PT Status, Type of Contract, and the Returns to Experience

Daniel Fernández-Kranz
IE Business School

Marie Paul
née Waller, Albert-Ludwigs-University Freiburg

Núria Rodríguez-Planas
Universitat Autònoma de Barcelona

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Abstract

Within the part-time (PT) literature, two results emerge. First, PT work is both a support and a trap for women's future careers, but these alternative roles apply to different groups of women. Second, a PT/full-time (FT) hourly wage differential persists (for a given level of human capital) even after controlling for time-invariant unobserved heterogeneity using a fixed-effects approach. However, such approach is limited as identification arises from those workers who switch from FT to PT work (or vice-versa), ignoring wage differences between different individuals with diverging employment histories. We begin with the same wage equation as previous studies. We follow Buchinsky, Fougère, Kramarz, and Tchernis, 2010, and explicitly model the participation/employment and PT status and type of contract mobility decisions, which, in turn, define the individual's experience in different segments of the labor market. We introduce into the wage equation a summary of the workers' entire career path. The three-equation system is estimated simultaneously using data from the Continuous Survey Working History (CSWH). We find no PT/FT wage gap for currently working PT in a permanent contract but a low return to experience in these jobs. For PT work in fixed-term contracts our results suggest a penalty for currently working PT. Working in PT jobs for several consecutive years amounts to a considerable penalty which is similar in both labor markets. Our study highlights the importance of accounting for individuals' diverging employment histories especially across segments of the labor market.

Key words: Fixed-term and permanent contracts, part-time employment, returns to experience of differential work histories, random effects models, MCMC

JEL classification: J16, J24, J31, J41, C11, C33

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I. Introduction

Academics, politicians and practitioners are becoming increasingly interested in flexible work arrangements, such as part-time (PT) work or fixed-term contracts. Supporters of such type of arrangements argue that they are helpful in reconciling market and family work, or that they work as stepping-stones from non-employment into full employment. Deterrents, in contrast, argue that this type of jobs harm the worker's career as they receive lower wages, fringe benefits, training, and chances of promotion.

Within this literature, many researchers have analyzed the hourly wage differences between PT (or fixed-term contract) workers and their full-time (FT) (or permanent contract) counterparts. Most frequently, to address unobserved heterogeneity problems, such as the fact that workers in flexible work arrangements may have different tastes and preferences about work than those with more standard types of jobs, authors use panel data and fixed-effects-'within' estimator (see Booth and Wood, 2008; and Connolly and Gregory, 2009; Fernandez Kranz and Rodriguez-Planas (*forthcoming*)). However, such approach is limited as identification arises from those workers who switch from standard employment to flexible work arrangements (or vice-versa), ignoring wage differences between different individuals with diverging employment histories. An alternative approach has been to analyze the returns to experience in PT employment (or fixed-term contracts)—see for example, Connolly and Gregory, 2009; Fouarge and Muffels, 2009; Green and Ferber, 2005; and Russo and Hassink, 2008. But then the employment history is assumed exogenous.

In this paper we consider the possible endogeneity of employment/participation, PT status, as well as, type of contract, and its potential effects on the estimated returns to both tenure and experience.[1] To address this key issue, we follow Buchinsky, Fougère, Kramarz, and Tchernis, 2010, (BFKT, hereafter) and use an econometric specification in which two

[1] As in Buchinsky *et al.*, 2010, participation is the same as employment, thus we use the two terms interchangeably throughout the paper.

key decisions individuals make each period are modeled: (1) employment in FT or PT work, non-employment (NE), and (2) type of contract (fixed-term versus permanent). In turn, these decisions influence the observed outcome of interest, namely wages. Within this model we revisit the issue regarding the magnitude of the returns to flexible work arrangements and offer new perspectives, in particular with regard to dynamics.

This model gives rise to three reduced-form decision equations: (1) a participation/FT/PT employment; (2) a contract-type decision; and (3) a Mincer's wage equation. We estimate the three equations simultaneously. In this approach, both experience (in both full- and PT work) and contract type are endogenized with regard to time-constant individual heterogeneity because they are direct outcomes of the employment and type of contract decisions, respectively. Hence, the model accounts for the potential selection biases that stem from these endogenous decisions, and thus allows one to consistently estimate the parameters associated with the wage function, including the PT and fixed-term contract wage differentials. In contrast to Buchsinky *et al.*, 2010, we have to make the assumption that idiosyncratic errors are independent across equations. In our context we need to add regressors on the current employment situation to the wage equation. Our model would not be identified if we added in addition correlation terms for the idiosyncratic error terms as our data does not provide the exclusion restrictions which would be needed in this case. To account for individuals' unobserved heterogeneity we include person-specific random effects in the estimated equations. Controlling for unobserved heterogeneity in this fashion creates important links between the individual's decisions and the outcome of interest (i.e., wage) that, if ignored, could lead to severe biases in the estimated parameters of interest.

Our approach offers a unified framework with which one can address and reexamine results that have been previously found in the literature. This is because the statistical assumptions that are adopted in this study–along with the fact that PT status, contract type and

experience are endogenized–incorporate the most important elements from many previous studies.

We use data from the Spanish Social Security records (the Continuous Survey Working History, CSWH hereafter) and estimate, first, the model for all women, and then, separately for two separate groups of women: mothers (including mothers to be) versus women who will never bear a child. This distinction is done because clearly the role of PT work as a family-work-reconciling device takes place mainly among women who care (or will care) of small children in Spain. To estimate the joint posterior distribution for the model's parameters, we adopt a Bayesian approach and employ Markov Chain Monte Carlo (MCMC) methods.

Our results suggest that conditional on the labor market history, there is no penalty of currently working PT for mothers in a permanent contract, but there is a penalty of currently working PT for childless in a permanent contract and for both groups of women of currently working PT in a fixed term contract. Experience is rewarded less in PT work than in FT work and these (low) returns are quite similar in the two segments of the labor market. But there is a remarkable exception: there is a particular strong penalty for childless women of working PT in the primary segment of the labor market. With regard to the PT literature these findings imply that it is important to disentangle the effect of currently working PT and of the employment histories. Part of the PT penalty can be explained by differential employment histories of part-time and full-timeworkers. Furthermore our results suggest that the low wages observed among fixed-term contracts are largely due to differences in experience between women working in the primary and secondary segment of the labor market. Our study also sheds light on several important factors leading to the apparent differences between our estimates and those obtained in the literature. First, our study highlights the importance of accounting for individuals' diverging employment histories especially across segments of the labor market. We demonstrate the need to explicitly control for past experience by type of

contract and PT status in the wage function to capture the overall effect of the worker's specific career path on her market wage. In particular, our results show that it is important to account very carefully for experience when attempting to estimate causal effects of different contract types. We find that the magnitude of the estimated returns changes markedly when we account for this factor, indicating that the PT employment history during the course of one's career is important for her wage trajectory. This factor may be accounted for in a methodological approach which uses cross-sectional identification in addition to "within" identification and which endogenizes different forms of experience. This can be done by explicitly modeling the employment, PT status and type of contract decisions while accounting for state dependence and duration dependence when modeling the contract type and the employment type. Second, we establish the need to account for time invariant unobserved heterogeneity prevailing in the participation and contract type and PT status decisions, as well as in the wage function. Finally, heterogeneity analysis by motherhood status reveals interesting differences.

The remainder of the paper is organized as follows. Section 2 provides a brief review of the literature, and Sections 3 presents some institutional background on flexible work arrangements in Spain. Section 4 discusses the econometric specification and estimation methodology. Sections 5 and 6 present the data and the empirical results, respectively. Concluding remarks are in Section 7.

II. Literature Review

Within the PT literature, the following two interesting results emerge. First, there are two leading patterns in transitions through PT work: the "maintenance" PT spell supporting continued labor market participations within basically a FT career; and the PT work forming part of an "exclusionary" cycle of weak labor market attachment. For the US, where the

incidence of PT work among prime-aged women is as low as in Spain but temporary contracts are practically inexistent, Blank, 1998, identifies two groups of PT female workers. For one group of women, PT work serves as temporary alternative to FT work, to which they return; while for another group, there is a constant churning in and out of employment through PT work. Similarly, for the UK, where PT work among prime-aged women is considerably more common than in Spain but fixed-term contracts are not as frequently used, Connolly and Gregory, 2010, find that PT work serves two different functions. Women whose past history predominantly involves FT work, possibly in conjunction with spells of PT work or non-employment, revert to FT work; while women whose labor market histories combine spells in PT with non-employment are unlikely to subsequently take up FT work. According to these authors, PT work is both a support and a trap for women's future careers, but these alternative roles apply to different groups.[2] Finally, for Spain, Fernández-Kranz and Rodríguez-Planas, 2010, find that two similar leading patterns in transitions through PT work coexist, with the "maintenance" spell being more common among women with permanent contract and the "exclusionary" pattern more frequent among those with fixed-term contracts.

Second, recent work finds that PT/FT hourly wage differential persists (for a given level of human capital) even after controlling for time-invariant unobserved heterogeneity using a fixed-effects approach. While the literature on the PT/FT hourly wage differential goes back to the 1970s, it is only recently that this research has taken advantage of the panel structure of newly available datasets and has addressed the problem of the omitted variable bias due to unobserved quality differences between PT and FT workers by estimating a fixed-effects-'within' estimator (see Booth and Wood, 2008; Connolly and Gregory, 2009;

[2] In contrast with Connolly and Gregory, 2010, O'Reilly and Bothfeld, 2002, find that PT work in the UK follows only an "exclusionary" pattern where it is interspersed with spells of non-employment. However, these authors followed workers over a five-year period, whereas Connolly and Gregory followed a birth cohort (born in 1958) until they were aged 42 in 2000.

Fernández-Kranz and Rodríguez-Planas (*forthcoming*)).[3] However, such approach is limited as identification arises from those workers who switch from standard employment to flexible work arrangements (or vice-versa), ignoring wage differences between different individuals with diverging employment histories. This is particularly concerning, because as discussed earlier, there seems to be two clearly diverging patterns of transitions into PT work leading to drastically different employment histories.[4]

To address this concern, in the present paper we focus on the dynamic aspect of the PT/FT hourly wage differential. More specifically we are interested in disentangling how much of the PT/FT hourly wage differential is explained by differences in PT and FT workers' employment histories versus differences in the current hourly wage differential between PT and FT workers. Moreover, our approach enables us to identify the effects of different employment histories (involving switches between employment states and the two segments of the labor market) on PT wages, and to analyze whether the PT/FT wage differential is more severe for those with a long history of working PT than for those who experienced lots of churning between PT, FT work, and fixed-term contracts. These differential effects are identified because we distinguish between observed and time-constant unobserved individual heterogeneity, on the one hand; and model the employment histories, state dependence and duration dependence in PT/FT employment, as well as, in type of contract in addition to the individual specific unobserved heterogeneity.

[3] Alternatively, Hirsch, 2005, uses multiple short panels with two observations per worker (one year apart) to estimate the effect of switching between FT and PT status on wage changes.

[4] As discussed in the introduction, an alternative approach has been to analyze the returns to experience in PT employment. But then the employment history is assumed exogenous or only "within" identification is used.

III. Spanish Institutional Background

Most of the existing literature on PT employment uses data from the Netherlands, UK, US, or Australia where PT accounts for a very high percentage of overall female employment and where the incidence of female temporary or fixed-term employment is relatively low. Table 1 shows statistics for some of these countries and illustrates that Spain is among the countries with a lower incidence of PT work combined with an extremely high incidence of fixed-term employment (OECD, 2008). These figures highlight that the role of PT employment in Spain may differ given the unique specificities of its labor market. Below, we discuss the institutional background on PT and fixed-term employment in Spain.

The two most common forms of flexible work arrangements (fixed-term contracts and PT work) have evolved quite differently in Spain over the last two decades. Both types of contracts were first regulated by law in 1984 with the objective of adding flexibility and promoting employment in a rigid labor market with stringent employment protection legislation and high levels of unemployment. While fixed-term employment soared, the growth in PT employment was modest, at most. As a result, since the early 1990s, fixed-term employment represents one third of the Spanish labor force (by far, the highest share among European countries), whereas the share of PT employment is below one tenth of the labor force (far from the EU average of 18%).

Fixed-Term Contracts

Prior to 1984, most contracts in Spain were permanent contracts. With such contracts, the costs of dismissing a worker were high (up to 45 days of wages per year worked if the worker appealed to Court and the dismissal was declared "unfair", with a limit of 24 months' wages).[5] In 1984, in a context of high unemployment and given that an across-the-board reduction of dismissal costs was politically unfeasible; the use of temporary contracts was

[5] Izquierdo and Lacuesta, 2006, and Galdón-Sánchez and Güell, 2003, estimate that between 72% and 75% of cases that arrived to court were declared "unfair" by Spanish judges.

liberalized. As such, fixed-term contracts for regular activities entailed much lower severance payments than permanent contracts (initially of 12 days per year worker, zero if the firm waited until expiration), and their termination could not be appealed to labor courts (in contrast with their permanent counterpart). However, temporary contracts could only be used up to a maximum of three consecutive years.

After the 1984 regulation change, fixed-term employment soared and, since the early 1990s, they have represented one third of the Spanish labor force. The surge of fixed-term contracts led to a dual labor market with workers with fixed-term contracts holding unstable, low protected and poorly paid jobs, while workers with permanent contracts enjoyed protection and presumably also higher wages. According to Bover and Gómez, 2004, between 1985 and 1994, over 95% of all new hires were employed through temporary contracts and the conversion rate from temporary to permanent contracts was only around 10%.

PT Work

PT work in Spain has traditionally been a second-best job. Prior to the 1980s, PT work in Spain heavily penalized workers in terms of higher social security costs and was only legal for certain types of workers considered at-risk of social exclusion, such as disabled workers, first time job seekers, workers over 40 years of age or with family responsibilities, and the long-term unemployed. The 1980 Worker's Act (*Estatuto del Trabajador)* removed the social security costs' penalty, and the 1984 reform eliminated the hiring restrictions and legalized (for the first time) the conversion from FT to PT contracts. However despite these changes, PT employment remained modest, hovering around 5% of the labor force. It is not until the mid-1990s with the economic crisis and the 1994 reform that aimed at increasing PT work as a flexible work arrangement by reducing the social rights of some PT workers (those working less than 12 hours per week or less than 48 hours per month), that PT employment rose to 8% of the labor force. Four years later, the 1994 changes were undone as the Spanish

PT regulation converged to that of the EU, but the share of PT work remained stable around 9% of the labor force--and far from the EU average.

PT work is not only less frequent among female workers in Spain than in the neighboring European countries, but it is also less stable (Buddelmeyer *et al.*, 2005; Muñoz de Bustillo Llorente *et al.*, 2008; and Fernández-Kranz and Rodríguez-Planas, 2010). Moreover, just as in the case of fixed-term contracts, there is much evidence supporting the hypothesis that those working in PT jobs in Spain are in secondary employment, as these workers are more likely to be working under a temporary contract than a permanent one, and they tend to be concentrated among certain industries—especially services, such as, retail sales, janitors, real estate, restaurants, education, and other social and personal services—, and low-skilled occupations, such as non-qualified occupations or non-professional white-collar jobs (Muñoz de Bustillo Llorente *et al.*, 2008). Finally, while PT work may be a cultural choice among Spanish women with permanent contracts—as 28.60 % of them report working PT because of family reasons (that is, taking care of their children or an elderly)—, it is quite unlikely that this is the case in the secondary market—since only 12.48 % of those working PT in fixed-term contracts report doing so to take care of their family members (1999-2009 Spanish Labor Force Survey, LFS hereafter). Because of this, our analysis will also de done for two separate groups of workers: those women who are mothers (or mothers to be) and those who will never bear a child as to identify how much of our findings are due to choosing PT because of non-market responsibilities (measured by presence of children in the household).

Wage Differences by Type of Contract or PT Status

The evidence on wage differences by type of contract or PT status has been scarce in Spain (mainly due to the lack of large databases containing individual information on wages until recently), and based on cross-sectional analysis. Given that wages are set by collective agreements and that these do not allow workers to be paid differently on type of contract, it

seems reasonable to think that employers do not discriminate against workers by type of contract. Despite this fact, several empirical studies find that permanent workers earn around 10% more, for men, and about 5% more, for women, after controlling for observed heterogeneity in personal and job-related characteristics and for selection into type of contract using an instrumental approach (Hernanz, 2002; and De la Rica, 2004).[6] Turning to the evidence on PT/FT wage differential, the evidence on wage differences between PT and FT workers in Spain has found that there is an 'unexpected' (in the light of the anecdotal evidence and job satisfaction indicators) wage *premium* to working PT (Pagán Rodríguez, 2007), or no effect (Pissarides *et al.*, 2005, and Muñoz de Bustillo Llorente *et al.*, 2008).[7] However, failure of correcting for unobserved heterogeneity and measurement problems raise caution before taking these estimates at face value—as acknowledged by Pissarides *et al.*, 2005. Recently, using the same data set than the one used in this paper, Fernández-Kranz and Rodríguez-Planas (*forthcoming*) have found that wage differentials between FT and PT workers persist (for a given level of human capital) even after controlling for individual- and firm-level heterogeneity (through fixed-effects) and measurement error in contractual hours. In addition, these authors find that the PT/FT wage differential is more pronounced among fixed-term contract workers.

[6] To identify participation into fixed-term versus permanent contract in Spain, Hernanz, 2002, uses gender, age, level of education, industry, public or private employer, firm size and region and working day duration (and occupation on the case of the estimation of the SES sample). De la Rica, 2007, uses age, tenure and education, controls for occupation (at one-digit) and the rate of fixed-term contracts by autonomous community. De la Rica, 2007, does not find evidence of selection into type of contract for females (while there is selection for males). Hernanz's estimates are not presented separately by sex, therefore we are unable to know whether her evidence of selection in the whole sample would hold when the analysis focuses on women.

[7] To identify participation into PT work in Spain, Pagán Rodríguez, 2007, uses age, level of education, marital status, number of children 5 years old or younger, number of children between 6 and 12 years old, region and household income. He finds evidence of sample selection among women working PT (but not among those working FT). Pissarides *et al.*, 2005, use family composition variables to identify participation into PT employment.

IV. Methodology

The Econometric Approach

We use a dynamic trivariate random effects panel data model which explicitly models transitions in between different contract-types and in between different employment states. The model consists of a wage equation and two selection equations: a contract-type equation and an employment-status equation. The wage equation is a linear equation of the log hourly wage implementing a flexible specification for the work history distinguishing FT employment, PT employment, and non-employment as well as permanent and fixed-term contracts. The contract type-equation is a dynamic probit equation modeling the decision between a permanent and a fixed-term contract for each period. The employment equation is a dynamic ordered probit equation modeling the decision between NE, PT, and FT. In this approach experience in FT and PT employment and in both contract types are endogenized with regard to time-constant individual heterogeneity, because they are the outcomes of the simultaneously estimated selection equations. Thus, the model accounts for potential selection biases that may evolve from the employment and contract decisions in current and past periods and allows estimating the returns to different employment patterns. Individual specific random effects in all three equations and their potential correlations across equations serve to capture time-constant unobserved heterogeneity and selection based on it. The idiosyncratic random components are assumed to be independent across correlations. This approach is methodologically similar to the approach BFKZ use to study the returns to seniority in the U.S. With regard to the idiosyncratic error terms our model is more restricted than their model as we assume independence of the three idiosyncratic error terms. This is a necessary assumption because we need to include a regressor on the current PT status into the wage equation. In this case we could only model dependences between the idiosyncratic errors if we had instruments involving variation over time and over individuals for both

selection equations which we do not have in our data. With regard to the employment equation we extend the model of BFKZ by introducing PT as a third employment status and then using an ordered probit model for the employment equation.

The Three-Equation Model

The wage equation is a random effects model of the natural logarithm of the deflated hourly wage of individual *i* in quarter *t* on the current PT status, the current contract status (FIX), the work history (H^W) and control variables:

$$(1) \qquad \ln W_{it} = \beta^W_0 + \beta^W_1 PT_{it} + \beta^W_2 FIX_{it} + \beta^W_3 H^W_{it} + \beta^W_4 x^W_{it} + \alpha^W_i + \varepsilon^W_{it}$$

The way the work history is modeled builds on the specification suggested by Light and Ureta, 1995.[8] It is summarized by a vector H^W which collects the employment status and the contract status each quarter from the year before the current period as far back as the sample allows in a very flexible way. This is done by defining an array of dummies for non-employment (NElag1 to NElag40), for PT employment (PTlag1 to Ptlag40) and for fixed-term contract status (FIXlag1 to FIXlag40). Such a dummy can equal zero for two reasons, either if the women is not in that state in the current year (thus if both the NE and the PT dummy equal zero the women would be in FT) or if the dummy refers to a year which is before the individual is observed in the sample. To distinguish these two reasons, an array of dummies to capture the time since the individual entered the sample is added. Some of the dummies are then added up to capture several quarters in one variable. We allow for some interactions between contract-type and employment-status. Control variables x^W include information on age, education, the age of the youngest child, unemployment rate and GDP growth in the region and well as the year and the season. For the exact specification of the model see the table with the results (table 4).

[8] Light and Ureta, 1995, use information on the share of weeks worked in each year in their model, but they do not distinguish between PT and FT work and by contract-types.

The contract equation is a dynamic random effects probit model of a fixed-term contract dummy on a PT dummy, the work history H^C and control variables:

(2) $FIX_{it} = \mathbf{I}(FIX_{it}^* > 0)$

$$FIX_{it}^* = \beta^C_0 + \beta^C_1 PT_{it} + \beta^C_2 H^C_{it} + \beta^C_3 x^C_{it} + \alpha^C_i + \varepsilon^C_{it}$$

The vector H^C and the vector x^C model the work history and the control variables following the same concept as in the wage equation, but specification search leads to a little different specification than in the wage equation. As H^C includes the first lag of the contract-type variable and also earlier lags, the contract-type equation allows for state dependence and duration dependence. Note that the work-history not only includes contract-history but also lagged PT and NE dummies. The wage equation and the contract-type equation are only estimated if the woman is employed in the respective period.

The employment equation is a dynamic random effects ordered probit model of the employment variable E (which equals 2 for FT, 1 for PT and 0 for NE, E* is the latent variable) on the work history H^E and control variables:

(2) $$E_{it}^* = \beta^E_0 + \beta^E_1 H^E_{it} + \beta^E_2 x^E_{it} + \alpha^E_i + \varepsilon^E_{it}$$

Work history and control variables are modeled following the same concept as in the other equations. Again work history includes the employment-status history as well as the contract-status history and state dependence and duration dependence are again allowed for. The initial conditions are modeled as suggested by Wooldridge, 2005; thus the values of the left-hand side variables for the initial period are added to the equations in each period.

All three equations include an individual unobserved effect α^W_i, α^C_i, α^E_i and these are assumed to follow a joint normal distribution $N(0,\Sigma)$ which allows for correlation among them. Thus the current employment status, the current contract-type, experience in FT and PT work as well as experience in a specific contract type are endogenized with regard to time-constant unobserved characteristics. The idiosyncratic errors are assumed to be independently normal distributed: $\varepsilon^W_{it} \sim N(0, \sigma^2)$, $\varepsilon^C_{it} \sim N(0,1)$, $\varepsilon^E_{it} \sim N(0,1)$.

Identification

Identification of this model depends on functional form restrictions and time-varying covariates (Hyslop, 1999). Identification could be supported by exclusion restrictions but in a dynamic context it is particularly difficult to find suitable exclusion restrictions because the variable or its effect has to change with the individual and with the time period (see for example Heckman and Navarro, 2007). Note that in our model identification comes from both the cross-sectional and the time-series dimensions. In the present application the time-series dimension supports identification as many women switch between states during the observation period and the cross-sectional dimension makes use of differences between different women who follow differential employment patterns over their working life. The latter is also an important reason why a fixed-effects estimator quickly reaches its limits when estimating the wage effect of work histories. Fixed-effects approaches only use the time-series dimension for identification ("within-identification"). Identification would solely rely on the wage changes women experience when they switch status. To obtain precise estimates of the wage effect of differential long-term histories may not be possible, because women experience only one employment history in their life, thus a fixed-effects estimator can only attempt to identify history effects from comparing wages at different stages of the history the women.[9] For these reasons we use a model which additionally makes use of cross-section variation (i.e. comparing women with different histories).

Estimation

To estimate the model, we use Bayesian Markov Chain Monte Carlo (MCMC) techniques. This approach avoids simulating integrals and allows a numerically robust estimation of the flexible model specification. The goal of this technique is to obtain a sample from the posterior distribution of the model parameters. From a classical perspective, the mean of the posterior distribution converges to the point estimator from maximum likelihood estimation

[9] As unobserved heterogeneity is important in this context, simply using OLS is not an alternative.

and the variance of the posterior distribution converges to the asymptotic variance of the point estimator in maximum likelihood estimation. Thus, the mean of the draws may be interpreted as the coefficient and the standard deviation as standard errors.[10] Conjugate but very diffuse priors are used. We obtain a sample of the posterior distribution of our model parameters by running 50,000 iterations of a Gibbs sampling algorithm. We monitor convergence by comparing the means at different stages of the chains. The first 5,000 iterations are discarded (the burn-in phase). We implemented the Gibbs sampler in Stata.

The analysis is first done for the whole sample, and then separately for whether women become mothers at some point in time or are childless women. The heterogeneity analysis by motherhood status responds to the fact that many mothers chose PT work as a means of reconciling family life and work, and thus, we want to disentangle how much of this story is driving our results.

V. Data and Descriptive Statistics

We use data from the 2006 wave of the Continuous Sample of Working Histories (hereafter CSWH), which is a 4% non-stratified random sample of the population registered with the Social Security Administration in 2006.[11] The CSWH consists of nearly 1.1 million individuals and provides the complete labor market history of the selected individuals back to 1967.[12] It provides information on: (1) socio-demographic characteristics of the worker (such as, sex, education, nationality, province of residence, number of children in the household

[10] See Chib, 2003, for a survey of MCMC methods for Panel data and Train, 2003, for an overview over important properties of MCMC estimators. Recent applications in labor economics are BFKT, Fitzenberger *et. al.*, 2010, Horny *et al.*, 2009, and Troske and Voicu, 2010.

[11] This includes any person that has either contributed to the Social Security or has received a pension or unemployment benefits from the Social Security during 2006. These data are extracted from administrative records from the Social Security, the Municipal Registry of Inhabitants, and the Spanish Internal Revenue Service. For a description of the CSWH and the sampling strategy, see Argimón and González, 2006.

[12] We have information on the dates the employment spell started and ended, but for those not working we do not know whether this is because they are studying, they are unemployed or another reason (the unemployed appear as such in the dataset only as long as they receive unemployment benefits from the Social Security but not if they are not eligible for such benefits). Therefore, we record spells of non-work as the time the person is not employed.

and their date of birth); (2) worker's job information (such as, the type of contract—fixed-term versus permanent contract—, the PT status, the occupation, and the dates the employment spell started and ended, and the monthly earnings); (3) employer's information (such as, industry—defined at the three-digits Spanish classification code or NACE—, public versus private sector—, the number of workers in the firm, and the location—at the province level). Although not reported in the CSWH, other variables such as working experience (in FT and PT work) and tenure can be easily calculated. In addition, information on the individual's education level, and the number and date of birth of children living in the household at the time of the interview (including own natural, adopted, step and foster children) is also available in the Spanish Municipal Registry of Inhabitants, which is matched at the person level with the Social Security records.

Because the CSWH does not have reliable information on type of contract prior to 1996, our analysis focuses on the years 1996 to 2006. However, we use information back to 1985 to calculate variables such as workers' experience and tenure. Therefore, although our PT/FT hourly wage differential analysis focuses on work histories from 1996 to 2006, we use information on the workers' employment experience (in FT and PT work) and tenure back to 1985.

Following Connolly and Gregory, 2009, and Fernández-Kranz and Rodríguez-Planas (*forthcoming*), we restrict our sample to women whose full labor market history to date can be observed. We focus our analysis on wage and salary workers, that is, we exclude from the analysis self-employed individuals.[13] We confine our selection to birth cohorts between 1961 and 1978. In addition, we restrict women in our sample to be aged between 24 and 45 years. The reason for dropping women younger than 24 years old is that we want to eliminate PT work by students. In addition, we confine our analysis to women living in households of five

[13] If the worker held more than one job, the analysis focuses on her main job, defined as the job in which the worker has a permanent contract—if she has one—, and in the case of multiple jobs with the same type of contract, the one for which the individual worked the largest number of days in a given year.

or fewer members (96.5% of the sample). The reason for restricting our attention to women 45 and younger living in households of five or fewer members is that we want to have accurate information on the number and age of children, which is unavailable in the CSWH but can be obtained from the information about the household composition as reported by the Spanish Municipal Registry of Inhabitants.[14] Finally, because we want to confine the analysis to women with a strong attachment to the labor force, we further restrict our sample to women who record at least three years in FT employment and who are observed for at least four quarters in the sample of analysis.[15]

This sample selection results in an unbalanced panel of 427,254 observations on 15,138 women, of which 2,157 (14.25%) are observed working PT at some point in time between 1997 and 2006. Following most of the European literature, we classify a worker working PT if she works 30 hours or less each week, and FT if she works 31 or more hours each week. We observe on average 34 quarters in our sample of analysis. Among the sample under study, we find that those working FT with permanent contracts represent 88% of the sample. In addition, the percentage of women working in PT with permanent contracts amounts to 3.4%. Furthermore, 7.7% of observations in our sample represent FT work in fixed-term contracts and 1.3% PT work in fixed-term contract (see first row of table 2). The remaining rows in table 2 indicate how many years of experience in the four states (FT Permanent, PT Permanent, FT Fixed-term, PT Fixed-term) the individuals have on average in the current period. Note that the average number of years already observed in the estimation sample in a given period is 5.43.[16] The data are divided in four groups, classified

[14] Although we know who lives in the household and their age, we do not know their relationship with the respondent in the CSWH. According to Lacuesta and Fernandez-Kranz, 2009, the information on family composition is reliable (relative to Census data) for the sub-population of women under 45 years old and for those living in small households. However, for older women and for women living in large households, the data becomes noisier as it is unclear whether the younger person in the household is a descendent or just a roommate.

[15] Furthermore we do not use the first period for our estimation but just for the construction of the state dependence variable.

[16] For a compact presentation, we use years in our tables, but we obtain the results based on a panel data set in quarters.

by PT status and type of contract.[17] The second row of table 2 shows that there are strong differences in the observed years of experience in between the different groups. Those who currently work FT under a permanent contract have on average 5 years of experience (observed in the estimation sample) in FT in a permanent contract, while those working FT under a fixed-term contract only have 1.7 years of experience in FT under a permanent contract. Those currently working PT under a permanent contract have on average 4.1 years of experience in FT under a permanent contract and those currently working PT in a fixed-term contract have on average 2.6 years of experience in that state. Also experience in the other three states differs a lot by the current employment status (rows three to five). These descriptives suggest that especially women working under a fixed-term contract have a very different labor market history than those working under permanent contracts, not only with regard to past contract states but also with regard the past PT states.

Table 3 presents descriptive statistics of key variables, again by PT status and by contract type (fixed-term versus permanent). When comparing the variables for women working in PT versus FT jobs, Table 3 shows that PT workers have lower (raw) hourly wages than FT workers given their the contract type. In addition, those working FT under a permanent contract have a higher hourly wage than those working FT under a fixed-term contract and those working PT under a permanent contract have higher wages than those working PT under a fixed-term contract.[18] However, this cannot be used as a reliable estimate of the pay penalty that a given woman would suffer if she changed from FT to PT status or if she changed contract status because women working PT (under a particular contract) are very different from those working FT (under a particular contract), as found in the subsequent rows of this table (and as we have already seen with regard to experience).

[17] Although one individual can appear under different categories in different waves of the panel, it should be noted that these four categories are mutually exclusive.

[18] Our measure of pay is hourly earnings, calculated as annual earnings excluding overtime divided by total contractual hours, deflated by the 2006 price deflator.

For instance, we observe that PT workers are less-educated, older and more likely to have children of all ages than FT workers. Overall the observed differences for PT versus FT workers hold across the two types of contract. Comparing those women who work FT under a fixed-term contract with those working FT under a permanent contract, table 3 indicates that those women working under a fixed-term contract are more likely to have children, they are younger and they are much more likely to hold a college or university degree. A comparison of those working under a fixed-term versus those working under a permanent contract within the PT workers reveals that those women working under a fixed-term contract are again more likely to have a child, but they tend to have lower education.

Compared to other datasets, our data has several advantages. First, the CSWH is a very large sample, which is important because PT work and switching from FT to PT (and vice-versa) is a relatively infrequent event, and more so when we focus the analysis in women strongly attached to the labor market. Second, the CSWH provides the complete labor market history for those women registered in the Social Security Administration in 2006, for up to 21 years. Although the restriction that information on type of contract is available reduces the length of women's earnings trajectories to up to ten years, it is still a non-negligible length of time. Third, it contains reliable information on monthly earnings, tenure, experience in FT and PT work, and change of employer, as the information comes directly from the payroll records. Measurement error due to recall bias or self-reporting for these key variables is minimized with this data set. Similarly, non-response is not an issue. Fourth, the dataset has rich information on individual characteristics, including education, age, ethnicity, marital status, and number and age of children in the household.

One of the short-comings of the CSWH is that it contains only information on individuals working in the formal sector. While Izquierdo *et al.*, 2010 find that, compared to the LFS, this dataset offers an accurate picture of the formal sector in Spain, Ramos Muñoz, 2007, find that some differences exists for youth, females and foreigners because the CSWH

does not account for the informal sector. For this to be a problem in our study, we would need to argue that the incidence of the informal sector is greater among PT workers than FT workers, which seems plausible. Unfortunately, there is little we can do about this. That said, it is important to keep in mind that the bulk of our analysis focuses on the time period 1996 through 2006, which was mainly a period of economic expansion, reducing the relevance of the informal sector. Most importantly, our analysis focuses on women strongly attached to the labor force (and these women tend to work in the formal sector of the economy).

Another short-coming is that we lack information on non-labor income data. Finally, another important short-coming of these data is that, instead of reporting actual hours worked, they collect information on contractual hours, which are likely to be consistently underreporting actual hours worked for PT workers relative to FT workers, leading to a differential measurement error in hours by PT status. An explanation for this is that employers have an incentive to underreport contractual hours to reduce their labor costs. Given that PT workers tend to be in more vulnerable situations than FT workers (Belous, 1989; Bardasi and Gornich, 2000; Connolly and Gregory 2008 and 2009; Manning and Petrongolo, 2008), and given the higher dispersion of hours worked among PT workers compared to FT workers in Spain (Muñoz de Bustillo Llorente *et al*., 2008), underreporting of contractual hours, albeit unlawful, seems to be a easier and more common practice for PT contracts than FT ones. Following Fernández-Kranz and Rodríguez-Planas (*forthcoming*) we use imputed effective hours (as opposed to contractual hours) to calculate the hourly wage using the same methodology.[19]

[19] To impute hours worked, we proceeded in the following manner. First, using the Spanish Time Use Survey, we regressed hours worked against contractual hours, a PT dummy, age, education, two-digit industry dummies and occupation dummies (the estimates are shown in the Appendix Table A.1). Then, using these estimated coefficients, we calculated fitted hours worked for the women in our CSWH sample.

VI. Results

We will begin by discussing our results on the wage effect of currently working in a certain employment state and then discuss the marginal and cumulative effects of experience and work histories in different employment states. In the second part of this section we turn to a discussion on how our results evolve by interpreting the variance parameters of our model and providing some sensitivity analysis. Tables 4 and 5 provide means and standard deviations (SD) of the posterior distribution of the model parameters that can be interpreted as coefficients and standard errors from a Maximum Likelihood estimation. But as the wage equation involves many dummy variables and several interactions, it is difficult to interpret the coefficients directly and we thus mainly refer to the cumulative and marginal effects of different employment states provided in table 6 to 11.

Wage Penalty to Current Employment Status

The first column of table 6 gives our estimates of the ceteris paribus wage effect of the current employment situation given the employment history. Results indicate that, on average, there is no *current* penalty for working PT under a permanent contract (conditional on the employment history). However, we observe a significant and large current hourly wage reduction (of up to 5%) for working PT under a fixed-term contract. These results corroborate those found by Fernández-Kranz and Rodríguez-Planas (*forthcoming*), who use a fixed-effect approach to estimate the penalties on current wages, and who find that the PT wage penalty is aggravated in the secondary labor market. However, we find that there is heterogeneity in these results (see first column in table 7 and 8). Indeed, we find that the PT wage gap is a little stronger among women with fixed-term contracts who eventually become mothers; whereas a PT/FT hourly wage differential among permanent-contract workers is mainly driven by the sample of childless women. The lack of a PT/FT hourly wage differential among mothers with permanent contracts is likely due to the fact that these

women are protected by law. In 1999, the Spanish Government implemented a law allowing all parents with children under seven to reduce their work-week schedule (and reducing their weekly salary proportionately). Most importantly, this law protected those workers with the reduced schedule against any layoff. While this law was binding in the primary labor market, it was not in the secondary as employers only had to wait for the contract to expire to let them go. Fernández-Kranz and Rodríguez-Planas, 2011a and 2011b, find evidence corroborating this asymmetry of the effectiveness of the law.

Wage Returns of Employment and Contract-type History

The parameter estimates of the wage equation with regard to employment status and contract-type in the past (see table 4) answer the question of how much it harms to have worked in a particular situation in the past (irrespective of the current work situation). For simplicity of exposure, we add tables 6 to 8 which present the estimated marginal returns of experience in a particular labor force state in a given period (relative to the baseline state, which is working FT with a permanent contract). Tables 9 to 11, in contrast, give the cumulative effects of working in a certain state in the current period plus in the year(s) before, for example in column three, the two years before. Thus the cumulative effects are summing up the marginal effects. Technically, the cumulative effects and their standard errors are obtained by calculating a draw of the posterior distribution of the cumulative effect of interest for each iteration of the Gibbs sampler and then describe the resulting draws by using their mean and their standard deviation. For example, the cumulative returns to working two years PT in a fixed-term contract is calculated by adding up the draws of the model parameter for working PT and for working in a fixed-term PT contract in the current quarter and in the last eight quarters taking interactions into account.

Our cumulative results in the second row of Table 9 indicate that there is a strong wage penalty of having worked part time in the current period and for the last year which increases with each additional year of PT work history. For instance, we find that women

who have been working PT in a permanent contract for one year *in addition to the current quarter* earn an hourly wage which is 12.2% lower than the wage of women who are comparable in observed and time-constant unobserved characteristics as well the rest of their employment history, but have not interrupted their FT career. If a woman has already been working PT for six consecutive years the wage difference to her FT counterpart has increased to -14.9%. Remember that, on average, there is no penalty of current PT work under permanent contracts, thus our results suggest that the estimated cumulative penalty is caused by experience in PT jobs being less rewarded than in FT jobs. Thus, in the primary labor market, there is no immediate penalty of switching to PT, but wages do not evolve in the way they do in FT jobs.

Looking at the effect of heterogeneity between mothers and childless women, Tables 10 and 11 show that the cumulative penalty is lower for mothers (10.7% after one year of PT history and 12.2% after six years), while it is higher for childless (16.7% after one year of PT history and 24% after six years). Note that only 3 percentage points of this difference can be explained by the penalty we have found for childless mothers currently working PT, thus also the returns to experience in PT under permanent contracts are lower for childless than for mothers. Thus, childless women are, overall, strongly penalized by working PT instead of FT in the primary labor market. In contrast, mothers are penalized less. Their wages do suffer from PT work, but mainly due to having worked PT in the last year.

Tables 6 to 8 use the same parameter estimates but depict them as marginal effects. Looking at the marginal effects makes it even more evident that there is a strong negative effect of experience in PT in a permanent contract. The marginal effects also indicate that, in many cases, it is the employment status in the last year (i.e. one to four quarters ago) that has a strong wage effect, thus wages seem to be influenced a lot by what a woman has done in the last year. In contrast, experience in earlier years has relatively small marginal effect. For mothers the marginal effects of remote years are even insignificant. Mothers in the primary

labor market are the group which is most likely to use PT work to reconcile labor market participation and family work. Our results suggest that their wages suffer in particular from having worked PT in the last year, while it does not matter so much what they did before. A possible explanation for the strong influence of experience in the last year in the primary labor market may be that in this labor market it is not possible to reduce hourly wages if a switch to PT occurs but that employers compensate for this in the following periods. The women do not experience an immediate wage reduction but stop moving up the wage scale for some time after having switched to PT.

Turning to the cumulative effects of PT in the secondary segment of the labor market (shown in Table 9), we get a similar picture with regard to the wage effects of working PT for several consecutive years as we observed among PT workers in the primary labor market. Our methodological approach allows us to compare the wages of women working PT in a fixed-term contract with those working FT in the primary or the secondary segments of the labor market, respectively, as we have endogenized the contract type decision. Note that typically the literature estimates heterogeneity in PT/FT wage differences given the average histories that come along with currently working in a particular segment of the labor market (see for example Fernández-Kranz and Rodríguez-Planas (*forthcoming*)). Our results suggest that a woman who has been working in PT in a fixed-term contract for one year in addition to the current quarter earns 13.1% less than a woman who has been working FT in a permanent contract but is comparable in observed and time-constant unobserved characteristics and has experienced the same employment history in earlier years. Having been working in PT fixed-term already for six year, the wage penalty is 16.3%. Thus, working PT under a fixed-term contract leads to a little higher cumulative wage penalty as under a permanent contract. But note that this cumulative effect is more negative due to the wage penalty of currently working PT under a fixed-term contract. It is this effect which drives the difference in between contract types in the cumulative penalties. PT experience alone is rewarded in a similar way

than on the primary segment of the labor market. In fact, table 6 with the marginal effects suggests that the marginal effects of experience in PT are similar in the primary and the secondary segments of the labor market, though there is some indication that on the primary segment of thelabor market the penalty of having worked PT in the last year is particularly strong.

Heterogeneity effects in Tables 10 and 11 reveal that the effect of having worked the current quarter plus one year PT in a fixed-term contract is -14.1% for mothers and -10.2% for childless; while the cumulative effect of having been working the current quarter plus six years PT in a fixed-term contract is -15.7% for mothers and -18.1% for childless (again compared to not having interrupted a FT career on the primary segment of the labor market). Thus it seems that for childless, it is in particular staying away from FT for a very long period that is additionally penalized.

Note that until here, we have always compared the effects to the reference situation of working FT in a permanent contract, as this is still the typical employment situation in Spain. As mentioned above, this comparison is different from the literature, in which the heterogeneity of the PT penalty by segment is studied without giving a causal interpretation to the effect of working PT in the primary as opposed to the secondary segment of the labor market. In other words, in the literature the effect of working PT in a certain contract type is the effect of working PT in this contract type given the decision to work in this contract type and the employment history that comes along with this contract type. Because we control for the employment history and endogenize it, subject to our model assumptions, we can give the contract-type effect a causal interpretation and thus compare all wage effects to FT work in the primary segment of the labor market. In addition, it is possible to look at the PT /FT wage differential within the secondary segment of the labor market. This differential is even a little larger than the differential discussed before, because we find a very small cumulative positive effect of working FT in a fixed-term contract as opposed to working FT in a permanent

contract (second last rows of table 9). The table with the marginal effects (table 6) suggests that experience in FT work in the secondary labor market is a little less rewarded then in the primary segment of the labor market, but there is a small wage premium of currently working in the secondary segment of the labor market. Note that this premium is conditional on the complete history and as usually those women with a less beneficial history select into the secondary segment of the labor market, this effect does in no way contradict that wages are lower in the secondary segment of the labor market. The fact that our results imply a less detrimental picture of the second labor market than most of the literature suggests, indicates that an important reason for lower wages in the secondary segment of the labor market is that those women working in the fixed-term contracts have a different labor market history and thus have collected less experience or a large part of their experience in jobs in which the returns to experience in them will be low. We will come back to this point in the section with the sensitivity analysis

Another advantage of our methodological approach is that we can not only compare the returns to experience in PT and FT, but also compare them to the wage effects of intermittent behavior. The last rows of table 6 to 11 show the penalty of non-employment in the past. A history of non-employment is much more penalized than working PT. Having interrupted employment for the last year goes along with a penalty of 14.5% (relative to permanent FT work) and after an interruption of six years the penalty is 35.3%. These results put the PT penalties into perspective by showing that while on the one hand experience in PT is less rewarded than in FT, on the other hand the effects are still much less detrimental than if a woman does interrupt employment.

To give a short summary of our discussion until here, we find that conditional on the labor market history, there is no penalty of currently working PT for mothers in a permanent contract, but there is a penalty of currently working PT for childless in a permanent contract and for both groups of women of currently working PT in a fixed term contract. Experience is

rewarded less in PT work than in FT work and these (low) returns are quite similar in the two segments of the labor market. But there is a remarkable exception: there is a particular strong penalty for childless women of working PT in the primary segment of the labor market. With regard to the PT literature these findings imply that it is important to disentangle the effect of currently working PT and of the employment histories. Part of the PT penalty can be explained by differential employment histories of PT and FE workers, a finding which is also supported by Hirsch, 2005, for the US and Paul, 2011, for Germany. With regard to the literature on fixed-term contracts, our results suggest that an important reason for lower wages in the secondary segment of the labor market is those women working in the fixed-term contracts having a different labor market history.

Parameter Estimate of the Employment and Contract-type Equations

Table 4 provides also means and standard deviations of the posterior distribution of the parameters that are not in the focus of our paper but some of them are worthwhile to look at. The direction of the remaining coefficients of the wage and the two reduced form equations are as expected and most of them are significant. In these two discrete choice equations we allow for state dependence which turns out to be very strong (see coefficients on the first lag of the outcome variables in table 4). Also the other coefficients relating to the employment history, i.e. the contract type dummies and the PT dummies referring to earlier periods are for the most part highly significant. This indicates that with regard to selection into fixed-term contract not only the contract history is very important but also the history with regard to FT, PT and NE decision in the past. Furthermore, for selection into PT, the history with regard to FT, PT and NE is important, but also the contract history.

Variance and Covariance Parameters

Table 5 provides means and standard deviations (SD) of the posterior distribution of the variance parameters (that can again be interpreted like coefficients and standard errors from Maximum Likelihood estimation). The estimated variance parameters suggest that an

important part of the variance is on the individual level in all three equations, indicating that time-constant unobserved heterogeneity is important. For the wage equation a very high part (83%) of the variance is on the individual level. This ratio is somewhat lower for the contract equation (23%) and the employment equation (29%). Moreover, there is a positive correlation between the unobserved individual effect of the wage equation and the employment equation. Thus, with regard to time-constant unobservables, those women who have unobservables leading to higher wages also tend to have unobservables leading to a higher employment propensity. The correlation of the unobserved individual effect of the wage equation and the contract equation is negative, so those unobserved individual characteristics that lead to a higher propensity to have a fixed-term contract tend to go along with those characteristics that lead to lower wages. Also the individual correlation between the employment equation and the contract equation is negative indicating that those time-constant unobserved individual characteristics that lead to a higher propensity to have a fixed-term contract tend to go along with those characteristics that lead to a lower employment propensity. All variance and covariance parameter are significant. Given the size and significance of the covariance parameters, the estimation results confirm that the heterogeneity on the individual level should not be neglected by for example estimating the wage equation separately or even using cross-sectional data.

The results from the separate analysis by whether the woman eventually becomes a mother or not are very similar to the aggregate ones confirming the relevance of unobserved heterogeneity. Perhaps worth highlighting is the magnitude of the individual correlation between the employment equation and the contract equation for childless women, suggesting that especially among this group those time-constant unobserved individual characteristics that lead to a higher propensity to have a fixed-term contract tend to go along with those characteristics that lead to a lower employment propensity.

Sensitivity Analysis

The results of the present study cannot be compared one-to-one to the previous literature on the PT penalty (mainly to Fernández-Kranz and Rodríguez-Planas, forthcoming, (FR hereafter), because this literature estimates different effects. First, the estimator of FR is a population average effect for the PT penalty of the women used by the estimator. Therefore, it would depend on the average elapsed duration in PT (in the sample the estimator uses for identification) which of our estimators of the cumulative effects the FR estimator had to be compared to. Second, if history has an effect, this effect is likely to be captured in the fixed effect in FR. With regard to this argument the FR estimate should be most comparable to our effect for the current PT penalty. Third, as discussed above, FR estimate the PT penalty separately by contract type, thus their estimator gives the effect of working PT in a particular segment of the labor market, given that the women works in this segment and given the employment history that goes along with working in this segment.

Despite those restrictions in comparing our findings to the literature, it is evident that our findings, as well as those of FR, suggest that there are penalties of currently working PT, but not for all groups in the labor market. Also the comparison to the fixed-term contract literature is limited, as on the one hand, we assume irrelevance of potential selection based on the idiosyncratic part of the error term, but on the other hand, we account very carefully for differential labor market histories by controlling for and endogenizing contract-type as well as experience in PT and FT work and potential interactions of those. In addition, our results only refer to women highly attached to the labor market, as described in section V. Nevertheless, it seems that our results indicate a somewhat less negative picture of wages in the secondary segment of the labor market than an important part of the literature suggests.

Remember that we use the three-equation RE model, because we want to flexibly account for the employment history in our model and therefore need to use cross-sectional identification (along with time-series identification which is necessary if unobserved heterogeneity is important). Because in a standard RE model one would need to assume

exogeneity of the current employment state and of experience with regard to unobserved heterogeneity, we need to use a three-equation model. Nevertheless, to see how our results evolve, it is instructive to re-estimate our results using restrictive, but more standard estimators. First, we estimate the same wage equation as before using a standard RE model with one equation. This model accounts for a time-constant unobserved individual effect, but assumes independence of the work history and the current employment and contract status with this random effect. As we have seen from our results on the variance parameters, such an assumption has been shown to be wrong, because we find strong and significant correlations between the random effects of the three equations. Nevertheless, when using the standard RE model instead of our three-equation model we get only little larger penalties than from our three-equation model (see table 12). The similarity of the results does not occur because the assumptions of the standard RE model are fulfilled, but because by chance the correlation of the random effect of the wage equation with the contract-type equation (which is negative) and the correlation with the employment equation (which is positive) and the correlation between the employment equation and the contract type equation (which is negative) partly happen to cancel out. Economically this means that the size of the positive unobserved selection into FT employment and the size of the negative selection into fixed-term contracts, given the link between them, given state dependence and given observed characteristics are such in our data that they partly offset each other. If and in how far this happens can of course not be anticipated ex ante and may be different in a different setting. Thus, this exercise should not be misinterpreted as evidence that the assumptions of a standard RE model are valid. What we can take away from this exercise is a hint that our results do not deviate from the literature because there were any problematic feature of our complex model. The exercise shows that nothing unexplainable happens.

Next we use pooled OLS to estimate our wage equation (see table 13). OLS results are strongly biased suggesting far too high PT penalties. Concerning the effect of working FT in

a fixed-term contract, already in the OLS estimation there is no negative effect of currently working in fixed-term contracts but the experience penalties are too large. It is thus indispensable to account for unobserved individual heterogeneity in particular with regard to the PT wage differential but also with regard to contract-type wage differentials.

Next we look in more detail at the issue that in our results women in fixed-term contracts do better (in PT as well as in FT) than in the previous literature summarized in the section on institutional backgrounds. The surprisingly good performance of women with fixed-term contracts working FT appears already (though to a little lower extend) when we estimate our model using OLS or a standard RE model, thus it is not (mainly) due to our methodological approach. But once we re-estimate the wage equation by POLS excluding the labor market history, we get a fixed-term contract penalty for FTrs of 16%. But if we add the covariates capturing PT and NE states in the past, this penalty reduces to 4% and adding the covariates capturing contract-history the penalty vanishes. This exercise highlights the importance of including the employment history with regard to NE, PT, and FT and the contract-history and interactions of these when estimating the wage effects of contract-types. A large part of the apparent fixed-term penalty seems to be driven by experience differences between women in the primary and the secondary segments of the labor market. But note that with regard to the returns to experience, fixed-term contracts are a bit worse than permanent contracts. Our study suggests that it is very important to clearly define if the estimated contract effect is a descriptive effect showing the wage differential of women working in the primary and the secondary segments of the labor market or a causal effect, i.e. reflecting the thought experiment of taking a women with an average labor market history and making her work in the secondary labor market. When attempting to estimate a causal contract type effect it turns out to be important to account for differences in the employment history, also with regard to PT work and intermittent behavior. Note that the relatively good

performance of fixed-term contracts does not only influence our results on FT fixed-term but also on PT fixed-term.

To sum up, sensitivity analysis shows that it is crucial to account for the history with regard to PT, FT and NE and with regard to contracts and interdependencies of these. We confirm that individual heterogeneity must be taken into account. Our methodology allows us to model history in a flexible way (while we do not have to assume exogeneity of the employment state and the history with regard to unobserved heterogeneity as in a standard RE model), because identification is supported by using cross-sectional comparison. In addition cross-sectional identification makes it possible to capture a situation in which wage paths may already diverge early in life

VII. Conclusion

In this study we re-examine wage differences between PT and FT workers and workers in fixed-term and permanent contracts. Employing a three-equation RE model, we use a different methodological approach than the previous literature which offers new perspectives. We follow Buchinsky, Fougère, Kramarz, and Tchernis, 2010, and explicitly model the participation/employment and PT status and type of contract mobility decisions, which, in turn, define the individual's experience in different segments of the labor market. We introduce into the wage equation a summary of the workers' entire career path. The three-equation system is estimated simultaneously using data from the Continuous Survey Working History (CSWH) and MCMC estimation.

Our results suggest that conditional on the labor market history, there is no penalty of currently working PT for mothers in a permanent contract, but there is a penalty of currently working PT for childless in a permanent contract and for both groups of women of currently working PT in a fixed term contract. Experience is rewarded less in PT work than in FT work

and these (low) returns are quite similar in the two segments of the labor market. But there is a remarkable exception: there is a particular strong penalty for childless women of working PT in the primary segment of the labor market. With regard to the PT literature these findings imply that it is important to disentangle the effect of currently working PT and of the employment histories. Part of the PT penalty can be explained by differential employment histories of PT and FE workers. Furthermore our results draw a less negative picture of wages of fixed-term contract work. Our results suggest that the low wages observed in the secondary segment of the labor market are largely due to differences in experience between women working in the primary and the secondary segments.

Our study also sheds light on several important factors leading to the apparent differences between our estimates and those obtained in the literature. First, our study highlights the importance of accounting for individuals' diverging employment histories especially across segments of the labor market. We demonstrate the need to explicitly control for past experience by type of contract and PT status in the wage function to capture the overall effect of the worker's specific career path on her market wage. In particular, our results show that it is important to account very carefully for experience when attempting to estimate causal effects of different contract types. We find that the magnitude of the estimated returns changes markedly when we account for this factor, indicating that the PT employment history during the course of one's career is important for her wage trajectory. This factor may be accounted for in a methodological approach which uses cross-sectional identification in addition to "within" identification and which endogenizes different form of experience. This can be done by explicitly modeling the employment, PT status and type of contract decisions while accounting for state dependence and duration dependence when modeling the contract type and the employment type. Second, we establish the need to account for time invariant unobserved heterogeneity prevailing in the participation and contract type and PT status decisions, as well as in the wage function. Finally, heterogeneity analysis by motherhood

status reveals interesting differences. As a next step we plan to allow for even more flexibility in the wage equation. This will allow us to calculate cumulative wage effects of additional employment patterns.

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Table 1
Incidence of Female PT and Fixed-Term Employment, OECD 2008

	Incidence of female PT employment	Incidence of female temporary employment
Australia	37.7%	5.9%
Belgium	33.8%	9.7%
Germany	38.6%	14.9%
The Netherlands	59.9%	20%
Norway	30.8%	11.1%
Spain	**21.1%**	**31.2%**
The United Kingdom	37.7%	6%
The United States	17.8%	4.2%

Table 2
Experience in Different Employment States (Means with Standard Deviations in Brackets)

	FT, Permanent	PT, Permanent	FT, Fixed-term	PT, Fixed-term
Dummy for respective employment status in t	0.876 (0.329)	0.0341 (0.181)	0.077 (0.267)	0.013 (0.111)
Years of experience in FT, Permanent observed since initial period of sample	5.011 (2.812)	4.137 (2.617)	1.668 (2.230)	2.597 (2.548)
Years of experience in PT, Permanent observed since initial period of sample	0.020 (0.242)	2.063 (1.862)	0.022 (0.241)	0.074 (0.433)
Years of experience in FT, Fixed-term observed since initial period of sample	0.104 (0.576)	0.227 (0.914)	2.457 (2.379)	0.381 (0.965)
Years of experience in PT, Fixed-term observed since initial period of sample	0.016 (0.206)	0.231 (0.774)	0.048 (0.312)	1.787 (2.104)
Year of FT experience in initial period	7.326 (3.774)	6.595 (3.268)	4.848 (2.613)	5.630 (3.020)
Years of PT experience in initial period	0.171 (0.602)	0.503 (1.038)	0.210 (0.576)	1.004 (1.409)

Table 3
Descriptive Statistics (Means with Standard Deviations in Brackets)

	FT, Permanent	PT, Permanent	FT, Fixed-term	PT, Fixed-term
Log hourly wage	2.386 (0.409)	2.046 (0.457)	2.277 (0.462)	1.806 (0.577)
Age from 23 to 29	0.072 (0.258)	0.013 (0.115)	0.102 (0.302)	0.028 (0.166)
Age from 30 to 34	0.315 (0.465)	0.271 (0.444)	0.389 (0.488)	0.219 (0.413)
Age from 35 to 39	0.393 (0.488)	0.447 (0.497)	0.309 (0.462)	0.429 (0.495)
Age from 40 to 45	0.219 (0.414)	0.269 (0.443)	0.120 (0.340)	0.324 (0.468)
Less than secondary education	0.322 (0.467)	0.455 (0.498)	0.315 (0.465)	0.599 (0.490)
Secondary education	0.410 (0.492)	0.379 (0.485)	0.230 (0.458)	0.269 (0.444)
More than secondary education	0.268 (0.443)	0.166 (0.372)	0.385 (0.487)	0.132 (0.338)
No child	0.486 (0.500)	0.251 (0.434)	0.582 (0.493)	0.342 (0.474)
Youngest child younger than 2 years	0.155 (0.362)	0.309 (0.462)	0.123 (0.328)	0.123 (0.329)
Youngest child 3 years	0.046 (0.209)	0.084 (0.277)	0.031 (0.173)	0.043 (0.202)
Youngest child 4 to 6 years	0.112 (0.315)	0.142 (0.349)	0.073 (0.259)	0.114 (0.318)
Youngest child 7 to 29 years	0.202 (0.401)	0.214 (0.450)	0.192 (0.394)	0.379 (0.485)

Table 4
Results (Means and Standard Deviations of Parameters)

	Whole Sample		Mothers (to be)		Childless	
	Mean	**SD**	**Mean**	**SD**	**Mean**	**SD**
	Wage Equation					
PT	-.0020	0.0036	-0.0073	0.0043	-0.0307	0.0069
PTlag1	-0.0634	0.0042	-0.0643	0.0050	-0.0696	0.0080
PTlag2 + PTlag3 + PTlag4	-0.0187	0.0012	-0.0167	0.0014	-0.0222	0.0022
PTlag5 + PTlag6 + PTlag7 + PTlag8	-0.0012	0.0008	-0.0024	0.0009	0.0037	0.0015
PTlag9 + PTlag10 + ... + PTlag40	-0.0014	0.0002	-0.0003	0.0002	-0.0055	0.0004
NElag1	-0.0306	0.0031	-0.0276	0.0038	-0.0368	0.0050
NElag2 +NElag3 +NElag4	-0.0383	0.0010	-0.0359	0.0012	-0.0435	0.0016
NElag5 +NElag6 +NElag7 +NElag8	-0.0256	0.0006	-0.0243	0.0008	-0.0282	0.0011
NElag9 +NElag10 + ... +NElag40	-0.0066	0.0002	-0.0049	0.0003	-0.0108	0.0004
FIX and FT	0.0290	0.0026	0.0350	0.0036	0.0221	0.0039
FIX and PT	-0.0508	0.0050	-0.0661	0.0062	-0.0059	0.0088
FIXlag1 and FTlag1	-0.0199	0.0029	-0.0280	0.0039	-0.0099	0.0043
FIXlag1 and PTlag1	0.0351	0.0054	0.0258	0.0067	0.0661	0.0095
FIXlag2 +FIXlag3 +FIXlag4	0.0018	0.0008	0.0021	0.0011	0.0017	0.0012
FIXlag5 +FIXlag6 +FIXlag7 +FIXlag8	-0.0004	0.0005	-0.0004	0.0007	-0.0002	0.0008
FIXlag9 +FIXlag10 + ... +FIXlag40	-0.0001	0.0001	0.0000	0.0002	-0.0003	0.0002
t=1	-0.0335	0.0017	-0.0303	0.0023	-0.0362	0.0025
t >1 & t <5	-0.0088	0.0004	-0.0080	0.0006	-0.0095	0.0006
t>4 & t < 9	-0.0018	0.0002	-0.0017	0.0002	-0.0019	0.0003
t	0.0061	0.0003	0.0061	0.0004	0.0058	0.0004
PT in t=0	-0.2029	0.0213	-0.1947	0.0244	-0.2502	0.0470
NE in t=0	-0.0461	0.0166	-0.0497	0.0202	-0.0648	0.0325
Experience in FT in t=0	0.0134	0.0009	0.0121	0.0012	0.0160	0.0016
Experience in PT in t=0	-0.0141	0.0036	-0.0125	0.0047	-0.0161	0.0060
FIX in t=0	-0.0261	0.0118	-0.0127	0.0173	-0.0442	0.0171
Experience in FIX in t=0	-0.0073	0.0126	-0.0334	0.0187	0.0277	0.0184
Age from 23 to 29	-0.0223	0.0019	-0.0205	0.0025	-0.0247	0.0029
Age from 30 to 34	-0.0043	0.0010	-0.0034	0.0012	-0.0054	0.0015
Age from 40 to 45	-0.0083	0.0011	-0.0061	0.0014	-0.0133	0.0017
Less than secondary education	-0.2185	0.0065	-0.2238	0.0085	-0.2153	0.0105
More than secondary education	0.3562	0.0063	0.3536	0.0086	0.3587	0.0100
No child	0.0192	0.0019	0.0231	0.0023	.	.
Youngest child younger than 2 years	-0.0022	0.0014	-0.0004	0.0016	.	.
Youngest child 3 years	-0.0054	0.0015	-0.0041	0.0017	.	.
Youngest child 4 to 6years	-0.0040	0.0012	-0.0032	0.0012	.	.
1997	0.0541	0.0095	0.0541	0.0138	0.0369	0.0151
1998	0.0443	0.0085	0.0451	0.0123	0.0293	0.0135
1999	0.0463	0.0074	0.0496	0.0108	0.0292	0.0118
2000	0.0343	0.0064	0.0369	0.0093	0.0203	0.0101
2001	0.0345	0.0053	0.0377	0.0077	0.0212	0.0085
2002	0.0264	0.0043	0.0298	0.0062	0.0145	0.0068
2003	0.0274	0.0033	0.0300	0.0047	0.0187	0.0052
2004	0.0163	0.0023	0.0196	0.0033	0.0084	0.0036
2005	0.0131	0.0014	0.0151	0.0019	0.0089	0.0022
April to June	-0.0072	0.0007	-0.0069	0.0009	-0.0071	0.0011
July to September	-0.0104	0.0008	-0.0102	0.0012	-0.0099	0.0013
October to December	-0.0115	0.0010	-0.0109	0.0014	-0.0111	0.0016
UR in province (deviation)	-0.0003	0.0001	-0.0005	0.0002	0.0001	0.0002
GDP growth in province (deviation)	0.0478	0.0162	0.0638	0.0211	0.0171	0.0256
Constant	2.1209	0.0089	2.1373	0.0123	2.1230	0.0137
	Contract Equation					
FIXlag1	3.6145	0.0417	3.6510	0.0520	3.5880	0.0529
FIXlag2 +FIXlag3 +FIXlag4	0.1213	0.0124	0.1248	0.0162	0.0984	0.0182
PT	0.4736	0.0382	0.3401	0.0485	0.7565	0.0662
PTlag1	-0.1947	0.0535	-0.1664	0.0694	-0.2641	0.0891
PTlag2 + PTlag3 + PTlag4	-0.0361	0.0193	-0.0038	0.0238	-0.0952	0.0334
NElag1	2.6181	0.0510	2.5723	0.0605	2.7272	0.0654
NElag2 +NElag3 +NElag4	0.2136	0.0150	0.2375	0.0191	0.1948	0.0241

FIX in t=0	0.4349	0.0700	0.3866	0.0971	0.5845	0.1079
Experience in FIX in t=0	0.5734	0.0719	0.6488	0.1023	0.5068	0.1097
t=1	0.3798	0.0454	0.4737	0.0578	0.2840	0.0624
t>1	0.0684	0.0094	0.0822	0.0125	0.0538	0.0132
Age from 23 to 29	0.2254	0.0503	0.2416	0.0730	0.2090	0.0764
Age from 30 to 34	0.0519	0.0270	0.0197	0.0380	0.0742	0.0406
Age from 40 to 45	-0.0731	0.0285	-0.0468	0.0345	-0.1116	0.0474
Less than secondary education	0.1333	0.0483	0.1587	0.0789	0.0979	0.1021
More than secondary education	0.0819	0.0580	0.0386	0.0923	0.1284	0.0858
No child	-0.0364	0.0381	-0.0070	0.0664	.	.
Youngest child younger than 2 years	-0.2537	0.0402	-0.2045	0.0505	.	.
Youngest child 3 years	-0.1233	0.0500	-0.0833	0.0579	.	.
Youngest child 4 to 6years	-0.0929	0.0387	-0.0661	0.0429	.	.
1997	-0.3862	0.0635	-0.3946	0.0796	-0.3817	0.0981
1998	-0.2500	0.0526	-0.2378	0.0711	-0.2572	0.0826
1999	-0.2943	0.0523	-0.2823	0.0700	-0.3112	0.0807
2000	-0.1213	0.0435	-0.1170	0.0593	-0.1113	0.0684
2001	-0.0537	0.0370	-0.0486	0.0512	-0.0533	0.0598
2002	0.0011	0.0346	-0.0268	0.0488	0.0480	0.0513
2003	0.0136	0.0322	0.0057	0.0422	0.0366	0.0485
2004	0.0771	0.0291	0.0666	0.0384	0.0938	0.0442
2005	0.0914	0.0261	0.0859	0.0354	0.1033	0.0408
April to June	0.1280	0.0227	0.1672	0.0299	0.0775	0.0339
July to September	0.0738	0.0220	0.1428	0.0291	-0.0175	0.0343
October to December	0.0634	0.0223	0.1200	0.0309	-0.0124	0.0340
UR in province (deviation)	0.0009	0.0036	0.0019	0.0046	0.0007	0.0053
GDP growth in province (deviation)	-0.2751	0.5310	-0.0251	0.7177	-0.5603	0.8614
Constant	-3.3298	0.0664	-3.4345	0.0876	-3.3555	0.0789
	Employment Equation					
PTlag1	-2.1835	0.0272	-2.1659	0.0317	-2.2191	0.0582
PTlag2 + PTlag3 + PTlag4	-0.0399	0.0079	-0.0474	0.0093	-0.0206	0.0151
NElag1	-3.0612	0.0213	-3.1133	0.0258	-2.9170	0.0429
NElag2 +NElag3 +NElag4	-0.0520	0.0068	-0.0498	0.0084	-0.0558	0.0117
FIXlag1 and FTlag1	-0.6840	0.0220	-0.6662	0.0269	-0.6998	0.0399
FIXlag1 and PTlag1	0.1132	0.0256	0.0874	0.0298	0.1840	0.0494
PT in t=0	-0.7528	0.0622	-0.7346	0.0696	-0.7866	0.1432
NE in t=0	-0.8479	0.0565	-0.8445	0.0613	-0.8454	0.1091
t=1	-0.1733	0.0276	-0.2268	0.0347	-0.0780	0.0466
t>1	-0.0347	0.0059	-0.0395	0.0077	-0.0254	0.0094
Age from 23 to 29	0.0446	0.0337	0.0152	0.0412	0.0602	0.0555
Age from 30 to 34	-0.0364	0.0146	-0.0355	0.0176	-0.0470	0.0260
Age from 40 to 45	0.1124	0.0165	0.1341	0.0198	0.0545	0.0307
Less than secondary education	-0.2190	0.0351	-0.2125	0.0429	-0.2292	0.0682
More than secondary education	0.0784	0.0404	0.0961	0.0541	0.0518	0.0691
Experience in FT in t=0	0.0226	0.0046	0.0200	0.0059	0.0287	0.0100
Experience in PT in t=0	-0.0164	0.0168	-0.0164	0.0214	-0.0102	0.0365
No child	0.0504	0.0255	0.1169	0.0373	.	.
Youngest child younger than 2 years	-0.4481	0.0225	-0.4340	0.0235	.	.
Youngest child 3 years	-0.1384	0.0261	-0.1254	0.0261	.	.
Youngest child 4 to 6 years	-0.0969	0.0205	-0.0865	0.0209	.	.
1997	0.1812	0.0400	0.2130	0.0479	0.1110	0.0761
1998	-0.1191	0.0296	-0.0916	0.0349	-0.1907	0.0545
1999	-0.1606	0.0274	-0.1346	0.0332	-0.2353	0.0492
2000	-0.2599	0.0244	-0.2162	0.0297	-0.3693	0.0458
2001	-0.3390	0.0221	-0.3133	0.0275	-0.4047	0.0409
2002	-0.3819	0.0202	-0.3281	0.0246	-0.4943	0.0377
2003	-0.3042	0.0186	-0.2779	0.0225	-0.3616	0.0344
2004	-0.3143	0.0171	-0.2866	0.0212	-0.3684	0.0318
2005	-0.2492	0.0160	-0.2345	0.0195	-0.2745	0.0297
April to June	0.0273	0.0121	0.0265	0.0146	0.0292	0.0218
July to September	0.0506	0.0122	0.0542	0.0151	0.0443	0.0215
October to December	0.0353	0.0119	0.0295	0.0142	0.0500	0.0211
UR in province (deviation)	0.0045	0.0021	0.0042	0.0026	0.0046	0.0037
Constant	3.7517	0.0498	3.7608	0.0572	3.6689	0.0799

Table 5
Results (Means and Standard Deviations of Variance Parameters)

	Whole Sample		**Mothers (to be)**		**Childless**	
	Mean	**SD**	**Mean**	**SD**	**Mean**	**SD**
	Variance and Covariance Paramters					
$Var(\alpha^W)$	0.1053	0.0012	0.1053	0.0016	0.1095	0.0020
$Var(\varepsilon^W)$	0.0212	0.0001	0.0219	0.0001	0.0202	0.0002
$Var(\alpha^C)$	0.3031	0.0177	0.3240	0.0220	0.3683	0.0268
$Var(\alpha^E)$	0.4091	0.0144	0.3672	0.0163	0.4831	0.0278
$Cov(\alpha^W, \alpha^E)$	0.0429	0.0027	0.0435	0.0032	0.0359	0.0046
$Cov(\alpha^W, \alpha^C)$	-0.0255	0.0033	-0.0276	0.0042	-0.0214	0.0051
$Cov(\alpha^C, \alpha^E)$	-0.0587	0.0104	-0.0517	0.0121	-0.0739	0.0166
$Var(\alpha^W) / (Var(\alpha^W) + Var(\varepsilon^W))$	0.8326	0.0017	0.8275	0.0023	0.8440	0.0026
$Var(\alpha^C) / (Var(\alpha^C) + 1)$	0.2325	0.0104	0.2445	0.0126	0.2689	0.0143
$Var(\alpha^E) / (Var(\alpha^E) + 1)$	0.2902	0.0072	0.2685	0.0087	0.3255	0.0126
$Corr(\alpha^W, \alpha^E)$	0.2065	0.0122	0.2212	0.0152	0.1563	0.0196
$Corr(\alpha^W, \alpha^C)$	-0.1426	0.0178	-0.1497	0.0220	-0.1065	0.0251
$Corr(\alpha^C, \alpha^E)$	-0.1665	0.0277	-0.1497	0.0334	-0.1753	0.0377

Table 6
Estimated Marginal Returns to a Particular Labor Force State: all*

Group	Marginal Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-5.281 (0.435)	-7.900 (0.515)	-0.637 (0.344)	-0.612 (0.088)	-0.612 (0.088)	-0.612 (0.088)	-0.612 (0.088)
PT Permanent	-0.201 (0.357)	-11.957 (0.417)	-0.473 (0.304)	-0.574 (0.081)	-0.574 (0.081)	-0.574 (0.081)	-0.574 (0.081)
FT fixed-term	2.895 (0.263)	-1.443 (0.296)	-0.163 (0.205)	-0.038 (0.051)	-0.038 (0.051)	-0.038 (0.051)	-0.038 (0.051)
NE	NA	-14.542 (0.329)	-10.228 (0.251)	-2.635 (0.099)	-2.635 (0.099)	-2.635 (0.099)	-2.635 (0.099)

*Year 0 is the current **quarter**, for the cumulative history the current quarter is included. "FT permanent" is the reference category. The returns are given in percent. Standard errors are as well multiplied by 100 and are given in brackets.

Table 7
Estimated Marginal Returns to a Particular Labor Force State: mothers

Group	Marginal Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-5.880 (0.556)	-8.230 (0.644)	-1.131 (0.424)	-0.125 (0.106)	-0.125 (0.106)	-0.125 (0.106)	-0.125 (0.106)
PT Permanent	0.731 (0.426)	-11.449 (0.497)	-0.973 (0.359)	-0.134 (0.096)	-0.134 (0.096)	-0.134 (0.096)	-0.134 (0.096)
FT fixed-term	3.502 (0.356)	-2.168 (0.402)	-0.158 (0.273)	0.009 (0.069)	0.009 (0.069)	0.009 (0.069)	0.009 (0.069)
NE	NA	-13.518 (0.410)	-9.713 (0.311)	-1.944 (0.121)	-1.944 (0.121)	-1.944 (0.121)	-1.944 (0.121)

Table 8

Estimated Marginal Returns to a Particular Labor Force State: childless

Group	Marginal Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-3.660 (0.693)	-6.516 (0.866)	1.389 (0.624)	-2.318 (0.168)	-2.318 (0.168)	-2.318 (0.168)	-2.318 (0.168)
PT Permanent	-3.074 (0.688)	-13.628 (0.808)	1.485 (0.582)	-2.205 (0.162)	-2.205 (0.162)	-2.205 (0.162)	-2.205 (0.162)
FT fixed-term	2.206 (0.392)	-0.489 (0.443)	-0.096 (0.310)	-0.113 (0.077)	-0.113 (0.077)	-0.113 (0.077)	-0.113 (0.077)
NE	NA	-16.738 (0.549)	-11.283 (0.430)	-4.309 (0.177)	-4.309 (0.177)	-4.309 (0.177)	-4.309 (0.177)

Table 9
Estimated Cumulative Returns to a Particular Labor Force State: all

Group	Cumulative Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-5.281 (0.435)	-13.181 (0.437)	-13.817 (0.431)	-14.429 (0.446)	-15.041 (0.478)	-15.653 (0.522)	-16.265 (0.577)
PT Permanent	-0.201 (0.357)	-12.159 (0.274)	-12.632 (0.251)	-13.206 (0.243)	-13.780 (0.262)	-14.354 (0.303)	-14.928 (0.357)
FT fixed-term	2.895 (0.263)	1.452 (0.230)	1.289 (0.241)	1.251 (0.256)	1.213 (0.280)	1.176 (0.310)	1.138 (0.345)
NE	NA	-14.542 (0.329)	-24.770 (0.342)	-27.405 (0.357)	-30.040 (0.397)	-32.675 (0.455)	-35.310 (0.526)

Table 10
Estimated Cumulative Returns to a Particular Labor Force State: mothers

Group	Cumulative Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-5.880 (0.556)	-14.111 (0.534)	-15.242 (0.527)	-15.367 (0.550)	-15.491 (0.591)	-15.616 (0.647)	-15.741 (0.714)
PT Permanent	0.731 (0.426)	-10.718 (0.328)	-11.691 (0.293)	-11.824 (0.285)	-11.958 (0.308)	-12.092 (0.356)	-12.225 (0.421)
FT fixed-term	3.502 (0.356)	1.334 (0.308)	1.176 (0.319)	1.185 (0.341)	1.194 (0.375)	1.202 (0.418)	1.211 (0.466)
NE	NA	-13.518 (0.410)	-23.231 (0.428)	-25.175 (0.446)	-27.119 (0.495)	-29.063 (0.565)	-31.007 (0.651)

Table 11
Estimated Cumulative Returns to a Particular Labor Force State: childless

Group	Cumulative Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-3.660 (0.693)	-10.176 (0.772)	-8.787 (0.768)	-11.105 (0.780)	-13.423 (0.827)	-15.741 (0.903)	-18.059 (1.003)
PT Permanent	-3.074 (0.688)	-16.702 (0.532)	-15.217 (0.494)	-17.422 (0.477)	-19.627 (0.514)	-21.832 (0.595)	-24.037 (0.705
FT fixed-term	2.206 (0.392)	1.717 (0.345)	1.621 (0.369)	1.508 (0.387)	1.395 (0.419)	1.282 (0.461)	1.170 (0.511)
NE	NA	-16.738 (0.549)	-28.021 (0.572)	-32.329 (0.601)	-36.638 (0.677)	-40.947 (0.786)	-45.256 (0.917)

Table 12

Sensitivity Analysis: Cumulative Returns Estimated by a Standard RE Model: all

Group	Cumulative Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-5.450	-13.417	-14.085	-14.721	-15.356	-15.992	-16.627
PT Permanent	-0.271	-12.260	-12.748	-13.334	-13.920	-14.506	-15.092
FT fixed-term	2.797	1.352	1.171	1.122	1.072	1.023	0.974
NE	NA	-14.707	-25.022	-27.767	-30.512	-33.258	-36.003

Table 13

Sensitivity Analysis: Cumulative Returns Estimated by a Pooled OLS: all

Group	Cumulative Returns (in %)						
	Years in that state						
	0*	1	2	3	4	5	6
PT fixed-term	-16.929	-30.743	-37.497	-39.729	-41.961	-44.193	-46.424
PT Permanent	-10.438	-25.352	-29.935	-30.436	-30.937	-31.437	-31.938
FT fixed-term	1.814	-1.999	-4.170	-5.901	-7.632	-9.363	-11.094
NE	NA	-19.401	-34.123	-38.792	-43.461	-48.131	-52.800