non-participation) to the intensive margin (average of the working alternatives when weighted by the choice probabilities), we find that the persistence rate is 0.76 and 0.63, respectively. In the next section, we use these concepts to analyze how each source of state dependence contributes to persistence.

[9] This type of measure has, for example, been frequently used in the literature on school mobility, see e.g., Dobson et al. (2000).

TABLE 2

Observed transition probabilities and persistence rates, over the period 2006-2007

2006 \ 2007	Non-participation	Short part-time	Long part-time	Full-time	Overtime
Non-participation	**81.7%**	15.6%	1.5%	0.9%	0.4%
Short part-time	2.0%	**76.3%**	16.4%	4.2%	1.0%
Long part-time	0.1%	6.5%	**77.7%**	13.5%	2.2%
Full-time	0.0%	1.5%	8.9%	**81.2%**	8.4%
Overtime	0.0%	1.0%	4.9%	31.6%	**62.5%**
Persistence rate*	**0.76**	**0.64**	**0.64**	**0.67**	**0.43**

*The persistence rate is defined as the number of individuals who are in the specific working time arrangement both periods divided by the number of individuals who have been present in the working time arrangement at least one of the two periods.

V Estimation Results and Simulations

In the following we describe the main estimation results, including how each source of state dependence stemming from preferences, opportunity constraints and adjustment costs, contributes to persistence. Next, we show how the estimated model can be used to simulate the path of adjustment, and how this path differs across individuals according to observed characteristics. Finally, we demonstrate that the path of adjustment is robust to alternative model specifications of both the wage rate and unobserved heterogeneity.

Estimated parameters and model fit

The model is estimated by the method of maximum likelihood, where we summarize the likelihood contribution over individuals, given by eq. (9), where $t = 0$ refers to the initial conditions in year 2003 and $t = T$ refers to year 2007. This essentially means that the parameters of the utility function, the job opportunity measure, adjustment costs and the probabilities of each unobserved type of individual are estimated simultaneously in order to maximize the probability of the observed path of labor market decisions. For each individual, the observed path of labor market decisions is given by the sequence of working time (5 categories) over the 4-year period

(2004-2007), such that $5^4 = 625$ different paths of working time are possible.[10] The estimated parameters are reported in Table 3.

The utility function turns out to be concave and increasing with respect to consumption and leisure for all individuals, and the job opportunity measure has the expected sign. We find evidence of individual heterogeneity in preferences and job opportunities. In particular, we find observed heterogeneity in preferences related to age and the presence of children, and in terms of unobserved heterogeneity in the job opportunity measure.

Since parameter estimates of $\mathbf{b}_2' = (b_{22}, b_{23}, b_{24}, b_{25})$ are negative and increasing in absolute value ($b_{2j+1} > b_{2j}$), individuals who worked long hours previous period value leisure less in the present period, which implies that there is positive state dependence in preferences. Moreover, significant negative estimates of $\gamma_2' = (\gamma_{22}, \gamma_{23}, \gamma_{24}, \gamma_{25})$ imply positive state dependence in labor market constraints, as individuals who worked in the previous period obtain a larger job opportunity measure, and are thereby more likely to participate in the labor market the current period. The effect is less strong if only working short part time the previous period (γ_{22}). The estimates of adjustment costs, (ac_1, ac_2), are positive and significant for both unobserved types. This suggests that altering the previous period's labor supply decision induces a loss, interpreted as adjustment costs.

To facilitate model predictions, we use the so-called empirical Bayes-method (Skrondal and Rabe-Hesketh, 2004; Train, 2009) to specify individual-specific probabilities (the posterior distributions) of belonging to each unobserved type, $k = (1, 2)$. The individual weights for each type are defined as $w_k = p_k \prod_t \varphi_k(h_t|h_{t-1}) / \sum_{j \in \{1,2\}} p_j \prod_t \varphi_j(h_t|h_{t-1})$.

Predictions of the estimated model reproduce the observed choices very well. We find close correspondence between observed and predicted outcomes, with respect to both the probability distributions of working hours and with respect to labor supply transitions over the estimation period (2004-2007).

To test out-of-sample predictions of the model is challenging, as the model is not designed to control for macroeconomic fluctuations and other institutional changes in the labor market. We therefore only conduct a simple out-of-sample comparison for the year 2008. Again we

[10] Data for year 2003 are used for initial conditions only.

find a close correspondence between model predictions and observed outcome, see Figure A.1 and Table A.2 in the Appendix.[11]

How each source of state dependence contributes to persistence

As already denoted, model estimates suggest that there are contributions to state dependence from all three components of preferences, labor market constraints and adjustment costs. However, standard statistical tests on the coefficients cannot provide useful information on the relative importance of the three components. We therefore suggest an alternative approach by measuring the degree of persistence caused by each component of state dependence. To do this, we compare the simulated persistence rate (defined in Section IV) obtained from three constrained models, where we sequentially leave out each component of state dependence from the full baseline model.

To be more precise, based on the estimated baseline model, we perform three simulations where we impose the following restrictions: (i) Past labor supply decisions have no impact on the preferences for leisure, i.e., replace b_{2j} by a constant average, $\bar{b_2}$; (ii) Past labor supply decisions have no impact on job opportunities, i.e., replace γ_{2j} by a constant average, $\bar{\gamma_2}$; (iii) Past labor supply decisions have no impact on adjustment costs, i.e., replace ac_1 and ac_2 by 0.

In Table 4 we report the persistence rate at the extensive and intensive margin for the baseline model, and for each alternative case, (i)-(iii). Note that the observed persistence rates reported in Section IV (0.76 and 0.63) are almost perfectly reproduced by the baseline model (0.77 and 0.61). The difference in the estimated persistence rate for each alternative case, compared to the baseline model, is labeled as 'loss' and can be interpreted as the contribution of the specific source of state dependence. The relative importance of each component of state dependence are ranked accordingly.

Overall we find that adjustment costs are the most important contributor to persistence, as the rates drop extensively from 0.77 to 0.38 and from 0.61 to 0.26 at the extensive and inten-

[11] We acknowledge that, as the year 2008 had no large changes in the tax schedule, a static model without state dependence also generate a reasonable out-of-sample fit in this case. An alternative test of model fit is to compare model predictions by quasi-experimental findings on adjustment to tax reforms, as seen in Thoresen and Vattø (2015). An approach along these lines is not within the scope of the present paper.

TABLE 3

Baseline model coefficients

	Parameter	Coefficient	Std. error
Probability distribution $(\alpha_0, b_0, ac, \gamma_0)$			
Probability, type 1	$p1$	0.4592***	0.0110
Probability, type 2	$p2$	0.5408***	0.0110
Preferences, consumption			
Constant (Scale 10^{-4}) , Type 1	α_{01}	1.0270***	0.0553
Constant (Scale 10^{-4}) , Type 2	α_{02}	1.1886***	0.0733
Exponent	α_1	0.6512***	0.0157
Preferences, leisure			
Constant (Scale $1/80$) , Type 1	b_{01}	5.7626***	0.2747
Constant (Scale $1/80$) , Type 2	b_{02}	6.2598***	0.2859
Exponent	β_1	-2.7513***	0.0610
Taste modifiers			
Age (Scale 10^{-1})	b_{11}	-0.6826***	0.0797
Age squared (Scale 10^{-2})	b_{12}	0.0793***	0.0089
Child(ren) under age 6	b_{13}	0.2069***	0.0208
Child(ren) under age 12	b_{14}	-0.1533***	0.0205
State dependence			
Choice 2 period t-1	b_{22}	-1.8407***	0.1178
Choice 3 period t-1	b_{23}	-2.6012***	0.1260
Choice 4 period t-1	b_{24}	-3.2183***	0.1386
Choice 5 period t-1	b_{25}	-3.4015***	0.1442
Adjustment costs			
Type 1	ac_1	0.6794***	0.0232
Type 2	ac_2	2.9297***	0.0530
Opportunity measure			
Constant, Type 1	γ_{01}	0.3952***	0.0860
Constant, Type 2	γ_{02}	1.7667***	0.1641
Labor market tightness	γ_{11}	-1.3875***	0.1695
Low education ($\leq$ 10 years)	γ_{12}	0.0399	0.0630
State dependence			
Choice 2 period t-1	γ_{22}	-0.6745***	0.1128
Choice 3 period t-1	γ_{23}	-3.7672***	0.2595
Choice 4 period t-1	γ_{24}	-3.2675***	0.3004
Choice 5 period t-1	γ_{25}	-3.0217***	0.4664
Opportunity density			
Full-time peak	$g(h_t)$	0.4307***	0.0136

Notes: The estimation sample contains 29,757 individuals. Initial working time has been included accordingly as state dependence, in order to estimate unobserved heterogeneity conditionally on the initial period observations.

TABLE 4

The relative contribution of state dependence from preferences, opportunities and adjustment costs

		Extensive margin			*Intensive Margin*		
		Rate	*Loss*	*Rank*	*Rate*	*Loss*	*Rank*
	Baseline model	0.77			0.61		
(i)	No state dependence in preferences	0.61	0.16	3	0.52	0.09	2
(ii)	No state dependence in opportunities	0.45	0.32	2	0.60	0.01	3
(iii)	No adjustment costs	0.38	0.39	1	0.26	0.35	1

Notes: The persistence rate is defined as the number of individuals who are in the specific working time arrangement in both periods (here: 2006 and 2007) divided by the number of individuals who have been present in the working time arrangement at least one of the two periods (see Section IV). This can be expressed in state and transition probabilities according to the following formula for the rate of persistence in non-participation: $Rate(NP) = P(NP_t|NP_{t-1}) / \left(1 + (1/P(NP_{t-1})) \cdot \sum_{X=\{SP,LP,FT,OT\}} (P(NP_t|X_{t-1}) * P(X_{t-1}))\right)$ where NP, SP, LP, FT and OT refer to the five working time categories. The rate of persistence for the other working time alternatives can be computed accordingly by replacing NP by SP, LP, FT or OT.

sive margin respectively, when this component is excluded. State dependence in job opportunities also contributes substantially to explain persistence at the extensive (participation) margin, where the rate is reduced from 0.77 to 0.45. Recall that this component of state dependence does not affect the intensive margin decisions by definition, see Section III. State dependence in preferences is on average less important, but the persistence rates are still significantly reduced both at the extensive and intensive margin by 0.16 and 0.09 respectively, when excluded from the model.[12]

Simulating the path of adjustment

Simulated responses to a general tax cut

We now use the estimated model to simulate (out-of-sample) the effect of a tax cut on the path of labor supply adjustment. We describe the path of adjustment under both a permanent and a transitory cut in the general tax rate, and compare the results to simulations from a static

[12] Note that a decomposition of the persistence rate is not possible because of non-linearities. We can therefore not add up the contribution of each component in order to obtain the persistence rate of the baseline model.

version of the model, without state dependence. The static model without state dependence is estimated on the same data, pooled over the period 2004-2007.

We describe the adjustment path by developments in labor supply elasticities. The elasticities are obtained by simulating the average predicted working time over individuals before and after a general tax cut applies. They can be interpreted as the percentage change in mean working hours when the net wage is increased by 1 percent. Further, the labor supply elasticity is decomposed into a participation elasticity and an elasticity conditional on participation, measuring the extensive and intensive margins, respectively.

Now, in order to estimate labor supply elasticities over time, we start out from the observed initial choices of labor supply, and let the model simulate the labor supply decisions for the subsequent years according to the formula $\sum_k w_k \prod_t \varphi_k(h_t|h_{t-1})$, where w_k is the individual weight of type k. We compare a reference path, with no tax cuts, to two alternative paths: one reflecting a transitory (one period) tax cut and the other reflecting a permanent (all periods) tax cut. Individual characteristics, including non-labor income and gross wage rate, as well as the tax schedule (apart from the tax cut) are kept constant in the simulations. Thus, the labor supply elasticities reflect the effect of the tax cut only.

As individuals are assumed to optimize on a period-to-period basis, they respond equally to a transitory and permanent tax cut in the first period.[13] Also for the transitory tax cut, the labor supply effect will last for more than one period, because of state dependence in preferences, labor market constraints and adjustment costs.

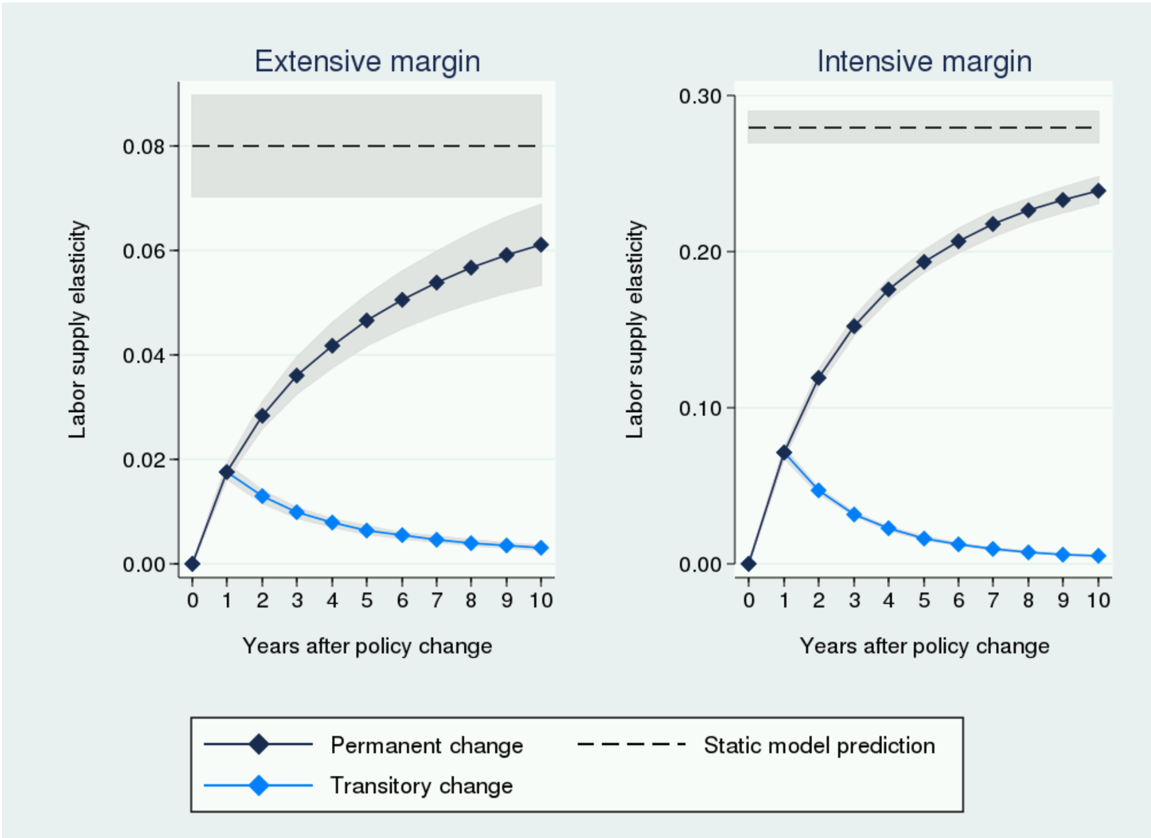
Figure 2 provides a graphical illustration of how the simulated labor supply elasticities at the extensive and intensive margins evolve over time. The intensive margin responses are of a larger magnitude than the extensive margin responses, but the transitions over time follow relatively similar paths.[14] At both the intensive and the extensive margins, the first year responses account

[13] In reality, individuals might respond differently already from the first period dependent on whether the change is expected to be transitory or permanent. However, in order to capture such behavior, one must allow individuals to form expectations about future tax changes, for instance in a life-cycle framework, such as Keane (2015).

[14] In general, one often finds that extensive margin responses are larger than intensive margin responses (see e.g. Heckman, 1993). However, elasticities depend on characteristics of the data sample because of non-linearities, see e.g. Dagsvik et al. (2014). High participation rates in the Norwegian context might explain the modest elasticities we find at the extensive margin.

FIGURE 2

Predicted labor supply elasticities over time



Notes: The shaded area represents the 95 percent confidence interval obtained by non-parametric bootstrapping, 50 repetitions.

for about 1/3 of the 10th year responses. The time until 90 percent of the 10th year outcome is achieved is about 7 years at the intensive margin and approximately 8 years at the extensive margin.

When comparing the elasticities obtained from our model to the prediction of a static model without state dependence, we see that the difference is especially pronounced in the short run. This suggests that state dependence has a considerable effect in the short run, by bringing down the first year effect to less than a third. It seems like our model eventually reach predictions from the static model in the long run, which suggest that state dependence, including adjustment costs, do not prevent individuals from adjusting in the long run.[15]

[15] See also Haan (2010) and Gelber et al. (2013) for similar conclusions.

Heterogeneity in lagged responses

The results reported so far are based on average effects over individuals. Now, we take a closer look at different subsets, based on observed characteristics of age and education level.

Remember that age is included as a taste modifier in the utility function, and the education level enters the specification of job opportunities. State dependence in preferences, opportunities or adjustment costs are not altered by observed characteristics of individuals, but the predicted path of responses might still be affected, as the weights assigned to the two unobserved types differ (w_k).

Given the latent class framework, we have two different unobserved types of individuals, each with a unique set of parameters $(\alpha_0, b_0, ac, \gamma_0)$, with estimates reported in Table 3. One type of individuals has larger adjustment costs ($ac_2 > ac_1$), and thus reacts considerably slower than those of the other type. This is reflected by predicted transition probabilities for each type in Table A.3 in the Appendix, where the diagonal elements (estimated probability of staying) are substantially larger for this type.

In Table 5, we report the short-run (1st year) and long-run (15th year) labor supply elasticities for these different groups of individuals. The ratio of short run responses in percent of long-run responses indicates how fast individuals adjust. We find that young and high educated women are more likely to be of the rapidly responding type, with lower adjustment costs and a larger job opportunity measure than older and lower educated females. This is reasonable, as young and high educated might be less tied-up by firm-specific skills, as well as mental costs and search costs associated with a new contract are lower. We see that the short- to long-run ratio is almost 10 percent points lower for females aged 50-62, than for young females aged 26-35. The ratio gap between high educated (university) and low educated (primary/ lower secondary) is about half the size.

TABLE 5

Heterogeneity in how fast individuals respond to tax changes

		Labor supply elasticities		
Heterogeneous groups		*Short-run**	*Long-run*†	*Ratio*‡
Age	25-35 years old	0.11	0.32	32.8%
	50-62 years old	0.08	0.36	23.3%
Education	High education	0.07	0.25	29.7%
	Low education	0.10	0.39	25.2%

$*$ First year responses, † 15th year responses, ‡ The ratio of short-run relative to long-run elasticities (in percent)
Notes: The labor supply elasticities refer to the total elasticities (extensive and intensive margin) wrt. the net wage rate. The heterogeneous groups are defined by the respective observed characteristics in the first estimation period.

Robustness checks

In the following we report the results of a number of robustness checks regarding unobserved heterogeneity, and state dependence in the wage rate. As more elaborate specifications of unobserved heterogeneity and state dependence are computationally demanding, we rely on a random subset (about 8,900 individuals) of the original estimation sample to obtain estimates. We do not find any substantial effects on the labor supply elasticities over time by increasing the number of unobserved individual types from two to three or four. The results are reported in Table 6.

In our main analysis we focused on state dependence in preferences, labor market constraints and adjustment costs, and used a pooled Heckman selection regression (Table A.1 in Appendix) to assign wage rates to each individual. However, state dependence and unobserved heterogeneity are possibly also present in wages (see e.g. Eckstein and Wolpin, 1989), and it might be a concern that accounting for this could change our main results regarding labor supply elasticities over time. We therefore test an alternative specification of the wage rate, where each individual has five different predicted wage rates depending on the previous period's choice of working time. Random effects are added to the predicted individual wage rates based on draws (30 draws per individual) from a normal distribution of the individual specific effects. We find that the implication of using this alternative specification of the wage rate, has little effect, as reported in Table 6.

TABLE 6

Robustness checks. More elaborate specifications of unobserved heterogeneity and state dependence in wages

	Labor supply elasticities					
	Extensive margin			*Intensive margin*		
	Short-run*	*Long-run*†	*Time*‡	*Short-run**	*Long-run*†	*Time*‡
Baseline model (2 types)	0.02	0.07	10	0.07	0.25	8
3 unobserved types	0.02	0.06	10	0.07	0.24	8
4 unobserved types	0.01	0.06	9	0.08	0.23	7
Alt. specification of wages	0.02	0.08	10	0.08	0.28	8

$*$ First year responses, † 15th year responses, ‡ Years until 90% of long run effect is reached

VI Concluding Remarks

The aim of this paper has been to develop a structural model that explicitly allows for possible sources of lagged responses in labor supply decisions. While controlling for individual heterogeneity (so-called spurious state dependence), we define three sources of (true) state dependence arising from preferences, labor market constraints and adjustment costs. These elements cause lagged responses, and consequently a gradual path of labor supply adjustment.

Our study is, to our knowledge, the first decision model to define adjustment costs, discussed by Chetty et al. (2011) and Chetty (2012), as a distinct source of state dependence. This enables us to analyze the effect of adjustment costs over time in a random utility framework. Moreover, we elaborate on earlier studies on state dependence in working time decisions, such as Haan (2010) and Prowse (2012), by defining state dependence more precisely from a tractable job choice framework. The benefit of our approach is twofold: first, it provides new insights into sources of state dependence, including adjustment costs and state dependence in labor market constraints, and secondly, it provides a more coherent structural decision model, which is better equipped to simulate the path of labor supply responses under policy changes.

The decision model is estimated as a multinomial logit model, where the dependent variable is the sequence of working time for each individual. Preferences and job opportunities can be identified by repeated cross-sectional variation, whereas transitions over time allow us to identify individual-specific unobserved heterogeneity and state dependence.

Whereas we define state dependence in preferences as a shift in the preferences for leisure, adjustment costs are characterized as a one-time fixed costs of altering working time from last period's decision. This means that state dependence in preferences may induce individuals to slowly react to incentive changes in a stepwise fashion, going from non-participation to part-time and further to full-time, as the preferences for leisure are influenced by previous working time. Adjustment costs on the other hand, simply capture that a change is costly, independently of in what direction the adjustment takes place or how radical it is. State dependence in job opportunities is identified as an additional lag in responses, when deciding to participate in the labor market after a period of non-participation.

Although each component of state dependence enters in a quite restrictive fashion, we find that the estimated model reproduces the correlations observed in data very well. In order to analyze the relative importance of the different sources of state dependence, we let the model simulate persistence over time when eliminating each estimated component of state dependence sequentially. We find that adjustment costs dominate, whereas state dependence in labor market constraints (job opportunities) is almost as important at the extensive margin. We also find evidence of state dependence in preferences, although to a lesser extent.

When simulating the path of adjustment to a general tax cut, we find that state dependence leads to considerable lags in labor supply responses. The estimated first year responses are brought down to about one third of the long-run effect. We find that the females reach the proximity of the full effect after about seven years. In the long run, simulated labor supply responses are close to predictions from a static version of the model without state dependence. This suggests that state dependence, including adjustment costs, do not prevent individuals from adjusting in the long run.

Despite a number of structural and simplifying assumptions, we nevertheless think this study adds to the understanding of underlying sources of lagged responses in labor supply decisions. Furthermore, we add to an emerging literature which suggest that individuals respond gradually rather than immediately to policy reforms.

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Appendix

Imputation of missing working time observations

The imputation of missing working time observations (about 30 percent) is done similarly to Pronzato (2015), by using observed information on annual labor income combined with predicted monthly wage income for a full-time job. The monthly wage income is predicted based on a set of individual characteristics (experience, field and level of education, land background and county). Cut-offs for each working time alternative are calibrated by adjusting the simulated to the actual distribution of working time for individuals observed in the wage statistics sample. We use the following cut-offs: Short part-time: 0.3-7.7 times predicted monthly wage, long part-time: 7.7-10.95, full-time: 10.95-15.35 and overtime: >15.35. Non-participants are defined as earning less than NOK 5,000 (about EUR 600) annually or less than 0.3 times the predicted monthly full-time earnings.

Assigning wage rate and wage income for each alternative choice

Table A.1 presents the wage regression, which is used to assign each individual a wage rate. The dependent variable in the regression is the hourly wage rate which is calculated by dividing contractual monthly pay by monthly contractual working hours. We use a (pooled) Mincer wage regression with year fixed effects, allowing for participation selection effects (Heckman, 1979).[16] We assign the predicted hourly wage rate of the Heckman regression to each individual. Wage income is constructed by multiplying the individual wage rate and the median working time (yearly) for each alternative. We deduct taxes by employing a tax simulator.

Out-of-sample predictions and predictions for each unobserved type of individuals

Figure A.1 and Table A.2 present out-of-sample predictions of the estimated model. Table A.3 provides in-sample predictions for the two unobserved types of individuals.

[16] Mills lambda is significantly positive which suggests that unobservables are both positively related to participating in the labor market and positively related to wages. This indicates that an ordinary least square regression (OLS) without correction for selection effects could be biased.

TABLE A.1

Wage regression, Heckman 2-Stage

	ln(Wage)		*Participation*	
	Coefficient	*Std. error*	*Coefficient*	*Std. error*
Constant	4.6834***	0.0068	1.5323***	(0.0653)
Experience	0.0164***	0.0003	0.0023	(0.0040)
Experience squared	-0.0003***	0.0000	-0.0008***	(0.0001)
Low education	-0.0968***	0.0022	-0.2082***	(0.0201)
High education	0.2791***	0.0016	0.4424***	(0.0207)
Non-Western immigrants	-0.1192***	0.0038	-0.9994***	(0.0234)
Residence in metropolitan area	0.0733***	0.0013	-0.0781***	(0.0142)
Year dummies (base: 2003)				
Year 2004	0.0459***	(0.0017)	0.0606**	(0.0189)
Year 2005	0.0820***	(0.0017)	0.1878***	(0.0193)
Year 2006	0.1243***	(0.0017)	0.2488***	(0.0196)
Year 2007	0.1768***	(0.0017)	0.3279***	(0.0201)
Education category (base: ‘unknown’)				
General	0.0206***	(0.0052)	0.6999***	(0.0323)
Human, Art	-0.0624***	(0.0055)	0.4348***	(0.0396)
Education	-0.0736***	(0.0054)	0.9016***	(0.0423)
Social, Law	0.0557***	(0.0062)	0.8511***	(0.0614)
Business	0.0284***	(0.0052)	0.8295***	(0.0329)
Technology	0.0677***	(0.0054)	0.8294***	(0.0387)
Health	-0.0809***	(0.0052)	1.0256***	(0.0331)
Primary	-0.0049	(0.0095)	0.3533***	(0.0787)
Service	-0.0267***	(0.0064)	0.6344***	(0.0499)
Exclusion restrictions				
No. of children under age 3			-0.1002***	(0.0206)
No. of children under age 6			-0.0879***	(0.0174)
No. of children under age 12			-0.3349***	(0.0098)
Net wealth (in NOK 10,000)			-0.0000**	(0.0000)
Partner’s net income (in NOK 10,000)			-0.0024***	(0.0001)
Mills Lambda	0.0273***	(0.0058)		
Number of observations	110,001			
Number of individuals	26,631			

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Notes: Net wealth is measured by the end of the year. Net wealth and partner’s net income are measured in current NOK (1 EUR≈8 NOK).

FIGURE A.1

Out of sample fit, choice probabilities, year 2008

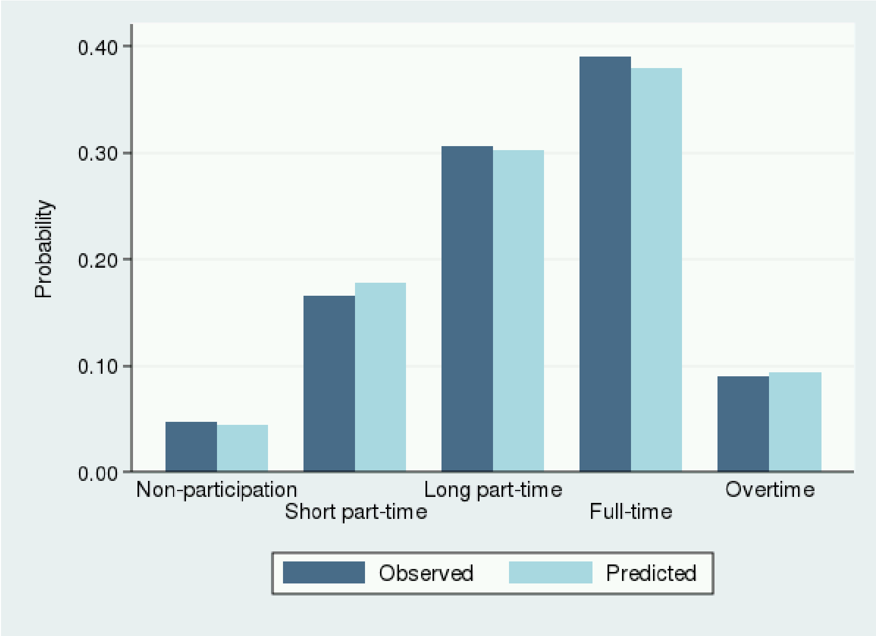

TABLE A.2

Out-of-sample fit, predicted and observed transition probabilities (in percent) 2007-2008

Predicted:					
2007 \ 2008	Non-participation	Short part-time	Long part-time	Full-time	Overtime
Non-participation	**83.7%**	12.0%	4.0%	0.3%	0.0%
Short part-time	1.9%	**76.2%**	14.9%	6.7%	0.3%
Long part-time	0.0%	8.5%	**74.7%**	14.4%	2.4%
Full-time	0.0%	3.3%	8.8%	**79.2%**	8.7%
Overtime	0.0%	3.9%	11.3%	27.9%	**56.9%**
Observed:					
2007 \ 2008	Non-participation	Short part-time	Long part-time	Full-time	Overtime
Non-participation	**84.8%**	13.0%	1.3%	0.9%	0.0%
Short part-time	2.5%	**76.5%**	16.1%	4.0%	0.9%
Long part-time	0.1%	7.2%	**77.9%**	12.4%	2.4%
Full-time	0.1%	1.4%	8.4%	**82.7%**	7.4%
Overtime	0.3%	1.3%	5.6%	37.0%	**55.8%**

TABLE A.3

In-sample predicted transition probabilities for the two unobserved types of women

Type	t+1 / t	Non-participation	Short part-time	Long part-time	Full-time	Overtime
1	Non-participation	**32.2%**	45.7%	19.9%	2.1%	0.0%
	Short part-time	2.8%	**49.9%**	30.1%	16.4%	0.9%
	Long part-time	0.1%	14.7%	**52.3%**	28.4%	4.5%
	Full-time	0.0%	6.9%	17.4%	**63.1%**	12.6%
	Overtime	0.0%	6.3%	17.5%	37.8%	**38.4%**
2	Non-participation	**94.3%**	4.0%	1.6%	0.1%	0.0%
	Short part-time	1.9%	**90.0%**	5.4%	2.2%	0.1%
	Long part-time	0.0%	2.9%	**92.8%**	3.9%	0.3%
	Full-time	0.0%	1.5%	3.7%	**93.7%**	1.1%
	Overtime	0.1%	3.4%	8.9%	14.2%	**73.5%**

Notes: The observed individual characteristics match the sample average in year 2006. The predicted transition from year 2006 to year 2007 is considered under the constraint that the choice environment is kept constant.

Statistics Norway

Postal address:
PO Box 8131 Dept
NO-0033 Oslo

Office address:
Akersveien 26, Oslo
Oterveien 23, Kongsvinger

E-mail: ssb@ssb.no
Internet: www.ssb.no
Telephone: + 47 62 88 50 00

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Statistics Norway