Labor Supply Inertia and Lagged Tax Responses: Effects of State Dependence and Adjustment Costs

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Abstract

Behavioral micro simulation models, estimated by static discrete choice, provide essential information to policy makers regarding labor supply effects of prospective tax-benefit reforms. However, due to the static framework it is not clear whether its predictions should be interpreted as immediate or long run expected outcomes. In the present study we estimate a simulation model which allows for a gradual adjustment in labor supply subsequent to a policy reform. In contrast to earlier studies, we explicitly model frictions in the labor market by distinguishing between state dependence in preferences and job opportunities, and fixed adjustment costs. We estimate the model on Norwegian administrative data for married and cohabiting women over a five years period. The simulation results suggest a considerable component of gradual adjustment, mainly driven by fixed adjustment costs. The immediate effect only accounts for about 1/3 of the long run effect, in which the predicted 'long run' effect is consistent with the static model predictions.

Keywords: labor supply transitions, inertia, state dependence, adjustment costs, income tax
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1 Introduction

Behavioral micro simulation models provide essential information to policy makers regarding labor supply effects of prospective tax-benefit reforms.[1] Discrete choice models are popular for this purpose, as they easily account for nonlinearities and possible non-convex budget sets implied by actual tax and benefit systems. The models are typically estimated on cross-sectional observations of individual's choice of working time and disposable income. It follows that these models are static in nature and therefore provide a static prediction for the labor supply effect or budgetary effect of altered tax changes. However, it is not clear whether this effect can be expected already the same year as the policy reform was implemented, or whether it should be considered as a long run prediction. This distinction is crucial, and it is therefore a concern among policy makers that the timing dimension for ex-ante predictions is not clarified.

From an economic point of view this concern is justified. One can think of several reasons for a gradual rather than immediate adjustment in labor supply after a change in the tax schedule, for instance due to inattention or information costs, habit persistence, costs associated with obtaining a new job position, or a lag in the observed outcome (working hours) of altered effort or search behavior. These issues are disregarded in a static model setting, and might therefore provide biased predictions.

There is indeed some evidence, although the literature is scarce, on gradual adjustment in labor earnings when evaluating a tax reform period ex-post. Holmlund and Söderström (2011) estimate an error correction model which demonstrates the presence of lagged responses in labor earnings when a new tax schedule applies.[2] Schroeder (2016) finds also delayed adjustment in working hours to a change in the net wage rate. Moreover, Gelber et al. (2014) present some recent evidence of gradual adjustment to changes in the US tax-benefit schedule by using information on the amount of bunching around kink points.

The ex-post evidence on gradual adjustment suggests a demand for an ex-ante tax-benefit simulation model which accounts for the timing of responses. In this study we follow a similar approach by (Haan, 2010) who estimated an intertemporal discrete choice model using German data.[3] Because of our rich model set-up we are able to go further than (Haan, 2010) in modelling the different mechanisms of lagged responses. Whereas (Haan, 2010) relied on interactions dummies for each combination of transitions, we instead model state dependence and adjustment costs explicitly. This is both an advantage for a simulation model to be used for analyzing prospective tax policies; in addition to that it provides essential information on the composition of different sources of lagged responses.

[1] The Norwegian behavioral microsimulation model LOTTE-Arbeid is such an example.

[2] Vattø (2015) adopts a corresponding reduced form framework analyzing Norwegian data over a tax reform period, and finds a similar pattern.

[3] Haan et al. (2012) and Haan and Uhlendorff (2013) build on the same type of intertemporal model as Haan (2010), in which the former improve the estimation techniques, and the latter provides an extension with distinction between voluntary and involuntary unemployment.

Based on a static discrete job choice model (Dagsvik and Jia, 2014), we develop a model which allows for frictions in the labor market, by incorporating fixed adjustment costs, together with state dependence in preferences for leisure and state dependence in job opportunities.

The model is estimated on data from Norwegian administrative registers over the period 2003-2008. We focus on the group of married and cohabiting women, because of more variation in working time, and also a wider specter of transitions over the time period. Moreover, married women are typically thought to be more responsive to changes in tax schedule and therefore provide an interesting group from a policy perspective.

We find a considerable component of gradual adjustment in female labor supply, in which the immediate effect only accounts for about 1/3 of the long run effect. In accordance to (Haan, 2010) we find that the long run effect corresponds to the static models predictions.

Further, our results suggest that fixed cost of adjustment, is a crucial source of gradual adjustment in labor supply responses. This fixed cost of adjustment might include the cost of negotiating a new contract with an employer, or the time and financial cost of job search (Gelber et al., 2014), but it might also reflect informational and psychological costs of even considering and recognizing the choice situation (Samuelson and Zeckhauser, 1988).

This paper is organized as follows. In Section 2, we present our modeling framework which we denote the intertemporal job choice model. In Section 3, we describe the data sources and organization of the data, followed by the empirical implementation of the model in Section 4. In Section 5 we present the estimation results including wage elasticities, tax simulations and the separate effect of state dependence and adjustment costs. Section 6 concludes the paper.

2 State dependence and adjustment costs as sources of gradual tax responses

Before turning to the more formal description of our model in the next section, we here introduce the concepts of state dependence and the special case of adjustment costs. We review some of the main contributions to this literature and discuss how state dependence and adjustment costs affect the dynamics of adjustment in a random utility model.

State dependence can broadly be defined as the causal effect of experiencing an event on an individual's evaluation of future choices (Heckman, 1981). For instance, individuals might be more willing to participate in the labor market if they did so last period, despite controlling for observed and unobserved individual heterogeneity. Such positive state dependence at the extensive margin (participation in the labor market) is found on US data (Hyslop (1999), Eckstein and Wolpin (1989)), Japanese data (Okamura and Islam, 2009), British (Prowse, 2012) and German data (Haan, 2010). At the intensive margin (working hours given participation) evidence seem to be somewhat more dispersed. WhileProwse (2012) and

Haan (2010) find significant state dependence also at the intensive margin, Cai (2010) reports no evidence of true state dependence for married Australian women after unobserved heterogeneity and serial correlation of transitory errors are controlled for.

In the presence of positive state dependence the immediate adjustment in working hours after a policy reform is reduced because individuals are still influenced by their previous optimal decisions. One can think of a share of individuals adjusting to their new optimal equilibrium each period, leading to an observed gradual adjustment in labor supply. In a random utility model this is enforced by an increased probability of transition to the new equilibrium, and a reduced probability of returning to the old equilibrium once adjusting. As time goes by a new stable equilibrium at the aggregated level is attained.

State dependence might originate from prices, preferences or constraints of future periods (Prowse, 2012), but it might also take a more particular form such as fixed costs of adjustment. In the following we review a recent literature on how optimization frictions such as adjustment costs are thought to prevent individuals from adjusting their labor supply in the short run.

Chetty (2012) introduced the concept of optimization frictions, which represent the utility loss agents can tolerate to deviate from the frictionless optimum. He argues that optimization frictions, such as adjustment costs and inattention, prevent individuals from responding to tax reforms.

Gelber et al. (2014) use evidence from the social security earnings test to measure adjustment frictions in earnings. They use information on the amount of bunching at kinks before and after policy changes in earnings incentives around the kinks, and find that the short run impact of changes in the effective marginal tax rate can be substantially attenuated.

They suggest that the information costs associated with navigating a new tax regime, the cost of negotiating a new contract with an employer, or the time and financial cost of job search could be possible sources of adjustment costs/optimization frictions.[4] In line with our random utility model, they interpret the frictions they observe as evidence of barriers to making an immediate adjustment in response to changes in incentives. They do not find that optimization frictions prevent individuals from adjusting in the long run.

Adjustment cost might also be related to the concept of status quo bias (Samuelson and Zeckhauser, 1988), in which individuals in experiments seem to choose status quo, such as maintaining the previous decision, disproportionally often. According to Samuelson and Zeckhauser (1988), an individual may retain the status quo out of convenience, habit, custom, because of fear or innate conservatism, or through simple rationalization, for instance by minimizing the psychological cost of even considering and recognizing the choice situation.

In the following sections we provide a model concept which disentangles between adjustment costs in

[4] Blundell et al. (2008) suggest that labor supply adjustment occur mainly through job shifts.

addition to two other sources of state dependence: State dependence in preferences for leisure and state dependence in job opportunities.

The presence of state dependence in preferences infer that preferences for leisure can be altered by experience. For instance, individuals who have worked full time or over time last period might obtain higher preferences for working as they have developed an attitude for working long hours, influenced by colleagues and peers.

State dependence in preferences might lead to a gradual adjustment also at the individual level if for instance it was optimal to choose a short part time job in the old policy regime and a full time job in the new tax regime, individuals might not recognize this immediately, and instead need to increase working hours gradually from short part time to long part time until realizing that a full time job is preferred.

State dependence in opportunities occur when individuals who previously did not participate in the labor market lack job opportunities the present period due to for example depreciation of human capital, reduced network, higher search costs or discouraged worker effects (see e.g. Hyslop, 1999).

State dependence in opportunities lead to a partly postponed adjustment from non-employment to an employed position in a random utility model. Although the individual prefer this transition, she might have few job options in the market due to non-participation the previous period. Once the transition occur, it is less probable to choose to go back to non-participation, so gradually the optimal adjustment will take place.

3 The model

In this section, we start with a brief review of the *static* job choice model followed by a discussion of the extensions necessary to analyze labor supply responses in an intertemporal setting, including state dependence, persistent unobserved heterogeneity and serial correlation in the error terms. The empirical implementation and specification of the model is described in Section 4.

3.1 The static job choice model

The static job choice model (applied by e.g. Aaberge, Dagsvik, and Strøm, 1995; Dagsvik and Strøm, 2006) builds on a conditional logit framework (see McFadden, 1974), similarly to the standard discrete choice labor supply model, in which individuals are assumed to choose among feasible combinations of leisure and disposable income; see for example van Soest (1995) and Creedy and Kalb (2005). The discrete choice model has practical advantages over the more traditional continuous labor supply models with marginal optimization, as it easily deals with possibly non-linear and non-convex budget sets (Blundell and MaCurdy, 1999). It is therefore a widely used tool to study labor supply, especially for the purpose of policy simulation using micro data.

Neither the continuous nor the discrete labor supply model are able to explain the peaks in full-time and part-time hours observed in the distribution of working hours. Most studies apply an ad hoc adjustment to account for this observation, such as in van Soest (1995), by introducing working hour specific (dis)utility elements. The job choice model offers, on the other hand, a theoretical rationale for this practice by introducing a concept of job opportunities; see e.g. Dagsvik, Jia, Kornstad, and Thoresen (2014); Dagsvik and Jia (2014). In this framework, labor supply behavior is viewed as an outcome of agents choosing among a set of job "packages", with additional constraints on the set of available jobs. The model allows the researchers to accommodate latent choice restrictions in a convenient way which leads to similar practical implementation as the standard discrete choice model.

The job choice model is specified as follows. At any given time period t, individuals are assumed to have preferences over a set of "jobs" where each job (indexed by $z = 1, 2, ...$) is characterized by disposable income $C_t(z)$, hours of work $h_t(z)$, and other non-pecuniary job attributes, such as the nature of the job-specific tasks to be performed, location of the workplace etc. Disposable income for a given job is defined as $C_t(z) = f_t(h_t(z)w_t(z), I_t)$, where $w_t(z)$ is the offered wage rate for the given job z, I_t is the non-labor income and $f_t(.)$ is the net-of-tax function. In the present study, we assume that the offered wage rate $w_t(z)$ does not vary across jobs for a given individual. The individual's utility of choosing job z at period t is represented as $U_t(C, h, z)$.

We assume that the set of available jobs to choose from is individual and time specific. In practice, we are not primarily interested in the job choice per se but rather in the wage and hours of work combination that follows from the job choice. Dagsvik and Jia (2014) show that when the utility function is additively separable, namely $U_t(C, h, z) = v_t(C, h) + \varepsilon_t(z)$, the labor supply probabilities can be specified in absence of detailed information about the choice sets. A variable for job opportunities is introduced which represents the number of available jobs for a given working time alternative h: $m_t(h)$. It can be shown that under suitable distributional assumption regarding the error terms $\{\varepsilon_t(z)\}$ as extreme value distributed, the probability that the worker chooses a particular job with hours of work h at period t, $\varphi_t(h)$ can be written as

$$\varphi_t(h) = \frac{exp(v_t(C, h))m_t(h)}{\sum_h exp(v_t(C, h))m_t(h)} \tag{1}$$

This expression is analogous to a multinomial logit model with representative utility $v_t(C, h)$, weighted by the number of available jobs $m_t(h)$. Since $m_t(h)$ is not observable, one need to estimate it simultaneously with the systematic part of the utility function $v_t(C, h)$. The number of non-working opportunities is normalized to one, i.e. $m_t(0) = 1$.

3.2 The intertemporal job choice model

In the following, we extend the static job choice model to study individual's labor supply behavior over time ($t = 1, ..., T$). Although individuals are still assumed to be myopic, and therefore optimize on a period to period basis, we allow for correlation in individuals' choices over time due to persistence in observed and unobserved heterogeneity and due to state dependence.

As already mentioned, state dependence exists when past experience has a casual effect on individual's choice behavior. In the case of labor supply, this could happen if past employment affects for example either preferences, the price of labor (in terms of wage), or labor market constraints (Heckman, 1981; Prowse, 2012). Following Haan (2010) and Hyslop (1999) we assume that past employment history influences current period's labor supply decision, only through last period's decision. In other words, we assume that state dependence follows a first order Markovian structure over time.

In contrast to earlier studies, we attempt to separate out the pure effect of adjustment costs and distinctively define state dependence in preferences and opportunities by utilizing the theoretical job choice framework.[5] Notice that last period's decision influences both current individual preferences as well as the current available job opportunities. This leads to that individual's preferences and opportunities might vary by last periods experience, although the objective preference and opportunity functions are time independent. In addition we allow for adjustment costs which is assumed to occur if an individual alters his/her choice of working time from last period.

Preferences can be affected by state dependence when past experience influences the taste of leisure and consumption, leading to intertemporally non-separable utility functions, $v_t(C,h) = v_t(C,h|h_{t-1})$,where h_{t-1} is last period's labor supply choice. Haan (2010) mention that preferences for leisure might change according to previous experience because of influence from colleagues and peers. Other reasons might be habit or a subjective perception of preferences for leisure which change according to experience. Uncertainty regarding other alternatives, regret avoidance and incomplete information can be reasons for both state dependence in preferences and the more specific effect of adjustment costs. Notice that possible preferences for "closer" rather than "distant" alternatives to status quo are related to what we define as state dependence in preferences.

Wages can, moreover, be affected by state dependence, as past working experience contributes to accumulation of job-related human capital, which thus induce an increase in wages or other types of compensation. We have tested for state dependence in wages in a robustness test. We do find some evidence that wages increase by past experience. However, to account for this does not have a significant impact on our results. Thus, in our main specification we have relied on a more standard Heckman selection regression for the wage rate without state dependence.

[5] Haan and Uhlendorff (2013) introduced state dependence in labor market rationing in a model which distinguish between voluntary and involuntary unemployment for males. In our model this correspond to that individuals might choose not to participate in the labor market due to preferences for leisure ("voluntary") or due to lack of job opportunities ("involuntary").

Labor market constraints or job opportunities can be affected through signaling and scarring effects from past employment, or as Hyslop (1999) argues by higher fixed cost of searching, for those who are currently unemployed. Moreover, smaller network or discouraged worker effects implies less job opportunities. One nice feature of our job choice model is that it explicitly introduces labor market constraints through $m_t(h)$. So we can directly model state dependence due to labor market constraints by allowing that $m_t(h) = m_t(h|h_{t-1})$.

We moreover define adjustment costs, $AC(h_t|h_{t-1})$, as a more specific kind of state dependence, where past experience defines status quo and thus a fixed cost related to adjustment if not maintaining previous choice alternative. Adjustment costs can be psychological, informational, related to time or monetary costs involved in changing "job" to another working time alternative. Adjustment costs are directly related to status quo bias, in which individuals for various reasons tend to maintain the reference path, for both rational reasons as mentioned above and more irrational reasons related to e.g. loss aversion and endowment effects. The periodic utility function is extended to account for the adjustment cost, i.e. $v_t = v_t(C, AC, h_t|h_{t-1})$. Thus, the joint probability for an individual to choose the labor supply sequence $h_1, h_2, \cdots, h_t$ is given by:

$$\varphi(h_1, h_2 \cdots h_T) = \prod_{t=2}^{T} \varphi_t(h_t|h_{t-1})\varphi_1(h_1), \tag{2}$$

where

$$\varphi_t(h_t|h_{t-1}) = \frac{exp(v_t(C, AC, h_t|h_{t-1}))m_t(h_t|h_{t-1})}{\sum_h exp(v_t(C, AC, h|h_{t-1}))m_t(h|h_{t-1})} \tag{3}$$

Under suitable assumptions (see section 4), we are able to distinguish state dependence in preferences, state dependence in opportunities and adjustment costs. To be more precise, the main restrictions we impose for identification is as follows: state dependence in preferences is related only to shifts in the continuous utility function, state dependence in opportunities captures the extensive margin, and adjustment costs are limited to status quo maintenance, and not allowed to depend on how far or in what direction a transition is made. At first sight, these restrictions seem quite restrictive. However, it turns out this provides a similar flexible pattern of state dependence as if we had estimated indicator variables for each transition as conducted in Haan (2010).

We further control for persistent unobserved heterogeneity both in preferences, opportunities and in adjustment costs. A related aspect is serial correlation in the error terms $\{\varepsilon_t(z)\}$ which arises as a result of random shocks which last longer than one period or unobserved individual characteristics that change slowly over time. Hyslop (1999) argues for instance that allowing for serial correlation in the error terms can correct for the endogeneity of non-labor income and fertility. A simple form of serial correlation can be introduced in the discrete choice model by allowing for choice specific permanent unobserved heterogeneity. Consider a simple representation of serial correlation structure where the error term consists of two parts; a permanent individual specific effect that does not change over time, and a transitory part. We will assume that the individual permanent effects for jobs with the same working time are constant. Namely: $\varepsilon_t(z) = a(h) + \zeta_t(z)$, where $\zeta_t(z)$ is i.i.d across time periods and jobs. If we assume that $a(h)$ is random across individuals and distributed according to a suitable distribution $f(.)$, we see immediately that the choice probability $\varphi_t(h_t|h_{t-1})$ follows a standard random effect approach. It follows that allowing for unobserved heterogeneity in the opportunity measure, $m_t(h|h_{t-1})$, can be interpreted as a special type of serial correlation in the error terms $\{\varepsilon_t(z)\}$.

4 Data and descriptive statistics

Before we probe deeper into the empirical specification of the model, we shall in this section describe the data sample and provide some descriptive statistics.

Our estimation sample on Norwegian married and cohabiting women over the period 2003-2007 is collected from a large panel data set of administrative registers.

To obtain information on working time and monthly wages we use "Wage statistics" which is based on administrative registration of employers reporting of their employees (Statistics Norway, 2006, 2009). The statistics are collected from a stratified selection of Norwegian firms with at least 3-5 employees (depending on industry). The statistics cover 50-60 percent of the employees in the private sector and 100 percent in public sector. In total, we have information on about 70 percent of Norwegian wage earners. The large number of employees included each year implies that we can utilize the panel dimension of the data source.

Information about annual incomes and tax return, family composition, number of children, education, etc. is obtained from the "Income Statistics for Persons and Families" (Statistics Norway, 2005) and linked to the Wage statistics, using unique personal identification numbers. Unemployed, self-employed, disabled persons and students are excluded from the sample. We further limit the sample to married and cohabiting women [6] aged 25-62 years. To avoid attrition/selection effects over time we impute information on working hours for those individuals not observed in the wage statistics (almost 30 percent) to match up with the sample of the income statistics.

To simplify the analysis computationally, we concentrate on a balanced ten percent random subsample (about 30,000 individuals).

The discretization of working time is obtained by dividing into five categories based on weekly hours of work: $h \in \{0, 1-19, 20-34, 35-40, 40+\}$. The categorization of individuals not observed in the wage statistics is done similarly to Pronzato (2015), by regressing the monthly wage of individuals in full time job on potential experience, field and level of education, origin, residence and time fixed effects, and using the predicted monthly wage together with observed annual wage income as basis

[6] We limit the analysis to couples where the husband's total pre-tax income level exceed yearly NOK 150,000 (about EUR 20,000) to abstract from various welfare transfers. This restriction affects less than five percent of the couples.

for categorizing. Individuals with observed annual income about 12 times the predicted monthly wage are for instance assumed to work full time. The cut-offs are calibrated by adjusting the simulated to the actual distribution of working time for individuals observed in the wage statistics sample.[7] Non-participants are defined as earning less than NOK 5,000 annually or less than 0.3 times the predicted monthly full time earnings.

Summary statistics of the main variables are presented in Table 1 for the estimation period 2004-2007 (2003 is used for initial conditions only).

Figure 1 depicts the observed share of individuals in each working time arrangement. It appears to be a trend towards increased labor market participation and a shift from short part time to full time work during the period of consideration. The observed pattern of increased working time is somewhat exaggerated by the balanced panel properties, but might be partly explained by a cyclical upturn in the economy in addition to a higher number of availabel kindergarten places.[8] Moreover, the 2006 Norwegian tax reform implied lower marginal tax rates for high income levels in addition to some smaller adjustments in the basic tax allowance.[9]

Table 2 presents the transition probabilities for year 2005 to 2006.[10] The diagonal entries of the matrix stand out and suggests strong persistence in individuals' labor supply over time. In accordance with earlier studies such as Haan (2010), we focus on annual transitions in contrast to monthly or quarterly transitions, as this suits our data better. In general, we expect that annual transitions contain less persistence than quarterly measures. Moreover, when working hours is averaged over a longer period, this contribute to less persistence than if it measured over a short period of time.[11]

Table A.1 describes the characteristics included as taste modifying variables in the utility function, $v_t(C,h)$, in addition to characteristics included in the specification of the job opportunity measure,

[7] The following annual income cut-offs corresponds to the observed distribution of working hours. 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.

[8] In our model framework we interpret this as increased number of (relevant) job opportunities. Both labor market tightness and available kindergarten capacity increased over the period 2004-2007 in all regions. The unemployment rate in Norway declined from 4.5 percent in 2004 to 2.5 percent in 2007, (Statistics Norway, 2013), and the daycare coverage of children aged 1-5 years increased from 72.1 percent in 2004 to 84.3 percent in 2007 (Statistics Norway, 2013). Since the rise in coverage might also partly reflect increased demand, we have instead relied on a regional measure of number of kids per employed as a proxy for free capacity in child care.

[9] We have deliberately chosen to estimate the model over a period with changes in the tax schedule in order to enhance the structural estimates of the intertemporal model, although cross-sectional variation among individuals in principle are sufficient to identify the model.

[10] The annual transition probabilities are similar for the other years. It is interesting to see that the entries of the transition matrix are of very similar magnitude as for German women reported by Haan (2010). This although the state probabilities for each choice alternative differ, with considerably higher probability of non-participation (27 percent) and less probability of full time (19 percent) in the German sample.

[11] Keep in mind that while our working time information originates from the wage statistics in September/October each year, the income information is gathered from the whole income year. So for individuals where we impute hours worked from the annual income information we might underestimate some of the variation relatively to the monthly data, because an individual who undertakes a transition from full time to non-participation within a year might be reported as working part time. For the majority of individuals, however, we rely on working hours and wage information from October, although other income (such as husbands income) originate from the whole year.

Table 1: Sample characteristics over time

	Nominal mean income (NOK)				
Variable	2004	2005	2006	2007	N
Labor income (reference month)	20,859	21,530	22,624	24,109	20,353
Labor income (annual)	247,564	260,500	277,688	299,909	29,848
Non-labor income	28,357	39,364	24,822	26,569	29,848
Husbands total income	515,546	580,154	515,479	573,527	29,848

$m_t(h)$. The taste modifying variables include age and indicator variables for having children under 6 years and children under 12 years in the household. Information about education,[12] and measures of labor market tightness and kindergarten capacity are included in the specification of the job opportunities. Regional labor market tightness is measured as vacancies divided by the number of unemployed. Regional available kindergarten capacity is proxied by the number of children (adjusted for age) per kindergarten employee. For both measures information from 19 main regions in Norway is utilized.

[12] Years of education are imputed for a small fraction of individuals with missing observations. A few individuals are coded with 0 years of education. In the specification of the job opportunities we instead utilize an indicator variable of low education defined as whether an individual has less than 10 years education or not.

Figure 1: Observed choice probabilities

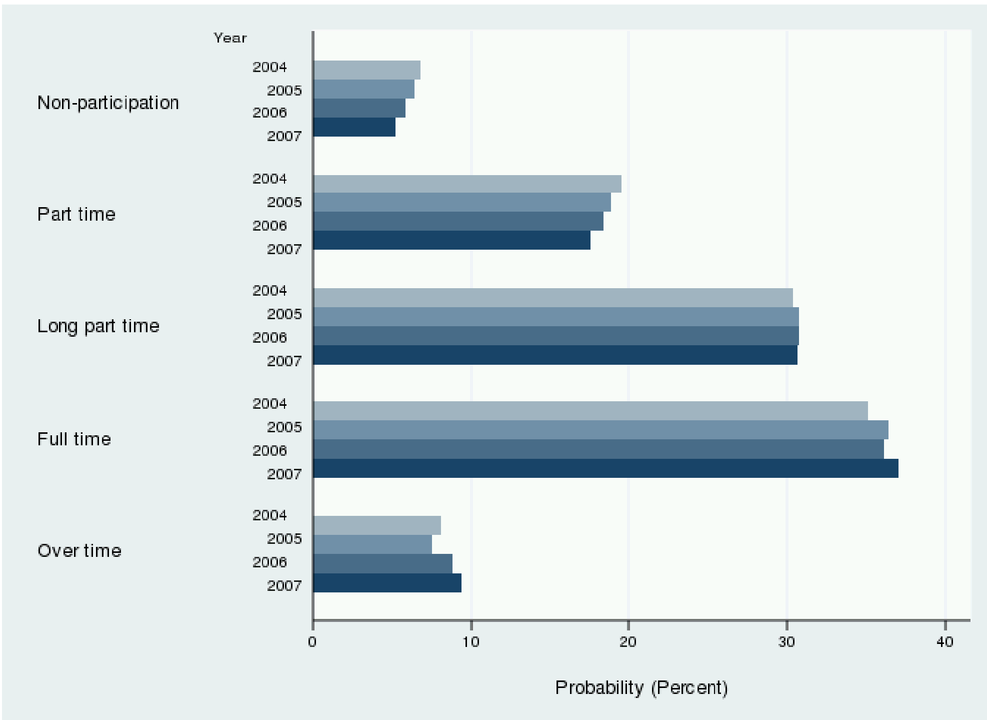


Table 2: Observed transition probabilities 2005-2006

		2006				
		Non-participation	Short part time	Long part time	Full time	Over time
	Non-participation	**81.9%**	15.8%	1.2%	0.9%	0.3%
	Short part time	2.7%	**76.0%**	16.5%	4.1%	0.7%
2005	Long part time	0.1%	7.7%	**77.3%**	12.7%	2.2%
	Full time	0.1%	1.6%	9.1%	**79.8%**	9.4%
	Over time	0.0%	1.9%	6.4%	31.1%	**60.7%**

5 Empirical specification

In this section, we discuss the empirical specification of the model. We specify in particular the functional form for the deterministic part of the utility function $v(C,h)$ and the opportunity measure $m(h)$.

We start by specifying how adjustment costs enters the utility function.

Adjustment costs

Adjustment costs is assumed to be directly associated with status quo maintenance, and is unrelated to how far or in what direction one moves. Thus, we impose a fixed cost (*ac*) for the working time transition (h_t, h_{t-1}) for all other alternatives, than the alternative which is equal to last periods choice (status quo).

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

This measure can be related to what Chetty (2012) denote optimization frictions (such as adjustment costs, inattention and status quo bias), in which the degree of friction is measured by the utility losses agents tolerate to deviate from the frictionless optimum.

The utility function for choosing job z in period t can now be written as

$$U_t(C,h,z|h_{t-1}) = v_t(C(z),h(z)|h_{t-1}) - AC(h(z);h_{t-1}) + \varepsilon_t(z)$$

Preferences

We assume that preferences $v_t(C,h|h_{t-1})$ can be described by a smooth time-independent utility function. As regards the systematic part of the utility function, we follow Dagsvik and Strøm (2006); Dagsvik and Jia (2014) and make use of a generalized Box-Cox function in leisure and disposable income. This functional form has the advantage of being strictly concave under some simple constraints on the parameters.[13] We follow Dagsvik and Jia (2014) and assume;[14]

$$v_t(C,h|h_{t-1}) = \alpha_0 \cdot \frac{(C_t - C_0)^{\alpha_1} - 1}{\alpha_1} + \beta_0(h_{t-1}) \cdot \frac{(\bar{h} - h_t)^{\beta_1} - 1}{\beta_1} \tag{4}$$

[13] Many researchers apply a polynomial (in most cases, a quadratic) specification which is quite flexible and linear in parameters. One problem with a polynomial functional form is that it is not globally monotone in consumption or leisure. See Dagsvik, Jia, Kornstad, and Thoresen (2014) for a more detailed discussion on this issue.

[14] An additional interaction term between consumption and leisure has negligible effect and is dropped.

Disposable income is defined as $C_t = f_t(h_t w_t, I_t)$, where w_t is the individual wage rate, I_t is the non-labor income and $f_t(.)$ is the net-of-tax function. The minimum or subsistence consumption level is set to $C_0 = 50,000 \cdot \sqrt{hh}$ where hh is the number of individuals in the household.[15] $\bar{h}$ is defined as 80 hours per week and h is working hours per week, so that $(\bar{h} - h)$ measures leisure time. The leisure coefficient, β_0, is defined to depend on observed individual characteristics, $\mathbf{x_t}$, and incorporate state dependence, such that last period's labor supply behavior affect current preferences for leisure.

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

State dependence in preference is accounted for by a set of five indicator variables for the individual's chosen working time alternative in period $t-1$, $\mathbf{I}(h_{t-1})$. A set of indicator variables for the initial period choice $\mathbf{I}(h_0)$ is also included to solve the initial condition problem following the method by Wooldridge (2005).[16]

Job opportunities

The job opportunity measure $m(h)$ can be separated in the following manner:

$$m(h|h_{t-1}) = \theta(h_{t-1})g(h)$$

We denote $g(h)$ the opportunity distribution of hours, which can be interpreted as the fraction of jobs available to the agent with offered hours of work equal to h, where $\sum_{h>0} g(h) = 1$. The parameter $\theta(h_{t-1})$ describes the total number of jobs available to the agent (relatively to non-participation as $m(0)$ is normalized to 1).[17]

The opportunity distribution of hours $g(h)$ is considered to stem from the demand side due to institutional restrictions as a result of centralized negotiations between labor unions and employers' associations. We assume that the opportunity distribution is uniformly distributed among working time alternatives, except for a possible peak (estimated within the model) for full time jobs. Moreover, we treat $g(h)$ as constant over time as institutional settings are relatively stable.

$$log(g_t(h)) = \begin{cases} g & if\, working\, full\, time \\ 0 & otherwise \end{cases}$$

[15] The results are insensitive to exact which level of subsistence level that is chosen.

[16] The initial condition problem refers to the fact that the initial period can not be considered independently of the unobserved individual effect. Wooldridge (2005) suggests to estimate the unobserved heterogeneity conditionally on the initial period observations. A similar strategy for the structure of state dependence can be found in e.g. Haan, Kemptner, and Uhlendorff (2012).

[17] Note that θ can be extended to include fixed cost of working; see Cogan (1981).

Regarding the size of the job opportunity set $\theta(h_{t-1})$, we assume it depends on regional labor market tightness, kindergarten availability (measured as an indicator of regional availability for households with children below 3 years) and education, in addition to last period's working time experience, included as inicator variables (again we include $\mathbf{I}(h_0)$ to control for initial conditions).[18]

$$log\theta_t(h_{t-1}) = \gamma_0 + \gamma_{11} \cdot tight_{rt} + \gamma_{12} \cdot K_{rt} \cdot ch3_t + \gamma_{13} \cdot edu_t + \gamma_2' \mathbf{I}(h_{t-1}) + \gamma_3' \mathbf{I}(h_0)$$

We interpret $\theta(h_{t-1})$ as a measure of the total number of perceived job opportunities in the market (relative to non-participation). Notice that it has an impact on the participation decision only.

Unobserved heterogeneity

Unobserved individual heterogeneity in the model is represented by allowing the parameters $\{\alpha_0, b_0, ac, \gamma_0\}$ to be random. We assume that these random variables follow a discrete distribution (see Heckman and Singer, 1984) and take values on k mass points with probability $p_k \in (0,1)$.[19] This leads essentially to a finite mixture model and has the advantage that unobserved individual characteristics can be handled flexibly without imposing a parametric structure. We estimate simultaneously the mass points and the corresponding probabilities. In principle we impose no restrictions on possible correlation structure between these random variables, see for example Haan (2010) for similar arguments.

In order to obtain mixing weights $(p_1, .., p_K)$ for each mass point, one estimate $q_2, .., q_K$ where $q_1$is normalized to *0* (see Haan, 2010).

$$p_k = \frac{exp(q_k)}{\sum\limits_{m=1}^{K} exp(q_m)}, \sum_{k=1}^{K} p_k = 1$$

With the assumptions described above, the joint likelihood for individual to choose the sequence $(h_1, h_2 \cdots h_T)$ can be written as

$$\varphi(h_1, h_2 \cdots h_T) = \sum_k p_k \prod_t \varphi_t^k(h_t | h_{t-1}) \tag{5}$$

with the conditional probabilities for each mass point given by eq. (3).

In Section 2, we argued that we can account for certain types of auto-correlation by random effects specifications. In particular, if we assume that γ_0 is a time invariant random effect which vary across

[18] To retain the possibility to simulate effects out-of-sample, we avoid including time and regional fixed effects and instead rely on measurable characteristics of labor market conditions.

[19] Another interesting approach to specify the bi-variate distribution is to use the so called "two-factor loading" models, see for example Haan and Uhlendorff (2013) for an application in labor supply modeling.

individuals, it also account for serial correlations in the error terms over time: $\varepsilon_t(z) = \gamma_0 + \zeta_t(z)$ if $z > 0$, and $\varepsilon_t(z) = \zeta_t(z)$ if $z = 0$.[20]

In our main analysis, we model the individuals' unobserved heterogeneity using a finite mixture with two mass points ($K = 2$), each characterized by an parameter vector (α_0^k, b_0^k, ac^k, γ_0^k) and mixing weight p_k where $k = 1,2$.

Model implementation

The dependent variable is the sequence of choices over the period 2004-2007 for each individual, where $5^4 = 625$ different paths are possible.

Individual disposable income ("consumption") for each working time alternative for each year is constructed by adding the predicted wage income (based on the median working time in each alternative) to capital income, transfers and husbands income. Taxes are simulated for the household according to the applicable tax schedule for each year, and deducted such that net disposable household income remains. Disposable income is inflated to 2007 level income for all years by using growth in median husbands labor earnings.

The individual wage rate is assumed to be independent of different working time alternatives and is predicted as follows. We divide contractual monthly pay by monthly contractual working hours and estimate a pooled Mincer wage regression with year fixed effects, allowing for selection effects for participation, according to Heckman (1979). The regression output is reported in the Appendix. In order to be consistent, and to alleviate systematic bias in the computed wage, we use the predicted wage rate instead of the constructed actual wage rate for all individuals.

The likelihood contribution for each individual is given by $\varphi(h_1, h_2 \cdots h_T)$ for the observed sequence $(h_1, .., h_T)$ and 1 for all other sequences. By maximum likelihood we estimate the parameters of the utility function, job opportunity measure and adjustment costs which best describe the observed patterns over the sum of individuals.

6 Results

In this section, we present the estimated results of our labor supply model, and check both in-sample and out-of-sample fit by comparing predicted and observed outcomes. Based on the estimated model, we derive wage and income elasticities, and demonstrate how the model can be used to analyze counterfactual tax reforms with a focus on dynamics of responses. Then, we look further into the concept of state dependence by decomposing the effect in preferences, opportunities and adjustment costs by

[20] The assumption may be restrictive since it implies that the serial correlations among working states are identical.

providing numerical examples and graphical presentations. Finally, we provide some robustness checks on alternative modeling of state dependence and unobserved heterogeneity.

Model estimates

The estimated model coefficients are presented in Table 3. The utility functions for both subgroups are concave and increasing with respect to both consumption and leisure (also when accounting for taste modifiers, state dependence and initial conditions in leisure).

We find significant effects of state dependence in preferences, where negative coefficients $b_{22} - b_{25}$ implies that individuals who worked more previous periods have ceteris paribus lower valuations for leisure in the current period.

The estimates of adjustment costs are positive and significant which suggests that altering the last periods decision of labor supply induces ceteris paribus a loss in utility.

The coefficients of the job opportunity measure indicate a positive effect of labor market tightness (number of vacancies divided by number of unemployed) and a negative effect of child care availability. Again we find significant effects of state dependence ($\gamma_{22} - \gamma_{25}$), which suggest that individuals who worked previous period have a larger set of job opportunities, and is ceteris paribus more prone to participate in the labor market.

Since we use a mixture model with two mass points in our main analysis,[21] we refer to the two subgroups as type I and type II individuals. The two (unobserved) subgroups of individuals are estimated to cover respectively 46% and 54% of the sample. Predicted transition matrices for each type separately reveal that type II has a much stronger tendency to retain their decision over time (see Table A.2 in Appendix).

[21] As two mass points might be considered restrictive, we have tried specifications with three and four mass points. When we introduce more mass points this barely improves the fit of the model, moreover, we do not find any significant differences in the predicted wage elasticities or tax simulations over time. In order to limit computational time, we therefore choose to present the two mass point case as our reference. Wage elasticity results for three and four mass points are reported at the end of this section.

Table 3: Parameter estimates of the labor supply model for women in couples

		Coefficient	Std error
Probability distribution $(\alpha_0, b_0, ac, \gamma_0)$			
Probability mass point 1	$p1$	0.4592***	(0.0110)
Probability mass point 2	$p2$	0.5408***	(0.0110)
Preferences, consumption			
Constant (Scale 10^{-4})			
Mass point 1	α_{01}	1.0297***	(0.0554)
Mass point 2	α_{02}	1.1718***	(0.0724)
Exponent	α_1	0.6523***	(0.0157)
Preferences, leisure			
Constant (Scale $1/80$)			
Mass point 1	b_{01}	5.7935***	(0.2764)
Mass point 2	b_{02}	6.2698***	(0.2867)
Taste modifiers			
Age (Scale 10^{-1})	b_{11}	-0.6840***	(0.0803)
Age squared (Scale 10^{-2})	b_{12}	0.0796***	(0.0089)
Children under age 6	b_{13}	0.2060***	(0.0209)
Children under age 12	b_{14}	-0.1539***	(0.0206)
State dependence			
Choice 2 period t-1	b_{22}	-1.8419***	(0.1182)
Choice 3 period t-1	b_{23}	-2.6062***	(0.1264)
Choice 4 period t-1	b_{24}	-3.2280***	(0.1391)
Choice 5 period t-1	b_{25}	-3.4107***	(0.1447)
Initial conditions			
Choice 2 period 2003	b_{32}	0.0276	(0.0573)
Choice 3 period 2003	b_{33}	-0.2733***	(0.0578)
Choice 4 period 2003	b_{34}	-0.5771***	(0.0606)
Choice 5 period 2003	b_{35}	-0.7918***	(0.0657)
Exponent	β_1	-2.7404***	(0.0610)

Adjustment costs			
Mass point 1	ac_1	0.6799***	(0.0229)
Mass point 2	ac_2	2.9285***	(0.0521)
Opportunity measure (Inverse)			
Constant			
Mass point 1	γ_{01}	0.2464**	(0.0955)
Mass point 2	γ_{02}	1.7172***	(0.1609)
Regional characteristics			
Labor market tightness	γ_{11}	-1.3309***	(0.1704)
Child care availability	γ_{12}	0.0345***	(0.0102)
Education			
Low education ($\leq$ 10 years)	γ_{13}	0.0841	(0.0648)
State dependence			
Choice 2 period t-1	γ_{22}	-0.6019***	(0.1160)
Choice 3 period t-1	γ_{23}	-3.6897***	(0.2616)
Choice 4 period t-1	γ_{24}	-3.1807***	(0.3027)
Choice 5 period t-1	γ_{25}	-2.9309***	(0.4693)
Initial conditions			
Choice 2 period 2003	γ_{32}	-1.0003***	(0.1207)
Choice 3 period 2003	γ_{33}	-0.9754***	(0.1543)
Choice 4 period 2003	γ_{34}	-0.4249*	(0.1897)
Choice 5 period 2003	γ_{35}	-0.0923	(0.3072)
Opportunity density			
Full time peak	g_4	0.4298***	(0.0136)
Number of individuals		29,757	

Model fit

We first evaluate the model's performance by looking at the in-sample fit of the model. The marginal distributions and labor supply transitions are predicted from the estimated model and then compared with observed measures from the same sample.

In order to predict individual's working time from the estimated model, we use the socalled empirical Bayes method (Skrondal and Rabe-Hesketh, 2004 and Train, 2009) to specify individual specific distributions (the posterior distributions) of the random effects. We define the individual weights for each type as

$$w_{ik} = \frac{p_k \prod_t \varphi_{it}^k (h_t | h_{t-1})}{\sum_j p_j \prod_t \varphi_{it}^j (h_t | h_{t-1})}.$$

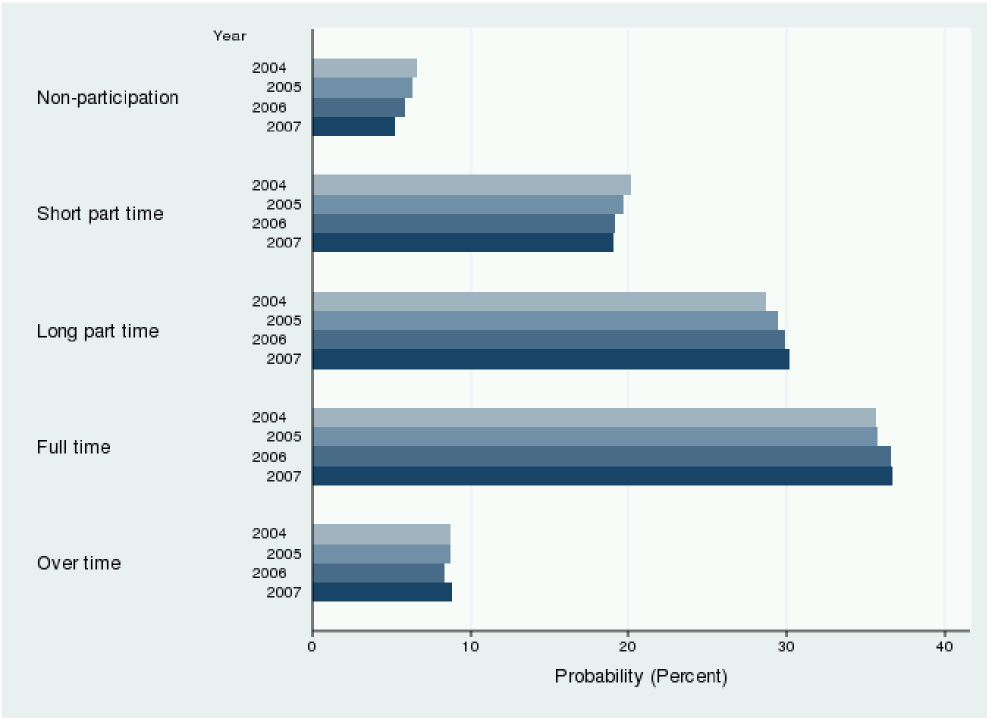
Figure 2 presents the aggregated predicted marginal distribution of working time, which is in close correspondance to the observed marginal distribution reported in Section 3, Table 1. Moreover, Table 4 presents the predicted labor market transitions from the period 2005-2006, which can be compared to the observed transition probabilities in Section 3, Table 2. We find that also the observed persistence is well captured by the model.

Further, we conduct a simple out-of-sample prediction test of the model for year 2008 (one year after the last estimation year). We use the estimated model parameters to predict the transition probabilities given the relevant tax schedule and relevant regional characteristics for year 2008. Both the marginal probabilities and the transition probabilities are again in close correspondance with the observed measures (see Table A.4 and Figure A.1 in Appendix). However, as there were no large changes in taxes from 2007 to 2008, a static specification provides similar out-of-sample marginal probabilities as our intertemporal model.

Table 4: Predicted transition probabilities 2005-2006

		2006				
		No work	Short part time	Long part time	Full time	Over time
2005	Non-participation	**83.0%**	12.8%	3.9%	0.2%	0.0%
	Short part time	2.9%	**75.8%**	14.7%	6.3%	0.3%
	Long part time	0.1%	8.5%	**75.1%**	14.2%	2.3%
	Full time	0.0%	3.1%	8.4%	**79.7%**	8.7%
	Over time	0.0%	3.7%	10.5%	26.2%	**59.6%**

Figure 2: Predicted choice probabilities



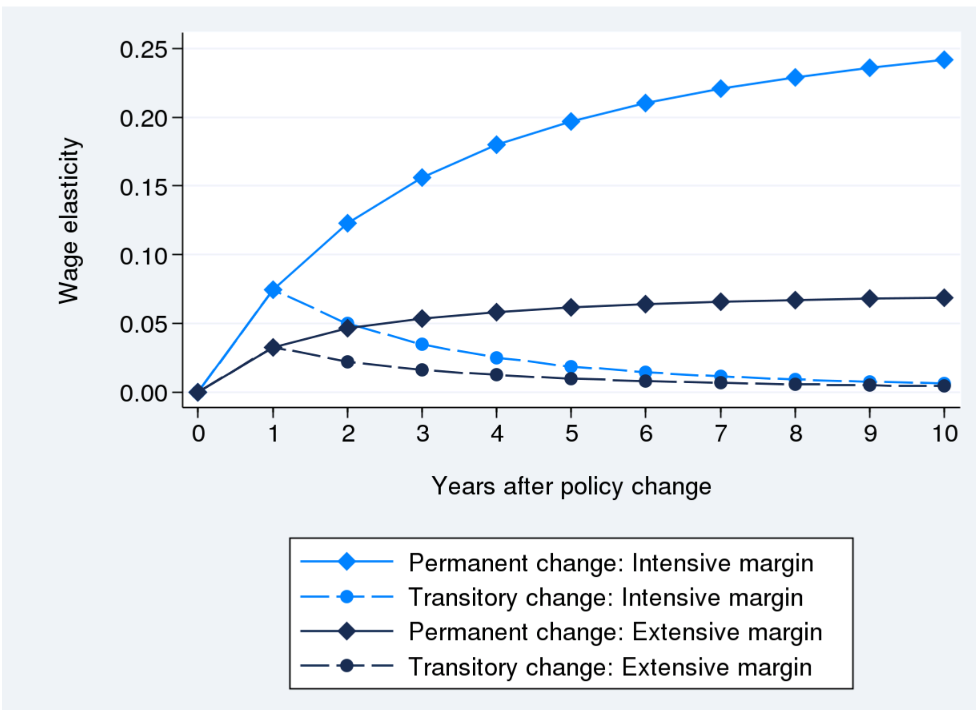
Wage elasticities

The wage elasticities are simulated by increasing the gross wage rate by ten percent and measuring the percentage change in predicted working hours for each individual over time; the average elasticity over individuals is reported. The analysis is conducted by starting with the predicted probability distribution for 2004 by the observed initial conditions and obtaining a probability distribution for the next year according to the following formula $\varphi_{it}(h_t|h_0) = \sum_{h_{t-1=j}} \varphi_{it}(h|h_{t-1}, h_0) \cdot \varphi_{it-1}(h_{t-1}|h_0)$. For each individual we compare a reference path with actual 2007-wage income to an alternative path with a transitory or permanent increase in the wage rate in year 2007. Prior to 2007 the two paths are identical, and then they diverge from 2007 ("1st year") onwards. The 2007 tax system and non-labor income is left unchanged for both paths in the subsequent years.

Figure 3 provides a graphical illustration of the wage elasticities over time (see also Table A.5 and A.6 in Appendix).

The results suggest that within the first year of a permanent change in the wage rate about 30 percent of the full labor supply response can be expected, whereas 90 percent of the response is reached after about 4 years at the extensive and 8 year at the intensive margin. Haan (2010) finds a similar pattern for German women, in which the first year responses correspond to almost 40 percent, and where 90 percent of the long run value is reached after about 4 years.

Figure 3: Predicted wage elasticities over time



Note: The estimated standard errors are very small (below 0.005) and therefore not explicitly reported. Be aware of that this impression of high preciseness of the elasticity estimate is somewhat artificial given that the estimated standard errors hinges on assumptions regarding the particular model specification.

Table 5: Share of short run relative to long run responses

General tax cut	1 st year responses	10 th year responses	1st year/10th year share
Small (1 p.p.)	+0.04 hours	+0.12 hours	34.4 %
Medium (5 p.p.)	+0.22 hours	+0.59 hours	36.7 %
Large (10 p.p.)	+0.43 hours	+1.16 hours	37.4 %

For comparison we also estimated a static model on pooled data over the period 2004-2007. The predicted wage elasticities from the static model (still the 2007 situation is used as reference) are 0.07 at the extensive and 0.28 at the intensive margin. Interestingly, the results from the static model are very close to the long run predictions from our intertemporal model. This suggests that the static model performs well in a longer time perspective, but might seriously overstate the responses to a tax reform in the short run.

Tax simulations of small versus large tax reforms

When fixed adjustment costs are crucial, one might see that the responses to larger tax changes are faster than the responses to small tax changes (see for example discussions in Chetty, Guren, Manoli, and Weber, 2011; Chetty, Friedman, Olsen, and Pistaferri, 2011). Indeed, we find this effect, although modest, presented in Table 5 by comparing 1st year responses and 10th year responses for a small, medium and large tax reform, where the general tax rate is cut by respectively 1, 5 and 10 percentage points.[22]

The separate effect of each source of state dependence on persistence

Previous studies such as Haan (2010) concentrated on a very flexible specification of state dependence, by estimating a parameter for each possible pair of working time transitions. One of the main contributions of our study is that we attempt to analyze the concept of state dependence more thoroughly, by separating out different sources of state dependence, and give them a structural interpretation. We have considered three different sources of structural state dependence in our framework: adjustment cost, state dependence in preferences, and state dependence in labor market opportunities. Moreover, we allowed for unobserved heterogeneity which generates so called "spurious state dependence". Given the estimated model, we would like to shed more light on the relative importance of these four different sources. Since the estimated model parameters related to state dependence are not directly comparable, we will depart from a measure based on the predicted labor supply transition matrices from period $t-1$ to period t : $P(h_t|h_{t-1})$.

[22] We treat husbands taxes as fixed, and do not alter the tax rates for joint taxation.

First, consider a benchmark case, where we eliminate all sources of structural state dependence in the estimated model. We impose the following restrictions: i. There are no adjustment cost, i.e. $ac = 0$; ii. Past labor supply behaviors have no impact on the preference parameters, i.e. replace $\beta_0(h_{t-1})$ by $E_{h_{t-1}}[\beta_0(h_{t-1})]$; iii. Past labor supply behaviors have no impact on labor market opportunity, i.e. replace $\theta_0(h_{t-1})$ by $E_{h_{t-1}}[\theta_0(h_{t-1})]$. For period t, the marginal choice probabilities becomes

$$\varphi_t(h_t) = \frac{exp(E_{t-1}(v_t(C,h_t|h_{t-1})))g_t(h)}{v_t(C,0|h_{t-1})(E_{h_{t-1}}[\theta_0(h_{t-1})])^{-1} + \sum_{h>0} exp(v_t(C,h|h_{t-1}))g_t(h)} \tag{6}$$

where $E_{t-1}(v_t(C,h_t|h_{t-1}))= \alpha_0 \cdot \frac{(C_t-C_0)^{\alpha_1}-1}{\alpha_1} + E_{h_{t-1}}[\beta_0(h_{t-1})] \cdot \frac{(L_t)^{\beta_1}-1}{\beta_1}$. We assume for simplicity that the choice environment does not vary over time, and as there is no state dependency, the marginal (conditional) labor supply choice probabilities are constant over time, i.e.$\varphi_t(h_t)$ =$\varphi_{t-1}(h_{t-1})$. Thus, the marginal probabilities in the benchmark case can be solved by iterations and the transition matrix P_{bench} follows.[23]

Now, consider three cases where only one source of state dependence is constrained (either i. ii. or iii). We denote the constrained transition matrices as $P_{con(source)}$ and the transition matrix of the full model as P_{full}.

To measure the relative contribution for each of the three sources of state dependence, we define

$$\kappa_{source}(h) = \frac{P_{full}(h_t|h_{t-1}) - P_{con(source)}(h_t|h_{t-1})}{P_{full}(h_t|h_{t-1}) - P_{bench}(h_t|h_{t-1})}$$

which lies in the interval $[0,1]$. When $\kappa_{source}(h) = 0$ one can disregard the specific source of state dependence (i,ii or iii) without deviating from the full model transition matrix ($P_{con(source)}(h_t|h_{t-1}) = P_{full}(h_t|h_{t-1})$), whereas when $\kappa_{source}(h) = 1$ the other sources of state dependence can be considered redundant as we are left with the same transition matrix as in a scenario with no state dependence altogether ($P_{con(source)}(h_t|h_{t-1}) = P_{bench}(h_t|h_{t-1})$).

The results for $h_{t-1} = h_t$, and conditional on the individual types (unobserved heterogeneity), are reported in table 6. Notice that the different sources of state dependence work jointly given the nonlinear nature of the model which means that the measures from different sources will typically not sum up to one. The cells for the non-participation alternative of type I can be read as follows: Adjustment costs alone contribute to 40% of the predicted persistence in non-participation. State dependence in opportunities accounts for 93%, whereas state dependence in preferences capture 88% of the persistence. The sum of contributions largely exceeds 1 which indicate that for the non-participating alternative there are strong interaction effects of disregarding all sources of state dependence.

Among the three sources, adjustment costs seem in general to be the most important one – especially for

[23] We use an "average" female and simulate the labor supply behavior using the model parameters in Table 3. The benchmark marginal choice probabilities for each type are solved using eq. (6). For type I individual: $(0.11\%, 12.39\%, 28.42\%, 45.11\%, 13.97\%)$; and for type II $(0.09\%, 22.93\%, 39.63\%, 33.26\%, 3.22\%)$.

Table 6: Share of persistence explained by each source of state dependence

	Non-participation	Short part time	Long part time	Full time	Overtime
Type I					
Adjustment costs	0.40	0.47	0.76	0.92	0.50
State dependence in preferences	0.88	0.73	0.28	0.08	0.67
State dependence in opportunities	0.93	-	-	-	-
Type II					
Adjustment costs	0.49	0.85	0.99	0.81	0.85
State dependence in preferences	0.22	0.10	0.02	0.06	0.51
State dependence in opportunities	0.45	-	-	-	-

type II individuals. State dependence in preferences play a lesser role, although some exceptions. State dependence in opportunities have a strong effect on the choice of participation, but almost no effect on the other working time alternatives. This is essentially due to the specification of the model where the distribution of available jobs across different working hours $g(h)$ is assumed to be independent of past choices.

So far, we have ignored the effect of unobserved heterogeneity as the above analysis was done conditional on the individual types. To investigate how much unobserved heterogeneity contribute to the observed persistence in our model estimation, we define the corresponding aggregated transition matrix $P_{aggregate}$ as follows :

$$\begin{aligned} P_{aggregate}(h_t|h_{t-1}) &= \frac{P_{aggregate}(h_t,h_{t-1})}{P_{aggregate}(h_{t-1})} \\ &= \frac{P_{type1}(h_t|h_{t-1})P_{type1}(h_{t-1})p + P_{type2}(h_t|h_{t-1})P_{type2}(h_{t-1})(1-p)}{P_{type1}(h_{t-1})p + P_{type2}(h_{t-1})(1-p)} \end{aligned}$$

where p is the share of type I individuals. We start from the benchmark cases for the two types of individuals, and construct the "benchmark aggregated transition matrix". As no structural state dependence is present, all the correlation in the labor supply behavior over time in the aggregated benchmark matrix is due to unobserved heterogeneity. As we can see from the aggregated transition matrix, Table A.8, there are some indications of correlations of labor market status over time (compared to the type-specific benchmark transition matrices which would have identical figures in each column), but not very strong – the diagonal elements do not stand out compared with others. Thus, we conclude

Table 7: Robustness check: Unobserved heterogeneity and state dependence in wages

	Wage elasticities					
	Intensive margin			Extensive margin		
	Short-run[a]	Long-run[b]	Time[c]	Short-run[a]	Long-run[b]	Time[c]
Reference (2 mass points)	0.07	0.25	8	0.04	0.07	6
3 mass points in $(\alpha_{0i}, b_{0i}, ac_i, \gamma_{0i})$	0.08	0.24	7	0.03	0.07	6
4 mass points in $(\alpha_{0i}, b_{0i}, ac_i, \gamma_{0i})$	0.08	0.24	7	0.03	0.06	6
State dep and unobs het. in wage rate	0.08	0.28	8	0.04	0.08	5

a First year responses, *b* 15th year responses, *c* Years until 90% of long run effect reached

that unobserved heterogeneity was not a major factor of the observed state dependence over time in our estimated model.

Robustness checks

In the following we provide a set of robustness checks which are estimated on a random subset of the estimation sample (in order to reduce computational time).

First, we check a set of alternative model specifications regarding state dependence and adjustment costs to test if for instance the effect of adjustment costs would be captured by the parameters for state dependence in preferences if not separately specified.[24] In Appendix, Table A.9 we report the log likelihood measure together with the Akaike Information criteria (AIC) and the Bayesian Information criteria (BIC) for the alternative model specifications. We find that the preferred specification corresponds to our baseline model.

In Table 7 we report robustness checks on wage elasticties regarding the number of mass points which is used to specify unobserved heterogeneity. We can see that introducing more mass points hardly affect the results. When introducing more than four mass points, we experience problems with convergence. In the literature one rarely see more than four mass point specifications.

In our main analysis we focused on state dependence in preferences and in the opportunity measure. However, state dependence might also be present in wages. The idea is that previous working time can be seen as an investment in (job-specific) human capital which increases productivity and leads to a rise in the hourly wage rate. We therefore conduct a robustness check of our main results by using

[24] Notice that we here estimate and test alternative more restrictive models, in contrast to last section where the idea was to switch off the effect of some parameters in the preferred specification.

an alternative specification on wage dynamics which takes both unobserved heterogeneity and state dependence in the wage rate into account.[25]

The wage regression is carried out by a random effects specification where last periods working hours is included as explanatory variable, see Table A.10. The random effects specification rests on the usual assumption that the individual unobserved effect is independent of the full set of explanatory variables. We allow for a quadratic form of last periods working hours and find a positive, concave relation as expected (although the effect is quite small). We add random effects to the predicted individual wage rates by drawing from the distribution of individual specific effects predicted from the model. 30 draws are picked for each individual.

The random effects in the wage rate for each individual is treated analogously to the random coefficients in the preference and opportunity measure specification, by maximizing the simulated likelihood defined as

$$lnL_i = ln\frac{\sum_{r=1}^{R} L_i(wage_r)}{R}$$

where $R = 30$.[26]

Notice that each individual now have five different wage rates dependent on last periods choice. We find that the implication of using this alternative specification of the wage rate relatively to the standard specification of a pooled Heckman selection model is small both with regards to model predictions and in terms of predicted wage elasticities over time.

7 Conclusion and discussion

Lagged labor supply responses are typically ignored, both with respect to a recent discussion of low responsiveness due to adjustment costs (Chetty etc), and in structural discrete choice models with an aim to simulate prospective policy analyses. We go further than previous studies on extending the discrete choice model to an intertemporal setting (Haan, 2010, Prowse, 2012) by departing from a more structural framework denoted the static job choice model (see Dagsvik and Jia, 2014) in order to distinguish between different sources of state dependence. In particular, we incorporate a more clear-cut fixed cost of adjustment (see e.g. Chetty, Friedman, Olsen, and Pistaferri, 2011; Chetty, Guren, Manoli, and Weber, 2011; Chetty, 2012).

We find, by using Norwegian wage register data on married and cohabiting women, that for a ten percent permanent increase in the individual wage rate, within the first year only about 30 percent of the full

[25] Remember that in our main specification we used a pooled Heckman selection regression.
[26] See e.g. Train (2009) for the asymptotic properties of maximum simulated likelihood.

labor supply response can be expected, whereas 90 percent of the response is reached after about 7 years. According to our analyses, a static model simulation replicate the ‘long run’ results and clearly overestimate predictions in the short run. By analyzing each sources of state dependence separately, we find that both state dependence in preferences and state dependence in job opportunities are crucial to explain labor supply persistence. However, according to our findings, fixed costs of adjustments seems to be the major part of what Heckman (1981) denotes as true state dependence.

Notice that we only consider a subgroup of the population in our empirical analysis such that the results might not correspond to the population of wage earners in total.[27] Still, we believe that the results for women in couples can provide an interesting example of the dynamics of labor supply responses in general.

Notice that our model is myopic since it assumes that individuals optimize on a period to period basis. The dynamics follow from the fact that this period’s decision will influence next period’s evaluation of preferences and opportunities. It follows that there is no scope for intertemporal adjustment at the individual level in our model, which is clearly a simplification (e.g., see Altonji, 1986; Imai and Keane, 2004), as one might think of intertemporal adjustment in working hours following the announcement of a tax reform, see e.g Blundell, Francesconi, and van der Klaauw (2011). A life cycle model is a more sophisticated framework to approach in this respect, but it is not clear to which degree individuals indeed are long-sighted rational, and it can be questioned how appropriate this type of model is to analyze relatively short sighted tax reforms. Moreover, as one might gain more reality in terms of a life cycle perspective, one looses the possibility for other details necessary to defend a model which can be used for policy makers in practice.

Although our model seem quite restrictive in how we have defined state dependence and adjustment costs, we attain a similar fit to observed transitions as for instance the model presented by Haan (2010) which is estimated on a much more flexible (and ad-hoc) specification of state dependence. The advantage of more structure is that we have a framework to capture the underlying preference parameters and to distinguish the sources of persistence. Keep however in mind that the assumptions regarding time independent unobserved heterogeneity and a special restrictive form of serial correlation might be strong and could influence the estimated effect of state dependence (see e.g. Prowse, 2012), although we have shown that the results are robust to including a more flexible specification of unobserved heterogeneity and robust to including state dependence and unobserved heterogeneity in the wage rate.

Notice that there might be further reasons for lagged responses to tax changes which are not explicitely discussed here, for instance, there might be a information lag, in which individuals learn the tax code after some time. Moreover as alerady mentioned, announced tax changes might have another pattern

[27] The same framework can be used for men and single women. However, women in couples differ from the other groups by a wider variation in working hours. As our data source is administrative and not based on survey information, the measure of working time is relatively crude, and only contains contractual working hours and registered overtime work. As the vast majority of Norwegian males work at least full time, we do not think it is much to gain to estimate the model jointly for the couple, although possible.

of responses than unannounced tax changes. To incorporate these issues in the simulation model is left for future research.

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Appendix

Examples of tax simulations

In the following, we demonstrate how the estimated model can be used for evaluating counterfactual policy reforms. In order to interpret the simulation results, we briefly describe the tax schedule for Norwegian wage earners in year 2007, which serves as the benchmark when conducting alternative tax policies.

Norway has a dual income tax system implemented through two separate tax bases denoted "ordinary income" and "personal income". Ordinary income consist of all taxable income from labor and capital after basic allowance and other tax deductions are subtracted, and is taxed by a flat rate of 28 percent. Personal income includes gross labor income and serves as the tax base for social security contribution (7.8 percent for wage earners) and a two tier surtax rate system. In 2007, the first surtax tier was applied at a rate of 9 percent to incomes above NOK 400,000, and the second tier of 12 percent applied to income in excess of NOK 750,000. Two separate tax classes apply to Norwegian residents. The majority are taxed individually in tax class 1 (more than 90 percent). Single parents are taxed in tax class 2, in which the "personal allowance" is doubled. Tax class 2 also applies to married couples if joint assessment is beneficiary (usually only one-income families).

We consider three counterfactual policy changes presented in Table A.11 with 2007 as a starting point and reference path. The first policy change is the abolishment of the second tax class, which in our context means an abolishment of the possibility of joint taxation for women in couples. The second policy change is an abolishment of the basic and personal tax allowance (corresponds to maximum NOK 100,800 in tax class 1 in 2007), and the third policy change we consider is the abolishment of the two-tier surtax schedule. We briefly comment on the results for each of the three policy changes.

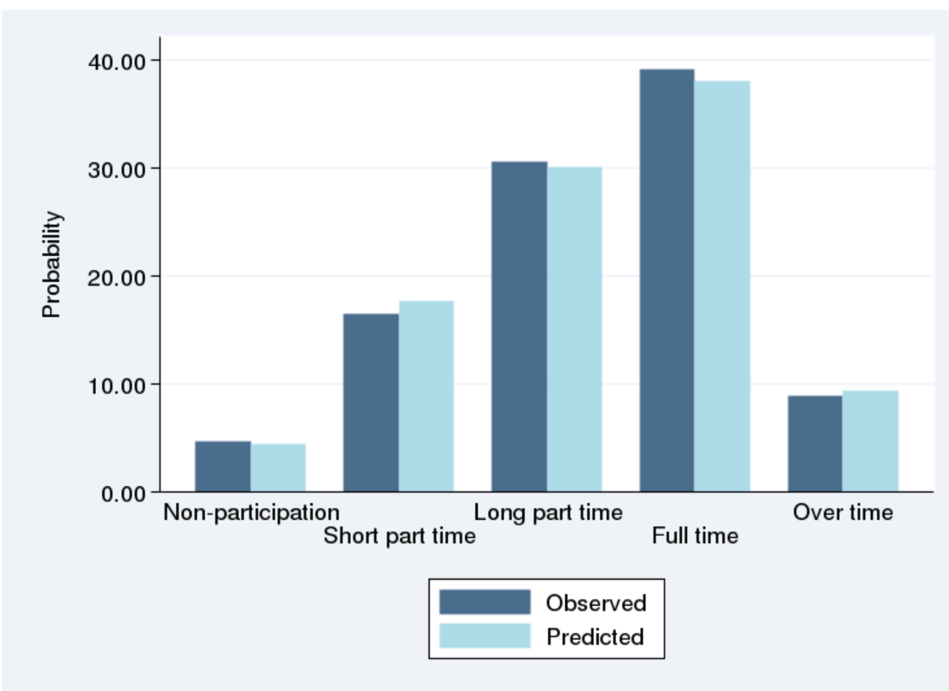
The first counterfactual policy change, the abolishment of joint taxation, leads to a less attractive alternative of non-participation for married women as the husband looses the additional personal allowance. As expected we see a negative predicted effect on the probability of choosing non-participation, and a correspondingly higher probability of longer part time and full time. The non-participation probability decreases by about 0.3 percentage points the first year and by about 0.8 percentage points in the long run. Predicted average working hours increase only slightly by 0.1-0.3 hours per week.

The second policy change, the abolishment of the basic tax allowance decreases the incentives to participate in the labor market. However, one special feature of the Norwegian basic tax allowance ("minstefradrag") is that it is increasing with earnings until a maximum level is reached (amounting to NOK 63,800 for wage earners in 2007). So, in currency amounts, only women working at least long part time or full time are fully profiting from the basic tax allowance. At the same time the tax policy affects husbands disposable income such that there is a negative income effect both from own and husband's disposable income which in itself is expected to increase the number of working hours. The results

Table A.1: Pooled sample characteristics

Variable	Mean	Std. Dev.	Min.	Max.
Age	44.2	8.9	25	62
Child(ren) under age 3	0.18	0.39	0	1
Child(ren) under age 6	0.30	0.46	0	1
Child(ren) under age 12	0.49	0.50	0	1
Years of education	12.1	2.6	0	20
Low education ($\leq$10 years)	0.40	0.49	0	1
Vacancies/Unemployed	0.28	0.20	0.07	1.1
Children per pre-school employee	6.93	0.43	5.75	8.2

Figure A.1: Out of sample fit, year 2008



suggest an increase in non-participation and short part time and a decrease in long part time and full time. We see that all changes are rather sluggish in particular for adjustments in participation and full time.

The last counterfactual policy change we consider, is an abolishment of the surtax schedule, leaving the tax schedule flat, apart from the features of the basic tax allowances and deductions. As expected, this induces an increase in overtime work, since a considerable proportion of the women would be in surtax rate position conducting a job characterized with overtime work. Depending on the wage rate, the choice of working full time would, for a certain fraction of the women, also be affected by the surtax rate abolishment. This effect of increased working hours appears surprisingly low. Instead, it seems like the effect through husband's income is stronger. The majority of husbands in our sample are affected by the abolishment of the surtax rates, which leads to a positive income effect for the women in our modeling framework. The model predicts an overall negative effect on the probability of working long part time and full time. Consequently there is a negative overall effect on working hours.

Similar analyses can be conducted for a range of counterfactual policy changes, these three analyzes serve primarily as examples of how the estimated model can be used for policy makers in practice.

Table A.2: Predicted transition probabilities for the two types of women with average characteristics

Type		Non-participation	Short part time	Long part time	Full time	Over time
I	Non-participation	32.9%	43.7%	20.7%	2.6%	0.0%
	Short part time	2.9%	46.5%	30.3%	19.0%	1.3%
	Long part time	0.1%	12.9%	49.8%	30.9%	6.2%
	Full time	0.0%	5.5%	15.2%	63.3%	16.0%
	Over time	0.0%	4.7%	14.3%	35.5%	45.4%
II	Non-participation	94.6%	3.7%	1.5%	0.1%	0.0%
	Short part time	2.1%	90.1%	5.4%	2.3%	0.1%
	Long part time	0.1%	2.9%	92.6%	4.0%	0.4%
	Full time	0.0%	1.5%	3.6%	93.8%	1.1%
	Over time	0.1%	3.1%	8.2%	13.5%	75.0%

We look at the labor supply dynamics of two hypothetical females who belong to different sub-groups, but are observably identical "average" with observed individual characteristics matching the sample average in year 2006. The choice situation is kept constant as in 2006, i.e. neither individual characteristics nor labor market institutions change over time.

Table A.3: Predicted choice probabilities

	2004	2005	2006	2007
Non-participation	6.6%	6.4%	5.9%	5.2%
Short part time	20.2%	19.7%	19.2%	19.0%
Long part time	28.7%	29.5%	29.9%	30.2%
Full time	35.7%	35.8%	36.6%	36.7%
Over time	8.8%	8.7%	8.4%	8.8%

Table A.4: Out-of-sample predicted and observed transition probabilities 2007-2008

	Predicted	2008				
		Non-participation	Short part time	Long part time	Full time	Over time
2007	Non-participation	**84.4%**	11.5%	3.7%	0.3%	0.0%
	Short part time	1.9%	**76.3%**	14.9%	6.7%	0.3%
	Long part time	0.0%	8.5%	**74.7%**	14.4%	2.3%
	Full time	0.0%	3.3%	8.7%	**79.3%**	8.6%
	Over time	0.0%	3.9%	11.3%	28.0%	**56.8%**
	Observed	2008				
		Non-participation	Short part time	Long part time	Full time	Over time
2007	Non-participation	**84.8%**	13.0%	1.3%	0.9%	0.0%
	Short part time	2.5%	**76.4%**	16.2%	4.0%	0.9%
	Long part time	0.2%	7.3%	**77.8%**	12.4%	2.4%
	Full time	0.1%	1.4%	8.4%	**82.7%**	7.4%
	Over time	0.3%	1.3%	5.8%	37.1%	**55.5%**

Table A.5: Predicted wage elasticities over time (ref. year 2007)

	Wage elasticities[a]			
	Transitory change		Permanent change	
	Intensive margin	Extensive margin	Intensive margin	Extensive margin
1st year	0.07	0.03	0.07	0.03
2nd year	0.05	0.02	0.12	0.05
3rd year	0.03	0.02	0.16	0.05
4th year	0.02	0.01	0.18	0.06
5th year	0.02	0.01	0.20	0.06
15th year	0.00	0.00	0.26	0.07
Pooled static	0.28	0.07	0.28	0.07

a The estimated standard errors are very small (below 0.005) and therefore not explicitly reported. Be aware of that this impression of high preciseness of the elasticity estimate is somewhat artificial given that the estimated standard errors hinges on assumptions regarding the particular model specification.

Table A.6: Predicted income elasticities over time (ref. year 2007)

	Income elasticities (Std error)			
	Transitory change		Permanent change	
	Intensive margin	Extensive margin	Intensive margin	Extensive margin
1st year	-0.02	-0.01	-0.02	-0.01
2nd year	-0.01	-0.01	-0.03	-0.01
3rd year	-0.01	-0.00	-0.04	-0.01
4th year	-0.01	-0.00	-0.04	-0.01
5th year	-0.00	-0.00	-0.05	-0.02
15th year	-0.00	-0.00	-0.06	-0.02

Table A.7: Pooled Heckman 2-Stage wage regression: Women in couples

	ln(Wage)		Participation	
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 origin	-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)
Constant	4.6834***	(0.0068)	1.5323***	(0.0653)
Exclusion restrictions				
No. children under age 3			-0.1002***	(0.0206)
No. children under age 6			-0.0879***	(0.0174)
No. children under age 12			-0.3349***	(0.0098)
Wealth (in NOK 10,000)			-0.0000**	(0.0000)
Husband'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			

Standard errors in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.8: The benchmark aggregated transition matrix

	Non-participation	Short part time	Long part time	Full time	Overtime
Non-participation	0.9%	22.0%	38.7%	34.3%	4.1%
Short part time	0.7%	19.6%	36.1%	37.0%	6.6%
Long part time	0.6%	18.9%	35.4%	37.8%	7.3%
Full time	0.5%	17.3%	33.6%	39.6%	9.0%
Overtime	0.3%	14.6%	30.8%	42.6%	11.7%

Table A.9: Alternative specifications of state dependence

	log L	AIC	BIC	Rank
Baseline model	-25753	51576	51825	1
No state dependence in preferences	-26387	52828	53020	2
No adjustment costs	-28839	57744	57978	3
No state dependence and no adjustment costs	-32935	65920	66098	4
Pooled static model	-48894	97822	97943	5

The baseline model equates to the main specification described in the previous sections, apart from the smaller sample size. The second model is specified without any state dependence in preferences. The third model is specified without adjustment costs, and the fourth model is further specified without any state dependence in preferences and opportunities, and without adjustment costs. It still includes initial conditions from period 2003 and unobserved heterogeneity. The last model is a pooled static model for the years 2004-2007, where the time dimension is completely ignored. The static model includes unobserved heterogeneity which varies both over individuals and over time.

Table A.10: Alternative wage regression: Random effects regression with state dependence

	ln(Wage)	
Hours worked $t-1$ (Scale 10^{-1})	0.0163***	(0.0028)
Hours worked $t-1$ squared (Scale 10^{-2})	-0.0005	(0.0005)
Experience	0.0126***	(0.0004)
Experience squared	-0.0002***	(0.0000)
Low education	-0.0650***	(0.0027)
High education	0.2223***	(0.0030)
Non-western origin	-0.1156***	(0.0067)
Residence in metropolitan area	0.0672***	(0.0023)
Year dummies (base: 2003)		
Year 2005	0.0306***	(0.0007)
Year 2006	0.0750***	(0.0007)
Year 2007	0.1270***	(0.0008)
Education category (base: "unknown")		
General	-0.0280**	(0.0089)
Human, Art	-0.0434***	(0.0096)
Education	-0.0736***	(0.0094)
Social, Law	0.0452***	(0.0109)
Business	-0.0037	(0.0090)
Technology	0.0146	(0.0093)
Health	-0.0556***	(0.0089)
Primary	-0.0122	(0.0160)
Service	-0.0287**	(0.0106)
Constant	4.7724***	(0.0112)
Variance over individuals (σ_u)	0.1531	
Mills Lambda	0.0273***	(0.0058)
Number of observations	72,990	
Number of individuals	21,638	

Standard errors in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table A.11: Simulation results - 3 hypothetical tax reforms

	Change in percentage points (relative to reference path)					
Abolish joint taxation	Non-participation	Short part time	Long part time	Full time	Over time	ΔWorking hours
1st year	-0.3	0.1	0.2	0.1	0.0	0.09
2nd year	-0.5	0.1	0.2	0.2	0.0	0.15
3rd year	-0.6	0.0	0.3	0.2	0.0	0.18
10th year	-0.8	-0.1	0.4	0.4	0.1	0.28

Abolish basic tax allowance	Non-participation	Short part time	Long part time	Full time	Over time	ΔWorking hours
1st year	0.2	0.9	-0.7	-0.4	0.1	-0.19
2nd year	0.3	1.4	-1.1	-0.7	0.0	-0.32
3rd year	0.4	1.8	-1.3	-0.9	0.0	-0.41
10th year	0.7	2.8	-1.9	-1.6	-0.1	-0.72

Abolish surtax schedules	Non-participation	Short part time	Long part time	Full time	Over time	ΔWorking hours
1st year	0.0	0.3	-0.2	-0.3	0.1	-0.05
2nd year	0.0	0.6	-0.3	-0.4	0.1	-0.09
3rd year	0.0	0.7	-0.3	-0.5	0.1	-0.12
10th year	0.0	1.1	-0.4	-0.8	0.2	-0.20