Fixed-term contracts, incentives and effort*

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Abstract

This paper focuses on labor market transitions and especially on those involving fixed-term contracts. Our contribution is twofold: first, we provide an accurate measure of labor market transitions; second, we analyze the potential incentive effect of fixed-term contracts on "effort". To deal with unobserved heterogeneity, we use a dynamic multinomial logit with fixed effects, following the method proposed by Magnac (2000). We construct an indicator of effort for fixed-term workers based on a "compared" weekly working time. Using French data, we find that a fixed-term contract provides significantly better perspectives than unemployment, but no evidence of any significant impact of exerting more effort on the probability of getting an open-ended contract.

JEL: J60, J24, J41

Key words: *Fixed-term contracts, incentives, effort, transitions.*

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Introduction

Like many other European countries, France enhanced in the mid-1980s the labor market flexibility by allowing employers to hire workers on a fixed-term basis. These short-term contracts have been widely used over the past two decades, e.g. "fixed-term contracts" but also temporary work (*interim* missions). Short-term contracts replace progressively usual open-ended contracts as common hiring device: in 2007, more than two over three French hiring contracts have a fixed duration. Short-term work accounts for more than 12% of the French labor workforce, which is a huge figure compared with what this proportion was twenty years before – almost zero. As a result, the emergence of fixed-term contracts as a massive phenomenon is often considered to be one of the striking features of the current French labor market after persisting unemployment.

This flexibility was motivated by recurrent criticisms on European labor markets' sclerosis (see Bentolila and Bartola, 1990); in fact, labor protections were created during the *Glorious Thirty* and would be now responsible for a slowdown in the job creation process. On the contrary, employment *at will* would help the US economy to keep on creating new jobs.

Yet labor market consequences for workers of short spells are still unknown. As stated by Booth, Francesconi and Franck (2002) in a seminal paper, the question is to know whether this kind of jobs are a *stepping stone* to permanent employment or a *dead end*. This is a difficult empirical question since individual heterogeneity is likely to matter a lot in transitions on the labor market. One must be able to correctly distinguish *state dependence* from *unobserved heterogeneity*, an old puz-

zle brought into the literature by Heckman (1981). To date, few papers dealing with short-term employment have tackled this issue, especially in the French case.

Reasons why temporary jobs turn into permanent jobs are not well-known either. On the one hand, these jobs are an obvious answer to demand shocks on the goods market. In that case, we should expect conversion of fixed-term into open-ended contracts to occur only when the transitory shock turns out to be permanent; and the employee's "quality" is not involved here. On the other hand, a fixed-term contract could be used like a screening device, as Engellandt and Riphahn (2005) suggest, and then the fixed-term contract – like a probationary period – would provide incentives: either the employee is making effort to send a good signal on his productivity and he should see his chances of getting a "promotion" increase, or he's shirking and sending some bad signal: then he is not hired at the end of the initial contract.

Using the recent French Labor Force Survey (LFS), we estimate a model of labor market transitions. To distinguish carefully state dependence from unobserved heterogeneity, we estimate a dynamic multinomial logit with fixed effects, following Magnac (2000). The French LFS provides very detailed quarterly information on labor market spells; the large sample size allows us to examine different labor market situations. It's therefore well suited to a transition analysis. Moreover, it has some useful information on actual weekly working time, which we use to construct our proxy for effort. To be more precise, we compare the weekly number of working hours for a given fixed-term worker with the distribution of working hours of other fixed-term workers in the same industry. We are then able to estimate

whether a fixed-term worker who exerts more effort is more likely *ceteris paribus* than the shirker to get his contract converted.

Our main conclusions are the following: (a) on the whole, a spell of fixed-term contract is likely to increase the probability of getting a permanent job and offers definitely better perspectives than unemployment to this regard; (b) but we don't observe evidence of any significant effect of exerting effort on such transition rates once unobserved heterogeneity is taken into account. This last conclusion is robust to various specifications for unobserved heterogeneity.

Note that these results come from reduced-form estimation and describe the current French labor market; they wouldn't necessarily hold in case of a structural change in the market. It is not however the issue we are willing to deal with.

The next section is devoted to a brief literature survey on short-term contracts. We describe the data and our measurement of effort in section 2. In section 3, we discuss the econometric method to identify state dependence, results of which are exposed in section 4. Last section presents some robustness checks.

1 Short-term contracts, labor demand and effort

1.1 Stepping-stone or dead-end?

Since the seminal contribution of Booth *et al.* (*ibid.*) that was comparing the situation of fixed-term workers with the one of permanent workers in the UK, several papers tried to characterize career paths of such workers. This is a nontrivial problem to tackle since individual characteristics are likely to matter a lot. The

question is, are these contracts truly a better opportunity than unemployment regarding future outcomes on the labor market? On Swedish data, Zijl, Van der Berg and Heyma (2004) conclude that fixed-term contracts reduce the unemployment duration and increase substantially the proportion of unemployed obtaining a permanent job within years following the first unemployment spell. Analyzing the labor market in Belgium, Cockx and Picchio (2009) find a stepping-stone effect. Gagliarducci (2005) looks at the impact of having recurrent spells of temporary jobs on the probability of finding a permanent job; on Italian data, he shows that nonlinear phenomena occur over time: this propension is high at the beginning but decreases after some time. People knowing such recurrent phases, especially when they are followed by unemployment spells, have a lower probability of finding a permanent job, as expected. Autor and Houseman (2005) use a quasi-experiment to find that temporary work has a negative impact on future employment.

Some papers paid attention to the French case; they were mainly focused on the evaluation of specific training programs, some examples are provided by Bonnal, Fougère, Sérandon (1997), Magnac (2000) and Havet (2006). Adopting a slightly different approach, Beffy, Coudin and Rathelot (2008) analyze labor market trajectories of temporary workers assuming that the labor market is divided *ex-ante* between *movers* and *stayers*.

1.2 The nature of short-term contracts

The use of such fixed-term contracts has been justified by the firm's need for being insured against risk. On the one hand, she faces uncertainty related to demand shocks. In an unstable economic environment, firms are using fixed-term contracts to face purely unexpected and transitory demand shocks (Givord and

Maurin, 2001): indeed, expecting dismissal indemnities imposed by law would prevent employers from offering open-ended contracts by fear of additional costs in case of low demand states. Changes in firm's organization schemes like the generalization of the *just-in-time* and the diffusion of new technologies that make employees closer substitutes each other would then account for this increase in demand for fixed-term work. This is also in relation with the assertion that too much labor protection has a depressing impact on the level of employment, an idea which has been defended among others by Bentolila and Bertola (1990).

On the other hand, the firm faces another source of uncertainty: she doesn't know the true quality of her employee (Lazear, 1995). The firm is not able to infer *ex-ante* his productivity, although she tries to gather the maximum of relevant information on it through job interviews. She might use a probationary period to screen the productivity: in practice, even permanent jobs do incorporate a probationary period at the end of which an employee can either be fired or hired. It is sometimes argued that fixed-term contracts could be a new type of probationary period.

In this setting, fixed-term workers would have incentives to make effort in order to get a permanent job – at least during the initial duration of the contract – because as rational agents employees expect their outcome on the labor market to depend on the signal they send. This signal however is a mix of intrinsic productivity and effort*. Riphahn (2004) wonders whether the labor protection encourages shirking. She finds some evidence of that, by measuring (non-)effort thanks to the absen-

*Notice that we assume that such workers actually want an open-ended contract. Indeed, according to a recent LFS, this is not the case for only one fifth of FTC workers – mostly students with summer jobs.

teeism rate (see also Ichino and Riphahn, 2005). Engellandt and Riphahn (2005) exploit the Swiss labor force survey and show that employees with a fixed-term contract do significantly more overtime work than people belonging to the permanent workforce. Moreover, the absenteeism is lower for temporary workers. There seems to be here some evidence of the existence of incentives (Anger, 2008). However, such an analysis is performed on cross-sectional data and could result from composition effects; more generally, it is likely to miss unobserved heterogeneity effects.

A dynamic analysis of transitions from fixed-term to permanent contracts should give more accurate answers to these questions, especially to the existence of an incentive effect. Meyer and Wallette (2005) implement such an idea and use similar proxies for effort than Engellandt and Riphahn. On Swedish data, they don't observe any impact of absenteeism on the transition rate from a fixed-term to an open-ended contract. Yet, the use of the absenteeism indicator casts doubts on the interpretation of it as a good measure for effort. Contrary to Engellandt and Riphahn, they don't see any cross-sectional evidence of differences in overtime work between temporary and permanent workers, which means in particular that overtime work behavior is likely to differ across countries; especially no conclusion can be drawn for now in the French case.

Finally, we do not discriminate among the two motivations for the use of short-term contracts: uncertainty towards future demand and with respect to employee's quality. We think there is no hope to provide any definite answer to that puzzle. First, it is probably useless to oppose both of these explanations since they can

easily be concealed with: employers may hire temporary workers because of a peak but also screen him and turn his contract into an open-ended one if the employee proved himself as a good worker. Second, effort measurements can be interpreted in both ways: a firm hiring short-term workers to face a boom of activity may also ask them to work more (more than a firm that would have hired them as probationary)... if the boom turns out to be persistent, the short-term contract is also more likely to end up with a permanent job.

1.3 Distinguishing state dependence from unobserved heterogeneity

Our empirical questions could be stated as: 1. how likely are FTC workers to get an open-ended contract compared with being unemployed? 2. does "effort" impact this chance of getting an open-ended contract?

From an econometric point of view, it could be rephrased as: do we observe state dependence in labor market transitions, that is, does past occupation affect the likelihood of occupying some given state on the labor market? As usual, this analysis is complicated by heterogeneity issues. Transitions in the labor market may be related to observed or unobserved individual characteristics. People with a fixed-term contract might have special characteristics that are likely to affect their career paths. Indeed, probationary periods used like a screening device lead to more selection (see Loh, 1994 and Marinescu, 2007). The basic problem for us will then consist in determining the true impact of contract and effort, which can be correlated to characteristics – observed or not – impacting the position on the labor market. For instance, one can think of people with no family constraints, who dispose of more available time, and who are therefore more likely to do over-

time work and to get an open-ended contract for the same reason.

To properly disentangle state dependence from unobserved heterogeneity, we follow the method suggested by Magnac (2000). He extends the conditional logit procedure proposed by Chamberlain (1984) to the multinomial case, allowing for different transitions to occur. This model incorporates fixed effects that constitute a satisfying nonparametric way of dealing with unobserved heterogeneity. One of the decisive advantage of this method is that it does not rely on some parametric assumption as random-effect models usually do.

2 Data

2.1 Source

We use the Labor Force Survey provided by the INSEE, the French National Statistical Institute. This survey consists in a rotating panel of a large sample (75,000 people). Quarterly data are available from 2002 to 2008. In theory, each individual is interrogated six consecutive quarters, the sample being renewed by sixth every quarter. The survey provides detailed information on individual characteristics and on career paths as well. It includes a precise description of the working time, the type and duration of the contract, and is therefore well-suited to our analysis. To go into more details, we know exactly how many hours an employee worked a given week, the so-called "reference week" – a week is assigned to every individual in the sample and this assignment is designed to correct for seasonal effects like holidays or public holidays. In most of cases, this variable is equal to 35 or 39 hours – legal durations. We will rely on this variable to construct our measure of effort (see *infra*). We know little about overtime work: this information is unfor-

tunately not available every quarter. To sum up, the main interest of this source for our purpose is to give precise and relevant information at the quarterly level that does not come from retrospective surveys, the quality of which is known to be potentially problematic (Magnac and Visser, 1999).

2.2 Description of the sample

In our final sample, we keep people aged from 18 to 60 who are interviewed six quarters consecutively. In fact, we need at least four observations per individual to identify and estimate the state dependence (see *infra*). To simplify, we use a balanced panel with six observations per individual; endogenous attrition could be an issue here since temporary workers are more likely to move than "insiders" and therefore to disappear from the original sample. We are left with more than 125,000 individuals, which is a quite large sample: it enables us to divide it into several categories. In first interrogation, concerning the labor market situation, 7% of individuals are unemployed and 22% are non-participant. Half of employees have a full-time permanent contract in the private sector. A small part of the population only has a short-term contract: $1,8\%$ has a fixed-term contract and $1,2\%$ works on an interim mission. Remember that the boom of short-term contracts is rather new, that's why it concerns only a specific part of the population in stock. More details are provided by Table 1.

Some general descriptive statistics can be found in Table 2. People with a fixed-term contract (FTC) are most of the time women, employees and work in services industry; interim missions on the contrary are occupied by men (e.g. factory workers). The younger you are, the shorter you get: one half of temporary workers are less than 30 while this proportion is only 15% among full-time permanent workers.

Distributions of short-term contracts' initial durations are displayed in Table 3 and exhibit large dispersion. For FTC, the median is equal to six months while interim's one is only one month (the mode is one week).

The weekly working time is displayed by Table 4 according to each type of labor contract. One shall notice that people with an open-ended contract work apparently more than temporary workers: they work $37,3$ hours a week on average *versus* $35,9$ for temporary workers (medians are respectively 35 and 37 hours). These insignificant differences vanish as soon as we consider non-executives. Thus, on French data, we don't observe any massive effect of having a fixed-term contract on the working time. Note also that these distributions are skewed to the left: for FTC workers, the last quartile is 39 hours while the first quartile is equal to the median (35 hours).

Our empirical question is to estimate the state dependence on the labor market, that is, the importance of past on current occupation. We will show therefore transition matrices from an origin to a destination state. More precisely, these matrices measure the intensity of transition; they are not gross frequencies, which brings only limited information: there is no absolute threshold above (resp. below) which a conversion rate is high (resp. low) enough to consider that FTC are a stepping-stone (trap) to stable employment... Instead, we consider here a relative indicator based on relative transition probabilities. To be more explicit, let fix some reference state 0. Then a measure for the intensity of transition from state

j to k may be given by

$$log\frac{\hat{\mathbb{P}}(y_{it}=j|y_{i,t-1}=k)/\hat{\mathbb{P}}(y_{it}=0|y_{i,t-1}=k)}{\hat{\mathbb{P}}(y_{it}=j|y_{i,t-1}=0)/\hat{\mathbb{P}}(y_{it}=0|y_{i,t-1}=0)}$$

which is the coefficient of a descriptive multinomial logit model of y_{it} on $y_{i,t-1}$ (see next section). δ_{kj} measures odds of being in state j rather than in the reference state, given that one comes from k rather than from the reference state. These coefficients can be compared within a row (odds of being in state j rather than in state j', given that one comes from k rather non from the reference state) or a column (odds of being in state j rather than in the reference state, given that one comes from k rather non from k').

In practice, thanks to our large sample, we consider 7 possible different states in the labor market. Non-participation aside, people can be unemployed, have a full-time open-ended contract, a full-time fixed-term contract, a full-time interim mission, a full-time other activity (civil servants, self-employed and subsidized jobs...) or a part-time job. Remember that it is worth putting part-time workers aside because our definition of effort will rely on the weekly working time. Concerning the definition of these states, one faces the well-known dilemma: either we aggregate different states in one category but then we miss specificities due to composition effects, or we stay at a rather detailed level and the interpretation is more tricky since the identification relies on few observations only, which is bad for the accuracy of the estimator.

Table 5 shows the matrix of intensity of transitions from one quarter to another. It exhibits a strong inertia measured by high diagonal terms, especially in the case of permanent workers. But on the whole it is not a *ceteris paribus* analysis and

this is the reason why we try to distinguish state dependence from unobserved heterogeneity in the following model.

3 Econometrics

3.1 A model of transitions on the labor market

We follow a standard approach in the job occupational choice literature inspired by McFadden (1974) class of models that postulate the existence of latent individual propensions for every different state on the labor market. The actual state is related to these unobserved variables thanks to the following relationship:

$$y_{it} = j \Longleftrightarrow y^*_{ijt} = Max_k \;\; y^*_{ikt}$$

if y^*_{ijt} denotes the latent utility of individual i for state j at time t and y_{it} the observed state taking its values in $\{0, 1, ..., J-1\}$.

As previously mentioned, a dynamic model is required for such an analysis: labor market histories are strongly time dependent. To simplify, we assume that this state dependence follows a Markov process of order 1, that is, current state depends on history through the last period only. In other words, y^*_{ijt} depends on $y_{i,t-1}$ like

$$y^*_{ijt} = \sum_{k=0}^{J-1} \delta_{kj} \mathbb{1}_{y_{i,t-1}=k} + \epsilon_{ijt}$$

For now, ϵ_{ijt} can be everything that is related to the individual: observed heterogeneity, idiosyncratic shock or permanent individual effect.

Parameters of interest measuring the state dependence include the vector δ. Indeed, the information on the frequency of the transitions from a given state to

another is given by δ_{kj}, a mobility index. δ_{jj} can be seen as the inertia parameter since it measures the persistence of the state j in the labor market.

To estimate the model easily with a parametric logit form, one also assumes that idiosyncratic shocks are extreme-value distributed. Without any explaining variable X_{jt} accounting for observed heterogeneity, one can write for instance:

$$\mathbb{P}(y_{it} = j | y_{i,t-1} = k; \delta) = \frac{e^{\delta_{kj} - \delta_{k0}}}{1 + \sum_{l \neq 0} e^{\delta_{kl} - \delta_{k0}}}$$

One can then interpret δ_{kj} as the logarithm of *odds ratios*:

$$\delta_{kj} = log \frac{\mathbb{P}(y_{it} = j | y_{i,t-1} = k) / \mathbb{P}(y_{it} = 0 | y_{i,t-1} = k)}{\mathbb{P}(y_{it} = j | y_{i,t-1} = 0) / \mathbb{P}(y_{it} = 0 | y_{i,t-1} = 0)}$$

Not all coefficients are identified; we shall normalize

$$\delta_{j0} = \delta_{0j} = 0 \quad \forall j = 0, ..., J-1$$

In practice, we choose unemployment to be the reference state, which we denote by 0. This normalization is innocuous but not meaningless since our future interpretations of δ_{kj} will all be related to this reference state. Transition coefficients have thus the following "relative" meaning: δ_{kj} measures the risk of being in state j a quarter later rather than being unemployed, given that one comes from state k rather than unemployment. The choice of unemployment as the reference state is ours: we are investigating the impact *per se* of spells of temporary work on the career path in comparison with "staying unemployed" for some part of the population.

A first naive estimator of mobility in the labor market consists in ratios using empirical counterparts of probabilities (observed frequencies), as presented in previous section (Table 5).

One wants the model to take individual heterogeneity into account by allowing people to differ in personal characteristics. It is well-known that such characteristics are very likely to influence the transition rate on the labor market. Obviously observed and unobserved heterogeneity can matter. Then there should be some permanent component accounting for all omitted variables and entering latent propensions to occupy different states in the labor market. Formally, one decomposes the error term ϵ_{ijt} into a permanent component that is individual- and state-specific α_{ij} and a pure idiosyncratic shock u_{ijt}:

$$\epsilon_{ijt} = \alpha_{ij} + u_{ijt}$$

The individual fixed-effect α_{ij} corresponds to the individual heterogeneity in propensions to occupy the state j. As usual, α_{i0} shall be normalized to 0 for identification issues.

Assuming now an extreme-value distribution for u_{ijt}, this model is a dynamic multinomial logit with fixed effects. The conditional probability that agent i is observed in state j at time t is now given by

$$\mathbb{P}(y_{it} = j | y_{i,t-1} = k; \alpha, \delta) = \frac{e^{\delta_{kj}+\alpha_{ij}}}{1 + \sum_{l \neq 0} e^{\delta_{kl}+\alpha_{il}}}$$

3.2 Identification and estimation

Magnac (*ibid.*) proposes an elegant method that provides identification and gives a natural estimator for these models. He shows that conditional on an appropriate statistics, fixed effects disappear from the conditional likelihood: thus, we don't need to estimate them and we avoid the problem of incidental parameters in panel data.

Formally,

$$y^*_{ijt} = \sum_{k=0}^{J-1} \delta_{kj} \mathbb{1}_{y_{i,t-1}=k} + \alpha_{ij} + u_{ijt}$$

where state dependence coefficients δ are still logarithms of odds ratios:

$$\delta_{kj} = log \frac{\mathbb{P}(y_{it} = j | y_{i,t-1} = k; \alpha) / \mathbb{P}(y_{it} = 0 | y_{i,t-1} = k; \alpha)}{\mathbb{P}(y_{it} = j | y_{i,t-1} = 0; \alpha) / \mathbb{P}(y_{it} = 0 | y_{i,t-1} = 0; \alpha)}$$

Magnac proves that a sufficient statistics is given by the initial state y_{i1}, the final state y_{iT} and the number of occurrences for each state that is observed between $t = 2$ and $t = T - 1$:

$$L(y_{i2}, ..., y_{i,T-1} | y_{i1}, y_{iT}, \{n_{ik} = \sum_{t=2}^{T-1} \mathbb{1}_{y_{it}=k}\}_{k=1,...,J}) = \frac{exp(\sum_{k,j,t=2,...,T-1} \delta_{kj} \mathbb{1}_{y_{i,t-1}=k} \mathbb{1}_{y_{it}=j})}{\sum_B exp(\sum_{k,j,t=2,...,T-1} \delta_{kj} \mathbb{1}_{b_{i,t-1}=k} \mathbb{1}_{b_{it}=j})}$$

with $B = \{(b_{i2}...b_{i,T-1}) / \forall k \quad \sum_{t=2}^{T-1} \mathbb{1}_{b_{it}=k} = n_{ik}\}$ being the set of labor market histories that are compatible with numbers of occurrences n_{ik}. Individual state-dependent fixed effects α_{ij} did vanish from this conditional likelihood.

The maximization of this conditional likelihood (CLE) provides a consistent and asymptotically normal estimator of the vector δ (Andersen, 1973).

The identification of the model and especially the state dependence coefficients comes from the comparison between the observed path and nontrivial permutations of it, that is, "equivalent" paths with the same numbers of occurrences between 2 and $T - 1$. In practice, at least 4 periods are required, since variation is needed conditional on the first, last period and a given number of occurrences. The "first" nontrivial case occurs with 4 periods and exactly one transition between the second and the third period. We dispose of 6 quarters. As a result, the identification relies on people whose state on the labor market varies between the

second and the fifth quarter: constant paths do not contribute to the likelihood. Therefore, excluding "stayers" from the estimation might bias inertia downwards.

Note that the model does not have any observable characteristic. However, fixed-effects account for all permanent characteristics and it is not obvious that one shall introduce time-varying covariates in this setting (see Honoré and Kyriazidou, 2003) on conditions for identification of such variables' effect).

To compare with the existing literature, one can think of a random-effect model that specifies a parametric distribution for unobserved terms α_{ij} conditional on the initial state and observable characteristics (Wooldridge, 2005). We investigate further in this direction and also consider a parametric specification (see *infra*) as a robustness check. Moreover, estimating the latter model on the full sample and on the restricted one gives some idea of the information that is lost. However, the parametric assumption is definitely an assumption that one does not want to make.

4 Results

4.1 Transitions between fixed-term and open-ended contracts

Table 6 displays the logarithm of odds ratios for different transitions, controlling for unobserved heterogeneity thanks to the fixed-effect model presented in last section. The (k, j) box is then equal to the logarithm of the transition rate between state k, rather than state 0 (unemployment), to state j, rather than state 0. The estimated matrix has a diagonal with high coefficients, which indicates persistence of states. Of course, this persistence is depending on the time interval and high

figures are not surprising since the time period is short. One can see that temporary workers are significantly more mobile than other active workers. As expected, open-ended contracts are the most protected as they exhibit a strong inertia and low transition rates. This persistence however is lower than the naive estimation of Table 5 that did not control for unobserved heterogeneity: this fact had already been noticed by Magnac (*ibid.*).

Restricting our attention to some transitions only, especially from a fixed-term contract or an interim mission to unemployment or to a permanent job, it seems that people with a fixed-term contract are $e^{1.16} = 3.2$ more likely to have an open-ended contract rather than be unemployed next period, in comparison with an unemployed. To some extent, the same is also true for temporary workers (interim) but the magnitude of the effect is lower (odds ratio is equal to 2.1). These results bring new evidence of the segmented labor market assumption and support the idea of a stepping stone to stable employment. However, the magnitude of this odds ratio is much lower than the one from a permanent job, which leads to qualify this diagnosis.

4.2 Restriction to movers' sample

At this point, one should discuss the econometric method. Remember that the estimation of the model taking unobserved heterogeneity into account relies on "movers" only. It leads us to drop out almost 83% of the sample and may limit our ability to extrapolate results to the whole population, especially if the state dependence is heterogeneous among individuals and the population is very heterogeneous itself – if "movers" differ much from "stayers" for instance.

A quick look at observable characteristics of subsamples (basically "movers" *versus* "stayers" for different labor contracts at the time of first interrogation) shows that significant differences do exist and concern especially unemployed, non-participants and permanent (all of them being most likely "stayers"). For instance, women, clerks and young people, as well as seniors, are over-represented in the sample of "movers" with open-ended contract. This is not the case with temporary workers where one observes no significant observable differences between "movers" *and* "stayers". However, the fact that naive estimates differ from fixed-effect estimates can be attributed either to unobserved heterogeneity or to the different composition of the two samples. That's why we show naive estimates in the "movers" sample only (Table 7). Surprisingly, compared with Table 5, differences are more striking not on but off the diagonal: persistence is naturally lower (but higher than the one from the model) but estimates of intensity of transitions between states are close to those obtained from the model. It is therefore difficult to disentangle what comes from the restriction to the "movers" sample and unobserved heterogeneity biases.

In case of individual-specific state dependence, our estimated coefficient is a weighted mean of individual coefficients and our model misses this effect. To tackle this issue of heterogeneous state dependence, that is, to allow for the causal effect of having a FTC spell on the future position on the labor market to be individual specific, one should estimate a random-coefficient model to allow δ to depend on i. There seems to be an interesting *caveat* here and we leave this for further research[†].

We briefly investigated how this estimated coefficient might vary among homoge-

[†]note that Beffy, Coudin and Rathelot, 2008 deal explicitly with heterogeneity of state dependence, but not with unobserved heterogeneity)

neous groups with identical given observable characteristics. For instance, we ran the model for women, men, less than 30, more than 50... only (see Tables 12 to 16). Surprisingly, we don't observe large differences between subsamples. Note however than young or old workers in interim have significantly less chances of getting an open-ended contract than people aged between 30 and 49 under the same regime.

4.3 Random-effect approach

To check for robustness, we specify another specification for unobserved heterogeneity, namely random effects, which adds a parametric assumption on the distribution of unobserved heterogeneity. It requires also to take the initial condition explicitly into account. Following Wooldridge (2005), and noting $\alpha_i = (\alpha_{i1}, ..., \alpha_{i6})'$ we specify

$$\alpha_i | y_{i1} \sim N(\alpha_0 + \sum_{k=1}^{6} \beta_k \mathbb{1}_{y_{i1}=k}, \Omega)$$

The estimation is performed by simulated maximum likelihood as the dimension of this multivariate distribution is six – the number of different states minus one, which corresponds to the normalization of the reference state. Indeed, the integral corresponding to the full likelihood – the integration of the conditional likelihood over the density of α_{ij} – has dimension six and can't be computed easily. We proceed to H draws with H being such that $\frac{H}{\sqrt{N}} \to +\infty$ to insure the SMLE to be consistent and asymptotically normal (Laroque and Salanié, 1993).

Results are given in Table 8. Estimates of state dependence are lower than naive ones but higher than those from the fixed-effect model, which makes us rather confident on the fact that the fixed-effect model distinguishes best state dependence

from unobserved heterogeneity (it is expected that naive estimation overestimates state dependence). The same diagnosis holds in regard of transitions from temporary to permanent contracts: FTC workers would have $e^{1.23} = 3.4$ more chances than unemployed to get an open-ended contract rather than being unemployed three months later. This ratio is equal to $e^{0.95} = 2.6$ for interim workers.

5 Incentives and effort: the more you work, the more you obtain?

5.1 A measure of effort

The empirical literature devoted to labor contracts and effort considers two proxies of effort, absenteeism and overtime work. One could also think of indicators for work during nights or week-ends, but they might concern specific industries or occupations.

In the French case the absenteeism rate is not so relevant since leave rights depend directly on how much employees have already been working. As a result, temporary workers have a low probability of owning such rights. Moreover, compensations for a sick leave could depend on the type of labor contract. In the case of a sick leave, the French welfare provides a daily indemnity that is less than the equivalent wage; there is a deductible for the first three days. But employees can get additional compensation, depending on collective agreements. The actual compensation rate could depend on industry, but also on experience and status. Employees with an open-ended contract are thus more likely to have a better coverage against health risks than temporary workers. Without any further information on the actual level of compensation rate for each individual, we decide

not to investigate further in this direction.

On the contrary, *working significantly more than others* is more likely to be a relevant measure of effort as a signal sent by the employee. An interesting measure of it that has been used is the fact of doing overtime work; in particular unpaid overtime work is a pure signal. *De facto* on cross-sectional data one observes that the quarterly transition rate to an open-ended contract is twice as high (14%) for FTC workers who do overtime work as it is for their colleagues doing no overtime work (7%). Surprisingly, the gap becomes non-significant when we look at unpaid overtime work. Unfortunately, overtime work has been measured at all interrogations since 2007 only, which makes the sample size shrink dramatically.

Instead, we could choose to compare the actual working time to a common legal duration, let say 35 or 39 hours a week. Yet, according to the occupation or the industry, usual working times might differ and we shall take it into account; in other words, this indicator does not measure the fact of "working more than siblings". That's why our favorite proxies for effort are relative indicators based on distributions of working times in the same industry and under the same regime. Each quarter we locate some FTC worker's working time in the distribution of working times of other FTC workers in the same industry. This gives us an idea of how much the FTC worker is working compared with others.

For instance, we define the fact of exerting an effort as being in the last quartile of the distribution. Similarly, "shirking" characterizes the fact of being in the first decile of the distribution[‡]. *A priori* this definition also accounts for absenteeism.

[‡]the first quartile is equal to the median, 35 hours (see *supra*)

Having a look at gross transitions shows that working significantly more increases slightly the probability of getting an open-ended contract the following quarter (10% – although the gap is not significant with previous 7%). Compared with overtime work, this indicator has more information since it is conditional on the industry and on characteristics of the job. Doing some overtime work might of course indicate a peak of the firm's activity. Our measure reduces some of the endogeneity since it controls for conjunctural and sectoral effects. Of course an individual heterogeneity remains: people working more may have some characteristics, e.g. motivation, that may also (positively) affect their probability of getting a permanent job.

5.2 Impact of effort *per se*

Our first question was to determine the potential impact on his future outcome, for a given temporary worker, to do extra effort. We have already explained that effort could be "working more than others". We now compare transition rates of fixed-term contract owners according to the fact that they exert some effort or not. That is, we split the FTC category in two: FTC effort and FTC non-effort. Interim is merged with the "other activity" state for obvious practical reasons. Once unobserved heterogeneity has been properly taken into account, one does not observe significant differences of getting an open-ended contract between FTC workers who exert an effort and those who do not.

Our measure is yet an imperfect proxy for effort but one has reasons to think that it is positively correlated with effort. One could worry of estimations depending too much on the effort definition; we show that this is not the case by considering alternative thresholds of the conditional distribution of weekly working times. For

instance, we use the top median and the last decile of the distribution instead of the last quartile. Results continue to hold in each of these alternatives. We also pay attention to the bottom of the distribution – "shirkers" – e.g. the first decile and we want to see whether these people are less likely to get an open-ended contract. This tends to be the case. Table 10 displays ratios of probabilities of being "promoted" according to several indicators: working more than 35 hours, 39 hours, doing overtime work or "working less than others" (a non-effort indicator). It seems that no matter which indicator is considered exerting extra effort has no significant impact on that outcome. This however does not mean that even "shirking" has no impact since we find a negative effect. It is worth saying a few words about the estimate related to overtime work: performing the estimation of the fixed-effect model on samples after 2007 only (remember that before 2006 overtime work is not trustable in the survey), we obtain that the ratio is 0.78 only for a FTC worker doing some overtime work *versus* a FTC worker doing no overtime work, that is, a negative effect. Remember however that doing overtime work doubled the gross transition rate from FTC to open-ended contracts, without taking unobserved heterogeneity into account. We insist yet on the fact that the size of the sample is too small to draw definite conclusions about it.

As a robustness check, one looks at the subpopulation of people under 30, the population that is most concerned by the temporary work phenomenon, which tends to become a mandatory step in labor market trajectories. One may think that for this reason the incentive effect is much stronger for them: many short-term jobs turn into permanent ones at the beginning of the career. Yet one does still not find any evidence of such an effect.

Conclusion

This paper revisits the analysis of labor market transitions in France. We use a large sample data set, the new French Labor Force Survey that enables us for instance to distinguish fixed-term contracts from interim missions. Moreover, such panel data allow the econometrician to take correctly unobserved heterogeneity into account. Doing so enables us to analyze perspectives that short-term contracts offer to their jobholders in terms of future positions on the labor market. Our estimations show that fixed-term contracts lead on average to a better outcome than unemployment, even though the access to a permanent job is far from being systematic. (Intensity of) transition matrices still exhibit strong inertia even after correcting from unobserved heterogeneity. An interesting extension would consist in strengthening the dynamic aspect of the model. Another topic that is worth being investigated is the heterogeneous state dependence (see for instance Beffy, Coudin and Rathelot 2008). A last puzzle is the estimation on the "movers"' sample.

Moreover, we discuss the question of incentives that is closely related to the nature of the contract. Our assumption is that if these contracts are used to provide employees incentives to exert supplementary effort, then people responding to these incentives should see their contracts converted more frequently *ceteris paribus*. Empirically this is true in a cross-sectional descriptive analysis but once unobserved heterogeneity is properly taken into account it does not hold anymore. Of course we can't go much further. First, our indicator is an imperfect measure of "working more". Second, doing overtime work is only one dimension of effort. On the whole the question we tackle here is the "probationary period" nature of a fixed-

term contract. To go further, it would be interesting to look at the timing of conversions into open-ended contracts (see Gagliarducci, *ibid.*, and Boockman and Hagen, 2005).

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Table 1: Description of the sample

Position	Number of individuals	Frequency (%)
Non-participants	27,736	22.0
Unemployed	8,553	6.8
Permanents *full time*	44,390	35.3
Civil servants *full time*	14,499	11.5
Fixed-term *full time*	2,237	1.8
Interim mission *full time*	1,548	1.2
Other temporary workers *full time*	2,510	2.0
Others *full time*	9,156	7.3
Part time	15,220	12.1
Total	125,849	100

Source: EEC 2002-2008, Sample of N=125,849 individuals interviewed 6 times aged from 18 to 60 in first interrogation

Table 2: Descriptive statistics

Contract	$Permanent$	$Fixed-term$	$Interim$
Women	0.35	0.52	0.28
	(0.002)	(0.01)	(0.01)
High-school degree	0.38	0.44	0.31
	(0.002)	(0.01)	(0.01)
15-29	0.15	0.49	0.47
	(0.002)	(0.01)	(0.01)
30-49	0.62	0.43	0.45
	(0.002)	(0.01)	(0.01)
50-60	0.23	0.08	0.08
	(0.002)	(0.01)	(0.01)
Workers	0.36	0.42	0.80
	(0.002)	(0.01)	(0.01)
Clerks	0.23	0.34	0.11
	(0.002)	(0.01)	(0.01)
Intermediates	0.26	0.17	0.08
	(0.002)	(0.01)	(0.01)
Executives	0.15	0.08	0.01
	(0.002)	(0.01)	(0.01)
Agriculture	0.02	0.06	0.00
	(0.001)	(0.005)	(0.00)
Manufacturing	0.30	0.19	0.46
	(0.002)	(0.01)	(0.01)
Construction	0.08	0.06	0.16
	(0.002)	(0.004)	(0.01)
Services	0.59	0.67	0.35
	(0.002)	(0.01)	(0.01)
Number of observations	$44,390$	$2,237$	$1,548$

Source: EEC 2002-2008. Sample of permanent, fixed-term and interim contracts, full time, first interrogation

Table 3: Initial duration (months) of fixed-term contracts – full time only

Contract	Mean	Std. dev.	Median	Mode	Number of observations
Fixed-term	6	6	5	6	2,237
Interim	3	7	1	0.25	1,548

Source: EEC 2002-2008, Sample of N=125,849 individuals interviewed 6 times aged from 18 to 60 in first interrogation

Table 4: Weekly working time – full time only

Contract	Mean	Std. dev.	p10	p25	p50	p75	p90	Nb. obs.
Permanent	37.3	9.6	28	35	37	40	48	37,979
Fixed-term (FTC)	35.9	9.3	27.5	35	35	39	45	1,699
Interim	34.5	7.8	24	35	35	39	40	1,408
Permanent without executives	36.5	9.1	28	35	36	40	45	32,119
FTC without executives	35.8	9	28	35	35	39	44	1,583
Interim without executives	34.4	7.8	24	35	35	39	40	1,392

Source: EEC 2002-2008, Sample of N=125,849 individuals interviewed 6 times aged from 18 to 60 in first interrogation (indicating the weekly working time)

Table 5: Intensity of transitions across states – Naive estimation

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	7.98	1.16	0.77	1.77	1.83	1.74
	(0.49)	(0.14)	(0.18)	(0.10)	(0.15)	(0.08)
FTC	2.54	3.86	0.87	0.83	0.69	1.03
	(0.12)	(0.09)	(0.18)	(0.14)	(0.24)	(0.10)
Interim	2.00	1.34	4.31	-0.01	0.36	0.40
	(0.14)	(0.15)	(0.09)	(0.22)	(0.31)	(0.14)
Part time	2.65	0.84	0.22	5.97	2.48	1.76
	(0.11)	(0.15)	(0.21)	(0.07)	(0.12)	(0.08)
Other	2.77	1.11	1.01	2.18	8.48	2.34
	(0.09)	(0.20)	(0.23)	(0.13)	(0.11)	(0.10)
Non-participation	1.31	0.81	0.17	1.22	2.79	4.52
	(0.10)	(0.09)	(0.13)	(0.07)	(0.09)	(0.04)

Source: EEC 2002-2008, Sample of N=125,849 individuals interviewed 6 times aged from 18 to 60, from first to second interrogation

Note: A FTC worker is $e^{2.54} = 12.7$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 6: Intensity of transitions across states – fixed-effect model

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	4.02	0.13	0.09	0.67	0.27	0.87
	(0.09)	(0.11)	(0.14)	(0.12)	(0.18)	(0.08)
FTC	1.16	1.77	-0.41	0.34	-0.10	0.48
	(0.09)	(0.06)	(0.12)	(0.10)	(0.14)	(0.07)
Interim	0.76	0.23	1.51	0.54	-0.12	0.11
	(0.12)	(0.10)	(0.07)	(0.14)	(0.21)	(0.10)
Part time	0.86	0.56	0.30	2.84	0.97	0.66
	(0.11)	(0.10)	(0.15)	(0.06)	(0.11)	(0.06)
Other	0.77	0.14	-0.06	0.49	3.28	0.73
	(0.15)	(0.14)	(0.20)	(0.11)	(0.10)	(0.08)
Non-participation	0.27	0.11	-0.06	0.27	0.77	1.60
	(0.09)	(0.07)	(0.09)	(0.06)	(0.08)	(0.04)

Source: EEC 2002-2008, Sample of N=20,996 "movers"

Note: A FTC is $e^{1.16} = 3.2$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 7: Intensity of transitions across states – Naive estimation, *"movers"* uniquement

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	6.19	0.92	0.45	1.05	0.90	1.16
	(0.13)	(0.16)	(0.21)	(0.16)	(0.27)	(0.11)
FTC	1.38	3.26	0.46	0.53	0.17	0.73
	(0.25)	(0.10)	(0.21)	(0.19)	(0.36)	(0.12)
Interim	1.27	0.94	3.74	-0.02	-0.14	0.19
	(0.29)	(0.17)	(0.11)	(0.27)	(0.46)	(0.16)
Part time	2.07	0.66	0.12	4.63	1.79	1.35
	(0.19)	(0.16)	(0.23)	(0.09)	(0.19)	(0.10)
Other	2.03	0.78	0.82	1.33	6.09	1.70
	(0.26)	(0.22)	(0.25)	(0.19)	(0.14)	(0.13)
Non-participation	1.31	0.67	-0.01	1.00	1.93	2.84
	(0.16)	(0.10)	(0.14)	(0.10)	(0.13)	(0.05)

Source: EEC 2002-2008, Sample of N=20,996 "movers"

Note: A FTC is $e^{1.38} = 4$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 8: Intensity of transitions across states – random-effect model

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	5.40	0.06	-0.02	0.93	0.52	1.09
	(0.07)	(0.09)	(0.11)	(0.08)	(0.12)	(0.06)
FTC	1.23	2.22	-0.45	0.34	-0.21	0.59
	(0.08)	(0.06)	(0.11)	(0.09)	(0.13)	(0.07)
Interim	0.95	0.24	1.90	0.57	-0.08	0.18
	(0.10)	(0.09)	(0.07)	(0.13)	(0.18)	(0.09)
Part time	1.31	0.55	0.37	3.63	0.89	0.77
	(0.08)	(0.08)	(0.13)	(0.06)	(0.09)	(0.06)
Other	0.99	0.12	-0.23	0.56	4.30	0.95
	(0.11)	(0.11)	(0.16)	(0.09)	(0.09)	(0.07)
Non-participation	0.30	0.16	-0.05	0.28	0.99	1.84
	(0.07)	(0.06)	(0.05)	(0.07)	(0.07)	(0.04)

Source: EEC 2002-2008, Full sample of N=125,849 individuals

Note: A FTC is $e^{1.23} = 3.4$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 9: Intensity of transitions across states – fixed-effect model. Impact of "working more"

	Permanent	FTC effort=0	FTC effort=1	Part time	Other	Non-participation
Permanent	4.08	0.06	-0.01	0.68	0.09	0.88
	(0.09)	(0.13)	(0.22)	(0.12)	(0.12)	(0.08)
FTC effort=0	1.20	1.76	1.57	0.29	-0.35	0.48
	(0.11)	(0.08)	(0.12)	(0.11)	(0.11)	(0.09)
FTC effort=1	1.29	1.47	1.77	0.32	-0.14	0.35
	(0.17)	(0.13)	(0.16)	(0.21)	(0.20)	(0.17)
Part time	0.90	0.55	0.39	2.85	0.47	0.65
	(0.11)	(0.11)	(0.20)	(0.06)	(0.09)	(0.06)
Other	0.74	0.19	0.30	0.28	2.13	0.32
	(0.10)	(0.10)	(0.18)	(0.09)	(0.06)	(0.06)
Non-participation	0.33	0.09	0.37	0.26	0.24	1.60
	(0.09)	(0.08)	(0.15)	(0.06)	(0.06)	(0.04)

Source: EEC 2002-2008, Sample of N=21,088 "movers"

Note: A FTC "working more" is $e^{1.29} = 3.6$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 10: From fixed-term to permanent contracts: impact of effort according to different indicators

Weekly working time $> 3^{rd}$ quartile	1.09
	(0.20)
Weekly working time $>$ 39 hours	0.95
	(0.17)
Weekly working time $>$ 35 hours	1.07
	(0.15)
Weekly working time $< 1^{st}$ decile	0.74
	(0.13)
Overtime work	0.78
	(0.30)

Source: EEC 2002-2008

Note: For a FTC, working more than 39 hours a week shifts by 0.95

the probability of getting an open-ended contract rather than being unemployed the quarter after

Table 11: Intensity of transitions across states – fixed-effect model, 18-29 only

	Permanent	FTC effort=0	FTC effort=1	Part time	Other	Non-participation
Permanent	4.04	0.04	-0.29	0.73	0.15	0.85
	(0.17)	(0.21)	(0.39)	(0.20)	(0.20)	(0.16)
FTC effort=0	1.09	1.66	1.54	0.25	-0.30	0.72
	(0.15)	(0.12)	(0.19)	(0.17)	(0.15)	(0.11)
FTC effort=1	1.30	1.37	1.36	0.02	-0.30	0.29
	(0.27)	(0.19)	(0.26)	(0.30)	(0.28)	(0.23)
Part time	0.84	0.51	0.31	2.50	0.51	0.94
	(0.18)	(0.16)	(0.29)	(0.11)	(0.13)	(0.10)
Other	0.55	0.08	-0.06	0.24	2.00	0.51
	(0.15)	(0.13)	(0.27)	(0.13)	(0.09)	(0.09)
Non-participation	0.28	0.19	0.28	0.37	0.50	1.81
	(0.14)	(0.11)	(0.21)	(0.09)	(0.08)	(0.06)

Source: EEC 2002-2008, Sample of N=7,705 "movers" aged from 18 to 29

Note: A FTC worker "working more" is $e^{1.30} = 3.6$ more likely...

Table 12: Intensity of transitions across states – fixed-effect model, 18-29 only

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	4.01	-0.02	0.15	0.73	0.32	0.85
	(0.17)	(0.19)	(0.24)	(0.20)	(0.32)	(0.16)
FTC	1.10	1.58	-0.37	0.21	-0.12	0.66
	(0.15)	(0.10)	(0.17)	(0.15)	(0.20)	(0.11)
Interim	0.57	0.16	1.47	0.32	-0.24	0.13
	(0.19)	(0.15)	(0.11)	(0.21)	(0.28)	(0.14)
Part time	0.83	0.45	0.25	2.51	1.02	0.96
	(0.18)	(0.15)	(0.20)	(0.11)	(0.16)	(0.10)
Other	0.64	-0.07	0.05	0.49	3.17	0.94
	(0.22)	(0.19)	(0.25)	(0.16)	(0.15)	(0.12)
Non-participation	0.26	0.19	0.07	0.38	1.06	1.84
	(0.14)	(0.10)	(0.12)	(0.09)	(0.11)	(0.06)

Source: EEC 2002-2008, Sample of N=7,676 "movers" aged from 18 to 29

Note: A FTC worker is $e^{1.10} = 2.7$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 13: Intensity of transitions across states – fixed-effect model, 30-49 only

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	4.06	0.24	0.18	0.64	0.23	1.02
	(0.13)	(0.17)	(0.20)	(0.18)	(0.28)	(0.12)
FTC	1.34	1.82	-0.50	0.43	-0.18	0.27
	(0.16)	(0.11)	(0.20)	(0.16)	(0.28)	(0.15)
Interim	1.04	0.24	1.57	0.63	0.14	0.20
	(0.19)	(0.17)	(0.11)	(0.23)	(0.37)	(0.18)
Part time	0.91	0.52	0.23	3.05	1.15	0.43
	(0.17)	(0.17)	(0.24)	(0.09)	(0.18)	(0.11)
Other	1.02	0.34	-0.12	0.48	3.38	0.68
	(0.25)	(0.28)	(0.36)	(0.18)	(0.16)	(0.16)
Non-participation	0.42	0.04	-0.22	0.22	0.50	1.50
	(0.13)	(0.14)	(0.17)	(0.10)	(0.14)	(0.06)

Source: EEC 2002-2008, Sample of N=8,466 "movers" aged from 30 to 49

Note: A FTC worker is $e^{1.34} = 3.8$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 14: Intensity of transitions across states – fixed-effect model, 50-60 only

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	4.30	-0.64	-0.17	0.82	0.31	0.67
	(0.28)	(0.52)	(0.58)	(0.32)	(0.44)	(0.19)
FTC	1.13	1.81	-0.02	0.13	1.47	0.78
	(0.39)	(0.24)	(0.49)	(0.43)	(0.82)	(0.31)
Interim	0.47	-0.18	1.40	1.41	-1.30	-0.63
	(0.55)	(0.48)	(0.32)	(0.64)	(1.20)	(0.58)
Part time	1.04	0.00	-0.17	2.96	0.55	0.22
	(0.34)	(0.42)	(0.78)	(0.17)	(0.35)	(0.18)
Other	1.08	0.30	-0.38	0.46	3.15	0.43
	(0.42)	(1.15)	(0.95)	(0.35)	(0.31)	(0.26)
Non-participation	0.37	0.20	0.08	0.06	0.37	1.47
	(0.26)	(0.30)	(0.49)	(0.19)	(0.25)	(0.09)

Source: EEC 2002-2008, Sample of N=3,805 "movers" aged from 50 to 60

Note: A FTC worker is $e^{1.13} = 3.1$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 15: Intensity of transitions across states – fixed-effect model, women only

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	4.14	-0.04	0.34	0.67	0.26	0.88
	(0.15)	(0.20)	(0.27)	(0.15)	(0.25)	((0.13)
FTC	1.19	1.68	-0.30	0.48	-0.35	0.45
	(0.16)	(0.10)	(0.20)	(0.13)	(0.22)	(0.11)
Interim	0.92	0.48	1.57	0.81	0.35	0.17
	(0.23)	(0.18)	(0.13)	(0.20)	(0.33)	(0.17)
Part time	1.05	0.47	0.50	2.84	0.88	0.62
	(0.14)	(0.13)	(0.20)	(0.(07)	(0.14)	(0.08)
Other	0.95	0.17	-0.08	0.49	3.14	0.72
	(0.22)	(0.21)	(0.36)	(0.14)	(0.14)	(0.12)
Non-participation	0.35	0.07	0.07	0.21	0.50	1.55
	(0.13)	(0.10)	(0.16)	(0.07)	(0.11)	(0.05)

Source: EEC 2002-2008, Sample of N=12,066 "movers"

Note: A FTC worker is $e^{1.19} = 3.3$ more likely than an unemployed to hold a permanent job three months later rather than being unemployed

Table 16: Intensity of transitions across states – fixed-effect model, men only

	Permanent	FTC	Interim	Part time	Other	Non-participation
Permanent	3.99	0.08	0.02	0.65	0.25	0.89
	(0.12)	(0.15)	(0.17)	(0.21)	(0.26)	(0.11)
FTC	1.23	1.74	-0.50	-0.16	-0.05	0.50
	(0.13)	(0.10)	(0.16)	(0.20)	(0.22)	(0.12)
Interim	0.68	0.07	1.47	0.30	-0.35	0.09
	(0.15)	(0.14)	(0.09)	(0.21)	(0.29)	(0.13)
Part time	0.58	0.59	0.14	2.88	1.05	0.78
	(0.20)	(0.18)	(0.22)	(0.12)	(0.19)	(0.13)
Other	0.67	0.02	-0.05	0.52	3.46	0.75
	(0.20)	(0.21)	(0.24)	(0.19)	(0.15)	(0.13)
Non-participation	0.30	0.18	-0.12	0.39	1.06	1.70
	(0.12)	(0.11)	(0.12)	(0.12)	(0.12)	(0.06)

Source: EEC 2002-2008, Sample of N=8,930 "movers"

Note: A FTC is $e^{1.23} = 3.4$ *more likely than an unemployed to hold a permanent job three months later rather than being unemployed*