Escola de Ciências Sociais e Humanas e ISCTE Business School

Essays on the Portuguese Labour Market: The Effects of Flexibility at the Margin

Marta Silva

Tese especialmente elaborada para obtenção do grau de

Doutor em Economia

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Resumo

Esta tese pretende estudar os efeitos da promoção da flexibilidade à margem, através da facilitação do uso dos contratos a termo, no mercado de trabalho português. Para tal, apresentam-se três ensaios onde se analisam empiricamente os efeitos de uma alteração de legislação que ocorreu em 2004 e que aumentou a duração legal máxima dos contratos a termo certo de três para seis anos. Dada a elevada segmentação do mercado de trabalho e a representatividade dos contratos a termo, considera-se que estes estudos podem contribuir para a definição de políticas públicas no futuro. A análise empírica é realizada utilizando a base de dados Quadros de Pessoal para o período compreendido entre 2002 e 2011.

Os resultados dos três ensaios sugerem que os contratos a termo podem desemprenhar papéis diferentes no mercado de trabalho, como processos de seleção e de ajustamentos quantitativos, e que isso deve ser tido em consideração quando se analisam os efeitos de reformas assimétricas da protecção ao emprego. Após controlar por diversas variáveis micro e macro, encontra-se evidência de que a possibilidade de utilizar o contrato a termo por um maior período de tempo teve efeitos negativos na probabilidade de conversão do contrato e contribuiu para aumentar a desigualdade salarial entre trabalhadores com contratos sem termo ou contratos convertidos em sem termo e trabalhadores que não obtiveram uma relação de emprego mais estável. Para além disso, mostra-se que esta alteração de legislação não se traduziu num aumento do crescimento do emprego e que teve inclusivamente um efeito negativo na criação de emprego com contratos a termo, devido principalmente ao efeito da não conversão dos contratos.

Esta tese sugere que se deve combater a segmentação do mercado de trabalho português e promover medidas para estimular a conversão dos contratos a termo em contratos sem termo.

Palavras-Chave: Contratos a termo, Legislação de Proteção ao Emprego, Qualidade do Par Trabalhador-Empresa, Fluxos de Emprego, Salário Relativo, *Endogenous Switching Regression Model*, Modelo de Vetores Autoregressivos, Regressão de Quantis

JEL Classification: J31, J41, J68, C21, C24, C33

Abstract

This thesis aims to study the effects of the promotion of flexibility at the margin in the Portuguese labour market through the facilitation of the use of fixed-term contracts. We present three empirical essays that assess the effects of the 2004 change in legislation that extended the maximum legal duration of fixed-term contracts from three to six years. Given the high labour market segmentation and representativeness of fixed-term contracts in the Portuguese labour market, we consider that these studies may contribute to the design of future public policies. The empirical analysis is conducted using the linked employer-employee database *Quadros de Pessoal* for the period between 2002 and 2011.

The results of the three essays suggest that fixed-term contracts may play different roles in the labour market, namely as screening devices and quantitative adjustment tools, and that this fact should be taken into account when the effects of asymmetric employment protection reforms are analysed. After controlling for several micro and macro variables, we find evidence that the extension of the fixed-term contract for a longer period had negative effects on the probability of conversion of the contract and contributed to increase the wage inequality between workers on permanent or converted fixed-term contracts and those that did not obtain a more stable employment relationship. Besides, we find evidence that this change in legislation did not contribute to increase employment growth and decrease the rate of creation of fixed-term jobs, especially due to the effects of the proportion of non-converted fixed-term contracts.

This research suggests that labour market segmentation should be tackled and policy makers should promote measures aiming to stimulate the conversion of fixed-term into open-ended contracts.

Keywords: Fixed-term contracts, Employment Protection Legislation, Worker-firm Match Quality, Job Flows, Relative Wage, Endogenous Switching Regression Model, Vector Autoregressive model, Quantile Regression

JEL Classification: J31, J41, J68, C21, C24, C33

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List of Abbreviations

CI	Confidence Interval
EPL	Employment Protection Legislation
FTC	Fixed-term Contract
GLS	Generalized Least Squares
GMM	Generalized Method of Moments
IMR	Inverse Mills Ratio
IRF	Impulse Response Function
LABREF	Labour Market Reforms database
MCMC	Markov chain Monte Carlo
OEC	Open-ended Contract
OECD	Organisation for Economic Co-operation and Development
OLS	Ordinary least squares
VAR	Vector Autoregressive Model

Chapter 1

Introduction

This thesis is composed by three essays that empirically study the effects of the promotion of flexibility at the margin through regulations on fixed-term contracts.

Since the 1980s, a large number of reforms on employment protection legislation (EPL) have been implemented in European countries, out of which more than half were two-tier reforms (Boeri, 2011). The use of temporary contracts[1] was facilitated in approximately two thirds of the OECD countries where EPL has changed during the 1990s, although this figure dropped to 50% at the beginning of the following decade (OECD, 2004). Therefore, labour market flexibility was achieved mainly by widening the asymmetries between temporary and regular contracts, for which the regulations suffered little changes (Kahn, 2010; OECD, 2004). This kind of labour market flexibility, introduced by reforms directed to a certain groups of workers such as temporary workers, has been referred to in the literature as flexibility at the margin[2]. Consequently, it is not surprising that regulations on temporary contracts account for a large share of the cross-country variability in the OECD EPL index (OECD, 2004) and that the index on temporary contracts shows greater within-country variance than that on regular contracts (Kahn, 2010). This type of policy has contributed to raise the proportion of temporary workers (Kahn, 2010; OECD, 2004) and gave rise to highly segmented labour markets.

The prevalence of this type of labour market reforms can be explained in the spirit of the insider-outsider theoretical framework (Lindbeck and Snower, 1988), such as in Saint-Paul (1993). EPL, and notably the type of contract, may influence

[1] Note that, according to OECD, temporary contract is a broader term than fixed-term contract, since it includes not only fixed-term but also temporary agency work. Both concepts are often used interchangeably in the literature but in the proposed analysis we focus specifically on fixed-term contracts.

[2] See, for example, Bentolila and Dolado (1994), Mertens et al. (2007), Sala et al. (2012).

workers' bargaining power[3]. This is such that high-tenured workers on regular contracts, protected by substantial labour turnover costs[4], receive higher wages and are insulated from the competition of fixed-term workers[5]. Consequently, policy makers may find social and political obstacles when they try to implement broader reforms promoting labour market flexibility (Saint-Paul, 1993). If incumbents represent the majority of voters, broad reforms reducing firing costs will be derailed since they are expected to reduce the utility of permanent workers due to a larger firing rate. Thus, according to Saint-Paul (1993), reforms that increase flexibility at the margin are more likely to gather the necessary political support and may be considered a transition phase until the necessary conditions to implement complete reforms are achieved. In this setting, the conversion clause of fixed-term contracts, establishing that the contract has to be terminated or converted to an open-ended contract after a certain time period, helps to build the political support to implement reforms at the margin by ensuring the continuous insider's representativeness.

The above mentioned suggests that the regulations on fixed-term contracts, such as their maximum legal duration and number of renewals, are important tools that governments may use to promote or reduce labour market flexibility (Saint-Paul, 1996). According to the Labour Market Reforms database (LABREF), some European countries have allowed the maximum duration of fixed-term contracts to be extended for an additional period of time over the last two decades, such as in Belgium in 2002, Romania in 2005, Latvia in 2006 and Romania, Slovakia, Czech Republic in 2011[6]. This type of regulations may have important effects on the perceived job security of temporary employees (OECD, 2014). Consequently, we believe that it is crucial to better understand the effects of facilitating the use of fixed-term contracts. This is especially relevant after the 2008 economic and financial downturn that led to a severe job crisis and gave rise to the idiosyncratic responses of the

[3]See Lindbeck and Snower (2002) and Bentolila and Dolado (1994).

[4]According to Lindbeck and Snower (1988), labour turnover costs may be divided in "production-related" (e.g. screening costs) and "rent-related" costs (e.g. severance pay and lack of insiders' cooperation with outsiders).

[5]Evidence on the wage differential between permanent workers and fixed-term workers can be found in Booth et al. (2002) and OECD (2014) for example.

[6]https://webgate.ec.europa.eu/labref/public/result.cfm (consulted on 31/05/2016).

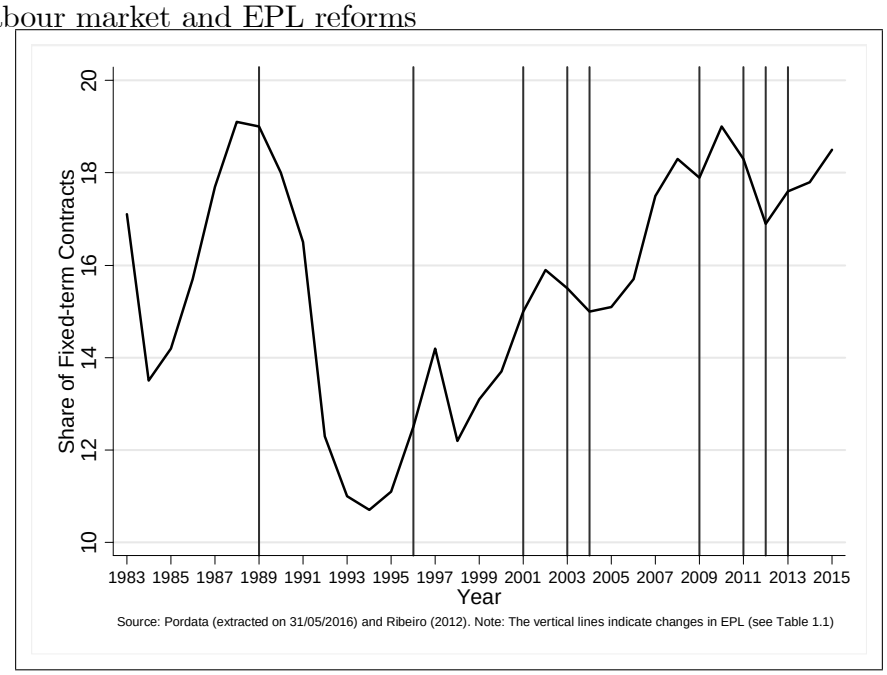
OECD labour markets (OECD, 2010). The effects of this crisis were particularly felt by temporary employees (OECD, 2010) who are used as a source of employment adjustment in countries such as Portugal, where the hiring and separation rates of temporary workers are more responsive to changes in employment than those of permanent workers (Carneiro et al., 2014).

Although the Portuguese unemployment rate showed an increasing trend since the early 2000s to 2014, it registered an unprecedented increase of approximately 67% between 2008 and 2012 (OECD, 2012). During the downturn, the adjustment of the Portuguese labour market was achieved more through employment reductions, which especially affected temporary contracts, than through wage cuts (Martins, 2016). This is mainly due to the fact that firms stopped hiring new employees and did not renew temporary contracts (Martins, 2016). This phenomenon is particularly relevant if we bear in mind that almost one fifth of total employees are on fixed-term contracts (Figure 1.1) and that this type of contract represented 80% of all new hires in the Portuguese labour market, between 2011 and 2012 (e.g. OECD, 2014). Hence, there is a non-negligible share of workers vulnerable to economic shocks and legislative changes.

Since 1976, when fixed-term contracts were firstly regulated in Portugal, important changes concerning their maximum number of renewals and, consequently, their maximum legal duration were implemented; the circumstances in which firms can hire workers on this type of contract and the severance payments due in case of dismissal were also altered (Ribeiro, 2012; de Campos, 2013). We summarize the main changes in EPL on fixed-term contracts in Table 1.1.

Although we cannot assume any causal effect between legislative changes and the proportion of fixed-term contracts from the simple observation of Figure 1.1, the latter seems to react with some lag to the changes in legislation described in Table 1.1. After the 1989 change in legislation that restricted the situations in which a worker could be hired under a fixed-term contract, the proportion of fixed-term employees sharply decreased and only started to recover in mid 1990s. Since then, we observe an increasing trend, disrupted since the 2008-9 downturn, as discussed above.

Figure 1·1: Proportion of fixed-term contracts in the Portuguese labour market and EPL reforms



According to Table 1.1, the maximum duration of the contract and number of renewals have been subject to at least four changes: in 2004[7], 2009, 2012 and 2013. From all of these legislative changes only that of 2009 intended to restrict the use of fixed-term contracts. Therefore, studying the effects of reforms increasing the maximum duration of fixed-term contracts is of special relevance to inform policy makers about their practical effects and contribute to the design of effective labour market policies.

For the above reasons, all the Chapters of this thesis focus on the Portuguese labour market and have as common thread to provide a complementary analysis of the effects of the 2004 change in legislation that extended the maximum duration of fixed-term contracts from three to six years. We focus in this change in legislation instead of the more recent ones not only because it was the one that most increased the maximum duration of the contract but also due to data availability constraints. We study the effects of the 2004 legislative change on four central labour market

[7]Although the extension of the contract was included in the 2003 Labour Code, it was only actually implemented on August 2004.

variables: the probability of conversion of fixed-term contracts, the wage growth experienced by fixed-term workers, job flows by type of contract and the ratio of permanent wages to fixed-term wages at the firm level. For that purpose, we use a rich database on the Portuguese labour market - *Quadros de Pessoal* - for the period between 2002 and 2011.

The present thesis fills a few gaps in this strand of literature. Firstly, to the best of our knowledge, this is the first comprehensive study of the effects of this type of legislative change in Portugal. Secondly, not enough attention has been given so far to the fact that workers and firms may be heterogeneously affected by the promotion of flexibility at the margin depending on the main use assigned to the fixed-term contract. We address this concern by allowing the effects of the legislative change to differ between fixed-term workers with converted and non-converted contracts and between firms with different wage policies regarding fixed-term contracts. This calls attention to the need to be careful when extrapolating some of the conclusions reached in studies for countries with different institutional characteristics.

In the remainder of this Chapter, we describe the way in which the thesis is organised and briefly present the motivation for each of the three essays that compose the thesis.

Table 1.1: Evolution of the Portuguese Employment Protection Legislation on Fixed-term Contracts: 1976-2013

Year	1976	1989	1996	2001
Categories	**Decret-Law N. 781/76, 28 October**	**Decret-Law N. 64-A/89, 27 February**	**Law N. 38/96, 31 August**	**Law N. 18/2001, 3 July**
Implemented in	October 1976	May 1989	August 1996	August 2001
admissibly (main situations)	To satisfy temporary needs and for casual and seasonal work	1) to replace a temporarily absent worker, 2) to satisfy an exceptional increase in firm's activity, 3) to satisfy seasonal or occasional activities, 4) to execute an occasional task or service, 5) to execute a temporary project, 6) to launch a new activity, firm or establishment, 7) to hire first job seekers or the long-term unemployed	The objective circumstance for using the fixed-term contract has to be stated	*idem*
successive contracts	n.a.	The worker can only be re-hired to the same job on a fixed-term contract 3 months after the fixed-term contract is terminated for reason not attributable to the worker	*idem*	The period in which re-hiring is forbidden increases to 6 months
preference in admission	n.a.	Workers on fixed-term contracts have preference in admission to identical functions under an open-ended contract. If this is violated, the employer has to pay the worker a compensation equal to one-half month's base wage.	*idem*	If the preference in admission clause is violated, the worker's compensation is increased to 6 months of base wage
equal treatment	n.a.	n.a.	n.a.	n.a.
training	n.a.	n.a.	n.a.	n.a.
maximum duration	3 years	3 years (exceptions: lauching a new activity, firm or establishment: 2 years)	*idem*	*idem*
maximum number of renewals	none	2	*idem*	*idem*
notice period	8 days	*idem*	*idem*	*idem*
severance payment	none	2 days of base wage for each complete month of duration	*idem*	3 days of base wage for each complete month of duration
unions	It cannot be modified by instrument of collective bargaining	*idem*	*idem*	Firms have to inform the workers' council and the union of the conclusion/extension/termination of fixed-term contracts
very short duration	The contract can be concluded for less than 6 months for specific temporary tasks	The contract can only be concluded for less than 6 months in the circumstances 1), 2), 3) and 4) described above. If violated: it is assumed that the contract was concluded for 6 months.	*idem*	*idem*

Source: All the referred legislation and Ribeiro (2012) and de Campos (2013). Notes: * Social security contributions can be augmented for firms with more than 15% of fixed-term contracts: 4/5 years of duration: 0,6%; 6 years: 1% (Law n. 35/2004, 29 July). n.a. stands for not applicable.

Year	2003	2009	2011	2012
Categories	**Labour Code (Law N. 99/2003, 27 August)**	**Labour Code (Law N. 7/2009, 12 February)**	**Law N. 53/2011, 14 October**	**Law N. 3/2012, 10 January**
Implemented in	December 2003 (the extension of the maximum legal duration was only implemented on August 2004 (Law n. 35/2004, 29 July))	February 2009 (regulation on the maximum duration only applies to new contracts)	November 2011 (only applies to new contracts)	January 2012 (only applies to fixed-term contracts created after 2009 that attains their maximum duration on 30 June 2013)
admissibly (main situations)	To satisfy temporary needs: *idem* + to indirectly substitute an absent employee	*idem* but fixed-term contracts can only be used to hire workers to lauching a firm or establishment with less than 750 workers.	*idem*	*idem*
successive contracts	The worker can only be re-hired to the same job after a period equal to one third of the total contract duration (Exceptions: new absence of the replaced employee, exceptional increase in firm's activity, seasonal activities and first job seekers)	*idem* but extended to the society to which the firm belongs	*idem*	*idem*
preference in admission	The worker has prefence in admission to an open-ended contract until 30 days after the termination of the contract. If violated: firm has to pay the worker three months of base wage.	*idem*	*idem*	*idem*
equal treatment	A fixed-term worker has the same rigths and the same obligations than a permanent worker.	*idem*	*idem*	*idem*
training	The employer has to provide training in a related field to workers on fixed-term contracts with more than 6 months of duration. The number of hours of training depends on the contract's duration: • <1 year: the number of hours of training is equal to 1% of the normal period of work • 1-3 years: 2% of the normal period of work • >3 years: 3% of the normal period of work	The employer has to provide training to workers on fixed-term contracts with more than 3 months of duration in the proportion of 35 hours of training/year	*idem*	*idem*
maximum duration	6 years *. (exceptions: launching of a new activity, firm or establishment and hiring long-term unemployed: 2 years. Hiring first job seekers: 18 months)	3 years. Exceptions: *idem*	*idem*	4,5 years
maximum number of renewals	introduces an additional renewal (Max.: 3 years, Min.: 1 year)	3	*idem*	introduced two additional renewals for contracts whose maximum duration is reached until June 2013 (Max.: 18 months; Min.: 1/6 of the contract's duration. The renewed contracts are only valid until 31 December 2014.
notice period	15 days	*idem*	*idem*	*idem*
severance payment	• Contracts with less than 6 months of duration: 3 days of base wage and seniority payment for each month of duration; • Contracts with more than 6 monts of duration: 2 days of base wage and seniority payment for each month of duration	*idem*	Contracts concluded before 1 November 2011: *idem* Contracts concluded after 1 November 2011: • 20 days of base wage and seniority payment for each year of tenure. Ceiling: the total compensation cannot be superior to 12 times the base wage and seniority payment or to 240 times the monthly minimum wage . The monthly wage and seniority payment cannot be superior to 20 times the monthly minimum wage (485€)	Before the extraordinary renewal: • Contracts with less than 6 months of duration: 3 days of base wage and seniority payment for each month of duration; • Contracts with more than 6 months of duration: 2 days of base wage and seniority payment for each month of duration. After the extraordinary renewal: 20 days of base wage and seniority payment for each year of tenure Ceiling: *idem*
unions	*idem*. It can be excluded and modified by instrument of collective bargaining (exceptions: hiring first job seekers and long-term unemployed)	*idem* (exceptions: *idem* + maximum duration and number of renewals)	*idem*	*idem*
very short duration	*idem*	introduction of a contract with very short duration (less than 1 week) for seasonal agriculture activities or turistic events. Do not have to be in writen form. The maximum duration of this type of contract is equal to 60 days per civil year. If violated it is assumed as a fixed-term contract of 6 month duration.	*idem*	*idem*

Year	2012	2013	2013
Categories	**Law N. 23/2012, 25 June**	**Law N. 69/2013, 30 August**	**Law N. 76/2013, 7 November**
Implemented in	August 2012	October 2013	November 2013
admissibly (main situations)	*idem*	*idem*	*idem*
successive contracts	*idem*	*idem*	*idem*
preference in admission	*idem*	*idem*	*idem*
equal treatment	*idem*	*idem*	*idem*
training	*idem*	*idem*	*idem*
maximum duration	*idem*	*idem*	5,5 years
maximum number of renewals	*idem*	*idem*	introduced 2 additional renewals for contracts whose maximum duration is reached until November 2015 (Max.: 12 months; Min.: 1/6 of the contract's duration). The renewed contracts are only valid until 31 December 2016
notice period	*idem*	*idem*	*idem*
severance payment	Until 31 October 2012: • Contracts with less than 6 months of duration: 3 days of base wage and seniority payment for each month of duration; • Contracts with more than 6 months of duration: 2 days of base wage and seniority payment for each month of duration After 31 October 2012: • 20 days of base wage and seniority payment for each year of tenure Ceiling: *idem*	Until 31 October 2012: • *idem*. Between 1 November 2012 and 30 September 2013: • *idem* After 30 September 2013: • 18 days (for the first 3 years) and 12 days (for the remainder of the contract duration) of base wage and seniority payment for each year of tenure Ceiling: *idem*	*idem*
unions	*idem*	*idem*	*idem*
very short duration	Written form is not necessary for very short term contracts with less than 15 days of duration in agriculture or turistic activities. The maximum duration of the contract has to be equal or lower than 70 days/year. If violated: *idem*	*idem*	*idem*

Structure and Main Goals of the Thesis

This thesis comprises three essays, each corresponding to one Chapter[8]. Chapter 2, entitled "Asymmetric labour market reforms and wage growth with fixed-term contracts: does match quality matter?"[9] addresses the impact of the legislative change on the wage growth experienced by workers on fixed-term contracts. We start by arguing that, since fixed-term contracts can play different roles in the labour market - they can be used as screening devices and as buffer stocks for permanent contracts - the effects on their wage growth should be heterogeneous depending on match quality. As argued by Jovanovic (1979), a worker-firm match is an "experience good", whose quality is only observed *a posteriori*. Fixed-term contracts may be considered an extension of the probationary period used to evaluate the quality of the match, which would be reflected in the conversion of the contract into an open-ended contract. We, therefore, test three hypothesis: i) whether extending the maximum duration of the contract influences the wage growth of fixed-term contracts; ii) whether converted fixed-term contracts are rewarded with higher wage growth, and iii) whether the extension of the contract penalizes more non-converted matches than good matches converted to an open-ended contract. In order to put these three hypothesis to test, we estimate an endogenous switching regression model, which allows us to study the wage growth of converted and non-converted fixed-term contracts separately, while correcting for the propension of the contract to be converted. In the second part of the analysis, we study the sources of the wage growth differential between both groups using a threefold decomposition, which allows us to estimate how this differential was affected in the period in which the change in legislation was in force. This essay is, to the best of our knowledge, the first empirical study of the effects of this type of employment protection reform on the wage growth differential between converted and non-converted fixed-term contracts. It contributes to the literature by presenting evidence that supports that

[8] Given this organisation choice and the fact that each chapters analyses different effects of the same change in legislation, there are a few repetitions throughout the Chapters.

[9] This Chapter is published as a Working Paper entitled: "Asymmetric labour market reforms and wage growth with fixed-term contracts: does learning about match quality matter?", Working Paper 15-04, Business Research Unit, ISCTE-IUL (co-authors L.F. Martins and H. Lopes).

This article was also submitted for publication on 30 June 2015 and is currently under revision.

if fixed-term contracts are used to assess the quality of the match, the effects of asymmetric employment protection reforms may be heterogeneous among workers on this type of contract.

We then turn to the study of how the change in legislation affected job flows by type of contract. In Chapter 3 - "Job Flows and Flexibility at the Margin" - we introduce a novel index of flexibility at the margin, built following Alexandre et al.'s (2010) methodology, and estimate a panel Vector Autoregressive (VAR) model. This study contributes to understanding the employment adjustment process that results from the promotion of flexibility at the margin by considering the effects of the latter on job flows by type of contract. Moreover, since we adopt a disaggregated unit of observation - resulting from the crossing of firms' sectors of activity, firm's region and firm's size and age classes - we are able to capture firm's heterogeneity, which has a relevant contribution to job flows dynamics (Davis and Haltiwanger, 1999). Finally, by introducing the index of flexibility at the margin, composed by the proportion of fixed-term contracts, the share of non-converted contracts and the average tenure of workers on fixed-term contracts, we are able to test the exogeneity of the index of flexibility at the margin and draw conclusions about not only how employment dynamics are affected but also whether and how they affect changes in EPL.

In Chapter 4, entitled "Wage inequality between permanent and fixed-term contracts: a firm-level analysis", we study how the within-firm relative wage between permanent and fixed-term contracts is affected during the period in which the fixed-term contract was allowed to be extended. Although the wage gap between fixed-term and open-ended contracts has been widely studied in the literature, we do not know much about the determinants of the relative wage at the firm level and how the latter is affected by the legislation extending the use of fixed-term contracts. This is mainly due to the lack of data on the average wage by type of contract at the firm level. We take advantage of a rich dataset and address this topic. However, since the within-firm wage gap may also reflect the way in which firms use fixed-term contracts, firms with larger wage differentials between both types of contract may react differently to the legislative change under study when compared to firms that do not pay wage premiums to open-ended contracts. Because of that, we estimate

a quantile regression with nonadditive fixed effects (Powell, 2016), to infer about the impact of the change in legislation on the conditional distribution of the wage ratio of permanent to fixed-term workers at the firm level. Finally, in an attempt to understand the importance that fixed-term contracts may have in the employment reallocation from the non-tradable to the tradable sector, we assess whether the effect of the change in legislation on the relative wage is homogeneous across firms in both sectors.

The general conclusions are discussed in Chapter 5, where we also present the main policy recommendations arising from the results of the empirical analysis conducted in the three essays.

Chapter 2

Asymmetric labour market reforms and wage growth with fixed-term contracts: does match quality matter?

2.1 Introduction

The productivity of a worker in a given firm depends on the quality of their match, which is learned over time by both parties (Jovanovic, 1979). The cost and the facility with which unproductive matches are terminated depend on the strictness of some labour market institutions, such as the EPL.

In recent years, EPL was reformed in some European countries in order to introduce some flexibility in the labour market mainly at the margin by relaxing the restrictions on the use of fixed-term contracts instead of reducing the protection of open-ended contracts (Kahn, 2010; Boeri, 2011). Prior evidence indicates that fixed-term contracts tend to bear the adjustment cost of legislation that widens the employment protection gap between open-ended and fixed-term contracts due to employment and wage levels. Namely, fixed-term contracts become less likely to be converted into permanent (Boeri, 2011; Centeno and Novo, 2012) and these workers may suffer a wage penalty resulting either from reforms that increase the protection of open-ended contracts (Centeno and Novo, 2014) or reforms that reduce the restrictions on the use of fixed-term contracts (Pérez et al., 2016).

Notwithstanding, previous contributions have neglected the fact that fixed-term contracts can play different roles in the labour market and, therefore, asymmetric employment protection reforms may have a heterogeneous impact. Following Jovanovic (1979), who classifies a worker-firm match as an "experience good", fixed-term contracts may play a crucial role by allowing firms to experiment different matches

before offering a permanent contract. Thus, if fixed-term contracts are used to extend the probationary period, their conversion into permanent contracts and the subsequent wage growth should reflect the performance of the match (Wang and Weiss, 1998; Loh, 1994). Good matches, i.e., matches that go from a fixed-term contract to an open-ended contract should be compensated through higher wage growth. They should also suffer less from the adverse impacts of reforms that widen the employment protection gap between fixed-term and open-ended contracts.

This article aims to provide further evidence of the impact of these institutional reforms by studying how they affect wage growth experienced by workers on fixed-term contracts given the learning process about match quality such contracts permit. To the best of our knowledge, this is the first empirical study on the sources of the wage growth differential between non-converted and converted fixed-term contracts and how it is affected by employment protection reforms that facilitate the use of fixed-term contracts. We focus on the change in the Portuguese EPL in 2004 that was subsequently overturned in 2009. This reform contributed to widening the protection gap between fixed-term and open-ended contracts by easing the restrictions on fixed-term contracts. More specifically, it introduced a third possible renewal of fixed-term contracts up to a maximum legal duration of 6 years and extended the conditions in which a fixed-term worker could be hired.

In order to test the abovementioned hypotheses, we use exceptionally rich Portuguese linked employer-employee data for the period 2003 to 2009 and estimate an endogenous switching regression model, similarly to Loh (1994) and Amuedo-Dorantes and Serrano-Padial (2007). This has the advantage of taking into account the possible selection bias arising from the fact that both the conversion and the wage growth of fixed-term contracts are simultaneously determined and affected by the learning process. Firstly, we test the significance and estimate the impact of the change in legislation on a proxy of match quality: the probability of conversion of fixed-term contracts into open-ended contracts. Secondly, we assess whether the change in legislation has a different impact on wage growth experienced by good matches, i.e., converted fixed-term contracts and non-converted fixed-term contracts. Thirdly, we study the sources of the wage growth differential between

those two groups using a threefold decomposition and evaluate how it is affected by the change in legislation.

Our results show that there is a statistically significant increase in the wage growth associated with the conversion of a fixed-term contract into a more stable employment relationship. Although the results suggest that match quality is negatively affected by employment protection reforms that ease the regulations on fixed-term contracts, the wage growth of good matches is less penalised by the change in legislation (-0.16 pp.) than that of non-converted fixed-term contracts (-0.55 pp.). We estimate that the implementation of this type of reform contributes to increase the wage growth differential between workers who remain on a temporary contract and those who receive an open-ended contract (15%). We argue that policy makers should tackle labour market segmentation given that asymmetric employment protection reforms that facilitate the use of fixed-term contracts may generate potential inefficiencies, such as the postponement of the conversion of the contract and the weakening of the link between this conversion and wage growth.

The next section characterises the Portuguese EPL and describes the change under analysis. Section 2.3 reviews some of the most relevant literature on the role of fixed-term contracts and briefly discusses the measurement of match quality. Section 2.4 presents the empirical approach and the dataset and Section 2.5 presents the main results obtained. Section 2.6 concludes.

2.2 The Portuguese Employment Protection Legislation

The Portuguese labour market is characterised by stringent EPL on regular contracts and by one of the largest employment protection gaps between temporary and open-ended contracts.

Fixed-term contracts were regulated in 1976 in the Portuguese labour market and their maximum legal duration was set at three years. In 1989 the situations in which a worker could be hired under a fixed-term contract were clearly defined and it was established that fixed-term contracts could only be renewed twice before reaching their maximum duration. This law also entitled the worker to receive a severance payment equal to two days for each month of work when the fixed-term

contract ends without conversion in a permanent contract[1].

Between 2003 and 2009, Portugal was the OECD country that most relaxed EPL (Venn, 2009). During this period, the main reform aimed to promote a more flexible labour market by easing the regulations on temporary contracts; meanwhile, the legislation on open-ended contracts was subject to little change. We examine the effect of the change to the legislation between 2004 and 2008 whereby the maximum duration and the situations in which fixed-term contracts could be used were extended. More specifically, the law introduced three changes: the possibility to renew the contract up to three times instead of just twice before reaching the maximum legal duration; the extension of the contract's maximum legal duration from three to six years; and the possibility to hire a worker on a fixed-term contract to satisfy temporary requirements at the firm level and notably to indirectly substitute an employee. The 2004 legislation also made it mandatory for firms to provide training for workers on fixed-term contracts of more than six months so as to bring their working conditions more in line with those on open-ended contracts. In 2009, the maximum legal duration of fixed-term contracts was restored to three years.

According to Eurostat, the proportion of temporary contracts in the total employment more than doubled between 1995 and 2009, reaching 22% in 2009. Given the growing representativeness of temporary contracts and the recurrent use of such changes in legislation on the Portuguese labour market, the impact of asymmetric employment protection reforms and, especially, how they affect workers on fixed-term contracts are major and current policy concerns.

2.3 Learning about Match Quality in Two-tier Systems

2.3.1 The Role of Fixed-term Contracts in Two-tier Systems

There is no consensus in the literature on the role of fixed-term contracts in the labour market. According to the segmented labour market theory, the labour market

[1] In 2004, the severance payment was equal to three days for each month of work for contracts with less than 6 months of duration and to two days for each month of work for contracts with more than six months of duration; this is not very different from the requirements in open-ended contracts i.e. 30 days per year of seniority. Nevertheless, for open-ended contracts, the administrative costs associated with a dismissal are significantly higher as discussed by Centeno and Novo (2014).

is composed of two segments characterised by distinct wage-setting behaviours and different non-pecuniary conditions. The primary segment offers higher wages, better working conditions and career progress and as Dickens and Lang (1985) highlight, tends to offer positive returns to schooling and experience, while the wage equation associated with the secondary segment is flat. Most fixed-term contracts are found in this secondary segment and suffer a non-negligible wage penalty relatively to open-ended contracts, for e.g. in France (Blanchard and Landier, 2002), Germany (Pfeifer, 2012; Hagen, 2002), Britain (Booth et al., 2002), Spain (Jimeno and Toharia, 1993) and Great Britain, Germany, France, Sweden and Portugal (Brown and Sessions, 2005). Similarly, using a French database of young workers, Blanchard and Landier (2002) warn that fixed-term contracts lead to high turnover rates even when good matches are formed as firms want to avoid the high firing costs associated with permanent contracts. Hence, workers on fixed-term contracts face a greater risk of becoming unemployed (McGinnity and Mertens, 2002) and being trapped in a chain of temporary contracts, as reported by Hagen (2002) for Germany and Gash and McGinnity (2007) for French female workers. Specifically, it is more difficult for women, youngsters and males with lower levels of education to escape from successive temporary jobs, since they are less likely to be promoted to permanent contracts (Alba-Ramírez, 1998). Fixed-term workers are therefore less likely to participate in training activities (Booth et al., 2002; Arulampalam et al., 2004).

As Bentolila and Saint-Paul (1992) predict, the introduction of temporary contracts may also make employment respond more to macroeconomic shocks. In other words, temporary workers can be used as a buffer stock that allows firms to respond to shocks more easily and at a lower cost by adjusting the employment level, especially downwards (Varejão and Portugal, 2007). This evidence is also supported by Boockmann and Hagen (2001) who argue that the probability of hiring on fixed-term contracts increases with positive fluctuations in product demand, measured by firm turnover, and with the level of employment protection of open-ended contracts.

Another strand of the literature explaining the role of fixed-term contracts rests on the screening hypothesis. Due to the existence of imperfect information, worker-firm matches are 'experience goods'(Jovanovic, 1979) and fixed-term contracts may

be used to assess the quality of the match before offering a permanent contract. Hence, fixed-term contracts may play a very important role by extending the probationary period and allowing firms to screen workers at a lower cost. This is documented by the high probability of fixed-term contracts to be converted into open-ended contracts reported for some countries, e.g. in France, one third of short-term contracts are converted at their maximum legal duration (Abowd et al., 1999), and in West Germany, nearly 40% of temporary contracts are converted within one year and most of them with the same employer (McGinnity and Mertens, 2002).

The use of fixed-term contracts as screening devices helps explain the heterogeneity of the pecuniary penalty associated with this type of contract and the catch-up with their permanent counterparts both in terms of wages and job stability as reported in the literature. Boockmann and Hagen (2008) find that the survival rate of German fixed-term contracts converges with that of open-ended contracts, although a match initiated with a fixed-term contract terminates more often in the two first years. Some authors using German data also argue that while the highest share of fixed-term contracts is found in the lower quartile of the wage distribution (Mertens and McGinnity, 2003)[2], the wage penalty of fixed-term contracts decreases as we move into higher quantiles (Mertens and McGinnity, 2003; Pfeifer, 2012; Mertens et al., 2007) and it is larger for matches lasting up to two years (Pfeifer, 2012); this supports the idea that there is a group of fixed-term contracts that faces a less severe pecuniary penalty. In line with Loh (1994) and Wang and Weiss (1998), if fixed-term contracts are used as screening devices, their wage may converge to the level of permanent contracts when converted and, therefore, they will experience higher wage growth (Sicilian, 1995). Accordingly, some authors such as McGinnity and Mertens (2002), for Germany, and Ruiz and Gomez (2009) and Amuedo-Dorantes and Serrano-Padial (2007), for Spain, find evidence that workers with fixed-term contracts experience higher wage growth than workers with open-ended contracts, especially those lasting more than one year and staying in the same job (Amuedo-Dorantes and Serrano-Padial, 2007) and those receiving an open-ended contract (Ruiz and Gomez, 2009). This steeper wage growth path is generally

[2]Note that the results in Mertens and McGinnity (2003) refer to West Germany only.

more marked in the case of female workers, whose wage penalty seems to be fully reversed due to the learning effect, measured by the accumulation of experience (Pavlopoulos, 2013), whereas males seem to suffer a more persistent wage penalty (Pavlopoulos, 2013; Booth et al., 2002). For example, Gash and McGinnity (2007) use a matching methodology to support this conclusion by showing that, unlike men, women on fixed-term contracts in West Germany experience higher wage growth than those on permanent contracts in the two years after being hired. Finally, Mertens and McGinnity (2003) argue that although only fixed-term contracts in the highest wage growth quartiles have a wage growth premium relatively to their permanent counterparts; fixed-term contracts in the lowest quartiles of the wage distribution are more likely to experience high wage growth.

As for the Portuguese labour market, there is some evidence indicating that fixed-term contracts are used as screening devices. Varejão and Portugal (2007) argue that even establishments with a stable employment level tend to hire more, rather than separate more from workers on temporary contracts, which means that some matches are continued and converted to permanent contracts. Similarly, Portugal and Varejão (2005) contend that a significant proportion of fixed-term contracts are converted into open-ended contracts, although workers on fixed-term contracts are more likely to switch jobs and become unemployed or inactive. In fact, the probability of conversion is low when the match is formed but tends to increase during the two first years of contract (Portugal and Varejão, 2009). The screening hypothesis is also supported by the fact that workers in longer employment relationships are less likely to move to another job (Portugal and Varejão, 2005).

Although fixed-term contracts can play different roles in the labour market, they tend to bear the adjustment cost of reforms that widen the employment protection gap between fixed-term and open-ended contracts. Using a difference-in-differences analysis, Centeno and Novo (2012) find that the extension of the employment protection of open-ended contracts to firms with 11 to 20 employees has not only increased the proportion of workers on fixed-term contracts but also their churning at firm level. Consequently, these workers also received lower wages, as reported in Centeno and Novo (2014). Thus, in a segmented labour market like that of Por-

tugal, fixed-term contracts may be used as a source of both wage and employment flexibility (Centeno and Novo, 2012; Centeno and Novo, 2014).

This paper focus on the impact of legislation reforms that facilitate the use of fixed-term contracts taking into account that this type of contract may be used to learn about match quality, which measurement is briefly discussed in Subsection 2.3.2.

2.3.2 The Measurement of Match Quality

There is robust evidence of a non-negligible impact of match quality on wages (Hersch and Reagan, 1990) and wage growth (Yamaguchi, 2010).

However, match quality contains various dimensions and can therefore be measured by several proxies. The job-search literature predicts that, after a match is formed, better alternative matches might appear which offer a higher wage than the worker's reservation wage. Therefore, the starting wage is a good proxy to measure match quality, and turnover is the mechanism used to form more efficient matches. Accordingly, some authors such as Gaure et al. (2012), Centeno and Novo (2006) and van Ours and Vodopivec (2008) use the starting wage as an *a priori* measure of match quality to study the impact of unemployment benefits on match quality.

Other authors classify a match as an "experience good", whose true value is only known *a posteriori* after experimentation (Jovanovic, 1979). Jovanovic's job matching hypothesis predicts that higher value matches endure and achieve higher wages while bad matches are terminated. According to this perspective, match quality can be measured by the duration of the employment relationship and by the wage growth. More specifically, tenure is used as a proxy of match quality by Centeno (2004) and Centeno and Novo (2006) to study the effects of unemployment insurance on match quality, by Allgood et al. (2012) to disentangle the impact of the expected match quality on the CEO's initial compensation and by Yankow (2009) to study the impact of match quality on job search behaviour in urban areas.

Finally, a few authors (e.g. Ferreira and Taylor, 2011), rely on subjective indicators of match quality based for example on worker's satisfaction and the will to switch jobs.

Given that the goal of the present analysis is to assess the impact of a change in the maximum legal duration of fixed-term contracts while taking the learning process about match quality into account, we classify a match as an "experience good" whose quality is measured *ex post*. However, tenure is not a suitable measure for our purposes since it would reflect not only the learning about match quality but also the direct impact of the reform on its upper bound. Therefore, we take the conversion of fixed-term contracts to permanent contracts and the subsequent wage growth as measures of match quality that reflect and incorporate the learning process.

In the next section, we present the econometric methodology that we find most suitable to assess the impact of the change in legislation taking into account the learning process about match quality.

2.4 Empirical Approach

According to Jovanovic (1979), a match needs to be experienced in order to evaluate its quality, which is a trial and error process. Therefore, fixed-term contracts could be an important tool to test different matches, learn about their quality and terminate the bad ones easily and at a lower cost.

Workers are matched with firms and they are given fixed-term contracts. The quality of the match is unobserved before the match is experienced:

$$Z^*_{mt} = w'_{mt}\omega + \Pi'_t\tau + D'_t\delta + \varepsilon_{mt}, \quad \text{m=1,..., M and t=1,...,T.} \tag{2.1}$$

It is assumed that Z^*_{mt} is a latent continuous random variable representing the match quality of a certain worker-firm pair m at period t. The total number of matches equals M and the total number of time periods equals T. As stated in equation 2.1, the value associated with a certain match m depends on a set of exogenous variables, w_{mt}, including the worker's characteristics (age and its square, tenure, gender, nationality, education, occupation) and firm's characteristics (dimension, region, sector of activity, share of fixed-term contracts[3] and capital ownership). Π_t

[3]We considered the one period lagged value of the share of fixed-term contracts, in order to

includes a set of year dummies to control for time effects and the annual unemployment rate to control for the business cycle. Since one of the purposes of the analysis is to evaluate how the change in legislation impacts on match quality, a variable D_t is also included, which is a regime dummy taking value zero in 2003 and 2009 and one in the remaining years of the sample in which the law was in force. The impact of the referred change in legislation is captured by δ.

Firms can hire a worker using a fixed-term contract up to a certain maximum legal duration, when the contract is automatically converted to permanent if the match is continued. Over time, both parties (worker and firm) learn about the value associated with the match and only good matches, i.e., matches yielding a positive value, are converted to permanent contracts since this type of contract is associated with higher labour turnover costs:

$$P_{mt} = I\left[Z^*_{mt} > 0\right]. \tag{2.2}$$

Thus, P_{mt} is a dummy variable taking value one when the match initiated with a fixed-term contract is converted into permanent between t-1 and t and zero when the match is continued but is not converted[4], which expresses the sign of the latent match quality. $I\left[.\right]$ is an indicator function assuming value one when the argument is true and zero otherwise. Thus, we assume that a good match is one that started with a fixed-term contract and was converted into a more stable employment relationship. Nevertheless, non-converted matches cannot be considered bad matches since the match is continued and the learning process may not yet be complete.

As Sicilian (1995) and Jovanovic (1979) argue, wage growth is a result of the learning process about match quality. Ceteris paribus, workers in good matches should experience higher wage growth than workers in low value matches. Accordingly, employment protection reforms could have an asymmetric impact on the wage growth of converted and non-converted matches. Since the marginal effect of the explanatory variables and the change in legislation is expected to differ, we should

account for endogeneity.

[4]Since the unit of observation is the worker-firm match and the wage is match specific, we only considered continuing matches in order to confine the study to the wage growth on the job rather than the wage growth resulting from job mobility.

distinguish between the wage growth of converted and non-converted matches:

$$W_{gt} = x'_{gt}\beta_g + \Pi'_t\tau_g + D'_t\delta_g + v_{gt} \; if \; P_{mt} = 1 \tag{2.3}$$

$$W_{bt} = x'_{bt}\beta_b + \Pi'_t\tau_b + D'_t\delta_b + v_{bt} \; if \; P_{mt} = 0, \tag{2.4}$$

a good match is represented by $g = 1, ..., G$ and a non-converted match by $b = 1, ...B$ over $t = 1, ..., T$ periods of time[5].

The wage growth experienced by good matches between t-1 and t (W_{gt}) is observed if the fixed-term contract is converted into a permanent contract between t-1 and t. Otherwise, we observe the wage growth of the matches that remained with a fixed-term contract between t-1 and t (W_{bt}). Since we intend to study the differences in the wage growth between these two groups, we introduce a set of independent variables, x_{gt} and x_{bt}, in order to ascertain the contribution of certain worker and firm characteristics. We are interested in obtaining the estimates of β and δ, representing the marginal impact of each covariate and the impact of the change in legislation on the wage growth of converted and non-converted matches respectively.

In such a scenario, where the sample is not random, using the standard OLS estimation would produce inconsistent estimates[6]. We adopt an endogenous switching regression model in order to tackle the problem arising from the simultaneous decision to convert the contract and the setting of the wage level and, thus, the non-random sampling, and consistently estimate the impact of the explanatory variables and the change in legislation. This type of model is an extension of the Heckman selection model (Heckman, 1979) in which both regimes are observable. Thus, assuming that the error term of the selection equation (ε_{mt}) is drawn from a standard

[5] Note that the total number of converted (G) and non-converted (B) matches corresponds to the whole sample dimension (M).

[6] $E(W_{gt}|P_{mt} = 1, x_{gt}, \Pi_t, D_t) \neq x'_{gt}\beta_g + \Pi'_t\tau_g + D'_t\delta_g$ and $E(W_{bt}|P_{mt} = 0, x_{bt}, \Pi_t, D_t) \neq x'_{bt}\beta_b + \Pi'_t\tau_b + D'_t\delta_b$ since $E(v_{gt}|P_{mt} = 1, x_{gt}, \Pi_t, D_t) \neq 0$ and $E(v_{bt}|P_{mt} = 0, x_{bt}, \Pi_t, D_t) \neq 0$.

normal distribution $N(0,1)$, while v_{gt} and v_{bt} follow a normal distribution $N(0,\sigma_g^2)$ and $N(0,\sigma_b^2)$ respectively, and that the switch is endogenous, i.e. v_{gt} and ε_{mt} and v_{bt} and ε_{mt} are significantly correlated, we follow the two-step procedure described by Maddala (1986, pp.223-228) in order to estimate the wage growth of both converted and non-converted matches[7]. The identification of the model is made not only through the assumption of joint normality but also by the exclusion of some covariates included in w_{mt}, from x_{gt} and x_{bt}. Specifically, we exclude two dummy variables accounting for less than 9 years of schooling, one dummy variable accounting for the activity sector of electricity production and distribution, and one dummy variable accounting for firm size of more than 401 employees. Thus, it is assumed that these variables only significantly affect the probability of conversion of fixed-term contracts and not the subsequent wage growth path[8].

As such, in the first step, equation 2.2 is estimated through maximum likelihood as a pooled[9] Probit regression in order to obtain the parameter estimates and compute the estimated inverse mills ratio. In the second step, a pooled generalised least square (GLS) estimator is used to estimate equations 2.5 and 2.6:

$$W_{gt} = x'_{gt}\beta_g + \Pi_t^{'}\tau_g + D'_t\delta_g + \sigma_g\rho_{g\varepsilon}\frac{\phi(w_{mt}^{'}\widehat{\omega} + \Pi_t^{'}\widehat{\tau} + D'_t\widehat{\delta})}{\Phi(w_{mt}^{'}\widehat{\omega} + \Pi_t^{'}\widehat{\tau} + D'_t\widehat{\delta})} + u_{gt}\ if\ P_{mt} = 1 \quad (2.5)$$

$$W_{bt} = x'_{bt}\beta_b + \Pi_t^{'}\tau_b + D'_t\delta_b - \sigma_b\rho_{b\varepsilon}\frac{\phi(w_{mt}^{'}\widehat{\omega} + \Pi_t^{'}\widehat{\tau} + D'_t\widehat{\delta})}{(1-\Phi(w_{mt}^{'}\widehat{\omega} + \Pi_t^{'}\widehat{\tau} + D'_t\widehat{\delta}))} + u_{bt}\ if\ P_{mt} = 0, \quad (2.6)$$

where ϕ and Φ represent the standard normal density function and the standard normal cumulative distribution function. $\frac{\phi(w_{mt}^{'}\widehat{\omega}+\Pi_t^{'}\widehat{\tau}+D'_t\widehat{\delta})}{\Phi(w_{mt}^{'}\widehat{\omega}+\Pi_t^{'}\widehat{\tau}+D'_t\widehat{\delta})}$ is the inverse mills ratio in the cases in which $P_{mt} = 1$ and $\frac{-\phi(w_{mt}^{'}\widehat{\omega}+\Pi_t^{'}\widehat{\tau}+D'_t\widehat{\delta})}{(1-\Phi(w_{mt}^{'}\widehat{\omega}+\Pi_t^{'}\widehat{\tau}+D'_t\widehat{\delta}))}$ for $P_{mt} = 0$. $\rho_{g\varepsilon}$ stands for the

[7] Although maximum likelihood is a more efficient estimation method, it may be computationally burdensome (Maddala, 1986, p.224) and the two-step estimation is a valid alternative.

[8] These exclusions are based on the estimation of the wage growth regression for the whole sample of fixed-term contracts (results available upon request).

[9] The model does not include unobserved match-specific heterogeneity since most variables have a lower within-variation than between-variation. In fact, converted matches appear only once in the database and approximately 64% of non-converted matches appear only twice in the sample. On average, each match is observed 1.7 times in the sample.

correlation coefficient between v_{gt} and ε_{mt} and $\rho_{b\varepsilon}$ for the correlation between v_{bt} and ε_{mt}. u_{gt} and u_{bt} are the disturbances with zero mean of the wage growth regression of converted and non-converted matches, respectively. Since we have unbalanced panel data, each match may be observed more than once and, as such, the hypothesis of independence across observations does not hold. Therefore, the variance-covariance matrix of the estimators is estimated taking into account the possible correlation of the error terms within matches by clustering observations at the match level, which simultaneously accounts for the existence of heteroskedasticity.

Given that the independent and dependent variables are always observed, if the match is either converted or not, and that some matches belong to both groups over the time period considered (18,5%), there may be efficiency gains accruing from the joint estimation of both wage growth regressions Maddala (1986, p.227). For this reason we estimate the following regression:

$$\begin{aligned} W_{mt} &= x'_{gt}\beta_g + \Pi'_{gt}\tau_g + D'_{gt}\delta_g + \sigma_g\rho_{g\varepsilon}\frac{\phi(w'_{gt}\widehat{\omega} + \Pi'_{gt}\widehat{\tau} + D'_{gt}\widehat{\delta})}{\Phi(w'_{gt}\widehat{\omega} + \Pi'_{gt}\widehat{\tau} + D'_{gt}\widehat{\delta})} + \\ & x'_{bt}\beta_b + \Pi'_{bt}\tau_b + D'_{bt}\delta_b - \sigma_b\rho_{b\varepsilon}\frac{\phi(w'_{bt}\widehat{\omega} + \Pi'_{bt}\widehat{\tau} + D'_{bt}\widehat{\delta})}{(1 - \Phi(w'_{bt}\widehat{\omega} + \Pi'_{bt}\widehat{\tau} + D'_{bt}\widehat{\delta}))} + u_{mt}, \end{aligned} \tag{2.7}$$

in which W_{mt} is the wage growth of fixed-term matches. All variables indexed by g assume their real values if the match was converted and are replaced by zero otherwise and the variables indexed by b assume their real values if the match was not converted and are replaced by zero otherwise. u_{mt} is the error term with zero mean.

The main parameters of interest are $\sigma_g\rho_{g\varepsilon}$, $\sigma_b\rho_{b\varepsilon}$, δ_g and δ_b. As previously stated, good matches are expected to be associated with a steeper wage growth. Thus, the switch is expected to be endogenous, i.e. the conversion of the fixed-term contract and the subsequent wage growth should be statistically correlated. It is also expected that good matches are less penalised by reforms that widen the employment protection gap between fixed-term and open-ended contracts if a learning process about match quality is in motion. In short, according to the hypothesis under analysis, it is expected that $\rho_{g\varepsilon} \neq 0$; $\rho_{b\varepsilon} \neq 0$ and $\delta_g > \delta_b$; $\delta_g, \delta_b < 0$.

2.4.1 *Quadros de Pessoal*

The analysis is based on *Quadros de Pessoal*, a Portuguese linked employer-employee database collected every year in October by the Ministry of Employment. *Quadros de Pessoal* is an exceptionally rich database suitable for develop the proposed analysis for several reasons. Firstly, it has a broad coverage and representativeness of the population since it is mandatory for all private firms with at least one wage-earner to provide information about the firm and all their employees. Secondly, given that the information is reported by the firm and is publicly available the measurement error of some variables (such as wages) is minimized. Thirdly, we can follow firms and workers over the years and easily identify the employer-employee matches, which are both assigned with a unique identification code.

This unique labour market database contains very detailed information on the worker, such as gender, age, tenure, education, skills, nationality, occupation, wages (base wage, overtime pay, regular and irregular benefits) and hours worked. Information about the contract type has been available since 2002. Firms are characterised in terms of their location, dimension, main economic activity, age and turnover.

The unit of observation is defined as the worker-firm match, observed from 2003 until 2009[10]. After correcting the time inconsistency in some variables such as education and gender (Cardoso, 2004), the data was filtered according to the following criteria (e.g. Cardoso et al., 2012). We only considered full-time workers with an open-ended or a fixed-term contract, aged between 18 and 65 years old, who earn more than 80% of the legal minimum wage each year[11] and less than 100.000 euros (at 2009 prices) and work less than 400 hours per month. Moreover, we excluded individuals employed in agriculture or fishery, firms operating abroad and International Organisations.

From this sample of workers, we restrict the analysis to all matches holding a

[10] We only considered data up to 2009 to avoid capturing the impact of the economic and financial crisis.

[11] This boundary corresponds to the minimum wage allowed for apprentices.

fixed-term contract in a certain year t-1 that were continued in t and either remained on a fixed-term contract or were converted into an open-ended contract. As a double check, we only considered fixed-term contracts with tenure at time t-1[12] lower than three years in 2003 and six years in the remaining years, in accordance with the legislation in force[13]. Finally, observations below the 2nd and above the 99th percentile of the wage growth distribution were excluded. After the exclusion of the missing data on relevant variables, we end up with an unbalanced panel of 702,242 different matches observed over a 7-year period, which corresponds to a total of 1,174,269 observations.

The worker's real wage is computed on an hourly basis and corresponds to the sum of the monthly base wage, regular benefits and overtime pay divided by the total hours worked (normal and overtime). The wage growth was calculated as the subtraction of the logarithms of real hourly wage over two consecutive years and is measured as a percentage. Real variables were computed using the Consumer Price Index (2012=100) and the business cycle is accounted for by the introduction of the annual unemployment rate reported by *Instituto Nacional de Estatística*. A brief description of the remaining variables is presented in Appendix A.

2.4.2 Descriptive Statistics

Table B.1 reports some summary statistics of the sample. Between 2003 and 2009, an average of 22.8% of fixed-term contracts were converted into open-ended contracts.

In the sample of fixed-term contracts, the average age of workers is 34 years, 45% are females and almost 8% are immigrants, although there is a higher share of non-native workers among non-converted fixed-term contracts. Workers on converted contracts are, on average, better educated than workers with non-converted fixed-term contracts. The former are also less concentrated in unqualified occupations (11%) than the latter group of workers (14%). The larger share of fixed-term contracts is observed in the services sector and in firms located in Lisbon and in the

[12]Note that firms report information annually in October. Thus, for the purposes of accuracy the exclusion is made using lagged tenure.

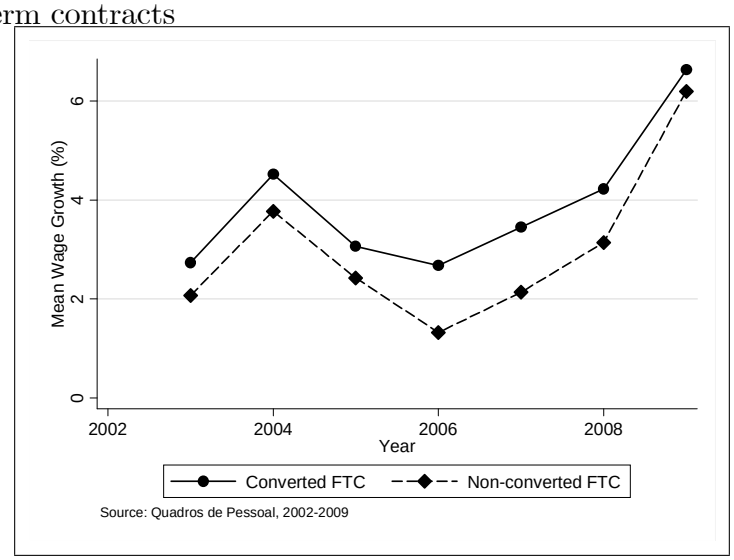
[13]While the 2004 change in legislation applies to all fixed-term contracts, the change introduced in 2009 only applies to newly created fixed-term contracts.

North region. Non-converted fixed-term contracts are found more in activity sectors and regions strongly affected by seasonality, such as construction and Algarve. It can be seen that more than 60% of workers on converted fixed-term contracts were converted in the two first years of tenure, while almost 50% of workers with non-converted fixed-term contracts have only one year of tenure. Although most fixed-term contracts are concentrated in firms with less than 100 employees, converted contracts are more represented in larger firms, notably in firms with more than 400 employees. Finally, on average, workers with converted fixed-term contracts receive higher raw hourly wages and experience higher wage growth, although there is not a significant difference in the supply of overtime hours between both types of contract.

Similarly to Mertens and McGinnity (2003), we compare the wage and the wage growth distributions of fixed-term contracts with the distributions in a sample of open-ended contracts. In line with their findings, although a greater proportion of fixed-term contracts is found in the lowest deciles of the wage distribution (Table B.2), they are also over-represented in both the lowest and the highest wage growth deciles, with nearly 25% of fixed-term contracts concentrated in the two highest wage growth deciles vs. 19% of open-ended contracts (Table B.2).

This preliminary evidence may indicate that an underlying learning process about match quality is associated with fixed-term contracts, which may be expressed by their conversion into open-ended contracts and their wage growth pattern. Figure 2·1 shows that the wage growth of converted fixed-term contracts is always higher than that of non-converted fixed-term contracts from 2003 until 2009, but the gap between them increased from 2005 until 2008, i.e. the period the change in the legislation was in force.

Figure 2·1: Hourly wage growth of converted and non-converted fixed-term contracts



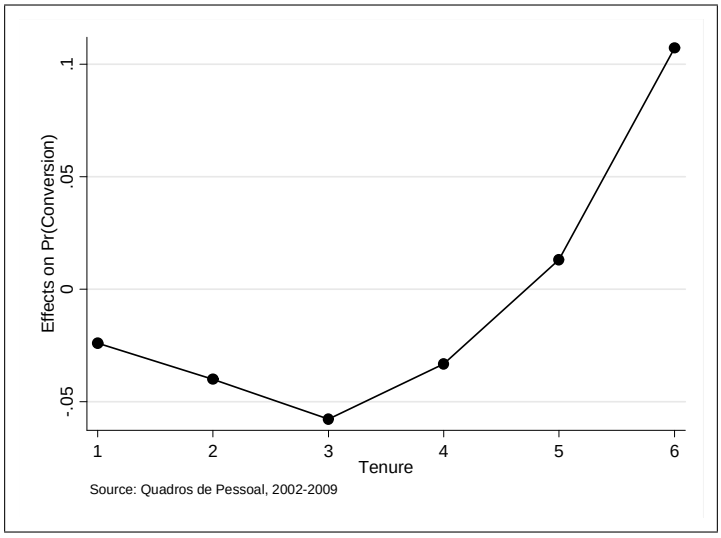
2.5 Empirical Results

2.5.1 Determinants of the Conversion of Fixed-term to Open-ended Contracts

In line with Boeri (2011) and Dolado et al. (2012), the results in Table 2.1 show that the change in Portuguese EPL that relaxed the regulations on fixed-term contracts had a negative and statistically significant impact, at a 99% confidence level, on the probability of a fixed-term contract being converted into an open-ended contract. In the years in which the change in legislation was in force, the probability of conversion was 3 percentage points lower (average marginal effect), ceteris paribus. Female fixed-term workers seem to be slightly more penalised by this type of reform than males, since the probability of conversion between 2004 and 2008 was 3.2 pp. lower for females and only 2.9 pp. lower for males (Table 2.1, columns (5) and (4), respectively). The results in Figure 2.2, based on the estimates in Table 2.1, column (2), indicate that this negative effect can be partly explained by the fact that the conversion of the contract during this period may have been postponed, especially at the end of the third year of the contract (-5.8 pp.). In fact, when the interaction between tenure dummies and the legislation dummy is considered, the average marginal effect of the change in legislation on the probability of conversion

is negative and statistically significant at a 5% significance level[14] in the first four years of contract.

Figure 2·2: Average marginal effect of the change in legislation at years of tenure



Tenure has a statistically significant and an inverse U-shaped impact on the probability of transition to an open-ended contract, increasing up to three years and decreasing thereafter; this is consistent with the evidence reported by Portugal and Varejão (2005) for Portugal and Güell and Petrongolo (2007) for Spain. This may indicate that, on average, the first years are crucial for firms and workers to assess the quality of the match.

As Bowlus (1995) argues, match quality is significantly affected by the business cycle and its behaviour depends on two opposite effects. During recessions, the increasing number of unemployed workers available to fill fewer job vacancies (congestion effect) negatively affects match quality despite the larger pool of available workers for firms to screen (agglomeration effect). Similarly to Bowlus (1995), we find evidence of a procyclical behaviour of match quality, proxied by the probability

[14] Standard error of all marginal effects are available upon request.

Table 2.1: Determinants of the conversion of fixed-term into permanent contracts

	(1)		(2)		(3)		(4)		(5)	
	Whole Sample		Whole Sample		Whole Sample		Males		Females	
VARIABLES	Estimates	AME	Estimates	AME	Estimates	AME	Estimates	AME	Estimates	AME
immigrant	-0.0994***	-0.0269	-0.0988***	-0.0267	-0.0994***	-0.0269	-0.115***	-0.0308	-0.0789***	-0.0215
	(0.00551)		(0.00551)		(0.00551)		(0.00702)		(0.00899)	
female	-0.0231***	-0.0064	-0.0233***	-0.0065	-0.0231***	-0.0064				
	(0.00299)		(0.00299)		(0.00299)					
4 years school	0.0604***	0.0156	0.0598***	0.0155	0.0604***	0.0156	0.0775***	0.0201	0.0343	0.0088
	(0.0136)		(0.0136)		(0.0136)		(0.0172)		(0.0223)	
6 years school	0.0729***	0.0189	0.0728***	0.0189	0.0729***	0.0189	0.0880***	0.0229	0.0529**	0.0137
	(0.0136)		(0.0136)		(0.0136)		(0.0172)		(0.0223)	
9 years school	0.123***	0.0327	0.123***	0.0326	0.123***	0.0327	0.123***	0.0325	0.128***	0.0340
	(0.0136)		(0.0136)		(0.0136)		(0.0172)		(0.0222)	
12 years school	0.180***	0.0487	0.180***	0.0486	0.180***	0.0487	0.178***	0.0479	0.183***	0.0497
	(0.0137)		(0.0137)		(0.0137)		(0.0175)		(0.0224)	
bachelor	0.188***	0.0511	0.188***	0.0509	0.188***	0.0511	0.180***	0.0484	0.190***	0.0517
	(0.0162)		(0.0162)		(0.0162)		(0.0215)		(0.0253)	
university	0.256***	0.0711	0.256***	0.0709	0.256***	0.0711	0.255***	0.0706	0.255***	0.0708
	(0.0145)		(0.0145)		(0.0145)		(0.0189)		(0.0233)	
tenure_1	-0.532***	-0.1466	-0.225***	-0.1539	-0.532***	-0.1466	-0.548***	-0.1515	-0.520***	-0.1425
	(0.00816)		(0.0153)		(0.00816)		(0.0113)		(0.0118)	
tenure_2	-0.0984***	-0.0313	0.223***	-0.0387	-0.0984***	-0.0313	-0.141***	-0.0447	-0.0518***	-0.0165
	(0.00816)		(0.0153)		(0.00816)		(0.0113)		(0.0118)	
tenure_3	0.174***	0.0592	0.521***	0.0524	0.174***	0.0592	0.187***	0.0641	0.156***	0.0524
	(0.00835)		(0.0158)		(0.00835)		(0.0116)		(0.0121)	
tenure_4	-0.0790***	-0.0253	0.226***	-0.0327	-0.0790***	-0.0253	-0.101***	-0.0325	-0.0537***	-0.0171
	(0.00909)		(0.0174)		(0.00909)		(0.0126)		(0.0131)	
tenure_5	-0.126***	-0.0397	0.0616***	-0.0503	-0.126***	-0.0397	-0.148***	-0.0469	-0.101***	-0.0317
	(0.0103)		(0.0208)		(0.0103)		(0.0143)		(0.0148)	
$\text{tenure}_1 \times leg$			-0.446***							
			(0.0179)							
$\text{tenure}_2 \times leg$			-0.467***							
			(0.0179)							
$\text{tenure}_3 \times leg$			-0.502***							
			(0.0185)							
$\text{tenure}_4 \times leg$			-0.444***							
			(0.0203)							
$\text{tenure}_5 \times leg$			-0.296***							
			(0.0240)							
managers	0.116***	0.0326	0.116***	0.0325	0.116***	0.0326	0.124***	0.0346	0.114***	0.0320
	(0.0134)		(0.0134)		(0.0134)		(0.0171)		(0.0221)	
experts	-0.0427***	-0.0113	-0.0423***	-0.0112	-0.0427***	-0.0113	0.00281	0.0007	-0.0747***	-0.0197
	(0.00740)		(0.00739)		(0.00740)		(0.0109)		(0.0103)	
technicians	0.0125**	0.0034	0.0123**	0.0033	0.0125**	0.0034	0.0294***	0.0079	-0.0108	-0.0029
	(0.00612)		(0.00611)		(0.00612)		(0.00798)		(0.0101)	
admin staff	0.0646***	0.0178	0.0645***	0.0177	0.0646***	0.0178	0.0732***	0.0200	0.0540***	0.0149
	(0.00570)		(0.00569)		(0.00570)		(0.00846)		(0.00804)	
salespeople	0.128***	0.0361	0.129***	0.0363	0.128***	0.0361	0.136***	0.0378	0.122***	0.0345
	(0.00499)		(0.00498)		(0.00499)		(0.00769)		(0.00672)	

(continued)

(continuation)

	(1) Whole Sample		(2) Whole Sample		(3) Whole Sample		(4) Males		(5) Females	
VARIABLES	Estimates	AME	Estimates	AME	Estimates	AME	Estimates	AME	Estimates	AME
craftsmen	0.0645***	0.0177	0.0641***	0.0176	0.0645***	0.0177	0.0693***	0.0189	0.0403***	0.0111
	(0.00556)		(0.00555)		(0.00556)		(0.00691)		(0.0116)	
mach operators	0.0188***	0.0051	0.0186***	0.0050	0.0188***	0.0051	0.0322***	0.0087	-0.0152	-0.0041
	(0.00569)		(0.00569)		(0.00569)		(0.00731)		(0.0109)	
age	0.00154	-0.0004	0.00160	-0.0004	0.00154	-0.0004	0.00433***	-0.0004	-0.00209	-0.0003
	(0.000984)		(0.000983)		(0.000984)		(0.00128)		(0.00156)	
agesq	-4.27e-05***		-4.35e-05***		-4.27e-05***		-8.34e-05***		1.36e-05	
	(1.30e-05)		(1.30e-05)		(1.30e-05)		(1.67e-05)		(2.10e-05)	
dimension_1	0.0213***	0.0056	0.0214***	0.0056	0.0213***	0.0056	0.0366***	0.0096	0.00140	0.0004
	(0.00522)		(0.00522)		(0.00522)		(0.00707)		(0.00777)	
dimension_2	0.0342***	0.0090	0.0341***	0.0090	0.0342***	0.0090	0.0567***	0.0150	0.00839	0.0022
	(0.00403)		(0.00403)		(0.00403)		(0.00554)		(0.00592)	
dimension_3	0.189***	0.0525	0.188***	0.0523	0.189***	0.0525	0.149***	0.0409	0.240***	0.0682
	(0.00443)		(0.00443)		(0.00443)		(0.00603)		(0.00660)	
dimension_4	0.376***	0.1114	0.375***	0.1110	0.376***	0.1114	0.313***	0.0906	0.457***	0.1389
	(0.00481)		(0.00480)		(0.00481)		(0.00655)		(0.00718)	
legislation	-0.106***	-0.0300	0.340***	-0.0297	0.290***	-0.0321	-0.103***	-0.0288	-0.111***	-0.0315
	(0.00423)		(0.0175)		(0.0920)		(0.00574)		(0.00628)	
unemrate	-0.0736***	-0.0205	-0.0665***	-0.0185	-0.0736***	-0.0309	-0.0777***	-0.0215	-0.0693***	-0.0194
	(0.00165)		(0.00168)		(0.00165)		(0.00223)		(0.00246)	
public capital	0.000897***	0.0002	0.000878***	0.0002	0.000897***	0.0002	0.00132***	0.0004	0.000368***	0.0001
	(8.72e-05)		(8.71e-05)		(8.72e-05)		(0.000125)		(0.000123)	
foreign capital	0.00146***	0.0004	0.00147***	0.0004	0.00146***	0.0004	0.00154***	0.0004	0.00133***	0.0004
	(4.58e-05)		(4.58e-05)		(4.58e-05)		(6.11e-05)		(6.97e-05)	
proportion_{t-1}	-0.00607***	-0.0017	-0.00608***	-0.0017	-0.00607***	-0.0017	-0.00617***	-0.0017	-0.00591***	-0.0017
	(5.18e-05)		(5.17e-05)		(5.18e-05)		(7.01e-05)		(7.73e-05)	
leg $\times unemrate$					-0.0521***					
					(0.0118)					
Constant	0.121***		-0.242***		0.121***		0.119***		0.108**	
	(0.0278)		(0.0311)		(0.0278)		(0.0368)		(0.0427)	
Region dummies	yes		yes		yes		yes		yes	
Sector dummies	yes		yes		yes		yes		yes	
Year dummies	yes		yes		yes		yes		yes	
Observations	1,174,269		1,174,269		1,174,269		642,813		531,456	
ll	-581255		-580836		-581255		-316559		-263983	
Pseudo-R-squared	0.0781		0.0788		0.0781		0.0763		0.0827	

Source: Quadros de Pessoal, 2002-2009. Notes: Probit regression with standard errors clustered in nmatch in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Base categories are: gender(male), education(<4 years of schooling), tenure(6/7 years), occupation(unqualified), dimension(<11), region(north), sector(services), year(y03,y08,y09). AME stands for Average Marginal Effects

of conversion. Fixed-term matches are less likely to be converted in periods of higher unemployment rates and the probability of conversion decreases by 2.1 pp. if the unemployment rate increases by 1 pp. (Table 2.1, column (1)), which may be explained by the firms' need for some downwards flexibility and to avoid high firing costs at times of economic distress, which is consistent with Varejão and Portugal's (2007) findings. Moreover, as Güell and Petrongolo (2007) predict, when unemployment increases, firms are less willing to convert fixed-contracts into open-ended contracts since workers are less likely to quit due to the worsening of outside opportunities.

Considering that policy makers tend to implement this type of reform when unemployment is rising (Saint-Paul, 1996), they may exacerbate the business cycle's negative impact on the probability of converting the contract. This is supported by the results presented in column (3) of Table 2.1, where the coefficient associated with the interaction term between the regime dummy reflecting the change in legislation and the current unemployment rate (leg x unemrate) is negative and statistically significant. Thus, in the years in which the legislation that widened the employment protection gap between open-ended and fixed-term contracts was in force, there was an increase in the adverse marginal effect of the current unemployment rate on the probability of conversion (from -2.2 pp. to -3.4 pp.)[15]. Although the direct impact of the change in legislation is positive and statistically significant at a 1% significance level when this interaction is considered, its overall marginal effect is still negative and statistically significant (-3.2 pp.).

Regarding workers' characteristics, the contracts of younger[16], male and more educated workers are more likely to be converted to permanent contracts. For example, workers with a university degree are 7.1 pp. more likely to have a contract converted into a more stable employment relationship than a worker with less than four years of schooling, ceteris paribus. There is also some evidence of discrimination against immigrant workers, especially in the case of male workers (on average, male immigrant workers are 3.1 pp. less likely to receive an open-ended contract, ceteris paribus).

[15] This result is robust to the use of alternative measures of business cycle, such as the unemployment rate at the start of the match.

[16] The average marginal effect of age is statistically significant at standard significance levels.

There is a greater tendency for fixed-term workers in management and sales occupations to be given an open-ended contract than workers performing unskilled tasks. This result was expected as fixed-term contracts are probably used less as a screening device for occupations requiring lower skill levels due to cost (Sicilian, 1995). Workers with fixed-term contracts matched either with smaller firms or firms with a larger percentage of fixed-term contracts have a slightly smaller probability of being given an open-ended contract. In fact, firms with a higher number of fixed-term contracts would be expected to have a greater need for flexibility and, thus, be less willing to change a fixed-term contract to a permanent one.

In Subsection 2.5.2 we distinguish between the wage growth of converted and non-converted fixed-term contracts and assess the impact of the change in legislation for both groups.

2.5.2 Wage growth of Converted and Non-converted Fixed-term Contracts

Assuming that firms and workers are not able to identify the true value of the match *ex-ante*, it is plausible that some matches start with fixed-term contracts and at a low wage level. However, as (Sicilian, 1995) argues, wage growth should reflect updated expectations of match quality. Therefore, while bad matches are terminated or remain with temporary contracts, good matches initiated with fixed-term contracts should experience higher wage growth and become a more stable employment relationship.

From the estimated coefficients associated with the inverse mills ratio (Table 2.2, column (1)) we can conclude that the error term of the selection equation and the error term of the wage growth regression for converted fixed-term contracts are positively and significantly correlated at a 99% confidence level, which supports the need to correct for the sample selection bias. Accordingly, unobserved factors increase the likelihood of a fixed-term contract being converted into an open-ended contract and lead to an above average wage growth. These results are in line with Sicilian (1995) and Loh's (1994) predictions, since there seems to be a non-negligible increase in wage growth associated with the conversion to a permanent contract that we estim-

ate to be equal to approximately 1.3 pp.[17]. Nevertheless, we find that workers on non-converted fixed-term contracts do not experience either a significantly lower or a higher wage growth than a random fixed-term worker would. It seems that the wage is only renegotiated when the contract is converted, which may be the result of the higher bargaining power gained on the conversion of the contract when the worker starts benefiting from higher employment protection levels. These results may also reflect the learning process about match quality associated with the use of fixed-term contracts or could be the result of the worker's integration in the firm's internal labour market.

Similarly to what we observe for the probability of conversion, the change in legislation also has a statistically significant and negative impact on the wage growth of fixed-term contracts. However, our findings indicate that not all fixed-term contracts are penalised evenly by the change in legislation. Although the change has a negative impact on the wage growth of both non-converted and converted fixed-term contracts (-0.55 pp. and -0.16 pp., respectively), the effect is statistically significant at a 1% significance level for the former group while only statistically significant at a 5% level for converted contracts. Besides there is evidence, at a 1% significance level, that the penalisation suffered by non-converted fixed-term contracts was greater than that of converted fixed-term contracts[18]. The renegotiation of wages between 2004 and 2008 may have been postponed as it was easier for firms to use fixed-term contracts for a longer period of time. Females in non-converted matches seem to be more affected by this type of change in legislation since they experience a significant decline in wage growth of approximately 0.71 pp. in the years the change was in force. The negative impact of the change in legislation on the wage growth of male workers does not seem to differ according to match quality[19] although it is only statistically significant for converted fixed-term contracts at a 95% confidence level.

It seems that the change in legislation affects the wage growth path of fixed-term contracts directly and indirectly through the link between contract's conversion and

[17] Evaluated at the sample mean inverse mills ratio

[18] The p-value of the Wald test of the equality of coefficients equals 0.0000.

[19] The p-value of the Wald test of the equality of coefficients equals to 0.2061.

Table 2.2: Determinants of the wage growth of non-converted and converted fixed-term contracts

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	0.305	1.050***	0.0438	0.556*	1.947***	0.0231	-0.183	-0.0451	0.7542
	(0.205)	(0.284)		(0.323)	(0.473)		(0.254)	(0.334)	
immigrant	0.138***	0.110	0.8114	0.156***	0.0589	0.5413	0.155**	0.124	0.8537
	(0.0443)	(0.105)		(0.0598)	(0.144)		(0.0654)	(0.152)	
female	-0.103***	-0.0537	0.3829						
	(0.0243)	(0.0500)							
9 years school	0.134***	0.347***	0.0034	0.232***	0.484***	0.0107	-0.0527	0.0600	0.2904
	(0.0295)	(0.0652)		(0.0408)	(0.0880)		(0.0414)	(0.0959)	
12 years school	0.531***	0.981***	0.0000	0.653***	1.138***	0.0001	0.338***	0.662***	0.0097
	(0.0363)	(0.0768)		(0.0514)	(0.106)		(0.0505)	(0.112)	
bachelor	1.132***	2.025***	0.0000	1.306***	2.335***	0.0002	0.869***	1.522***	0.0108
	(0.0870)	(0.163)		(0.132)	(0.240)		(0.116)	(0.223)	
university	1.138***	2.378***	0.0000	1.390***	3.069***	0.0000	0.908***	1.681***	0.0000
	(0.0585)	(0.117)		(0.0904)	(0.176)		(0.0774)	(0.160)	
$tenure_1$	0.377***	0.113	0.1710	0.323***	-0.375	0.0253	0.495***	0.768***	0.2521
	(0.0721)	(0.178)		(0.110)	(0.289)		(0.0940)	(0.218)	
$tenure_2$	0.346***	0.644***	0.0318	0.218**	0.282	0.7648	0.518***	1.075***	0.0019
	(0.0626)	(0.125)		(0.0924)	(0.195)		(0.0842)	(0.159)	
$tenure_3$	0.106	0.287**	0.2085	0.125	0.229	0.6325	0.0965	0.486***	0.0377
	(0.0693)	(0.125)		(0.104)	(0.189)		(0.0918)	(0.162)	
$tenure_4$	-0.140**	-0.0912	0.7567	-0.176*	-0.311	0.5716	-0.0890	0.185	0.1815
	(0.0705)	(0.141)		(0.104)	(0.215)		(0.0951)	(0.182)	
$tenure_5$	0.0675	-0.120	0.2963	-0.121	-0.466*	0.2054	0.282***	0.287	0.9829
	(0.0814)	(0.162)		(0.120)	(0.247)		(0.110)	(0.209)	
managers	0.679***	2.390***	0.0000	0.571***	2.288***	0.0000	1.048***	2.830***	0.0001
	(0.121)	(0.249)		(0.157)	(0.317)		(0.193)	(0.420)	
experts	0.542***	1.055***	0.0003	0.860***	1.479***	0.0050	0.497***	0.988***	0.0084
	(0.0610)	(0.126)		(0.0983)	(0.194)		(0.0776)	(0.166)	
technicians	0.851***	1.499***	0.0000	0.900***	1.593***	0.0000	0.962***	1.729***	0.0001
	(0.0505)	(0.108)		(0.0681)	(0.149)		(0.0794)	(0.169)	
admin staff	0.650***	1.301***	0.0000	0.593***	1.458***	0.0000	0.857***	1.453***	0.0000
	(0.0442)	(0.0934)		(0.0722)	(0.149)		(0.0565)	(0.122)	
salespeople	0.151***	0.607***	0.0000	-0.132**	0.211	0.0358	0.409***	0.953***	0.0000
	(0.0375)	(0.0853)		(0.0655)	(0.146)		(0.0443)	(0.103)	

(continued)

(continuation)

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
craftsmen	0.00460	0.447***	0.0000	0.134**	0.723***	0.0000	-0.293***	-0.346**	0.7841
	(0.0414)	(0.0934)		(0.0537)	(0.125)		(0.0757)	(0.171)	
mach. operators	-0.467***	-0.0117	0.0000	-0.323***	0.359***	0.0000	-0.427***	-0.524***	0.5689
	(0.0429)	(0.0939)		(0.0589)	(0.132)		(0.0690)	(0.152)	
age	-0.224***	-0.233***	0.5117	-0.287***	-0.316***	0.1514	-0.164***	-0.141***	0.2617
	(0.00714)	(0.0134)		(0.00995)	(0.0186)		(0.0102)	(0.0192)	
agesq	0.00211***	0.00212***	0.9716	0.00272***	0.00296***	0.3633	0.00164***	0.00123***	0.1261
	(9.21e-05)	(0.000177)		(0.000127)	(0.000241)		(0.000133)	(0.000258)	
dimension_1	0.367***	0.379***	0.9071	0.310***	0.291**	0.8862	0.476***	0.555***	0.5814
	(0.0371)	(0.0901)		(0.0520)	(0.123)		(0.0521)	(0.132)	
dimension_2	0.430***	0.395***	0.6387	0.458***	0.326***	0.1865	0.415***	0.553***	0.1923
	(0.0278)	(0.0666)		(0.0391)	(0.0903)		(0.0391)	(0.0967)	
dimension_3	0.269***	0.207***	0.3899	0.234***	0.0852	0.1405	0.320***	0.296***	0.8135
	(0.0332)	(0.0624)		(0.0456)	(0.0884)		(0.0481)	(0.0878)	
legislation	-0.547***	-0.159**	0.0000	-0.419***	-0.264**	0.2061	-0.711***	-0.0276	0.0000
	(0.0387)	(0.0759)		(0.0547)	(0.110)		(0.0540)	(0.104)	
unemrate	1.293***	1.164***	0.0001	1.175***	0.998***	0.0005	1.425***	1.343***	0.0569
	(0.0152)	(0.0305)		(0.0219)	(0.0473)		(0.0211)	(0.0399)	
public capital	-0.00103	0.0134***	0.0000	0.00438***	0.0167***	0.0000	-0.00655***	0.0100***	0.0000
	(0.000792)	(0.00130)		(0.00117)	(0.00190)		(0.00107)	(0.00178)	
foreign capital	0.00595***	0.00105	0.0000	0.00410***	0.000419	0.0123	0.00856***	0.00279**	0.0001
	(0.000515)	(0.000854)		(0.000716)	(0.00123)		(0.000742)	(0.00119)	
proportion_{t-1}	-0.00213***	-0.0177***	0.0000	-0.00370***	-0.0230***	0.0000	0.000679	-0.0103***	0.0000
	(0.000690)	(0.00183)		(0.00105)	(0.00286)		(0.000885)	(0.00233)	
Constant	-2.037***			0.193			-4.766***		
	(0.232)			(0.342)			(0.310)		
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,174,269			642,813			531,456		
Adjusted R-squared	0.033			0.030			0.041		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. 4 years school, 6 years school, dimension_4 and electricity excluded for identification purposes. Base categories are: gender(male), education(< 9 years school), tenure(6/7 years), occupation(unqualified), dimension(<11, >400), region(north), sector(electricity and services), year(y03,y08, y09). Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

wage growth (IMR x legislation[20], Table 2.3). Specifically, in the years the legislation was in force, this link was weakened for both types of contract, especially for non-converted fixed-term contracts. When this interaction is considered, the direct impact of the change in legislation on the wage growth of non-converted matches remains negative and statistically significant, but for converted fixed-term contracts it becomes positive although not statistically significant for males. In fact, the indirect penalisation of the change in legislation on wage growth is especially relevant to explain the negative but non-significant association between the probability of conversion and the subsequent wage growth observed for females with a converted contract (Table 2.2, column (3)). In the years the legislation was not in force, females with converted fixed-term contracts experienced a statistically significant (at a 10% significance level) increase in wage growth of approximately 0.78 pp.

The results also indicate that human capital variables have different returns for converted and non-converted fixed-term contracts (Table 2.2, column (1)). For both types of match, the returns on education are increasing, but they are always higher (at a 1% significance level) for converted fixed-term contracts, especially for levels of higher education. For example, for converted fixed-term contracts, a worker with a university degree experience a 2.4 pp. higher wage growth than a worker with less than nine years of schooling, while for non-converted matches this increase is only equal to 1.1 pp., ceteris paribus.

Moreover, as Amuedo-Dorantes and Serrano-Padial (2007) argue, the duration of the contract plays an important role in the explanation of the wage growth path and the evidence gathered shows that the moment at which the contract is converted has important implications. Workers only experience a significant higher wage growth than that at the end of the contract if their contracts are converted in the second or third year of tenure, while they face a wage growth penalisation of approximately 0.14 pp. in the fourth year of tenure if the contract is not converted.

The effects of workers' idiosyncratic characteristics, such as nationality, age and gender, are not statistically different in converted and non-converted matches at

[20] This interaction intends to assess if the link between contract's conversion and wage growth is different in the period in which the change in legislation was in force, as suggested in Semykina and Wooldridge (2010) for example.

Table 2.3: Direct and indirect impact of the change in legislation on the wage growth of non-converted and converted fixed-term contracts

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	1.363***	1.326***	0.9278	1.607***	1.941***	0.6104	1.098***	0.635*	0.3411
	(0.228)	(0.302)		(0.357)	(0.494)		(0.287)	(0.362)	
IMR x legislation	-1.743***	-0.594***	0.0000	-1.643***	-0.226	0.0000	-2.151***	-1.166***	0.0017
	(0.156)	(0.165)		(0.222)	(0.234)		(0.214)	(0.225)	
legislation	-1.175***	0.544***	0.0000	-1.001***	0.00634	0.0012	-1.504***	1.323***	0.0000
	(0.0687)	(0.208)		(0.0968)	(0.299)		(0.0949)	(0.282)	
Constant	-1.652***			0.624*			-4.344***		
	(0.237)			(0.351)			(0.317)		
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,174,269			642,813			531,456		
Adjusted R-squared	0.033			0.030			0.041		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. 4 years school, 6 years school, dimension_4 and electricity excluded for identification purposes. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

standard significance levels (Table 2.2). Ceteris paribus, apart from contract conversion, older workers experience lower wage growth and the rate at which the wage growth decreases slows with age up to about 53 and 55 years for non-converted and converted fixed-term contracts, respectively. Although native and female workers with non-converted fixed-term contracts experience lower wage growth rates on average, the wage growth rate of converted fixed-term contracts does not seem to be significantly affected by gender or nationality.

The highest wage growth rate is experienced by managers when the contract is converted and by technicians when the contract is not converted. It is also worth noting that machine operators experience lower wage growth than unqualified workers if the contract is not converted (-0.47 pp.) and they experience neither a statistically significant higher or a lower wage growth if the contract is converted; this may be due to the use of fixed-term contracts to screen matches for this occupation which requires specific training.

Contrary to what is reported for the conversion probability, the wage growth of fixed-term contracts seems to be countercyclical; this can be explained by the fact that during recessions firms separate from a higher share of less-educated and, thus, low-wage workers, while maintaining the employment relationship with high-wage earners.

2.5.3 Wage Growth Differential Between Non-converted and Converted Fixed-term Contracts

The aim of this subsection is to identify the main sources of the wage growth differential between non-converted and converted fixed-term contracts. To that end, we adopt a threefold decomposition initially proposed by Winsborough and Dickenson (1971). We start by decomposing the mean wage growth differential ($\overline{W}_{bt} - \overline{W}_{gt}$) into endowment and coefficient effects, and the interaction of both using the estimates of equations 3 and 4 without correcting for selectivity:

$$\begin{aligned}\overline{W}_{bt} - \overline{W}_{gt} &= [(\overline{x}_{bt} - \overline{x}_{gt})'\widehat{\beta}_g + (\overline{\Pi}_{bt} - \overline{\Pi}_{gt})'\widehat{\tau}_g + (\overline{D}_{bt} - \overline{D}_{gt})'\widehat{\delta}_g] + \\ &\quad [\overline{x}'_{gt}(\widehat{\beta}_b - \widehat{\beta}_g) + \overline{\Pi}'_{gt}(\widehat{\tau}_b - \widehat{\tau}_g) + \overline{D}'_{gt}(\widehat{\delta}_b - \widehat{\delta}_g)] + \\ &\quad [(\overline{x}_{bt} - \overline{x}_{gt})'(\widehat{\beta}_b - \widehat{\beta}_g) + (\overline{\Pi}_{bt} - \overline{\Pi}_{gt})'(\widehat{\tau}_b - \widehat{\tau}_g) + (\overline{D}_{bt} - \overline{D}_{gt})'(\widehat{\delta}_b - \widehat{\delta}_g)] \ ,\end{aligned} \tag{2.8}$$

where the first term in square brackets on the right hand side of the equation is the endowment effect, i.e., the part of the differential due to differences in the characteristics between non-converted and converted fixed-term matches. The second term in square brackets is the coefficient effect and corresponds to the part of the differential due to differences in the remuneration of characteristics between both groups. Finally, the third term corresponds to the interaction between the endowment and coefficient effects. These effects are computed using converted fixed-term contracts as the reference group. The results of the estimated differential are in Table 2.4.

On average, workers in good matches experience a higher wage growth than workers on non-converted fixed-term contracts, which is in line with our initial predictions. The observed average wage growth associated with good matches is equal to 4.02%, while workers with non-converted fixed-term contracts experience an average wage growth of 3.20% between 2003 and 2009. Thus, the mean wage growth differential between non-converted and converted fixed-term contracts is equal to -0.81 pp., which is statistically significant at standard significance levels.

The results show that more than 91% of this differential is due to differences in the remuneration of characteristics between both types of match. Good worker-firm matches not only appear to be better rewarded for their characteristics but also to have better endowments. Both effects and their interaction are statistically significant at standard significance levels.

However, as shown in the previous subsection, it is important to account for the selectivity bias arising from the correlation between the conversion of the contract and the subsequent wage growth. To do so we decompose the mean selectivity corrected wage growth differential (equations 5 and 6) by adapting Reimers's (1983) methodology and estimating the contribution of the selectivity effect equals to: $-\widehat{\sigma}_b\widehat{\rho}_{b\varepsilon}\frac{\phi(w'_{mt}\widehat{\omega}+\Pi'_t\widehat{\tau}+D'_t\widehat{\delta})}{(1-\Phi(w'_{mt}\widehat{\omega}+\Pi'_t\widehat{\tau}+D'_t\widehat{\delta}))} - \widehat{\sigma}_g\widehat{\rho}_{g\varepsilon}\frac{\phi(w'_{mt}\widehat{\omega}+\Pi'_t\widehat{\tau}+D'_t\widehat{\delta})}{\Phi(w'_{mt}\widehat{\omega}+\Pi'_t\widehat{\tau}+D'_t\widehat{\delta})}$. According to the results in Table 2.4, the selectivity effect is statistically significant and widens the wage growth differ-

Table 2.4: Threefold decomposition of the wage growth differential

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	3.203	4.016	-0.813***	-0.247***	-0.742***	0.175***	no
contribution (%)				-30.32%	-91.17%	21.49%	
legislation				0.000	-0.144***	0.000*	
contribution (%)				0.01%	-17.72%	0.05%	
Overall	3.203	4.016	-0.813***	-0.418***	0.255	0.346***	-0.997***
contribution (%)				-51.39%	31.38%	42.55%	-122.54%
legislation				0.000	-0.121***	0.000*	
contribution (%)				0.02%	-14.86%	0.04%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

ential. When this effect is considered, the contribution of the characteristics (endowment effect) to the wage growth differential increases to over 51% and the coefficient effect no longer contributes significantly to explaining the wage growth gap. If converted fixed-term matches had the characteristics of non-converted matches, they would experience a decrease of approximately 0.42 pp. in their wage growth rate, ceteris paribus.

Finally, we focus on the specific contribution of the variable accounting for the change in legislation to the wage growth differential. As expected, the endowment effect of the change in legislation does not significantly contribute to the wage growth differential since the reform applies to all fixed-term contracts. Instead, the change in legislation contributes to increase the wage growth differential through the coefficient effect ($\overline{D}'_{gt}(\widehat{\delta}_b - \widehat{\delta}_g)$). We estimate that almost 18% of the gap in the sample period is attributed to the way both groups were affected by the reform that eased the regulations on fixed-term contracts. This contribution slightly drops to approximately 15% when the selectivity effect is taken into account.

2.5.4 Robustness Analysis

The first robustness check consists of assessing the sensitivity of the results to different wage definitions and we therefore re-estimate the second stage of the model and the threefold decomposition using alternative and stricter wage definitions.

In Table B.3, we present results in which overtime pay is excluded and the wage is defined as the sum of base wages and regular benefits. The results seem to be robust to this alternative wage definition since not only do good matches experience an

increase in wage growth of approximately 1.2 pp. at the time of conversion but, when the change in legislation was in force, converted fixed-term contracts seem to have experienced a non-significant and lower wage growth penalisation (-0.007 pp.) than non-converted fixed-term contracts (-0.47 pp.)[21]. When the overtime pay component is excluded from the wage definition, this wage growth penalisation associated with the increase in the protection gap between the two types of contract is slightly lower, which may indicate that workers are also penalised by the payment for these hours or may reduce the amount of overtime hours worked when their employment protection level decreases. The results of the wage growth decomposition are also robust: the selectivity effect and the coefficient effect of the change in legislation contribute to increase the wage growth differential between non-converted and converted fixed-term contracts and are statistically significant at a 5% and 1% significance levels, respectively (Table B.4).

Further, we repeat the analysis considering only the growth of the hourly base wages (Table B.5). We find that the results are not robust to this wage definition, since the wage growth penalty associated with the change in legislation does not seem to differ significantly between converted and non-converted fixed-term contracts[22]. However, this stricter wage definition still allows us to conclude that workers in good matches are rewarded with a higher growth rate of base wages, especially female workers. We also find that the selectivity effect and the change in legislation no longer contribute significantly to explain the wage growth differential between both types of matches (Table B.6). This result is not surprising since base wages are more restricted by institutions such as minimum wage and collective bargaining and, thus, there is less scope to the effect of the change in legislation to be heterogeneous according to match quality.

Since the construction sector is strongly influenced by seasonality and where the share of non-converted fixed-term contracts is higher than that of converted fixed-term contracts, the sensitivity of the results to the exclusion of this sector should also be assessed (Table B.7 and Table B.8). The results were quite similar to

[21] The p-value of the Wald test of the equality of coefficients is equal to 0.0000.
[22] The p-value of the Wald test of the equality of coefficients is equal to 0.5580.

those discussed in Subsection 2.5.2, with the exception that when construction is not included in the estimation not only do workers experience an increase in wage growth when the contract is converted (1.56 pp.), but also a wage growth penalisation when the contract is not converted (-0.26 pp.); both are statistically significant at a 1% significance level. Workers in good matches are significantly less affected by the change in legislation than workers with non-converted contracts, although its impact becomes slightly more negative for both groups. The contribution of the selectivity effect and the change in legislation to the wage growth differential remains negative and statistically significant at standard levels (Table B.9).

The employment protection gap between fixed-term and open-ended contracts increased more in firms with 11 to 20 employees, since the 2004 change in legislation also increased the employment protection of open-ended contracts for this cohort (Centeno and Novo (2012) and Centeno and Novo (2014) study the impact of this change in excess worker turnover and wages, respectively). By excluding these firms from the sample, we find that the results discussed in the previous subsections are robust and not exclusively driven by them (Table B.10, Table B.11 and Table B.12).

Finally, the 2004 Labour Code revision introducing the change in legislation under study also introduced a penalisation in the social security contribution for firms that have more than 15% of the total employees on a fixed-term contract with more than four years' duration. After converting the contract to a permanent one, the firm can benefit from a reduction in the social security contribution. Since firms with a higher proportion of these contracts may also have had an incentive to convert them, in Table B.13, we present the results of the Probit model estimation considering the interaction between the one period lagged value of the proportion of fixed-term contracts and the legislation dummy. In firms with a higher proportion of fixed-term contracts, the probability of converting the contract was less penalised by the legislation that facilitated their use. However, this effect is negligible which may indicate that this type of measure promoting the conversion of fixed-term contracts is less effective when accompanied by measures increasing the flexibility on their use. The impact of the change in legislation on the wage growth of non-converted contracts remains statistically significant at standard levels and is more negative in

firms with a higher proportion of fixed-term contracts (Table B.14). As expected, in these firms the legislation contributes more to increase the wage growth differential between both types of match (Table B.15).

2.6 Conclusions and Policy Implications

Fixed-term contracts can play different roles in the labour market and therefore be unequally affected by asymmetric reforms that increase the employment protection gap between fixed-term and open-ended contracts. Our results show that it is relevant to consider match quality to assess how a reform that facilitates the use of fixed-term contracts affects their wage growth.

By estimating an endogenous regime switching model, we find that the 2004 change in the Portuguese EPL that eased the regulations on fixed-term contracts had a negative impact on match quality, measured by the probability of conversion of fixed-term contracts. However, we find evidence that not all fixed-term contracts are evenly affected by this type of reform. Not only is the conversion of the contract associated with a non-negligible increase in wage growth, but the wage growth experienced by workers in good matches, i.e., with converted fixed-term contracts, also seems to be less penalised by the asymmetric reform. In fact, in the years when the change in legislation was in force, workers on converted fixed-term contracts seem to have experienced a lower wage growth penalisation (-0.16 pp.), than those on non-converted fixed-term contracts (-0.55 pp.). Moreover, the change in legislation also had an indirect negative impact on the wage growth of both types of match, especially for non-converted fixed-term contracts, through the link between the conversion of the contract and the wage growth; this draws attention to the potential negative externalities of this type of employment protection reform. On average, we find that the change in legislation contributed to increase the wage growth differential between non-converted and converted fixed-term contracts in approximately 15%, ceteris paribus.

This paper aims to contribute to the ongoing discussion about the role of fixed-term contracts in the labour market and the impact of reforms that ease regulations

on their use. We argue that the burden of the adjustment of this type of reform is not spread homogeneously among workers on fixed-term contracts. On average, less than one fourth of fixed-term contracts are converted in open-ended contracts in the Portuguese labour market and employment protection reforms that facilitate their use generate potential inefficiencies by penalising and delaying the access of workers on fixed-term contracts to a more stable employment relationship. This may entail negative effects on labour productivity and human capital acquisition, since workers on this type of contract experience higher turnover rates (Centeno and Novo, 2012) and participate less in training activities (Booth et al., 2002) than workers on open-ended contracts. Our results also show that this type of reform contributes to increase the wage inequality between workers on converted fixed-term contracts and those who were not able to exit temporary employment. Tackling labour market segmentation may help to reduce inequality among workers. The future research agenda should assess the impact of the introduction of a single contract with increasing severance payments (Bentolila, Dolado and Jimeno, 2012), which could contribute to increase employment duration and decrease unemployment (Pérez and Osuna, 2014). Futher research also needs to be conducted in order to conclude about the impact of employment protection reforms, namely on employment level and non-pecuniary aspects of the employment relationship, such as the likelihood of promotion to a higher occupational level within the firm. Indeed, conversion to an open-ended contract may also be associated with access to career ladders, which would further amplify the negative impact of asymmetric reforms that increase the employment protection gap between fixed-term and open-ended contracts.

Appendix A

Description of Variables

Worker's characteristics:

- Nationality: 1 dummy variable- immigrant (1 if immigrant and 0 if native),
- Gender: 1 dummy variable- female (1 if female and 0 if male),
- Education: 7 dummy variables- less than 4 years of schooling; 4 years of schooling; 6 years of schooling; 9 years of schooling; 12 years of schooling; Bachelor degree and University education,
- Age: continuous variable measured in years,
- Tenure: 7 dummy variables- $tenure_1$ (1 year), $tenure_2$ (2 years), $tenure_3$ (3 years), $tenure_4$ (4 years), $tenure_5$ (5 years), $tenure_6$ (6/7 years),
- Occupation (Portuguese Classification of Occupations 2010): 8 dummy variables- managers, experts, technicians, administrative staff, salespeople, craftsmen, plant and machine operators, unqualified workers.

Firm's characteristics:

- Dimension: 5 dummy variables- $dimension_0$ (1-10 employees), $dimension_1$ (11-20 employees), $dimension_2$ (21-100 employees), $dimension_3$ (101-400 employees), $dimension_4$ (>400 employees),
- Region: 7 dummy variables- North, Lisbon, Algarve, Centre, Alentejo, Azores, Madeira,
- Sector of activity: 6 dummy variables- extractive industries, manufacturing, electricity production and distribution, construction, public administration, services,

- Share of fixed-term contracts: 1 continuous lagged variable (proportion$_{t-1}$) in percentage of total number of employees,
- Capital Ownership: 2 continuous variables- share of foreign capital in percentage and share of public capital in percentage.

Appendix B

Tables

Table B.1: Descriptive statistics

Variables	Non-converted FTC	Converted FTC	Whole Sample
Female (%)	45.04	46.00	45.26
Immigrant (%)	8.15	5.98	7.65
age (years)	34.10	33.03	33.86
	(9.83)	(9.27)	(9.71)
Education (%)			
<= 1st cycle	17.51	13.85	16.67
2nd cycle	19.69	17.14	19.11
3rd cycle	26.05	25.74	25.98
secondary education	23.32	26.62	24.07
bachelor degree	2.33	2.79	2.43
college	11.11	13.84	11.73
Tenure (%)			
1	47.30	28.99	43.13
2	26.43	33.74	28.10
3	12.60	22.97	14.96
4	7.13	7.72	7.26
5	3.87	3.63	3.82
6	1.94	1.92	1.94
7	0.73	1.03	0.80
Occupation (%)			
Managers	1.20	1.34	1.23
Experts	8.68	9.52	8.87
Intermediate-level technicians	10.59	11.05	10.70
Administrative staff	13.86	16.29	14.42
Sellers	22.59	24.99	23.14
Craftsmen	17.08	14.26	16.44
Plant and Machine Operators	12.22	11.71	12.10
Unqualified workers	13.77	10.84	13.10
Sector of Activity (%)			
Extractive Industries	0.38	0.36	0.38
Manufacturing	20.04	21.06	20.28
Electricity	0.40	0.54	0.43
Construction	14.12	9.59	13.08
Public Administration	2.22	0.37	1.80
Services	62.84	68.09	64.03
Region (%)			
North	28.61	27.42	28.34
Lisbon	36.36	42.12	37.68
Algarve	6.39	4.27	5.90
Alentejo	4.78	4.22	4.65
Centre	19.31	17.11	18.81
Azores	1.70	1.85	1.74
Madeira	2.85	3.01	2.88
Firm's Dimension (%)			
<=10	26.47	17.56	24.44
11 to 20	12.04	9.43	11.45
21 to 100	31.21	27.25	30.31
101 to 400	17.52	21.68	18.47
>=401	12.76	24.07	15.34
real wage (log)	1.50	1.59	1.52
	(0.41)	(0.43)	(0.42)
wage growth (%)	3.20	4.02	3.39
	(11.10)	(11.92)	(11.29)
overtime (hours)	2.09	2.33	2.15
	(9.26)	(9.66)	(9.36)
Observations	906,442	267,827	1,174,269

Source: Quadros de Pessoal, 2002-2009. Notes: FTC stands for fixed-term contract. Standard deviations in parentheses.

Table B.2: Distribution of open-ended and fixed-term contracts by wage and wage growth decile (%)

Wage Decile	OEC	FTC
1	9.48	12.74
2	9.43	13.04
3	9.32	13.55
4	9.55	12.40
5	9.74	11.38
6	9.89	10.58
7	10.11	9.44
8	10.21	8.91
9	10.83	5.62
10	11.45	2.33
Wage Growth Decile	OEC	FTC
1	9.84	10.85
2	10.93	9.20
3	9.52	8.43
4	10.18	9.07
5	10.32	8.33
6	10.19	8.97
7	10.10	9.50
8	9.86	10.73
9	9.80	11.08
10	9.27	13.84
Observations	6,211,944	1,174,269

Source: Quadros de Pessoal, 2002-2009. Notes: OEC stands for open-ended contract and FTC stands for fixed-term contract (converted and non-converted).

Table B.3: Determinants of the growth rate of base wages and regular benefits of converted and non-converted fixed-term contracts

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	0.260	0.997***	0.0437	0.614*	2.066***	0.0161	-0.294	-0.271	0.9575
	(0.202)	(0.281)		(0.317)	(0.467)		(0.251)	(0.331)	
legislation	-0.465***	-0.00721	0.0000	-0.320***	-0.0831	0.0488	-0.653***	0.0882	0.0000
	(0.0382)	(0.0752)		(0.0538)	(0.109)		(0.0537)	(0.103)	
Constant	-2.178***			-0.0712			-4.763***		
	(0.229)			(0.338)			(0.308)		
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,174,269			642,813			531,456		
Adjusted R-squared	0.035			0.032			0.042		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the growth rate of base wages plus regular benefits. 4 years school, 6 years school, dimension_4 and electricity excluded for identification purposes. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

Table B.4: Threefold decomposition of the wage (base wages and regular benefits) growth differential

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	3.212	4.033	-0.821***	-0.254***	-0.745***	0.178***	no
contribution (%)				-30.91%	-90.78%	21.68%	
legislation				-0.000	-0.167***	0.000*	
contribution (%)				0.00%	-20.31%	0.05%	
Overall	3.212	4.033	-0.821***	-0.409***	0.129	0.343***	-0.883**
contribution (%)				-49.85%	15.65%	41.71%	-107.52%
legislation				-0.000	-0.144***	0.000*	
contribution (%)				0.00%	-17.56%	0.05%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.5: Determinants of the growth rate of base wages of converted and non-converted fixed-term contracts

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	-0.352	2.295***	0.0000	0.971***	2.301***	0.1044	-1.395***	2.666***	0.0000
	(0.255)	(0.441)		(0.358)	(0.615)		(0.346)	(0.593)	
legislation	-0.964***	-1.009***	0.5580	-0.803***	-0.734***	0.5157	-1.185***	-1.361***	0.1002
	(0.0305)	(0.0678)		(0.0422)	(0.0946)		(0.0436)	(0.0953)	
Constant	-4.395***			-1.374***			-7.614***		
	(0.235)			(0.329)			(0.320)		
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,174,269			642,813			531,456		
Adjusted R-squared	0.068			0.063			0.080		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. Dependent variable is the growth rate of base wages. Immigrant, 4 years school, 6 years school, 9 years school and lisbon excluded for identification purposes. The control variables included are female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

Table B.6: Threefold decomposition of the base wage growth differential

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	2.909	3.788	-0.879***	-0.319***	-0.792***	0.231***	no
contribution (%)				-36.23%	-90.08%	26.31%	
legislation				0.001*	-0.034	0.000	
contribution (%)				0.09%	-3.85%	0.01%	
Overall	2.909	3.788	-0.879***	-0.389***	-0.838	0.427***	-0.078
contribution (%)				-44.29%	-95.35%	48.55%	-8.91%
legislation				0.001*	-0.007	0.000	
contribution (%)				0.10%	-0.79%	0.00%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.7: Impact of the change in legislation on probability of conversion of the contract without firms in construction sector

VARIABLES	Whole Sample Estimates	Males Estimates	Females Estimates
legislation	-0.103***	-0.0957***	-0.109***
	(0.00451)	(0.00640)	(0.00636)
Constant	0.0941***	0.0643	0.107**
	(0.0305)	(0.0431)	(0.0432)
Region dummies	yes	yes	yes
Sector dummies	yes	yes	yes
Year dummies	yes	yes	yes
Observations	1,020,643	502,932	517,711
ll	-514529	-256471	-257611
Pseudo-R-squared	0.0799	0.0778	0.0834

Source: Quadros de Pessoal, 2002-2009. Notes: Probit regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, industry dummies and year dummies.

Table B.8: Determinants of the wage growth of converted and non-converted fixed-term contracts without firms in construction sector

	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	0.689***	1.302***	0.0732	1.121***	2.267***	0.0320	-0.0709	0.0279	0.8211
	(0.195)	(0.261)		(0.301)	(0.409)		(0.253)	(0.331)	
legislation	-0.646***	-0.246***	0.0000	-0.574***	-0.419***	0.2303	-0.720***	-0.0340	0.0000
	(0.0412)	(0.0779)		(0.0618)	(0.115)		(0.0546)	(0.105)	
Constant		-1.713***			1.158***			-4.760***	
		(0.237)			(0.357)			(0.312)	
Region dummies		yes			yes			yes	
Sector dummies		yes			yes			yes	
Year dummies		yes			yes			yes	
Observations		1,020,643			502,932			517,711	
Adjusted R-squared		0.035			0.032			0.041	

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. 4 years school, 6 years school, dimension_4 and electricity excluded for identification purposes. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

Table B.9: Threefold decomposition of the wage growth differential without firms in construction sector

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	3.241	4.082	-0.841***	-0.229***	-0.795***	0.183***	no
contribution (%)				-27.21%	-94.53%	21.74%	
legislation				0.000	-0.154***	0.000	
contribution (%)				0.01%	-18.32%	0.02%	
Overall	3.241	4.082	-0.841***	-0.484***	0.900***	0.369***	-1.625***
contribution (%)				-57.61%	107.04%	43.83%	-193.26%
legislation				0.000	-0.131***	0.000	
contribution (%)				0.01%	-15.57%	0.02%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.10: Impact of the change in legislation on probability of conversion of the contract without firms with 11-20 employees

VARIABLES	Whole Sample Estimates	Males Estimates	Females Estimates
legislation	-0.103***	-0.0989***	-0.107***
	(0.00448)	(0.00609)	(0.00663)
Constant	0.167***	0.178***	0.140***
	(0.0296)	(0.0393)	(0.0453)
Region dummies	yes	yes	yes
Sector dummies	yes	yes	yes
Year dummies	yes	yes	yes
Observations	1,039,830	566,851	472,979
ll	-517749	-280464	-236574
Pseudo-R-squared	0.0834	0.0823	0.0872

Source: Quadros de Pessoal, 2002-2009. Notes: Probit regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, industry dummies and year dummies.

Table B.11: Determinants of the wage growth of converted and non-converted fixed-term contracts without firms with 11-20 employees

	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	0.229	1.090***	0.0206	0.334	1.937***	0.0105	-0.0568	0.114	0.6949
	(0.206)	(0.289)		(0.327)	(0.492)		(0.253)	(0.333)	
legislation	-0.571***	-0.176**	0.0000	-0.467***	-0.311***	0.2161	-0.706***	-0.0219	0.0000
	(0.0408)	(0.0783)		(0.0579)	(0.114)		(0.0569)	(0.107)	
Constant		-1.914***			0.269			-4.554***	
		(0.243)			(0.361)			(0.324)	
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,039,830			566,851			472,979		
Adjusted R-squared	0.034			0.031			0.042		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. 4 years school, 6 years school, 9 years school and dimension_4 excluded for identification purposes. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

Table B.12: Threefold decomposition of the wage growth differential without firms with 11-20 employees

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	3.195	4.019	-0.824***	-0.290***	-0.743***	0.209***	no
contribution (%)				-35.18%	-90.19%	25.37%	
legislation				0.000	-0.141***	0.000*	
contribution (%)				0.02%	-17.11%	0.04%	
Overall	3.195	4.019	-0.824***	-0.483***	0.240	0.418***	-0.999***
contribution (%)				-58.65%	29.15%	50.80%	-121.30%
legislation				0.000	-0.118***	0.000	
contribution (%)				0.02%	-14.28%	0.04%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table B.13: Impact of the change in legislation on probability of conversion of the contract for firms with a higher proportion of fixed-term contracts

	Whole Sample		Males		Females	
VARIABLES	Estimates	AME	Estimates	AME	Estimates	AME
legislation	-0.122***	-0.0301	-0.131***	-0.0290	-0.114***	-0.0315
	(0.00659)		(0.00890)		(0.00983)	
proportion_{t-1}	-0.00630***	-0.0017	-0.00659***	-0.0017	-0.00595***	-0.0017
	(9.11e-05)		(0.000125)		(0.000134)	
$\text{proportion}_{t-1} \times leg$	0.000320***		0.000588***		6.18e-05	
	(0.000104)		(0.000143)		(0.000153)	
Constant	0.132***		0.139***		0.110**	
	(0.0280)		(0.0371)		(0.0429)	
Region dummies	yes		yes		yes	
Sector dummies	yes		yes		yes	
Year dummies	yes		yes		yes	
Observations	1,174,269		642,813		531,456	
ll	-581250		-316550		-263983	
Pseudo-R-squared	0.0781		0.0763		0.0827	

Source: Quadros de Pessoal, 2002-2009. Notes: Probit regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, industry dummies and year dummies. AME stands for Average Marginal Effects

Table B.14: Impact of the change in legislation on wage growth for firms with a higher proportion of fixed-term contracts

	(1)			(2)			(3)		
	Whole Sample		Wald Tests	Males		Wald Tests	Females		Wald Tests
VARIABLES	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value	Non-converted FTC	Converted FTC	p-value
inverse mills ratio	0.319	1.076***	0.0403	0.561*	1.953***	0.0230	-0.162	-0.0246	0.7551
	(0.205)	(0.284)		(0.323)	(0.474)		(0.254)	(0.334)	
legislation	-0.200***	-0.103	0.4251	-0.125	-0.118	0.9688	-0.362***	-0.154	0.2190
	(0.0595)	(0.108)		(0.0847)	(0.155)		(0.0822)	(0.150)	
proportion_{t-1}	0.00234**	-0.0169***	0.0000	0.000125	-0.0205***	0.0000	0.00516***	-0.0124***	0.0000
	(0.000922)	(0.00228)		(0.00137)	(0.00354)		(0.00122)	(0.00298)	
$\text{proportion}_{t-1} \times leg$	-0.00629***	-0.00123	0.0115	-0.00535***	-0.00343	0.4982	-0.00632***	0.00294	0.0009
	(0.000850)	(0.00182)		(0.00121)	(0.00257)		(0.00117)	(0.00255)	
Constant	-2.208***			0.0223			-4.893***		
	(0.233)			(0.344)			(0.312)		
Region dummies	yes			yes			yes		
Sector dummies	yes			yes			yes		
Year dummies	yes			yes			yes		
Observations	1,174,269			642,813			531,456		
Adjusted R-squared	0.033			0.030			0.041		

Source: Quadros de Pessoal, 2002-2009. Notes: GLS regression with standard errors clustered in nmatch in parantheses. *** p<0.01, ** p<0.05, * p<0.1. 4 years school, 6 years school, dimension_4 and electricity excluded for identification purposes. The control variables included are immigrant, female, education dummies, tenure dummies, occupation dummies, age, agesq, firm's dimension dummies, region dummies, unemployment rate, capital ownership, lagged share of FTC, sector dummies and year dummies. Wald test of equality of the coefficients between converted and non-converted fixed-term contracts.

Table B.15: Threefold decomposition of the wage growth differential without firms in construction sector

Sample	E[Wb|p=0]	E[Wg|p=1]	Differential	Endowments	Coefficients	Interaction	Selectivity
Overall	3.203	4.016	-0.813***	-0.247***	-0.742***	0.175***	no
contribution (%)				-30.31%	-91.25%	21.56%	
legislation				0.000	-0.028	0.000	
contribution (%)				0.01%	-3.44%	0.01%	
proportion_{t-1}				-0.134***	0.531***	0.136***	
contribution (%)				-16.45%	65.33%	16.76%	
$\text{proportion}_{t-1} \times leg$				-0.003	-0.089***	-0.022***	
contribution (%)				-0.41%	-10.95%	-2.76%	
Overall	3.203	4.016	-0.813***	-0.419***	0.249	0.350***	-0.994***
contribution (%)				-51.49%	30.61%	43.03%	-122.15%
legislation				0.000	-0.001	0.000	
contribution (%)				0.02%	-0.15%	0.00%	
proportion_{t-1}				-0.186***	0.735***	0.189***	
contribution (%)				-22.82%	90.36%	23.18%	
$\text{proportion}_{t-1} \times leg$				-0.003	-0.092***	-0.023***	
contribution (%)				-0.33%	-11.29%	-2.85%	

Source: Quadros de Pessoal, 2002-2009. Notes: Threefold decomposition with normalized results and converted fixed-term contracts as the reference group. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Chapter 3

Job Flows and Flexibility at the Margin

3.1 Introduction

One of the main priorities in the European agenda is to stimulate job creation in order to achieve the Europe 2020 target of 75% employment among the working age population. Since the 1980s, many European countries have attempted to foster job creation and fight unemployment by boosting labour market flexibility and especially by easing the regulations on fixed-term contracts. This type of policy helped increase temporary employment (OECD, 2004), especially in labour markets characterised by stringent employment protection on open-ended contracts, notably in Portugal, Spain and France (Dolado et al., 2002; Bentolila and Saint-Paul, 1992; Blanchard and Landier, 2002), thus, contributing to greater labour market segmentation.

Other effects are reported in the literature. Extending the use of fixed-term contracts in labour markets characterised by high employment protection levels for open-ended contracts may not contribute to lowering unemployment but may in fact have perverse effects (Blanchard and Landier, 2002; Cahuc and Postel-Vinay, 2002; Bentolila, Cahuc, Dolado and Barbanchon, 2012). Also, this type of change in legislation is shown to increase employment volatility because it has more marked effects on the dynamics than on the stock of employment and unemployment (Boeri, 2011). Therefore, focusing only on the aggregate unemployment rate can mask very different dynamics in terms of job and worker flows (Portugal and Blanchard, 2001). Although this type of reforms is expected to boost both job creation and job destruction, the literature offers no clear prediction about which effect dominates over the other.

In this paper, our aim is to assess the impact of a change to the legislation in Portugal in 2004 that extended the maximum legal duration of fixed-term contracts

from three to six years. This is an important reform as it targeted both new and existing fixed-term contracts but its effects are yet to be analysed. Portugal is an interesting case study since it has the most stringent employment protection on regular contracts of all OECD countries and only mild employment protection on fixed-term contracts. Given the findings of previous studies referred above, we focus our analysis on the effects of this change in legislation on job flows. The impact on the job flows of fixed-term contracts is expected to be greater than on open-ended contracts, not only due to the target of the change in legislation but also because of the higher adjustment costs of open-ended contracts; we therefore assess the effect of fostering flexibility at the margin on job flows by contract type. In addition to providing a better understanding of the adjustment process that follows a change in EPL, the impact of flexibility at the margin is not underestimated because the focus is not only on aggregate job flows.

Empirical evidence shows that job creation and job destruction occur simultaneously exhibiting a persistent behaviour in a one-year period (Davis and Haltiwanger, 1992*a*) and that flexibility at the margin is promoted when unemployment is rising (Saint-Paul, 1996). A flexible econometric methodology must be used that can account for the fact that flexibility in the use of fixed-term contracts may cause and also be caused by job flows. A multivariate dynamic framework, such as the reduced-form vector autoregression (VAR) model, is suitable to study the impact of this EPL reform on job flows rates. Moreover, a VAR model is known to be a valid approach to account for the heterogeneity between firms' characteristics and their interdependences (Haltiwanger et al., 2013) without imposing strong restrictions (Canova and Ciccarelli, 2013). Finally, this empirical strategy makes it possible to formally test for the exogeneity of changes in the flexibility in the use of fixed-term contracts.

In this paper, we present an empirical within-country analysis of the impact of an asymmetric labour market reform that facilitates the use of fixed-term contracts on job creation and destruction by type of contract. To the best of our knowledge, this is the first empirical study of the impact of this change in legislation on job flows. Our analysis focuses on the period 2003-2011 and uses a rich longitudinal

database containing information on every private employer and their employees, thus allowing the calculation of job flows by type of contract, and the region, sector of activity, size and age of the firm. Our study contributes to the literature in three ways: i) it is conducted at a country level, therefore avoiding the criticisms of cross-country studies due to comparability of variables and diversity of institutional settings including EPL itself; ii) it uses disaggregated data, which allow us to capture the heterogeneity among firms and to draw a more detailed picture of job flows; iii) it uses a new index of labour market flexibility at the margin, which has the advantage of overcoming some drawbacks of the OECD EPL index and of capturing the degree of enforcement of the changes in legislation.

The results show that in the period of analysis, an increase in the flexibility at the margin Granger causes an increase in the destruction of fixed-term jobs and a larger and more lasting decrease in the creation of fixed-term jobs; meanwhile it indirectly promotes a substitution of open-ended contracts with fixed-term contracts. Our main findings show that during this period, in which the main change in legislation was to extend the maximum duration of fixed-term contracts, increased flexibility fosters the maintenance of existing fixed-term contracts rather than the creation of new ones and that there is a recomposition of job flows. These results are mainly explained by the proportion of non-converted fixed-term contracts and more associated with sectors with negative employment growth, since we conclude that the promotion of flexibility at the margin does not seem to significantly affect job flows in expanding sectors.

The paper is organised as follows. In the next section, we briefly describe the EPL in Portugal, focusing in particular on the legislation on fixed-term contracts. Section 3 provides a literature review of the effects of flexibility at the margin. In Section 4 we present the database and the empirical methodology. The results are described and discussed in Section 5 and the conclusions drawn in Section 6.

3.2 The Portuguese Labour Market

According to the OECD EPL index, since 1985 Portugal has had the most stringent empoyment protection legislation on regular contracts in the OECD and despite the

implementation of policies promoting overall flexibility, employment protection for regular contracts is still high. By contrast, Portugal has mild employment protection on fixed-term contracts. Additionally, Portugal has the third highest proportion of temporary contracts in Europe; they represent approximately one fifth of total employment, mainly due to the high proportion of fixed-term contracts.

Since the introduction of fixed-term contracts in 1976, the restrictions on their use have been relaxed substantially. In fact, fixed-term contracts may be contributing to increased job flows at the intensive (firm's expansion and downsizing) and extensive margins (new firms) as firms are able to hire a new worker on a fixed-term contract to substitute an absent employee, to satisfy seasonal or occasional activities, in response to an exceptional increase in firm's activity and to launch a new activity, firm or establishment. Fixed-term contracts can also be used to hire first job seekers or the long-term unemployed. In 1989, it was established that when a fixed-term contract ends without being converted to a permanent contract, the worker is entitled to severance pay. It is, however, much easier and less costly to dismiss workers on fixed-term contracts than on open-ended contracts. Although the financial component of the dismissal costs are only slightly higher for open-ended contracts[1], there are procedural costs involved, from which fixed-term contracts are exempted, such as the notifications of the workers council and unions and the possibility to appeal to courts (Centeno and Novo, 2014).

Between 2003 and 2009, Portugal was the OECD country that most reduced the employment protection strictness (Venn, 2009). It is in this period that one of the most important changes was made facilitating the use of fixed-term contracts, but the legislation on open-ended contracts remained practically unchanged. Between 2004 and 2008, a third extraordinary renewal of fixed-term contracts was permitted so that the contract could have a maximum duration of six instead of only three years. Additionally, thereafter workers could be hired on fixed-term contracts to indirectly substitute an absent employee.

[1]For open-ended contracts, the severance pay is equal to one month of base wage and seniority pay for each year of tenure, while workers on fixed-term contracts are entitled to receive a compensation equal to three (two) days of base wage and seniority pay for each month of tenure if the contract duration is lower (higher) than six months.

From 2009 to 2011, fixed-term contracts could still be renewed three times, but the maximum duration of new contracts was reset to three years. The use of fixed-term contracts was facilitated again in 2012 when fixed-term contracts could be subject to two additional renewals and could last for a maximum duration of 4 and a half years.

Herein, we study the effects of the legislation facilitating the use of fixed-term contracts on job flows in the Portuguese labour market, focusing on the period between 2003 and 2011.

3.3 Employment Protection Legislation and Job Flows

The existing theoretical models give an ambiguous prediction of the effects of EPL on aggregate employment and unemployment, since firms operating in a country with stringent employment protection reduce not only job destruction but also job creation as this allows them to avoid high dismissal costs in case of separation (e.g. Addison and Teixeira, 2003). As stated in Boeri's (2011) survey, most studies find EPL has a more significant impact on unemployment and employment flows than on their stock. Some empirical studies find that higher employment protection levels are associated with lower job flow rates (Gómez-Salvador et al., 2004; Kugler and Pica, 2008; Messina and Vallanti, 2007) while others find no significant impact (Avdagic, 2015).

Since broader employment protection reforms that decrease employment protection levels for regular contracts are politically difficult to implement (Saint-Paul, 1996), most European countries increased labour market flexibility by easing the regulations on fixed-term contracts (OECD, 2004; Boeri, 2011; Kahn, 2010). As Blanchard and Landier (2002) argue, the introduction of fixed-term contracts with lower dismissal costs than permanent contracts may fail to reduce unemployment since it encourages higher job turnover rates for this type of contract. This kind of reform may indeed fail to lower unemployment levels because the higher job creation may not offset the higher job destruction generated by the lower incentives to convert fixed-term into open-ended contracts. This is especially relevant in labour markets characterised by high employment protection levels for permanent contracts. The

introduction of flexibility at the margin tends to boost the proportion of fixed-term contracts in total employment (Eslava et al., 2014), increase employment volatility (Boeri and Garibaldi, 2007; Bentolila and Saint-Paul, 1992) and job turnover of fixed-term contracts (Aguirregabiria and Alonso-Borrego, 2014), and foster the replacement of permanent with fixed-term contracts (Boeri and Garibaldi, 2007; Aguirregabiria and Alonso-Borrego, 2014).

A number of theoretical frameworks have been used to simulate the impact of introducing or easing the use of fixed-term contracts (see for example Dolado et al. (2002) for a review of the most important models). The first strand of studies such as Cahuc and Postel-Vinay (2002), Sala et al. (2012), Dolado et al. (2007) extends the Mortensen and Pissarides's (1994) matching framework with endogenous job creation and destruction by including worker heterogeneity.

The results of Cahuc and Postel-Vinay (2002) support Blanchard and Landier's (2002) prediction since they show that facilitating the use of fixed-term contracts promotes an increase in job creation and job destruction, but the latter effect is stronger in countries with high employment protection levels for permanent contracts. There is evidence that this type of measure has an adverse impact on unemployment as workers on fixed-term contracts suffer from high job turnover when the dismissal of workers on regular contracts implies high costs for firms. However, as Dolado et al. (2007) argue, reducing the firing costs of less productive workers boosts job creation and may be more effective in reducing unemployment than reducing the firing costs of high productivity workers or of all workers simultaneously, although it increases the wage inequality between the two types of workers. Sala et al. (2012) show that stricter rules regarding the renewal or the duration of fixed-term contracts increasing the probability that a fixed-term contract ends may have different effects than increasing the restrictions on hiring workers on temporary contracts. They find that although both types of measure may help reduce the overall job destruction rate in a typical European labour market, the destruction of temporary jobs increases in the former case because firms avoid converting the contract to a permanent one. However, this last effect is overweighed by the lower proportion of fixed-term contracts.

The second strand of studies draw upon a dynamic labour demand model, for example Cabrales and Hopenhayn (1997), Boeri and Garibaldi (2007), Aguirregabiria and Alonso-Borrego (2014), Bentolila and Saint-Paul (1992). Cabrales and Hopenhayn (1997) advocate that the implementation of the 1984 Spanish reform easing the regulations on fixed-term contracts did not help significantly boost the average labour demand although it increased its volatility mainly due to the greater volatility of job creation; on the other hand Aguirregabiria and Alonso-Borrego (2014), find evidence that the introduction of fixed-term contracts increased employment by 3.5%. The transitional dynamics of two-tier reforms are described by Boeri and Garibaldi (2007). These authors argue that there is a honeymoon effect as this type of reform has a positive impact on job creation but job destruction is bounded to the stock of temporary contracts. Therefore, there is a transitory increase in employment, mostly due to the creation of temporary jobs. The substitution of open-ended with fixed-term contracts contributes to these dynamics, which are sluggish but stronger if fixed-term contracts are close substitutes for open-ended contracts (Bentolila and Saint-Paul, 1992).

Aguirregabiria and Alonso-Borrego (2014) argue that a comprehensive reform that halved the firing costs of both types of contract would have a similar positive impact on employment, but would also increase the proportion of permanent employment and have a larger positive impact on labour productivity. In turn, Alonso-Borrego et al. (2005) use a general equilibrium model and find that eliminating fixed-term contracts in the Spanish labour market would lead to a reduction in the unemployment rate and in the average unemployment duration because the job destruction rate would fall and households would exert more search effort because of their preference for permanent jobs. On the other hand, eliminating firing costs would increase the job destruction rate and lower the households' search effort, which would lead to an increase in the unemployment rate but also to labour productivity gains.

The lack of a significant impact on unemployment following the introduction or facilitating of the use of fixed-term contracts may be due to their effect on relative wages of permanent and fixed-term contracts and wage rigidities. Bentolila

and Dolado (1994) predict that a higher proportion of fixed-term contracts in total employment may help boost job creation due to the lower average labour costs, but this effect may be reversed given the lower adjustment costs of fixed-term contracts and if workers on open-ended contracts gain more bargaining power and push their wages up. In fact, empirical evidence shows that the proportion of fixed-term contracts is associated with higher wages for permanent contracts (Jimeno and Toharia, 1993; Bentolila and Dolado, 1994). Moreover, Güell (2000) and Güell and Rodríguez Mora (2010) argue that the introduction of fixed-term contracts may not have a positive impact on the employment level if the minimum wage is too high, since the increase in job creation is not enough to overcome the increase in job destruction if the wages of temporary contracts are not flexible enough. This effect is more marked for low-skilled workers (Güell, 2000; Güell and Rodríguez Mora, 2010).

Empirical evidence shows that employment becomes more responsive to shocks when firms are able to use fixed-term contracts, which suffer the burden of adjustment (Bentolila and Saint-Paul, 1992; Boeri and Garibaldi, 2007; Eslava et al., 2014). For the Italian labour market, Cappellari et al. (2012) estimate that this type of reform causes a small but statistically significant fall in aggregate employment and does not cause an increase in the growth rate of employment with fixed-term contracts. Notwithstanding, the effect of this type of reform is heterogeneous over the business cycle; it is more relevant in recessions than in expansions (Bentolila and Saint-Paul, 1992) and it depends on the role that fixed-term contracts play in the labour market[2]. By calibrating a matching model and assuming that match quality is unobserved *a priori*, Faccini (2014) shows that there may be welfare gains in using fixed-term contracts to learn about match quality and that increasing their share helps reduce unemployment. This occurs because when the proportion of fixed-term contracts increases, the larger creation of temporary jobs outweighs the higher temporary job destruction. According to this author's findings, the increase in the

[2]For example, Amuedo-Dorantes and Malo (2008) argue that temporary contracts are simultaneously used as a buffer stock for permanent contracts and as a screening device. The creation of temporary jobs outweighs that of permanent jobs when establishments expect employment gains in the short-run and employment gains and losses in the long-run. Establishments destroy fewer temporary than permanent jobs if they expect employment gains in the short-run, but destroy more temporary jobs if they expect employment losses in the same time horizon.

maximum duration of fixed-term contracts from two to three years does not produce very meaningful effects, since the screening process is not significantly improved.

In fact, the employment protection gap between fixed-term and open-ended contracts may help explain the divergence in the unemployment pattern between European countries (Bentolila, Cahuc, Dolado and Barbanchon, 2012). For example, according to Bentolila, Cahuc, Dolado and Barbanchon (2012), if Spain had decreased the gap between both types of contract and had adopted the French EPL, unemployment would have increased much less after the 2007 financial crisis.

The literature has so far focused mostly on the Spanish experience and notably on the effects of the 1984 change in legislation that removed restrictions on the circumstances in which firms could hire workers on fixed-term contracts. Several studies find that the proportion of temporary contracts increased after the reform (Dolado et al., 2002; Aguirregabiria and Alonso-Borrego, 2014), while the evidence on the impact on unemployment rate is not clear (e.g. Cabrales and Hopenhayn, 1997). However, fixed-term contracts also play an important role in the employment adjustment process in the Portuguese labour market (Centeno and Novo, 2012; Varejão and Portugal, 2007). The probability of an establishment adjusting its employment increases with the proportion of the workforce on fixed-term contracts and it starts adjusting employment by using fixed-term contracts, especially to make downward adjustments (Varejão and Portugal, 2007). As Centeno and Novo (2012) report, when the employment protection level for open-ended contracts is extended to firms with 11 to 20 employees, the excess worker turnover of fixed-term contracts in these firms increases, but there is no significant effect on the turnover of open-ended contracts. These authors also find a positive impact of that legislation change on the proportion of fixed-term contracts, which signals the high degree of substitution between both types of contract.

3.4 Data and Empirical Approach

3.4.1 *Quadros de Pessoal*

The data source used to compute job flows is *Quadros de Pessoal*, a Portuguese longitudinal database collected every year in October by the Ministry of Employment. This is a very important and detailed data source on the Portuguese labour market and has been used to study job flows by Centeno and Novo (2012) and Carneiro et al. (2014), for example. The survey is mandatory for all firms with at least one employee and therefore the sample is representative of the population under study. On average, it surveys approximately 300,000 firms employing over 2 million workers every year.

Quadros de Pessoal assigns a unique identification code to the firm and all its employees allowing the identification of worker-firm matches. This database has information on the firm's size, region[3], main economic activity, year of constitution, turnover and capital ownership. *Quadros de Pessoal* also provides information at the establishment level, namely on region, sector of activity and size. As we do not have information on the age of establishments, which is reported in the literature as a crucial variable to explain the heterogeneity in job flows (Davis and Haltiwanger, 1999; Haltiwanger et al., 2013), we compute job flows using the firm as the unit of observation. Finally, very detailed information is also available on each firm's employee, notably nationality, gender, age, education, skills, occupation, wage, hours worked and contract type[4].

The unit of observation j in this study is a narrow sector resulting from the crossing of 14 sectors of activity, 5 regions, 6 age and 9 size groups[5] at the firm level (as in Gómez-Salvador et al., 2004; Fuchs and Weyh, 2010). This allows us to construct panel data and to capture the heterogeneity in job flows across firms with different characteristics. A firm's entry is defined as the year in which the firm

[3]Since 2010 we have only had information for mainland Portugal and Madeira. As such, we restrict the whole analysis to mainland Portugal.

[4]Information on contract type is only available from 2002.

[5]The size group is defined using the current firm size, which equals the employment average between t and t-1. This methodology circumvents the size distribution fallacy described in Davis et al. (1996).

first reports a positive number of employees and a firm's exit is identified as the last year the firm reports a positive employment level[6]. We have an unbalanced panel of narrow sectors $j = 1, ..., 2315$[7] observed in $t = 2003, ..., 2011$. We pool all observations for each narrow sector and obtain a micro panel with a total of 13,856 observations.

3.4.2 Measurement of Job Flows

We define and measure job flows following Davis and Haltiwanger (1999). The job creation at time t (JC_t) is defined as the sum of the positive employment changes ($\Delta E > 0$) that occurred in all firms i expanding or entering the market between t-1 and t, while job destruction at time t (JD_t) corresponds to the sum of the negative employment changes ($\Delta E < 0$) in all firms i that contract or exit the market between t-1 and t.

In order to translate that measure into growth rates, JC_t and JD_t are divided by a size measure corresponding to the simple average of employment level between t-1 and t[8].

The job creation and destruction rates of group j, at time t result from the size-weighted sum of job creation and destruction growth rates of firms i belonging to j. Therefore, denoting the employment level of firm i belonging to group j at time t by $E_{i,j,t}$ and the size of group j at time t as $Z_{j,t} = 0,5(E_{j,t} + E_{j,t-1})$, the job creation and destruction rates for each group j are given by:

$$JCR_{j,t} = \sum_{\forall i} \frac{\Delta E_{i,j,t}}{Z_{j,t}} \; if \; \Delta E_{i,j,t} > 0 \tag{3.1}$$

$$JDR_{j,t} = \sum_{\forall i} \frac{|\Delta E_{i,j,t}|}{Z_{j,t}} \; if \; \Delta E_{i,j,t} < 0 \tag{3.2}$$

[6] The computations for aggregate job flows are based on the 2000-2011 sample in order to avoid overestimating the contribution of entries and exits to job flows. However, since information on contract type is only available since 2002, this is the start year to compute job flows by contract type.

[7] Note that we consider less than 3780 (14x5x6x9) narrow sectors because some of them are not observed for more than one time period.

[8] Note that this procedure has the advantage of accommodating firms' entries and exits.

Employment growth for each j is measured as the ratio between the employment change reported by the firm between t and t-1 and the average employment between t and t-1.

We also compute job flow rates by type of contract. The creation rate of fixed-term (permanent) jobs is defined as the ratio between the sum of positive employment changes under fixed-term (permanent) contracts in all firms that increase the number of fixed-term (permanent) contracts and the average fixed-term (permanent) employment between t and t-1. The destruction rate of fixed-term (permanent) jobs is defined as the ratio between the sum of negative fixed-term (permanent) employment changes in all firms that decrease the number of fixed-term (permanent) contracts and the average fixed-term (permanent) employment between t and t-1.

3.4.3 Measurement of Flexibility at the Margin

As stated in Addison and Teixeira's (2003) survey, measuring employment protection level is a challenging task and there are several approaches. Employment protection is conventionally measured by the OECD EPL index (Kahn, 2010; Messina and Vallanti, 2007) and more recently by subjective indexes (Di Tella and MacCulloch, 2005). Although the OECD EPL index captures labour market rigidity and can disentangle the contribution of legislation on regular, fixed-term and temporary agency contracts and collective agreements, it is subject to a few caveats. Namely, it does not account for the enforcement of legislation and may not capture all changes in legislation, since they may not be sufficient to change the country's ranking. The index is constructed at the country-level, but EPL may vary across firms and workers' characteristics such as firm's size and worker's age, respectively (Dolado et al., 2007). Also, a change in legislation easing the regulations on fixed-term contracts may impact across sectors and firms depending on their human resources practices (Portugal and Varejão, 2009). To overcome some of these limitations, Alexandre et al. (2010) propose an overall labour market flexibility index computed for each sector of activity. This index results from multiplying three standardised components following a logistic distribution: the proportion of workers not covered by a collective agreement, the proportion of workers with a part-time job and the

proportion of workers earning higher than the minimum wage.

Following Alexandre et al.'s (2010) methodology, we propose an alternative index of flexibility on the use of fixed-term contracts computed for each narrow sector j which has the advantage of reflecting how firms and employers react to changes in legislation promoting flexibility at the margin. Additionally, it captures asymmetries in flexibility- through the use of fixed-term contracts- between sectors of activity and firms with different sizes and ages and located in different regions.

In constructing this index, we focus on the main dimensions that previous studies have find to be affected by the promotion of flexibility at the margin. Namely, our index contains three dimensions: the proportion of fixed-term contracts, the non-conversion rate of fixed-term to open-ended contracts[9] and the average duration of the fixed-term employment relationship. Empirical evidence shows that countries that increase the employment protection gap between fixed-term and open-ended contracts experience a significant increase in the share of fixed-term contracts (Kahn, 2010; OECD, 2004; Centeno and Novo, 2012), especially those with stricter legislation on open-ended contracts such as the Portuguese case. Making it easier to use fixed-term contracts also places the adjustment burden on workers with this type of contract since they experience higher turnover rates and their contracts are less likely to be converted to open-ended contracts (Boeri, 2011)[10]. Finally, the extension of the maximum legal duration of fixed-term contracts has a direct impact on the average duration of the contract. Note that this index also has the advantage of capturing the links between regulations on both types of contract, since the components of the index are also affected by the stringency of regulations on regular contracts. As in the Centeno and Novo's (2012) study, when the employment protection for regular contracts is extended to firms with 11 to 20 employees, the proportion of fixed-term contracts in these firms and the rate of non-converted fixed-term contracts increase.

We plot the weighted average of these three variables in Figure C.1 in the Ap-

[9] This variable is defined as the share of workers that had a fixed-term contract in $t-1$ and that still have a fixed-term contract in t with the same or a different employer.

[10] In Chapter 2, our results show that when the legislation under study was in force, the probability of conversion of a fixed-term contract fell by approximately 3 pp.

pendix. We can observe an upward trend in the proportion of fixed-term contracts that is only reversed with the onset of the 2009 economic and debt crisis, mostly due to the high destruction rate of fixed-term jobs (see subsection 4.5). The average tenure of workers on fixed-term contracts increased over the sample period, reaching the maximum level in 2009. This is accompanied by an increase in the proportion of non-converted fixed-term contracts. In 2010, however, there was a marked decline in the average length of contract and in the non-conversion rate, which we attribute to the sharp drop in the proportion of fixed-term contracts and to the fact contracts drawn up in 2004 had to end or be converted to open-ended contracts in 2010.

The composite index results from the product of these three components, which follow a functional form F(.):

$$flex_{jt} = [F(\alpha)+F(f_{1,jt})].[F(\beta)+F(f_{2,jt})].[F(\delta)+F(f_{3,jt})],\ j = 1, ..., J \text{ and } t = 1, ..., T, \tag{3.3}$$

where j denotes the narrow sector and t denotes time measured in years. $f_{1,jt}$ is the standardised[11] share of workers on fixed-term contracts in each narrow sector j in period t, $f_{2,jt}$ is the standardised duration of fixed-term contracts in each narrow sector j in period t and $f_{3,jt}$ is the standardised share of workers with non-converted fixed-term contracts in each narrow sector j in period t. The parameters α, β and δ are the correction terms that ensure that the index is bounded and may be interpreted as the relative weight of each component in the index of flexibility at the margin. Although Alexandre et al. (2010) assume that $\alpha = \beta = \delta = 0$, we believe that some components may be more important sources of adjustment than others for labour market agents and, thus, we consider that they can be weighted differently in the index.

In order to estimate α, β and δ, we solve Equation 3.3 where $flex_{jt}$ is proxied by $legislation_{jt}$, which is equal to one if the narrow sector benefited from the 2004 change in legislation that eased regulations on fixed-term contracts, i.e., has at least

[11] Each measure in each year is standardised by subtracting the mean value and dividing by the standard deviation.

one worker, in year t, with a fixed-term contract and tenure of more than three years. Therefore, we use an *a posteriori* measure of the implementation of the most important change in legislation that extended the maximum legal duration of fixed-term contracts in Portugal in the period under analysis and estimate the following model using a Generalised Least Square estimator[12]:

$$\begin{aligned}(legislation_{jt} - f_{1,jt}.f_{2,jt}.f_{3,jt}) &= \lambda_0 + \lambda_1 f_{1,jt} + \lambda_2 f_{2,jt} + \lambda_3 f_{3,jt} + \\ &\quad \lambda_4 f_{1,jt}.f_{2,jt} + \lambda_5 f_{2,jt}.f_{3,jt} + \lambda_6 f_{1,jt}.f_{3,jt} + \varepsilon_{jt},\end{aligned} \tag{3.4}$$

where $\alpha = \lambda_5$, $\beta = \lambda_6$, $\delta = \lambda_4$. We have a straightforward interpretation of the impact of the three components in our index of flexibility at the margin, since the weights are computed such that the index reflects the effective impact of the 2004 change in legislation in the period under analysis[13].

Simlarly to Alexandre et al. (2010), we assume that the functional form F(.) is a logistic distribution:

$$\begin{aligned}flex_{jt} &= (\frac{\exp(\alpha)}{1+\exp(\alpha)} + \frac{\exp(f_{1,jt})}{1+\exp(f_{1,jt})}).(\frac{\exp(\beta)}{1+\exp(\beta)} + \\ &\quad \frac{\exp(f_{2,jt})}{1+\exp(f_{2,jt})}).(\frac{\exp(\delta)}{1+\exp(\delta)} + \frac{\exp(f_{3,jt})}{1+\exp(f_{3,jt})}),\end{aligned} \tag{3.5}$$

Thus, the weights (see Table 3.1) are given by: $\begin{bmatrix} \frac{\exp(\widehat{\alpha})}{1+\exp(\widehat{\alpha})} \\ \frac{\exp(\widehat{\beta})}{1+\exp(\widehat{\beta})} \\ \frac{\exp(\widehat{\delta})}{1+\exp(\widehat{\delta})} \end{bmatrix} = \begin{bmatrix} \frac{\exp(\widehat{\lambda}_5)}{1+\exp(\widehat{\lambda}_5)} \\ \frac{\exp(\widehat{\lambda}_6)}{1+\exp(\widehat{\lambda}_6)} \\ \frac{\exp(\widehat{\lambda}_4)}{1+\exp(\widehat{\lambda}_4)} \end{bmatrix} = \begin{bmatrix} 0.5704 \\ 0.5374 \\ 0.4512 \end{bmatrix}$.

From these results, we can conclude that the component that most contributed to flexibility at the margin between 2003 and 2011 was the proportion of fixed-

[12] We account for the possible correlation of the errors within the observations of the narrow sector.

[13] Note that we abstract from the functional form of $f_{1,jt}$, $f_{2,jt}$ and $f_{3,jt}$, such that the weigths are independent of the functional form F(.).

Table 3.1: Estimation of the weights of the index of flexibility at the margin

VARIABLES	legislation_dep
f1_j	-0.477***
	(0.0687)
f2_j	0.0598
	(0.156)
f3_j	-0.0167
	(0.0673)
f1f2	-0.196
	(0.242)
f2f3	0.284***
	(0.103)
f1f3	0.150*
	(0.0824)
Constant	0.578***
	(0.0517)
Observations	19,526
R-squared	0.160

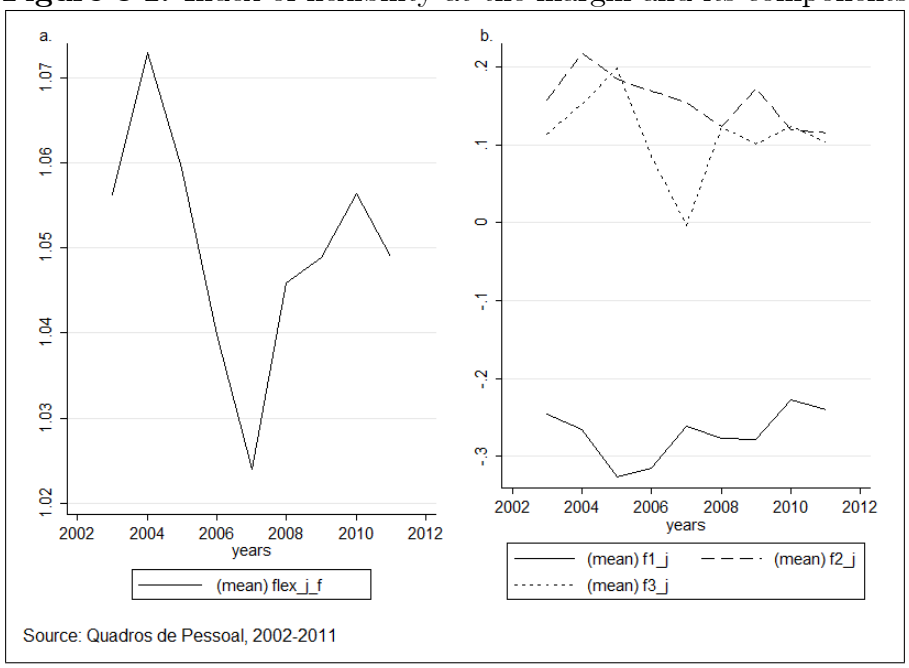
Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors clustered in narrow sectors in parantheses. ***p<0.01, ** p<0.05, * p<0.1.

term contracts, followed by the average duration of the contract and, finally, by the proportion of non-converted fixed-term contracts. Note also that the standardised proportion of non-converted fixed-term contracts is associated to a negative coefficient, which indicates that there is a value up to which this variable does not contribute positively to the index. These results are not surprising given the change in legislation during this period. The possible impact of this change is bounded to the stock of fixed-term contracts that can have their duration extended, such that it was potentially more beneficial for sectors with a larger proportion of fixed-term contracts. At the same time, the average duration of the contract is expected to increase, which is dependent on the proportion of contracts that remain as fixed-term from one year to the other and not converted to permanent contracts.

Panel a. of Figure 3.1 shows the evolution of the index over time (weighted by the share of employment in the narrow sector). In panel b., we plot the behaviour of the standardised components of the index of flexibility at the margin. The aggregate index of flexibility at the margin co-moves with the changes in legislation between 2003 and 2011. Notably, flexibility at the margin increased in 2004 in line with the decrease in the rigidity in the OECD EPL index for fixed-term contracts. In 2007, there was a drop in the flexibility at the margin mostly driven by the non-conversion

rate; this is not surprising as it was when 3 year contracts renewed in 2004 had to be terminated or converted in open-ended contracts.

Figure 3·1: Index of flexibility at the margin and its components



3.4.4 Econometric Methodology

We estimate a panel-data vector autoregression model to study the dynamic relationship between job flows and flexibility at the margin. The model is specified as follows:

$$y_{jt} = \Gamma_0 + \Gamma_1 y_{j,t-1} + \Psi_j + \Omega_t + e_{it},\ j = 1, ..., J \text{ and } t = 1, ..., T, \tag{3.6}$$

where j denotes the narrow sector and t denotes time measured in years. All variables in $y_{jt} = [flex_{jt}, jobdft_{jt}, jobcft_{jt}, jobdoe_{jt}, jobcoe_{jt}]$ are treated as endogenous. $flex_{jt}$ is the index of flexibility at the margin, $jobdft_{jt}, jobcft_{jt}$ are the destruction and creation rates of fixed-term jobs and $jobdoe_{jt}, jobcoe_{jt}$ are the destruction and creation rates of permanent jobs, respectively. We also control for unobserved group heterogeneity (Ψ_j) and year fixed-effects (Ω_t). By including year fixed-effects we account for aggregate and business cycle shocks. e_{it} is a random disturbance usually

defined as an innovation or shock to each variable in the system.

As Canova and Ciccarelli (2013) argue, when J is large and T is fixed, a GMM approach is a consistent estimation method. In order to estimate the model, the methodology introduced by Love and Zicchino (2006) is adopted[14]. We first remove the year fixed-effects by subtracting the cross-sectional mean of each variable in each period and eliminate the sector fixed effects by applying the Helmert transformation (Arellano and Bover, 1995), which is obtained by subtracting the mean of the subsequent values of each variable[15]. In order to select the model and moment conditions, and since we estimate a just-identified model, we rely on the coefficient of determination (Abrigo and Love, 2015). We therefore select a model with one lag and one-lagged variables as instruments, which we estimate using a GMM estimator.

A panel-VAR model also provides an adequate framework to test if the index of flexibility at the margin Granger causes and/or is Granger caused by job creation and job destruction. One variable is said to Granger cause the other if the coefficients of the lagged values of that variable are statistically different from zero and they therefore help predict this other variable. To draw conclusions on the causality between the variables in the system, we perform a Wald test on the lagged values of the variables in each equation.

After concluding about the stability of the parameters, we turn our attention to the orthogonalised impulse response functions that show how job creation and job destruction behave at $t + h$, where $h = 1, ..., 6$, periods after a shock in the index of flexibility at the margin. In order to identify that shock we use a Cholesky decomposition. The first variable in the Cholesky ordering is the index of flexibility on the use of fixed-term contracts. A shock to this variable in the narrow sector captures all forces that increase the flexibility in the use of fixed-term contracts, namely legislation reforms. This variable is followed in the ordering by job flows. Job destruction is included before job creation because the latter takes longer (not only to open a vacancy but to screen the potential candidates) than job destruction, which is documented by the higher volatility of job destruction than job creation

[14] We thank Inessa Love for making the STATA code available (Abrigo and Love, 2015).

[15] This procedure has the advantage of circumventing the loss of information of first differencing in unbalanced panels.

(Mortensen and Pissarides, 1994). We put the job flow rates for fixed-term contracts first in the Cholesky decomposition because empirical evidence shows that, since fixed-term contracts have lower adjustment costs, they are used by firms to adjust the employment level (Varejão and Portugal, 2007) and are a source of employment flexibility (Centeno and Novo, 2012), while firms are less willing to create and destroy permanent jobs given the higher firing costs they entail[16]. This specific ordering allows the index of flexibility at the margin to impact not only with a lag but also contemporaneously in the other variables in the system, while shocks in the destruction and creation of fixed-term and open-ended jobs do not reflect a variation in the index of flexibility at the margin. This assumption seems reasonable since it takes time for policy makers to understand the economic aggregate conditions and to create new legislation and put it in force in response to those conditions. The 95% confidence error bands of the impulse response functions are calculated using Monte Carlo simulations with 1000 repetitions.

Finally, we compute the forecast error variance decompositions for a 6-year time horizon, which gives information about the contribution of each variable's shocks to the variation in the other variables in the system.

3.4.5 Descriptive Statistics

We start by looking at the behaviour of job flows in the Portuguese labour market between 2003 and 2011. Over this period, on average, the job destruction rate (12.1%) was higher than the job creation rate (11.7%); therefore, there was net job destruction of approximately 0.5%. In panel a. of Figure C.2, we can distinguish two periods. From 2004 until 2008, there was net job creation in the Portuguese labour market, but from 2009 until 2011, a period of deep economic, financial and debt crisis, the net job destruction was equal to 5%. From 2003 until 2008, on average, the job creation rate was equal to approximately 12.8%, while the job destruction rate was about 11%. From that moment on, we observe a decrease in the rate at which jobs are created but also an increase in the rate of destruction. In fact, we can see that the net job destruction in the crisis period resulted especially from the marked

[16] This is a common assumption in theoretical models, such as in Cahuc and Postel-Vinay (2002).

increase in the destruction and the reduction in the creation of fixed-term jobs (panel b. of Figure C.2). This evidence is in line with that reported for OECD countries (OECD, 2010, pp.21-22) and confirms that fixed-term contracts are used as an adjustment mechanism in the Portuguese labour market (e.g. Varejão and Portugal, 2007), especially during the most recent crisis (Carneiro et al., 2014). Whereas, on average, 29.9% of fixed-term jobs were created and 22.4% were destroyed before the crisis, from 2009 onwards these figures were 23% and 35%, respectively. Therefore, the margin of adjustment relied on the destruction of temporary jobs. As expected, permanent job flows were lower than temporary job flows. On average, the creation of permanent jobs decreased from 12.4% to 11.5% when the economic and financial crisis started, while the permanent job destruction increased from 11.6% to 13.8%.

Additionally, we compute job flows by sector of activity, region, size and age of the firms (Tables D.1, D.2, D.3 and D.4, respectively).

Similarly to Davis et al. (1996), Centeno et al. (2007) and Lane et al. (1996), we find that the magnitude of job flows is larger in smaller and younger firms. We can also see that job dynamics differ during the 2009 crisis in line with firms' characteristics. As in other OECD countries, net job destruction was highest in construction, manufacturing and extractive industries (OECD, 2010, pp.20-23) and smaller and younger firms were also in a more vulnerable position and, thus, suffered a larger adjustment.

These results show that it is very important to consider the heterogeneity in job flows between sectors of activity, regions, size and age categories. Thus, the use of a disaggregated unit of analysis, such as the narrow sector j is justified. The main variables are described and summarised in Table D.5 in the Appendix.

3.5 Results

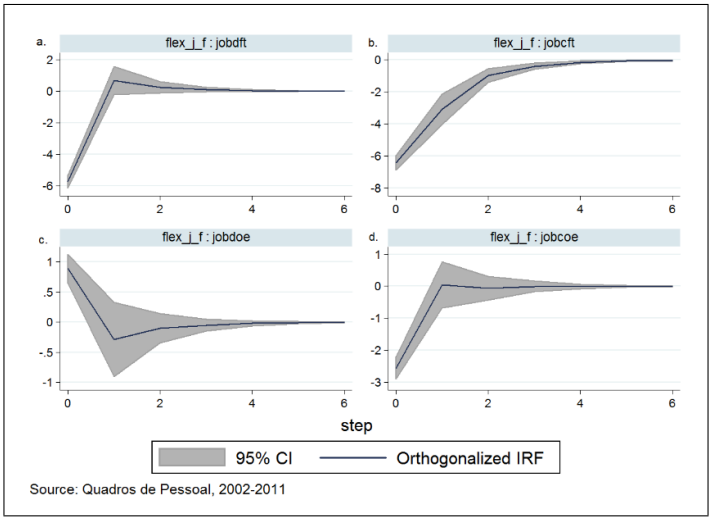
3.5.1 Job Flows and Flexibility at the margin

The results of the estimation of Equation 3.6 are depicted in Table D.6 in the Appendix. The dynamic relationship between the variables in the system is captured in one lag and the one-time lagged values of the variables in the system are adequate

instruments for each variable. After concluding that the system is stable, we are able to compute the impulse response functions (Figures 3.2-3.4)[17].

Ceteris paribus, a one standard deviation shock in the index of flexibility at the margin Granger causes a reduction in the rate at which fixed-term jobs are created in the three periods ahead and a slight increase in the rate at which they are destroyed in the two following years (panels b. and a. of Figure 3.2, respectively). The latter effect is only statistically significant at a 5% significance level and is smaller in magnitude in absolute terms: the fixed-term job destruction rate increases by approximately 0.7pp., while the creation rate of fixed-term jobs decreases by approximately 3pp. in the year after the shock in flexibility at the margin. We can therefore observe that the effects of the promotion of flexibility at the margin on fixed-term job flows tend to dissipate quickly. This may be explained by the fact that a lower (higher) creation (destruction) rate of fixed-term jobs also fosters a lower (higher) destruction (creation) rate of fixed-term jobs, counteracting the direct effect of the shock in the index of flexibility at the margin.

Figure 3·2: Response of job flows to a shock in flexibility at the margin

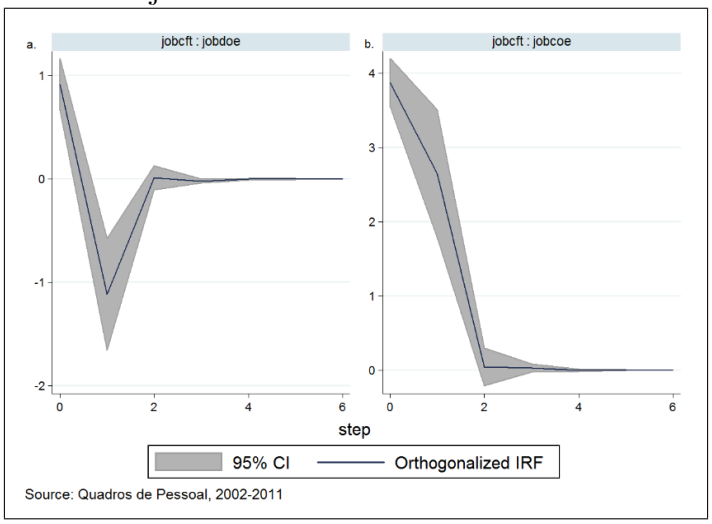


Although there is no direct causal effect of a shock in the index on permanent job flows, we find an indirect effect of the index of flexibility on the substitution

[17] The impulse response functions not shown are available upon request.

elasticity between fixed-term and open-ended contracts. The lower creation rate of fixed-term jobs, caused by a shock in flexibility at the margin, Granger causes an increase in destruction of permanent jobs and a larger decrease in their creation in the following year (Figure 3.3).This argument is in line with Bentolila and Saint-Paul's (1992) predictions and Centeno and Novo's (2012) empirical evidence, who find a high degree of substitution between the two types of contract.

Figure 3·3: Response of permanent job flows to a shock in the creation of fixed-term jobs

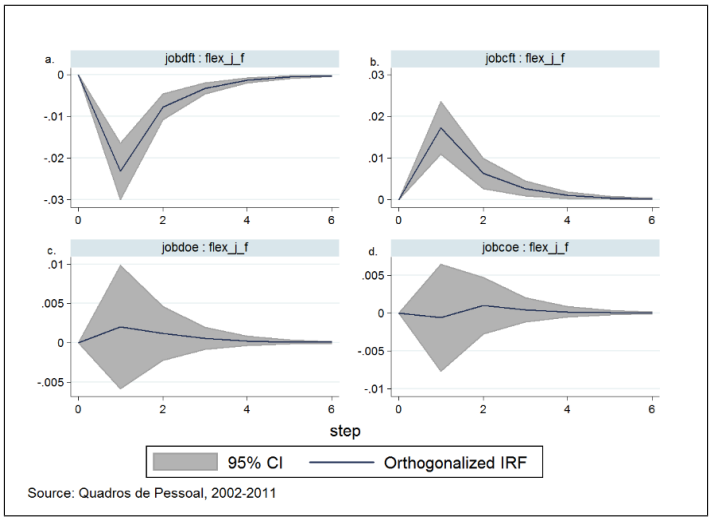


The dynamic behaviour between the index of flexibility and job flows shows that easing the regulations on fixed-term contracts, namely by increasing their maximum legal duration, seems to stimulate the maintenance of existing contracts instead of the creation of new ones and masks a recomposition of the two types of contract in the workforce. The positive response of destruction of jobs with fixed-term contracts to flexibility at the margin was expected (Sala et al., 2012). And although the negative effect on the creation of fixed-term jobs is at odds with theoretical predictions, it is also found in Faccini (2014) who shows that extending the maximum duration of fixed-term contracts leads to lower overall job creation in comparison with a reform that allows a higher acceptance rate of this type of contract.

Given the results stated above, a panel VAR model seems to be adequate to assess the impact of flexibility at the margin on job flows by type of contract, since

there is a dynamic behaviour between the index and fixed-term job flows. Our results support the hypothesis that labour market performance, namely job creation and destruction, triggers employment protection reforms that increase the use of fixed-term contracts, and that reverse causality should be taken into account. Notably, there is evidence that, at the standard significance levels, the index of flexibility at the margin is Granger caused by the creation and destruction of fixed-term jobs (Figure 3.4). A shock in the creation of fixed-term jobs fosters an increase in the index of flexibility at the margin, while a shock in the destruction of fixed-term jobs has a negative effect on that index. These effects last for about 4 years and were expected not only because of the way the index was constructed but also as the flexibility at the margin is direct and positively affected by the pool of available fixed-term workers. The level of flexibility is also higher when the contract can be extended (for example, by allowing additional renewals or longer maximum legal duration) and thus there is less destruction of fixed-term jobs.

Figure 3·4: Response of flexibility at the margin to a shock in job flows

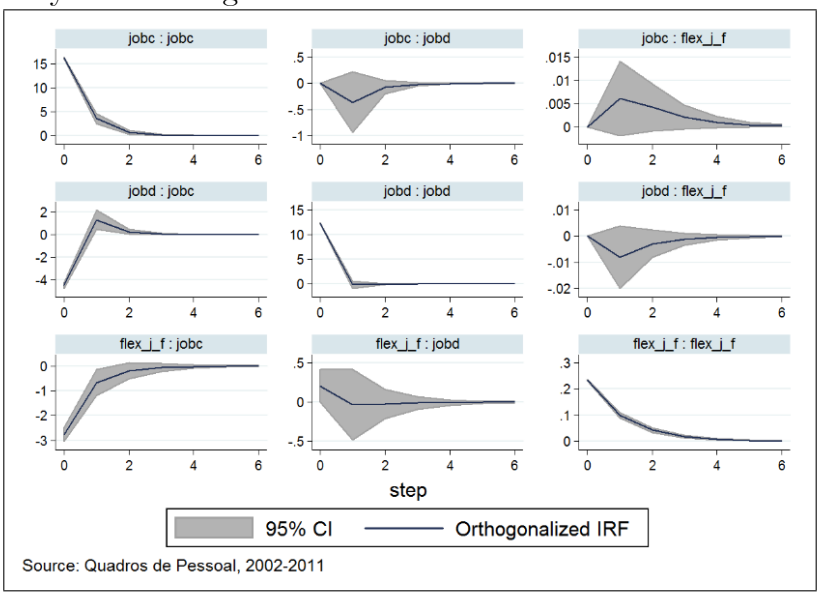


By estimating the panel VAR model using the overall job flows instead of job flows by contract type[18] (Figure 3.5), we can conclude that we might underestimate

[18]For this model and according to the information criteria, we select a model with one lag and the one-time lagged values of the variables as instruments.

the effect and fail to capture the dynamics associated with the implementation of this type of reform by focusing only on aggregate job flows. According to these estimates, there is no causal link between overall job flows and the index of flexibility at the margin. Thus, we find that this type of policy may in fact have a perverse effect as Blanchard and Landier (2002) predict and fail to improve labour market functioning. Namely, in line with Cappellari et al.'s (2012) findings, improved flexibility at the margin does not seem to significantly boost aggregate job creation and employment growth.

Figure 3·5: Impulse response functions of aggregate job flows and flexibility at the margin



Finally, the results of the variance decomposition are presented in Table 3.2. In line with the results discussed above, the index of flexibility at the margin helps to explain a non-marginal variation in the destruction and creation of fixed-term jobs one period ahead, approximately 4.9% and 5.2%, respectively. What is interesting to note is that the contribution of the index to the variation in the creation of fixed-term jobs increases over time and after 6 years is equal to approximately 6.4%. In turn, in the following period, 12.5% of the variation in the creation of fixed-term jobs is explained by their destruction rate. This may be due to the fact that workers on fixed-term contracts experience higher job turnover rates and that some fixed-term

Table 3.2: Variance decompositions

	s	flex_j_f	jobdft	jobcft	jobdoe	jobcoe
flex_j_f	1	1	0	0	0	0
jobdft	1	0.049	0.951	0	0	0
jobcft	1	0.052	0.125	0.822	0	0
jobdoe	1	0.003	0.020	0.004	0.973	0
jobcoe	1	0.016	0.00007	0.036	0.045	0.904
flex_j_f	6	0.985	0.009	0.005	0.00009	0.00002
jobdft	6	0.050	0.946	0.004	0.0004	0.0006
jobcft	6	0.064	0.130	0.803	0.0001	0.002
jobdoe	6	0.004	0.020	0.009	0.967	0.0001
jobcoe	6	0.015	0.0005	0.050	0.060	0.874

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

contracts are used for churning while others are used for screening purposes. It is also worth mentioning that the creation of fixed-term jobs explains approximately 5% of the creation of open-ended jobs in the 6 periods ahead.

3.5.2 Components of the Index of Flexibility at the Margin and Job Flows

In order to identify the sources of the results discussed in the last subsection, we re-estimate the model substituting the index of flexibility with each of its components: the proportion of fixed-term contracts, the average duration of the contract and the share of non-converted fixed-term contracts[19]. The resulting impulse response functions are presented in Figures 3.6, 3.7 and 3.8, respectively.

The one period ahead negative causal effect of a shock in flexibility at the margin on the creation of fixed-term jobs is explained by the shock in the proportion of fixed-term contracts (-2.9 pp.) and in the proportion of non-converted fixed-term contracts (-6.4 pp.). This is quite intuitive since the more contracts that remain fixed-term, the lower the need to replace them and create new fixed-term contracts. On the other hand, a shock in the average duration of the contract has no causal impact on the creation of fixed-term jobs.

An increase of one standard deviation in the average duration of fixed-term con-

[19] Selection criteria for the model and moment conditions hold for all models.

Figure 3·6: Response of job flows to a shock in the proportion of fixed-term contracts

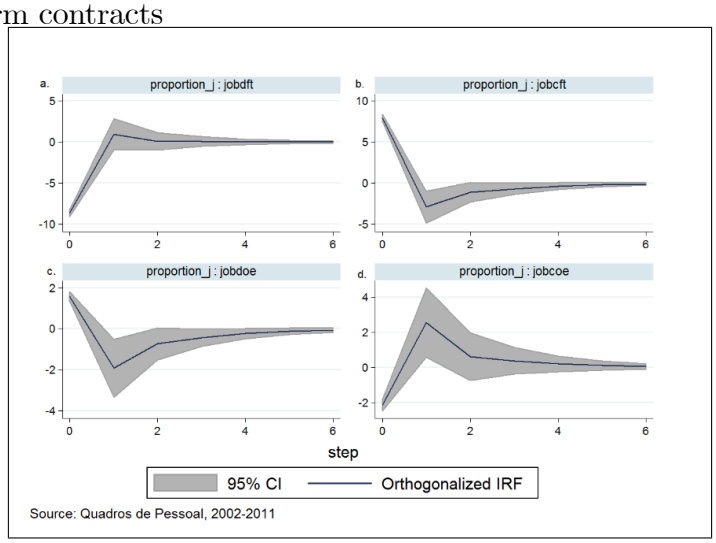
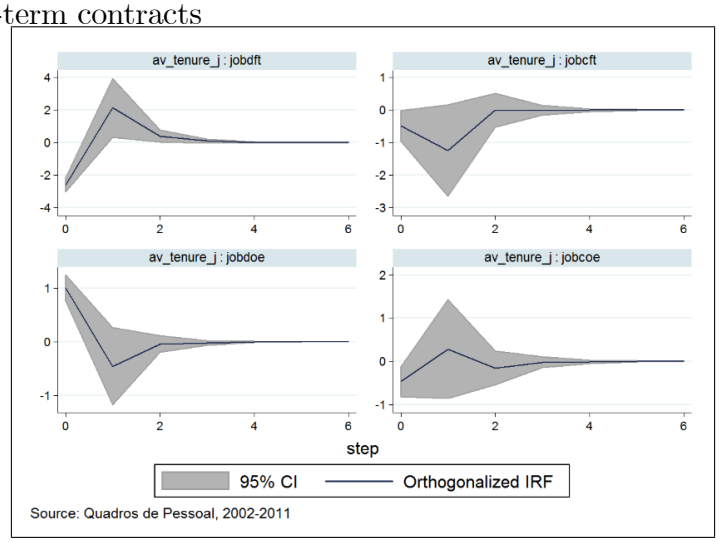


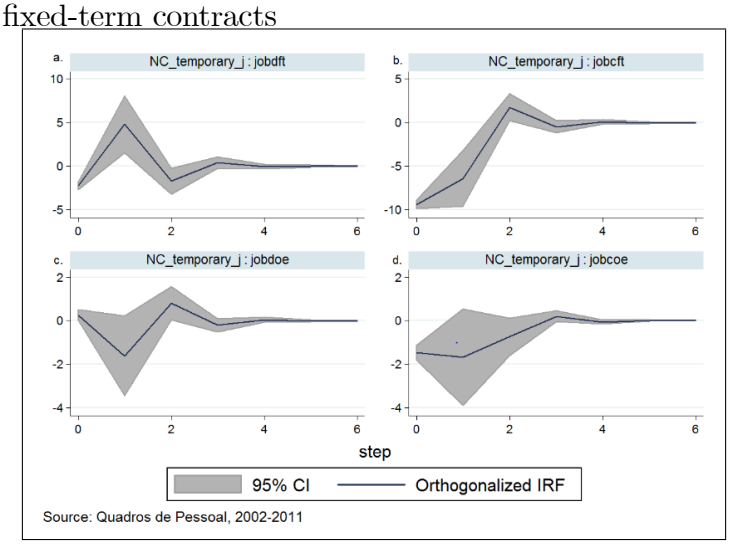
Figure 3·7: Response of job flows to a shock in the average duration of fixed-term contracts



tracts and in the proportion of non-converted contracts Granger causes an increase in the destruction of fixed-term jobs (2.1 pp. and 4.8 pp. in the period after the shock, respectively), which is mostly explained by the fact that the maximum duration and number of renewals is legally established in this type of contract and, once reached, more contracts have to be terminated.

Innovations to the share of non-converted contracts produce the largest and most lasting effects (4 years) on the job flows of fixed-term contracts. Therefore, we can argue that the maintenance of the contracts seems to be explained mainly by the shock in the non-conversion rate of fixed-term contracts, which may indicate and reflect the low or inefficient incentives for the conversion of fixed-term into open-ended contracts in the period in which their maximum duration was allowed to increase[20].

Figure 3·8: Response of job flows to a shock in the non-conversion rate of fixed-term contracts



Finally, it is worth mentioning that the destruction of open-ended jobs is negatively affected in the period following a shock in either the proportion of fixed-term contacts or in the share of non-converted fixed-term contracts; this may support the hypothesis of fixed-term contracts being used as a buffer for open-ended contracts.

[20]Note that, during the period under analysis, there was a simultaneous change in legislation that introduced a penalty in social security contributions for firms with more than 15% of employees on fixed-term contracts, which could be recouped if the contracts were converted.

The shocks in the index of flexibility at the margin and in the proportion of fixed-term contracts are quite persistent as the increase in these variables only vanishes after 4 years; on the other hand, the average duration of the contract and the proportion of non-converted fixed-term contracts are the least persistent variables (Figure C.3). This result is interesting as shocks in these two last variables fade out in 2 years and the maximum number of renewals as well as the additional duration of the contract was set at three years.

Regarding the variance decompositions, the proportion of non-converted fixed-term contracts is the component that explains the highest share of the variation in job creation (15.7%), while the variation in job destruction is mainly explained by the proportion of fixed-term contracts (11.2%) in the 6 periods ahead (Table D.7 and D.8, respectively). In fact, the proportion of fixed-term contracts helps explain 9.1% of the variation in the creation of fixed-term jobs and 3% of the variation in the destruction and creation of open-ended jobs after 6 years. In turn, the average duration of the contract only explains 1.7% of the variation in the destruction of fixed-term jobs in the 6 periods ahead and a marginal percentage of the variation in the remaining variables (Table D.9).

3.5.3 Employment growth pattern and the impact of flexibility at the margin on job flows

The effects of the changes in legislation that ease regulations on fixed-term contracts may be differentiated according to the stage of the firm's business cycle (Bentolila and Saint-Paul, 1992) and job flows in shrinking sectors may be more negatively affected by EPL than job flows in growing sectors (Messina and Vallanti, 2007). For that reason, we re-estimate the model for narrow sectors with negative and positive employment growth separately (Figures 3.9 and 3.10, respectively).

In expanding sectors, we find that an innovation to the index of flexibility at the margin does not Granger cause job flows. Conversely, in narrow sectors with negative employment growth, a one standard deviation in the index of flexibility at the margin causes an increase in the destruction and a decrease in the creation of fixed-term jobs for a 4-year period.

Figure 3·9: Response of job flows to a shock in flexibility at the margin in shrinking sectors

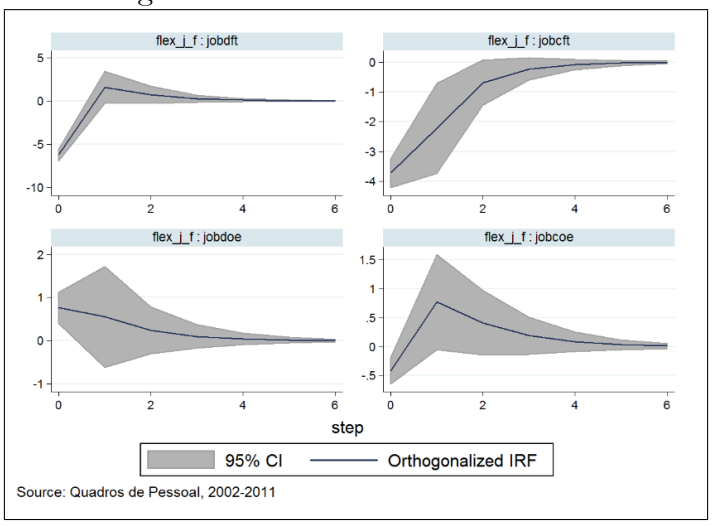
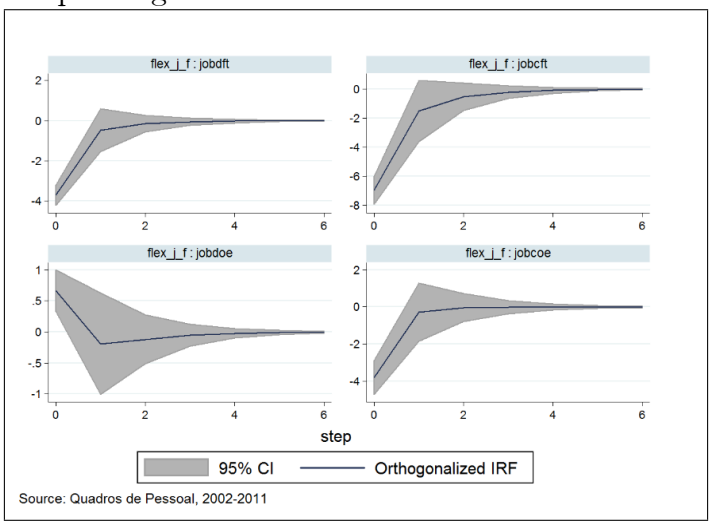


Figure 3·10: Response of job flows to a shock in flexibility at the margin in expanding sectors



Looking at the variance decompositions for both samples, we can conclude that the index of flexibility at the margin explains a non-negligible higher share of the destruction rate of fixed-term jobs in shrinking sectors (7.4%) than in expanding sectors (5.4%) in the 6 periods ahead (Table D.10 and D.11, respectively). It is also relevant to note that the destruction of fixed-term jobs explains a slightly higher percentage of the variation in job creation of fixed-term jobs in shrinking sectors than in expanding sectors (8.9% and 7.1% one period ahead, respectively), while the destruction of open-ended jobs explains a higher share of the permanent creation rate in expanding than in shrinking sectors (4.7% and 1.8%, respectively, in the following year).

These results show that the promotion of flexibility at the margin is especially important for firms that are reducing jobs, since those firms adjust the composition of job flows when a shock arises in the previous period. This is quite intuitive and is in line with the findings of Varejão and Portugal (2007) that show that the proportion of fixed-term contracts has a larger effect on the probability of establishments adjusting employment downwards rather than upwards. This pattern may signal that firms use fixed-term contracts longer and avoid converting them when they are in a downturn, which is in fact evidence that fixed-term contracts are also used as buffer stocks for open-ended contracts in the Portuguese labour market.

Therefore, the mean effect of flexibility at the margin described in subsection 3.5.1 is explained by the behaviour of shrinking rather than expanding sectors.

3.5.4 Robustness Analysis

We rely on the Cholesky decomposition to identify the impulse response functions. Although the ordering of the variables was based on stylised results found in the literature, we test the robustness of our findings by changing the ordering. Namely, we reversed the ordering putting the flexibility index last (Figure C.4 and C.5). The main conclusions remain about the impact of an innovation to job flows: there is less creation and more destruction of fixed-term jobs. Moreover, a negative (positive) response on the index of flexibility at the margin is observed following a shock in the destruction (creation) of fixed-term jobs. Since the results are qualitatively

unchanged, we are confident that they are not driven by the ordering of the variables in the system.

The results are also unchanged if we exclude the agriculture sector, which is under sampled in the data source used. The same also occurs if we compute the index of flexibility at the margin with the exogenous weights as proposed by Alexandre et al. (2010) (Figure C.6 and C.7).

As discussed by Dolado et al. (2002), the relative wage of permanent and fixed-term contracts may have an impact on the new equilibrium proportion of fixed-term contracts. The introduction and facilitation of the use of fixed-term contracts may indeed increase the bargaining power of workers with permanent contracts, especially in countries with high employment protection levels (Bentolila and Dolado, 1994). As such, we include the relative wage of permanent workers (the ratio between the average log hourly wage[21] of workers on permanent contracts and the average log hourly wage of workers on fixed-term contracts) in the model and conclude that the results remain qualitatively unchanged (Figures C.8, C.9 and C.10). Although the relative wage is not caused by the innovations to job flows, a shock in the relative wage Granger causes higher temporary job destruction, at a 5% significance level. This is in line with Dolado et al.'s (2002) argument, since if workers on permanent contracts can negotiate higher wages, firms may want to adjust employment by relying more heavily on fixed-term jobs that entail lower dismissal costs so as to tackle the higher wage bill. This effect may be more marked in sectors with higher union power.

Finally, we assess the robustness of our results to the inclusion of employment growth calculated for each narrow sector as a proxy for the business cycle at the sector level, which captures the idiosyncratic behaviour, namely the phase of the cycle and the future prospects of each unit j. The results in Figures C.12 and C.13 are similar to those discussed above. It is also stressed that although a shock in flexibility at the margin has a negative effect on employment growth, there is no causal relationship between the index of flexibility at the margin and employment growth (Figure C.11).

[21] Wage is equal to the sum of base wage, regular pay and overtime pay.

3.6 Conclusion

Existing theoretical models predict that increasing flexibility at the margin by easing the regulations or introducing fixed-term contracts has a positive impact on both job creation and job destruction although it is not known *a priori* which effect dominates. The net effect on job flows depends on the initial conditions and the institutional environment, such as the stringency of employment protection on open-ended contracts. Empirical studies that examine the effects of promoting flexibility at the margin are hence called for.

This paper provides such a study by analysing the effects on job flows of the 2004 Portuguese change in legislation that extended the maximum legal duration of fixed-term contracts from three to six years.

We use a new index of flexibility at the margin composed of three components: the proportion of fixed-term contracts, the average duration of the contract and the non-conversion rate of fixed-term contracts, and we estimate a reduced form panel VAR, which does not require strong assumptions, to examine the dynamic relationship between flexibility at the margin and job flows by type of contract.

Our results show that a panel VAR model is a suitable methodology to study the impact of flexibility at the margin on job flows and that it is important to disaggregate job flows by type of contract to capture a more accurate and detailed picture. We find that in a segmented labour market like that of Portugal, policy makers may not be able to boost overall job creation and increase employment if they adopt changes in legislation that ease regulations on fixed-term contracts. Our results show that an innovation to the index of flexibility at the margin has a negative impact on the creation rate and a positive impact on the destruction rate of fixed-term jobs. The effect on the creation rate of fixed-term jobs is of a greater magnitude in absolute terms and lasts for three years, while the impact on the destruction of fixed-term jobs fades out after two periods. The results also show evidence of a recomposition of fixed-term and open-ended contracts in the workforce since we identify that the promotion of flexibility at the margin has the indirect effect of increasing the destruction rate and causing a larger reduction in the creation rate

of open-ended jobs.

We therefore provide new evidence on the magnitude and duration of the effects of shocks in flexibility at the margin on job flows in the Portuguese labour market. We also conclude that the direction and magnitude of these effects is explained predominantly by the shock in the proportion of non-converted fixed-term contracts. This is not surprising since firms are expected to use fixed-term contracts longer instead of creating new ones when the maximum legal duration is increased. Finally, our results evidence a heterogeneous effect of flexibility at the margin between shrinking and expanding sectors as a statistically significant impact is only observed on job flows in sectors that are reducing jobs.

The conclusions stated above carry important policy implications. We advocate that effective labour market policies should be developed that aim to promote the conversion of fixed-term into open-ended contracts to counteract the negative effects of measures extending their use.

One of the limitations of this study is the fact that we use annual data. As Portugal and Blanchard (2001) argue, EPL is shown to have a less significant impact on job flows when studied using lower frequency data. It would be interesting to conduct this type of analysis on quarterly or monthly data. On the other hand, gross job flows show high one-year persistence (Davis et al., 1996; Davis and Haltiwanger, 1992*b*) and, for the purposes of the paper, annual data allows us to capture persistent changes instead of transitory job flows. This dataset may, however, underestimate job flows, especially for fixed-term contracts of short duration.

Future work is planned to complement the analysis by considering other institutional variables, such as employment protection for regular contracts, unemployment benefits, wage bargaining and union density, for example, that can help understanding the effects of employment protection reforms easing the regulations on fixed-term contracts (Bertola and Rogerson, 1997; Avdagic, 2015). It would also be interesting to focus on the asymmetric impact of shocks in the index of flexibility at the margin in recessions and expansions. The dataset used encompass only three years of crisis (2009-2011), which is not enough to draw meaningful and robust conclusions about business cycle asymmetries.

Appendix C

Figures

Figure C.1: Proportion of fixed-term contracts, non-conversion rate and average duration of the contract

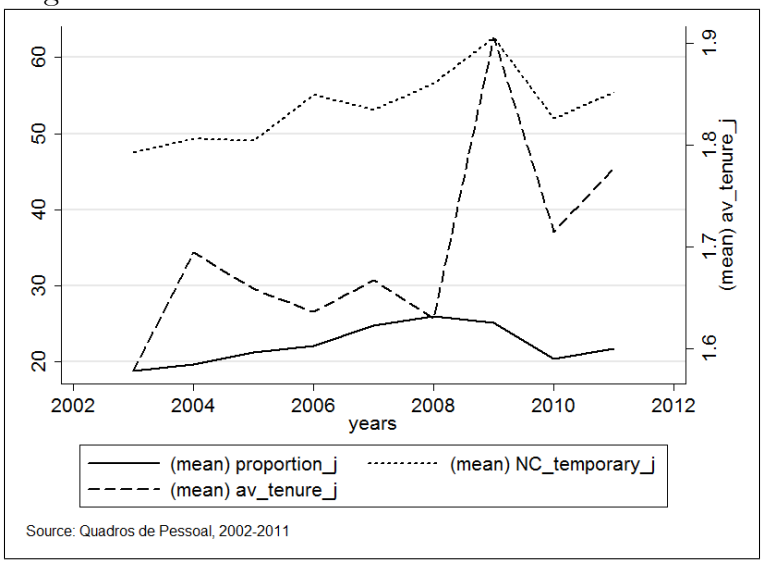


Figure C.2: Job flows

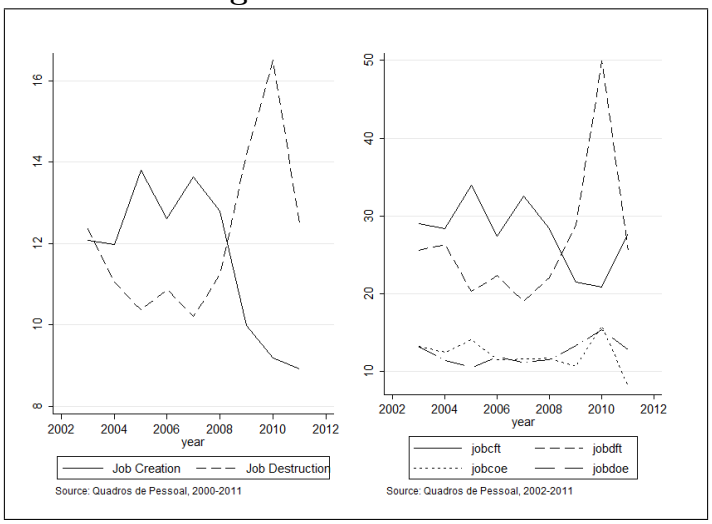


Figure C.3: Persistence of a shock in flexibility at the margin and its components

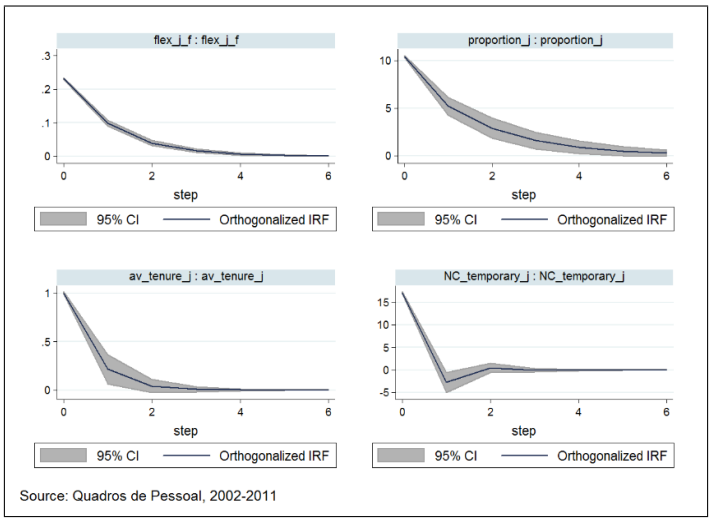


Figure C.4: Response of job flows to a shock in the index of flexibility at the margin with the index last in the Cholesky decomposition

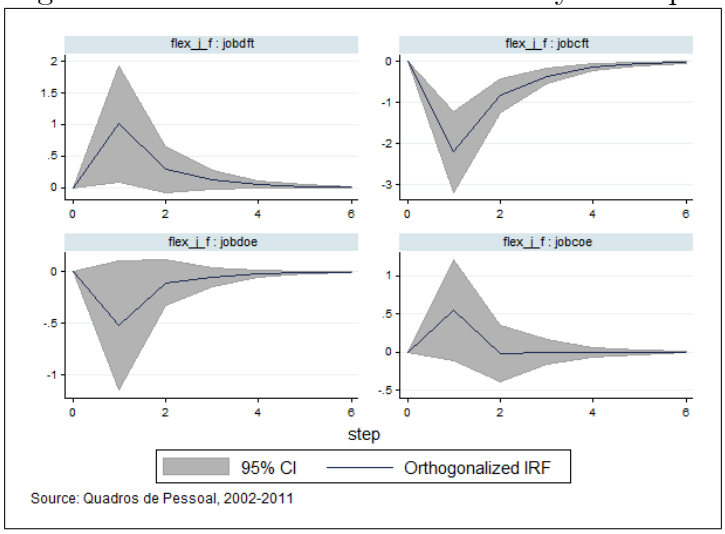


Figure C.5: Response of the index of flexibility at the margin to a shock in job flows with the index last in the Cholesky decomposition

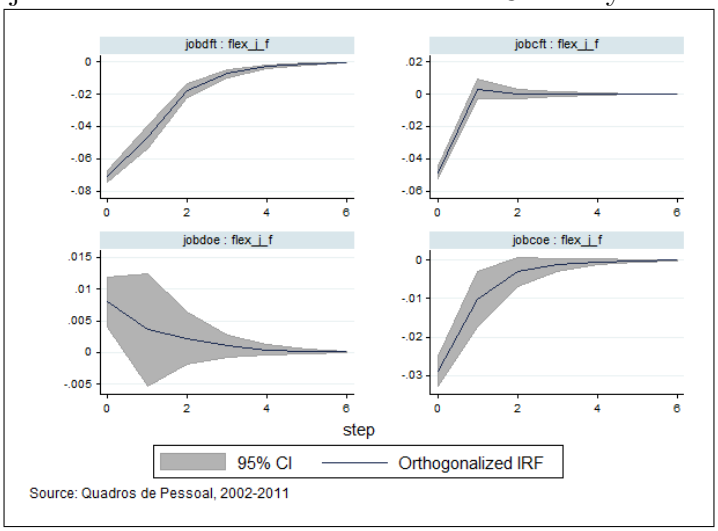


Figure C.6: Response of job flows to a shock in the index of flexibility at the margin with exogenous weights

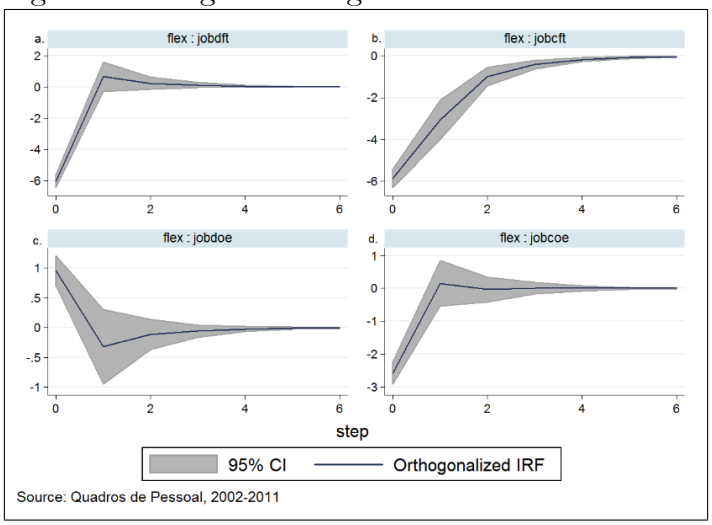


Figure C.7: Response of the index of flexibility at the margin with exogenous weights to a shock in job flows

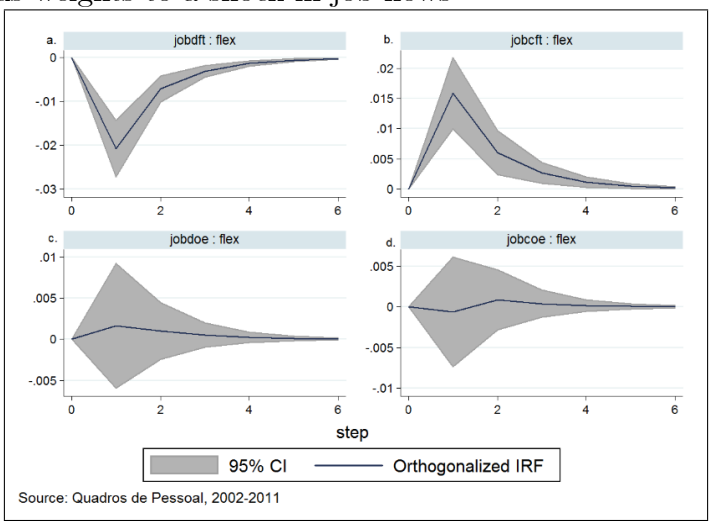


Figure C.8: Dynamic relationship between flexibility at the margin and relative wage

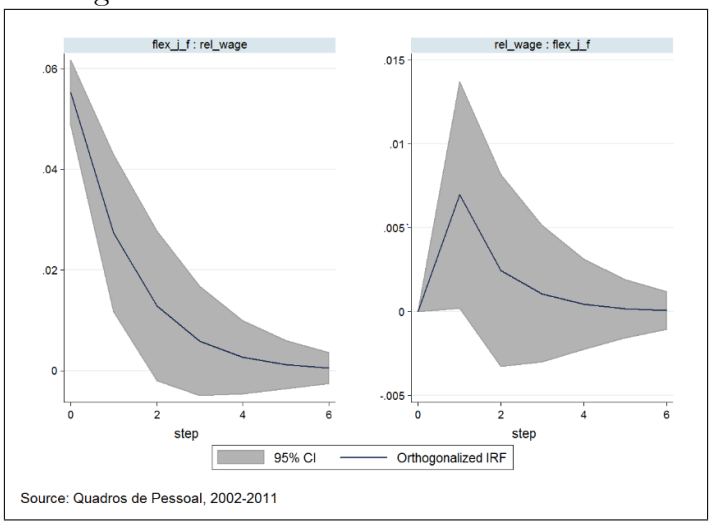


Figure C.9: Response of job flows to a shock in relative wage

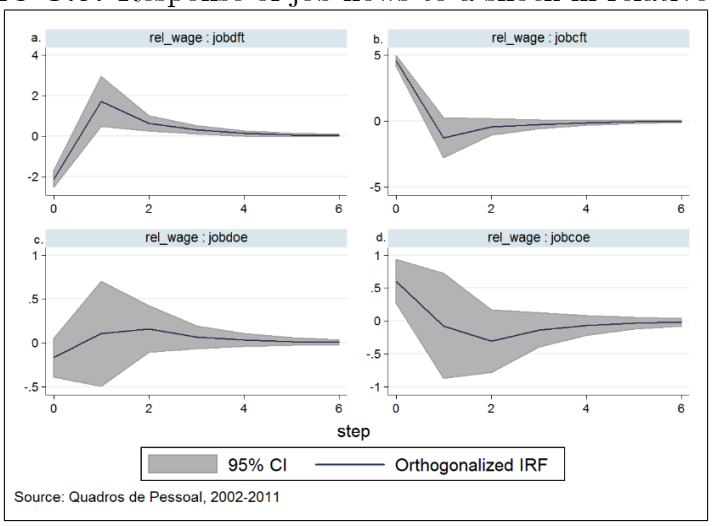


Figure C.10: Response of relative wage to a shock in job flows

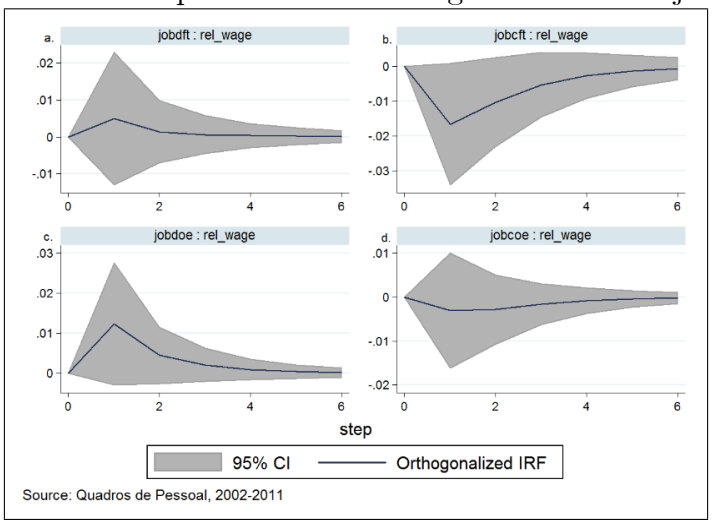


Figure C.11: Dynamic relationship between flexibility at the margin and employment growth

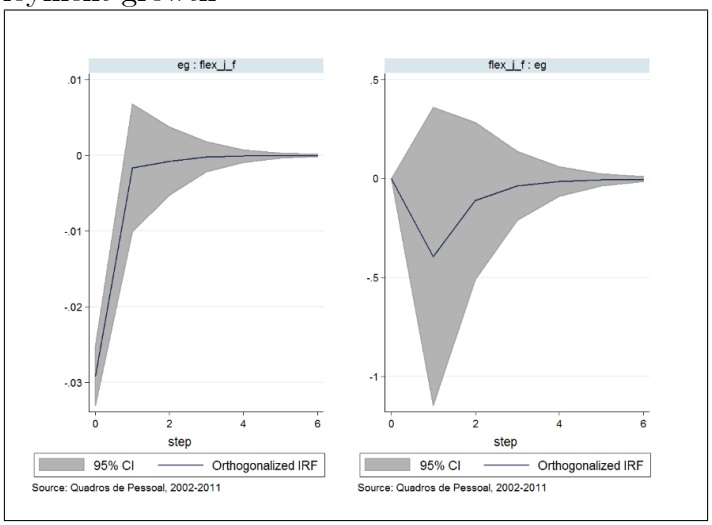


Figure C.12: Response of job flows to a shock in the index of flexibility at the margin including employment growth

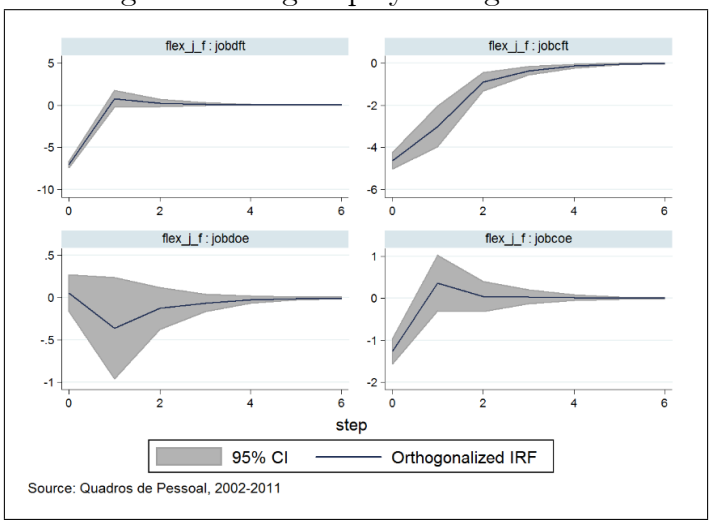
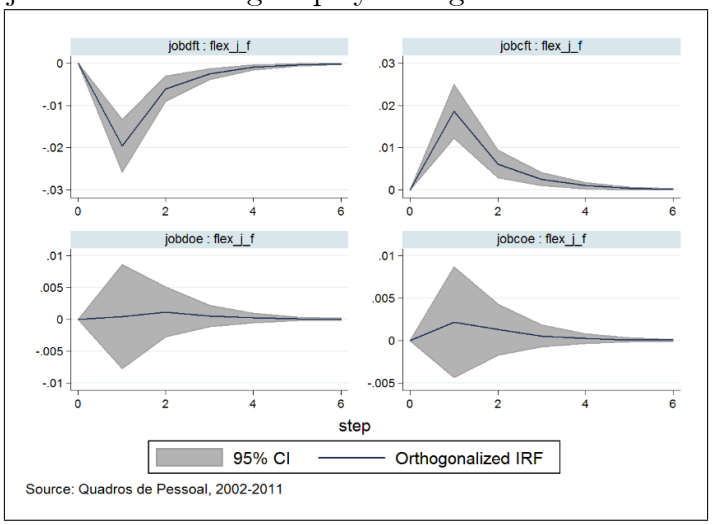


Figure C.13: Response of the index of flexibility at the margin to a shock in job flows including employment growth



Appendix D

Tables

Table D.1: Job flows by sector of activity

Sector	Job Creation		Job Destruction		Net Job Creation	
	2003-2008	2009-2011	2003-2008	2009-2011	2003-2008	2009-2011
extractive industries	8.90	6.17	12.30	13.37	-3.41	-7.20
manufacturing	8.07	6.75	10.38	12.53	-2.32	-5.77
electricity production and distribution	6.33	5.23	7.01	5.11	-0.67	0.11
construction	16.84	11.50	16.34	23.39	0.50	-11.88
wholesale and retail	12.02	9.08	10.28	12.89	1.74	-3.81
lodgind and restaurants	15.93	12.22	12.51	16.52	3.42	-4.31
transport	8.92	6.74	8.55	10.37	0.36	-3.63
financial activities	6.23	4.85	5.92	4.54	0.31	0.31
property	19.29	12.82	12.19	16.19	7.10	-3.38
education	18.94	8.43	9.34	22.35	9.60	-13.93
health and social work	13.75	9.24	5.29	7.14	8.46	2.10
collective. social and personal services	15.56	11.57	10.32	13.44	5.24	-1.87

Source: Quadros de Pessoal, 2000-2011. Notes: Agriculture and Public Administration were ommitted from the table because of lack of sample coverage.

Table D.2: Job flows by region

Region	Job Creation		Job Destruction		Net Job Creation	
	2003-2008	2009-2011	2003-2008	2009-2011	2003-2008	2009-2011
North	13.29	9.75	11.71	14.79	1.57	-5.04
Algarve	17.92	10.96	13.04	21.31	4.87	-10.35
Centre	12.11	8.98	10.58	14.08	1.53	-5.10
Lisbon	12.28	9.05	10.29	13.60	1.99	-4.55
Alentejo	14.86	11.20	12.86	16.97	1.99	-5.76

Source: Quadros de Pessoal, 2000-2011.

Table D.3: Job flows by firm size category

Firm size category	Job Creation		Job Destruction		Net Job Creation	
	2003-2008	2009-2011	2003-2008	2009-2011	2003-2008	2009-2011
]0,4]	27.14	17.97	22.41	30.57	4.73	-12.60
[5,9]	14.38	11.05	13.38	16.81	1.00	-5.76
[10,19]	12.76	9.40	11.61	15.39	1.15	-5.99
[20,49]	10.80	8.50	9.69	13.74	1.10	-5.24
[50,99]	9.77	7.22	8.28	11.54	1.49	-4.33
[100,249]	8.87	7.02	7.49	10.00	1.38	-2.97
[250,499]	9.27	7.09	8.12	8.94	1.15	-1.85
[500,999]	8.20	7.69	6.35	8.39	1.85	-0.70
[1000,+[	6.76	4.97	4.75	5.43	2.01	-0.46

Source: Quadros de Pessoal, 2000-2011.

Table D.4: Job flows by firm age category

Firm age category	Job Creation		Job Destruction		Net Job Creation	
	2003-2008	2009-2011	2003-2008	2009-2011	2003-2008	2009-2011
1	58.98	51.96	14.67	22.31	44.32	29.66
[2,5]	17.50	12.89	16.99	25.15	0.51	-12.26
[6,10]	10.70	8.12	13.69	18.03	-2.99	-9.91
[11,20]	7.86	5.60	9.35	12.39	-1.49	-6.80
[21,50]	5.55	4.56	8.75	11.19	-3.20	-6.63
[51,+[	5.95	3.20	7.14	10.11	-1.19	-6.90

Source: Quadros de Pessoal, 2000-2011.

Table D.5: Descriptive Statistics

Variable	Description	Unweighted Mean	Weighted Mean
flex_j_f	Index of flexibility at the margin in sector j	1.056	1.050
		(0.371)	(0.013)
proportion_j	Proportion of fixed-term contracts in sector j	27.864	22.241
		(20.839)	(2.418)
av_tenure_j	Average duration of fixed-term contract in sector j	1.469	1.696
		(1.485)	(0.092)
NC_temporary_j	Proportion of non-converted fixed-term contracts in sector j	51.518	53.507
		(22.909)	(4.421)
jobcft	Creation rate of fixed-term jobs in sector j	38.427	32.324
		(45.612)	(4.776)
jobdft	Destruction rate of fixed-term jobs in sector j	29.864	24.796
		(34.598)	(7.499)
jobcoe	Creation rate of open-ended jobs in sector j	24.292	13.982
		(39.700)	(2.266)
jobdoe	Destruction rate of open-ended jobs in sector j	14.590	11.869
		(21.393)	(1.182)

Source: Quadros de Pessoal, 2002-2011. Note: Weights are defined as the share of employment in each narrow sector j. Standard deviations in parentheses.

Table D.6: Dynamic relationship between job flows and flexibility at the margin

Response to	Response of flex_j_f(t)	Jobdft(t)	Jobcft(t)	Jobdoe(t)	Jobcoe(t)
flex_j_f(t-1)	0.427***	4.733**	-10.347***	-2.435	2.569
	(0.022)	(2.213)	(2.252)	(1.494)	(1.581)
jobdft(t-1)	-0.001***	0.0004	0.095***	-0.013	-0.002
	(0.0001)	(0.019)	(0.028)	(0.012)	(0.013)
jobcft(t-1)	0.001***	0.069***	-0.001	-0.043***	0.085***
	(0.0001)	(0.016)	(0.021)	(0.011)	(0.016)
jobdoe(t-1)	0.00012	-0.042	0.039	-0.058**	0.205***
	(0.0003)	(0.036)	(0.039)	(0.029)	(0.040)
jobcoe(t-1)	-0.00003	-0.032	0.067**	0.007	0.079***
	(0.0002)	(0.020)	(0.027)	(0.019)	(0.030)
Observations	13,856				
Narrow sectors	2,315				

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors in parantheses. *** p<0.01, ** p<0.05, * p<0.1.

Table D.7: Variance decompositions for the model with the proportion of non-converted fixed-term contracts

	s	NC_temporary_j	jobdft	jobcft	jobdoe	jobcoe
NC_temporary_j	1	1	0	0	0	0
jobdft	1	0.004	0.996	0	0	0
jobcft	1	0.150	0.102	0.748	0	0
jobdoe	1	0.000	0.015	0.001	0.984	0
jobcoe	1	0.008	0.001	0.038	0.047	0.906
NC_temporary_j	6	0.980	0.014	0.005	0.000	0.000
jobdft	6	0.005	0.989	0.005	0.000	0.001
jobcft	6	0.157	0.110	0.731	0.000	0.002
jobdoe	6	0.000	0.015	0.006	0.979	0.000
jobcoe	6	0.011	0.002	0.047	0.062	0.877

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

Table D.8: Variance decompositions for the model with the proportion of fixed-term contracts

	s	proportion_j	jobdft	jobcft	jobdoe	jobcoe
proportion_j	1	1	0	0	0	0
jobdft	1	0.111	0.889	0	0	0
jobcft	1	0.080	0.048	0.872	0	0
jobdoe	1	0.011	0.029	0.000	0.961	0
jobcoe	1	0.011	0.000	0.063	0.038	0.888
proportion_j	6	0.994	0.002	0.004	0.001	0.000
jobdft	6	0.112	0.885	0.002	0.000	0.001
jobcft	6	0.091	0.053	0.854	0.000	0.002
jobdoe	6	0.030	0.029	0.003	0.939	0.000
jobcoe	6	0.027	0.000	0.071	0.051	0.851

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

Table D.9: Variance decompositions for the model with the average duration of fixed-term contracts

	s	av_tenure_j	jobdft	jobcft	jobdoe	jobcoe
av_tenure_j	1	1	0	0	0	0
jobdft	1	0.010	0.990	0.000	0.000	0.000
jobcft	1	0.000	0.093	0.906	0.000	0.000
jobdoe	1	0.004	0.017	0.001	0.978	0.000
jobcoe	1	0.001	0.001	0.045	0.047	0.906
av_tenure_j	6	0.996	0.003	0.000	0.000	0.001
jobdft	6	0.017	0.979	0.003	0.001	0.001
jobcft	6	0.002	0.101	0.894	0.000	0.002
jobdoe	6	0.005	0.017	0.006	0.972	0.000
jobcoe	6	0.001	0.001	0.058	0.062	0.878

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

Table D.10: Variance decompositions for shrinking sectors

	s	flex_j_f	jobdft	jobcft	jobdoe	jobcoe
flex_j_f	1	1	0	0	0	0
jobdft	1	0.070	0.930	0	0	0
jobcft	1	0.043	0.089	0.868	0	0
jobdoe	1	0.004	0.011	0.020	0.965	0
jobcoe	1	0.003	0.011	0.008	0.018	0.960
flex_j_f	6	0.983	0.010	0.004	0.001	0.001
jobdft	6	0.074	0.907	0.015	0.003	0.001
jobcft	6	0.058	0.087	0.840	0.006	0.009
jobdoe	6	0.006	0.014	0.021	0.947	0.012
jobcoe	6	0.015	0.015	0.036	0.048	0.885

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

Table D.11: Variance decompositions for expanding sectors

	s	flex_j_f	jobdft	jobcft	jobdoe	jobcoe
flex_j_f	1	1	0	0	0	0
jobdft	1	0.054	0.946	0	0	0
jobcft	1	0.058	0.071	0.871	0	0
jobdoe	1	0.004	0.000	0.001	0.994	0
jobcoe	1	0.019	0.014	0.040	0.047	0.880
flex_j_f	6	0.991	0.002	0.005	0.001	0.001
jobdft	6	0.054	0.931	0.009	0.002	0.003
jobcft	6	0.061	0.073	0.866	0.000	0.001
jobdoe	6	0.004	0.000	0.004	0.991	0.001
jobcoe	6	0.018	0.014	0.051	0.052	0.864

Source: Quadros de Pessoal, 2002-2011. Notes: Percent of variation in the row variable explained by column variable in 1 and 6 periods after.

Chapter 4

Wage inequality between permanent and fixed-term contracts: a firm-level analysis

4.1 Introduction

It is important to understand not only how job and worker flows have been affected by the widespread implementation of employment protection reforms facilitating the use of fixed-term contracts in Europe but also what type of jobs are being created (Blanchard and Landier, 2002). Relative wages may be considered "price signals" that, along with "quantity signals" (eg. vacancy posting), drive employment reallocation across sectors of activity (Moghadam and Pissarides, 1989). Thus, the relative wage of permanent contracts may be an important determinant of their relative demand (Bentolila and Saint-Paul, 1992).

The present paper contributes to the literature by empirically studying the effects of a Portuguese change in legislation that facilitated the use of fixed-term contracts on the relative wage between workers on open-ended contracts and workers on fixed-term contracts. Most studies on the wage gap between both types of contract have been conducted using worker level data (e.g. Elia, 2010; Bosio, 2014; Ordine and Rose, 2016). This is partly due to the lack of available data on wages by type of contract at the firm level (e.g. Bentolila and Saint-Paul, 1992; Bentolila and Dolado, 1994; Benito and Hernando, 2008; Cipollone and Guelfi, 2006; Aguirregabiria and Alonso-Borrego, 2014). The unit of analysis of this study is the firm, since the firm's wage policy and employment flexibility may be interrelated (Suleman et al., 2014). We consider that the intra-firm wage differential between both types of contract may reflect the main use that firms assign to fixed-term contracts.

Chapter 2 highlights that acknowledging that firms may use fixed-term contracts

for different purposes is relevant to draw conclusions about the effects of flexibility at the margin. As mentioned, fixed-term contracts may play two primary roles. They allow firms to screen matches at a lower cost, especially in labour markets with high employment protection on regular contracts and low probationary periods (Faccini, 2014). The convergence between the wages of fixed-term workers and those of permanent workers has been considered in the literature as evidence of this screening effect (Booth et al., 2002; Amuedo-Dorantes and Serrano-Padial, 2007). Alternatively, fixed-term contracts may allow firms to adjust employment after an economic shock (Pfeifer, 2012). As argued by Pfeifer (2012), workers less likely to being integrated in the firm's internal labour market would suffer a larger wage penalty than workers at the top of the workers' wage distribution. Moreover, the higher the wages paid to workers on permanent contracts the higher the probability that a worker is hired under a fixed-term contract and the less likely is the conversion of the contract (Portugal and Varejão, 2009). Therefore, we consider that the relative wage of permanent to fixed-term workers may reflect the way in which firms primarily use fixed-term contracts[1]. Firms using fixed-term contracts primarily as a buffer stock may pay higher relative wages to permanent workers as argued by Pfeifer (2014). Alternatively, firms using fixed-term contracts primarily for secreening may have lower intra-firm wage differentials between both types of contract (Mertens and McGinnity, 2003). The effects of flexibility at the margin may, thus, be heterogeneous for firms with different wage policies with respect to fixed-term contracts.

Finally, this paper addresses an issue disregarded in the literature so far, i.e., the role played by fixed-term contracts in promoting employment reallocation towards the tradable sector. For price-taker's firms in the tradable sector, labour adjustment costs are potentially more binding than for firms in the non-tradable sector, given their exposure to international competition. Therefore, these firms may take special advantage of the use of this type of contract to reduce the firm's wage costs. This analysis is especially relevant for countries like Portugal, in which labour markets

[1]Note that the wage gap between permanent and fixed-term contracts has been considered a measure of labour market segmentation by some authors such as Osuna (2005).

are characterised by nominal wage rigidity and stringent employment protection on regular contracts. As documented by Reis (2013) and OECD (2012), there was a decline in the manufacturing employment share between 2000 and 2006 in Portugal and an increase in the employment share in nontradable sectors, especially pronounced in wholesale and retail trade, which was also one of the sectors that displays a large decline in productivity.

We use a rich longitudinal database on the Portuguese labour market to estimate a panel quantile regression model with nonadditive fixed effects (Powell, 2016). This is a useful econometric methodology to test whether the effect of the 2004 change in legislation and whether the tradable differential are heterogeneous over the relative wage distribution.

This paper contributes to the literature in two main ways. Firstly, to the best of our knowledge, this is the first study on the impact of asymmetric employment protection reforms on the distribution of the within-firm wage gap between permanent and fixed-term contracts. Secondly, we provide original empirical evidence on whether the effect of a larger EPL gap on the intra-firm wage differential is heterogeneous between tradable and non-tradable sectors.

Our results suggest that the 2004 change in legislation increased wage inequality between fixed-term and open-ended contracts in those firms at the median and top quantiles of the relative wage distribution by 1.6% and 1%, respectively. This result is in line with Elia (2010), who argues that fixed-term contracts are a new source of wage inequality. We also find evidence that although the relative wage is not differently affected by the change in legislation in firms in the tradable sector, these firms have higher within-firm wage inequality between both types of contract at the top of the conditional relative wage distribution. The relative wage of permanent contracts is 1.2% and 2.3% larger, at the 75th and 90th quantiles, respectively, in firms in the tradable sector. Finally, our results suggest that fixed-term and permanent workers are close substitutes in firms at the bottom and at the median of the relative wage distribution and are imperfect substitutes in firms where there is higher wage inequality between both types of contract.

The paper is organised as follows. Section 2 provides a brief characterisation of

the Portuguese labour market and describes the change in legislation under study. In section 3, we review the relevant literature and section 4 describes the dataset and presents the econometric methodology. Finally, sections 5 and 6 discuss the main results and draw conclusions about the policy implications of our results, respectively.

4.2 The Portuguese Labour Market

The Portuguese labour market is characterised by strong rigidity of nominal wages[2] and stringent employment protection on permanent contracts. Wages are determined under collective bargaining, usually held at the sector or occupation level (Marques et al., 2010). Until 2011, collective agreements were frequently extended to all the workers and firms in a sector of activity by the Government through extension clauses (Martins, 2015; Marques et al., 2010). Therefore, wages in the Portuguese labour market are bound by the national minimum wage and by the wage floor resulting from the wage bargaining (Portugal, 2008).

In such a setting, characterised by nominal wage rigidity, low inflation, and low scope for firms to adjust employment due to the high adjustment costs entailed by open-ended contracts, fixed-term contracts can be important instruments for firms to adjust employment. The reforms promoting flexibility at the margin and the strict employment protection on open-ended contracts led to an accentuated increase in the proportion of fixed-term contracts in Portugal over the last decades and in 2012, workers on fixed-term contracts represented approximately 20% of total workforce.

After the introduction of fixed-term contracts in 1976, the regulations on their use suffered several changes. One of the most relevant changes occurred in 2004, when an extraordinary renewal was introduced and the maximum duration of fixed-term contracts was extended from three to six years. This change in legislation also allowed firms to hire workers on fixed-term contracts to indirectly substitute an absent employee. In 2009, this legislative change was overturned for new contracts and the maximum legal duration of a fixed-term contract was re-established at three

[2]Nominal wage cuts are not allowed by Portuguese legislation (Portugal, 2008).

years. Although the employment protection of open-ended contracts remained quite stable during this period, the employment protection gap between fixed-term and open-ended contracts increased also in firms with 11 to 20 employees due to the increase in the procedural costs incurred in permanent worker's dimissal for firms of this size (Centeno and Novo, 2012).

Finally, although the Labour Code of 2003 introduced an equal treatment clause between workers on fixed-term and open-ended contracts, there is evidence of a wage premium for the average permanent worker (Silvério, 2015).

4.3 Literature Review

Fixed-term contracts may help firms to reduce wage costs (e.g. for the Netherlands, see Kleinknecht et al., 2006). In fact, the literature usually finds that fixed-term contracts receive, on average, lower wages than open-ended contracts (e.g. Jimeno and Toharia, 1993; Blanchard and Landier, 2002; Hagen, 2002), especially at the bottom of the worker's wage distribution for countries like Italy (Bosio, 2014) and Germany (Mertens and McGinnity, 2003; Mertens et al., 2007; Pfeifer, 2012).

As predicted by Boeri (2011), stricter EPL on open-ended contracts increases the wage premium of permanent over temporary contracts. However, this effect may depend on labour market tightness (Cao et al., 2010). If the job finding and vacancy filling probabilities remain constant (no search externalities), increasing firing costs on permanent contracts raise both the relative wage of workers on permanent contracts and the proportion of temporary contracts. However, if there are search externalities, the effects on wage inequality are negligible, because the proportion of temporary contracts decreases although the relative wage of permanent workers increases.

The promotion of flexibility at the margin may also have important consequences for wage setting (Jimeno and Toharia, 1993; Bentolila and Dolado, 1994) by increasing wage pressure and wage dispersion (Dolado et al., 2002). As predicted by insider-outsider theory, the existence of high labour turnover costs for insiders increases their bargaining power (Lindbeck and Snower, 1988). Bentolila and Dolado (1994) build upon this theory and argue that the strength of permanent contracts

in wage bargaining is reinforced by the existence of fixed-term contracts with lower firing costs, especially in countries where the employment protection on open-ended contracts is high. According to these authors, wages can be affected by the proportion of temporary contracts through the buffer and bargaining effects[3]. The higher proportion of fixed-term contracts can increase the wage growth experienced by workers on permanent contracts since the latter may be used as a buffer stock for permanent contracts and decrease the probability of a worker on a permanent contract being dismissed. The proportion of temporary employees may also increase the bargaining power of permanent workers if the latter threaten to not cooperate with the former. However, the bargaining power may also be reduced due to the lower strike frequency of temporary workers. The authors find empirical evidence of both the buffer and bargaining effects and estimate that the proportion of fixed-term contracts has a positive impact on the wage growth experienced by workers on open-ended contracts and a negative impact on the average labour cost per worker. This latter effect results from the fact that workers on fixed-term contracts suffer a wage penalty (Bentolila and Dolado, 1994), possibly due to wage discrimination, as pointed out by Jimeno and Toharia (1993). The results of Bentolila and Dolado (1994) hold for a panel of Spanish firms and a panel of sectors of activity in Denmark, France and Germany. However, the effects are more significant for countries with strong employment protection for open-ended contracts, which suggests that the bargaining effect is lower in countries such as Denmark.

Therefore, the introduction of fixed-term contracts in the Spanish labour market may have initially contributed to decrease the wage drift[4], due to the wage penalty suffered by workers on this type of contract. However, the increasing share of fixed-term contracts may have also contributed to strengthen the insiders' bargaining position and eventually contributed to increase the wage drift (Bentolila and Dolado, 1994). This empirical observation is in line with Jimeno and Toharia's (1993) findings of a positive impact of the proportion of fixed-term contracts in the

[3] Bentolila and Dolado (1994) also mention a composition effect when the researcher is not able to observe the wage by type of contract.

[4] The wage drift is the difference between the wage growth defined in collective bargaining and the observed average wage growth.

bargained wage growth. Besides, the use of fixed-term contracts as an adjustment tool in firms where unions have a higher coverage and bargain for higher wages for open-ended contracts is also documented by Dolado et al. (2002). These authots estimate that the proportion of fixed-term contracts on total employment is negatively affected by the proportion of workers covered by a union.

Job matching models also provide some insights about how the relative wage between permanent and fixed-term contracts may be affected by asymmetric employment protection reforms. Dolado et al. (2007) find that decreasing the firing costs of low-productivity workers does not have an expressive effect on average wages but increases wage inequality between high and low-productivity workers [5], while a reform reducing the firing costs for high-productivity workers shrinks the wage gap. Masui (2013) distinguishes the wage setting mechanism for open-ended and temporary contracts in a job matching model - the wages of workers on open-ended contracts are determined through collective bargaining and workers on fixed-term contracts receive a wage equal to their reservation wage. According to this model, the wage of permanent workers is positively (negatively) affected by the firing costs on permanent (temporary) contracts. The effects of firing costs of temporary workers on their wages depend on the average productivity and wages of workers on permanent contracts and on the contract's conversion.

Recent empirical evidence on the wage effects of a larger employment protection gap between open-ended and fixed-term contracts is scarce. Elia (2010) estimate that the introduction of new forms of temporary contracts (and the ease of the use of the apprenticeship contract [6]) in Italy increased the wage gap between permanent and temporary contracts between 8.2% and 10% in the short and long-run due to the reduction in temporary worker's wage. This effect is larger for skilled workers, which as Elia (2010) argues may indicate that fixed-term contracts are used as screening devices. Likewise, Bosio (2014) finds that this effect is larger for high

[5] As measured by the ratio of wages of high productivity workers and wages of low productivity workers.

[6] The circumstances allowing to hire a worker on a fixed-term contract were broadened and the requisites for an apprenticeship contract, such as age limit and certification of qualifications, were extended.

skilled workers at the bottom of the wage distribution. There is evidence, however, that the introduction of new forms of temporary contracts in Italy, had a negative effect on the wage of new permanent workers (Ordine and Rose, 2016). For Italy, there is also evidence that a 1% hiring of fixed-term contracts is valued by firms as equivalent to a reduction in permanent's workers' wage between 1.3% and 2.8%, which may explain the considerable increase in fixed-term contracts in Italy between 1995 and 2003 (Cipollone and Guelfi, 2006).

In the Portuguese labour market there is empirical evidence that fixed-term contracts facilitate employment adjustment (e.g. Varejão and Portugal, 2007) but also serve as screening devices (e.g. Portugal and Varejão, 2009). Previous literature finds that fixed-term workers suffer the burden of a larger employment protection gap between open-ended and fixed-term contracts. Higher employment protection for open-ended contracts increases excess worker turnover (Centeno and Novo, 2012) and reduces the wage of workers on fixed-term contracts (Centeno and Novo, 2014). On the other hand, extending the maximum duration of the fixed-term contract reduces and postpones the conversion of the contract and has a negative impact on the wage growth experienced by workers on this type of contract. However, and in line with the screening hypothesis, this negative effect is especially felt by workers whose contract was not converted to an open-ended contract (see Chapter 2).

Finally, reference should be made to the heterogeneous effects of fixed-term contracts on average wages and employment among firms and sectors of activity (Bentolila and Dolado, 1994; Kleinknecht et al., 2006). In services, where the proportion of fixed-term contracts is larger, the buffer and bargaining effects are less accentuated than in the manufacturing sector (Bentolila and Dolado, 1994). However, to the best of our knowledge, there are no studies addressing whether the effects of reforms increasing the employment protection gap between fixed-term and open-ended contracts are asymmetric between tradable and non-tradable sectors. Moreover, we find no empirical evidence of the impact of changes in EPL on the distribution of the relative wage between open-ended and fixed-term contracts at the firm level.

4.4 Empirical Methodology

4.4.1 *Quadros de Pessoal*

The analysis is conducted using *Quadros de Pessoal*, an administrative linked employer-employee database collected every year in October. This exceptionally rich database on the Portuguese labour market provides information on every employer in the private sector and their employees.

Quadros de Pessoal provides information on firm's sector of activity, location, constitution date, turnover and total number of employees. Moreover, it allows us to match firms and workers and, therefore, to characterise the firm's workforce in terms of personal characteristics (gender, education, nationality, age) and occupation level, tenure and type of contract. It also includes detailed information on wages and since the information is reported by the firm and is made publicly available, misreporting and error measurement are reduced.

We restrict the analysis to firms operating in mainland Portugal. Moreover, similarly to Cardoso et al. (2012), we only consider full time workers with a fixed-term or an open-ended contract, aged 18 to 65, earning at least 80% of the minimum wage[7]. Outliers in the worker's wage distribution, i.e., worker's observations below the 2nd and above the 99th percentile of the wage distribution, are also excluded from the analysis.

The unit of observation is the firm. After all the exclusions, we are left with an unbalanced panel of 162,480 different firms observed between 2002 and 2011[8], which corresponds to a total of 541,983 observations.

The real wage is defined in an hourly basis and equals the sum of base wages, overtime pay and regular benefits. This variable was deflated using the Consumer Price Index (2012=100). We then calculate the average wage[9] paid to workers on permanent contracts and to those on fixed-term contracts for each firm and year

[7] Workers with missing wage or type of contract, working more than 400 hours/month and with normal period of work lower than 120 hours are also dropped.

[8] Information on the contract type is only available since 2002. We restrict the analysis until 2011 to avoid capturing the effects of other changes in legislation, namely another extension of the maximum duration of fixed-term contracts occured in 2012.

[9] We use this location measure similarly to Edo (2016).

and compute the logarithm of their ratio.

The variable capturing the change in legislation varies over firms and time and it is defined as follows. It is a dummy variable taking value one between 2004 and 2008 and, after that, for firms that have a worker on a fixed-term contract hired until February 2009 [10]. The dummy takes value zero in 2002, 2003 and for those firms only employing workers on fixed-term contracts hired after February 2009.

The tradable sectors are defined as those that are exposed to international competition but the classification of firms in tradable and nontradable sectors is not trivial (Dwyer, 1992) and several methodologies are addressed in the literature. In this study we follow the classification in Catarino et al. (2006)[11]. Therefore, the tradable sector comprises the following two-digits sectors of activity: i) agriculture and fisheries; ii) extractive industries; iii) manufacturing; iv) transports; v) financial activities; vi) property. The non-tradable sector includes the remaining sectors, i.e.: i) production and distribution of electricity and water; ii) construction; iii) wholesale and retail; iv) lodging and restaurants; v) public administration; vi) education; vi) health and social work; vii) collective, social and personal services.

4.4.2 Econometric Methodology

Firms with different wage policies with respect to fixed-term and permanent contracts may use the former for different purposes, namely as screening and adjustment devices. Thus, studying only the mean effects of flexibility at the margin on the relative wage may provide a misleading and incomplete picture. Quantile regression methods allow us to estimate the impact of the 2004 change in legislation over the conditional quantile functions of the relative wage between permanent and fixed-term contracts (Koenker and Hallock, 2001). We therefore estimate the following panel quantile regression model with firm fixed effects:

[10] Note that the change in legislation was overturned only for new contracts in 2009.
[11] These authors use the classification criteria developed by Dwyer (1992).

$$Q_{w_{ft}}(\tau|x_{ft}, E_{ft}, Leg_{ft}, Trad_{ft}) = \beta(\tau)x_{ft} + \delta(\tau)E_{ft} + \sigma(\tau)Leg_{ft} + \kappa(\tau)Trad_{ft} + \gamma_t, \tag{4.1}$$

$$w_{ft} = Q_{w_{ft}}(\tau|x_{ft}, E_{ft}, Leg_{ft}, Trad_{ft}) + U^*_{ft}. \tag{4.2}$$

$Q_{w_{ft}}(\tau|x_{ft}, E_{ft}, Leg_{ft}, Trad_{ft})$ is the conditional τth quantile of w_{ft}, where $\tau \epsilon (0, 1)$. w_{ft} is the logarithm of the ratio between permanent and fixed-term contracts' wages ($\log(\frac{W_{OEC}}{W_{FTC}})$) at firm f and time t[12]. The covariates x_{ft} includes firm's size and firm's age dummies, the share of public and foreign capital, the proportion of workers by gender and immigrant status, education and occupation levels and the average age (in logs) of workers employed by the firm[13]. We also include in the model three treatment variables and their interactions: the logarithm of the ratio of permanent to fixed-term number of contracts ($E = \log(\frac{E_{OEC}}{E_{FTC}})$), a dummy representing the change in legislation (Leg_{ft}) and a dummy for tradable sectors ($Trad_{ft}$). Since the effects of the ratio of permanent to fixed-term workers on relative wages may have been shifted when the employment protection gap between both types of contracts widened, we include the interaction between the relative employment and the legislation dummy. In fact, this interaction term may capture the buffer effect (Bentolila and Dolado, 1994) and the composition effect of the change in legislation on the relative wage. Finally, we also allow the effect of the change in legislation to differ between firms in tradable and non-trabable sectors, by including the interaction between the legislation and tradable dummies.

The model includes both time fixed effects (γ_t), which allows the distribution of the relative wage to change over the years and firm fixed effects (α_f). U^*_{ft} is an unknown function of a firm fixed effect (α_f) and the disturbance term (ε_{ft}). Note that in this setting it is important to include firm fixed effects because it accounts

[12] We use the log of the variables to obtain the (semi-)elasticities of the dependent variable with respect to the independent variables. The results remain qualitatively unchanged if we did not transformed the ratio of relative wages.

[13] Note that we did not include region or sector dummies because they show low within variation in the data, which would result in lower efficiency in the fixed effects estimation.

for unobserved firm's characteristics, such as the relative productivity of permanent contracts that may be arbitrarily correlated with the treatment and control variables. This allows us to control for possible unobserved differences between those firms that benefited from the higher flexibility at the margin and those that did not and may help to mitigate possible concerns about the selection of firms with at least one fixed-term and one open-ended contract.

As Koenker (2004) states, it is not feasible to estimate a fixed effect for each quantile τ, thus the fixed effect has a location-shift effect on the conditional quantile function of the relative wage. Most quantile regression models for panel data, such as that introduced in Koenker (2004), include additive fixed effects (Powell, 2016). This technique, however, does not allow for the coefficients to vary with the fixed effect and their interpretation is modified since the distribution turns out to be $(w_{ft} - \alpha_f)|E_{ft}, Leg_{ft}, Trad_{ft}, x_{ft}$ instead of $w_{ft}|E_{ft}, Leg_{ft}, Trad_{ft}, x_{ft}$[14](Powell, 2016).

To overcome this issue, we implement the estimation method for quantile regression with nonadditive fixed effects proposed by Powell (2016). Identification is achieved through the within-firm variation.

As Powell (2016) argues, this method has several advantages: i) it allows the estimated coefficients to vary with some function of the fixed-effect and the disturbance term (U^*_{ft}); ii) it allows to mantain the interpretation of the coefficients as in standard cross-sectional quantile regression and iii) this method is consistent for samples with a low number of time periods, since the fixed effects are not estimated.

The model is estimated using a Generalised Method of Moments (GMM) estimator with two moment conditions as stated in Corollary 3.2 in Powell (2016). The standard errors are clustered at the firm level[15].

According to the literature presented in Section 3, we expect that, in the period in which the change in legislation increasing the employment protection gap between

[14]In this setting, we would estimate the effect of the treatment variables in the conditional distribution of the difference between relative wage and the unobserved characteristics of the firm.

[15]Note that we cannot compare the estimates over different quantiles without correcting the standard errors. We intend to do so by bootstrapping the standard errors, which is a time consuming exercise. Nevertheless, there is at least one variable for which the estimated coefficients change signs and remain statistical significant over the distribution of the relative wage, which evidences the suitability of the quantile regression.

both types of contract was in force, the wage premium paid to the average permanent worker increases, i.e. the wage ratio increases in the higher quantiles of the conditional distribution of $\log(\frac{W_{OEC}}{W_{FTC}})$. Moreover, we expect that this effect is smaller for firms with a higher ratio of permanent to fixed-term contracts (i.e., we expect that the interaction between relative permanent employment and the legislation dummy to be negative). Finally, we expect a higher wage inequality between workers on fixed-term and open-ended contracts in tradable sectors (i.e., a positive sign for the coefficient associated with the dummy for tradable sectors) and that these sectors are more affected by the change in legislation (i.e., a positive sign for the interaction of the tradable dummy with the legislation dummy), especially for the higher end of the distribution.

4.4.3 Descriptive Statistics

Figure 4·1 presents the graphical representation of the empirical quantiles of the relative wage distribution (in logs) and Table 4.1 presents the characteristics of the firms located at the mean and at the 10th, 50th and 90th quantiles of the empirical distribution of the relative wages (in logs).

Figure 4·1: Quantile plot of the relative wage

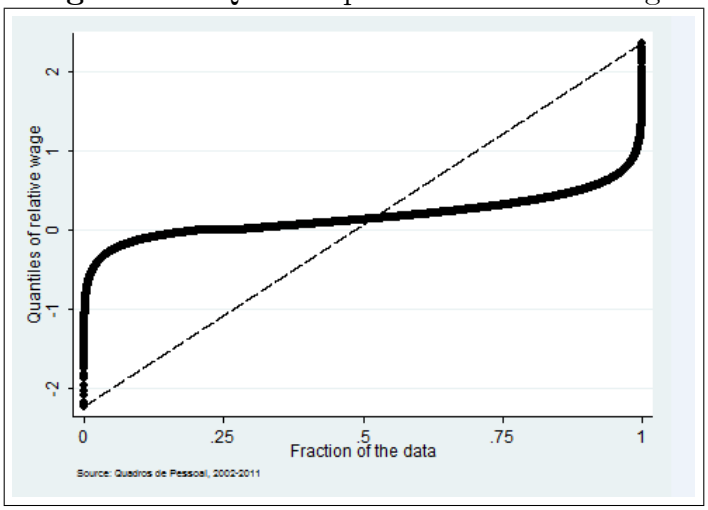


Table 4.1: Descriptive statistics

Variables	Mean	p10	p50	p90
wage_{FTC}	4.39	5.06	4.04	4.39
	(1.95)	(2.09)	(1.50)	(1.60)
wage_{OEC}	5.29	4.39	4.57	7.32
	(2.68)	(1.82)	(1.70)	(2.67)
relative employment (logs)	0.61	0.40	0.68	0.82
	(1.17)	(1.20)	(1.09)	(1.26)
E_{FTC}	5.88	3.92	5.91	7.58
	(24.28)	(10.42)	(20.03)	(30.03)
E_{OEC}	16.31	8.78	14.09	29.48
	(124.89)	(62.05)	(48.70)	(255.65)
Firms' characteristics				
dimension (number of employees)	31.35	17.72	27.79	50.20
	(194.57)	(72.60)	(83.64)	(349.28)
age (number of years)	15.81	13.77	16.09	17.20
	(20.61)	(19.53)	(18.74)	(17.76)
tradable sectors (%)	36.92	34.43	35.96	41.57
	(48.26)	(47.52)	(47.99)	(49.29)
public capital (%)	0.49	0.30	0.31	0.83
	(6.63)	(5.32)	(5.37)	(8.64)
foreign capital (%)	3.08	1.58	1.98	6.51
	(16.71)	(12.07)	(13.43)	(23.76)
Employment composition				
females (%)	43.97	41.60	42.86	42.80
	(34.74)	(34.40)	(35.38)	(30.99)
immigrants (%)	6.13	6.24	6.22	4.90
	(15.99)	(16.82)	(15.21)	(13.54)
average workers' age (log)	3.60	3.60	3.60	3.61

(continued)

(continuation)

Variables	Mean	p10	p50	p90
	(0.17)	(0.18)	(0.17)	(0.16)
Average workers' education				
less than high school	67.74	68.85	74.35	54.62
	(32.35)	(32.50)	(29.04)	(32.63)
high school	21.91	20.82	18.80	28.56
	(24.89)	(25.13)	(22.86)	(24.82)
bachelor degree	1.97	1.86	1.35	3.34
	(6.69)	(6.86)	(5.18)	(8.28)
college degree	8.38	8.47	5.50	13.48
	(17.56)	(18.06)	(14.07)	(20.53)
Average workers' occupation				
managers	4.58	4.12	3.18	7.30
	(11.23)	(11.37)	(9.36)	(12.96)
experts	6.36	6.95	4.07	9.63
	(16.72)	(17.80)	(13.51)	(19.15)
intermediate-level technicians	10.33	9.81	8.26	15.44
	(19.51)	(19.53)	(17.10)	(21.48)
admnistrative staff	14.12	13.99	12.05	20.00
	(21.78)	(21.73)	(20.06)	(24.07)
sellers	21.18	19.09	22.34	15.62
	(33.05)	(31.33)	(33.48)	(27.02)
qualified workers from agriculture	1.51	1.49	1.37	1.24
	(10.50)	(10.41)	(9.88)	(9.30)
craftsmen	19.77	22.70	23.94	14.71
	(31.19)	(32.66)	(32.88)	(25.96)
machine operators	10.10	10.09	11.19	6.84
	(23.21)	(23.11)	(23.90)	(16.77)

(continued)

(continuation)

Variables	Mean	p10	p50	p90
unqualified workers	11.35	11.02	12.93	8.45
	(21.10)	(21.00)	(21.57)	(16.44)
Total observations	541,983	5,419	5,420	5,420

Source: Quadros de Pessoal, 2002-2011. Notes: Standard deviations in parantheses.

At the mean of the relative wage distribution, the logarithm of the ratio between permanent and fixed-term wages is approximately 0.16 log points with a standard deviation of 0.29 log points and, therefore, there is evidence of a wage premium of 0.16 log points for the average worker on an open-ended contract. For instance, workers on fixed-term contracts receive, on average, 4.39€/hour, while workers on open-ended contracts receive approximately 5.29€/hour. Moreover, the average firm employs approximately six workers on fixed-term contracts for each sixteen workers on open-ended contracts. Firms have an average dimension of 31 employees and are approximately 16 years old. Almost 37% of the firms in the sample are in the tradable sector. Female workers represent approximately 44% of the total firm's workforce and 6% are immigrants. Finally, firm's workforce is characterised by a low level of education (approximately 68% of total workforce have less than high school education and only 8% have college education) and is allocated in low qualified occupations (approximately 41% of the workforce are salespeople and craftsmen).

However, there is a substantial heterogeneity in firm's characteristics and employment composition over the distribution of the ratio between permanent and fixed-term contracts' wages. The distribuition of the relative wage appears to be symmetric (Figure 4·1). Firms in the 10th quantile pay a slightly higher wage to workers on fixed-term than to workers on open-ended contracts (approximately -0.14 log points). At the median of the relative wage distribution, the wage penalty experienced by fixed-term contracts is equal to 0.12 log points, which is below the corresponding figure at the mean. In the 10% of firms that offer a higher relative wage to permanent contracts the wage gap is approximately 0.51 log points. This

is due to the higher wage for the average worker on an open-ended contract at the higher percentiles of the distribution, since the wage of the average fixed-term worker is in line with that at the mean of the distribution.

The firms in the bottom of the distribution are also smaller (on average 18 workers) and slightly younger (approximately 14 years of activity) and have a higher proportion of the total workforce on fixed-term contracts than firms at the top quantiles of the distribution. At the 10th quantile of the relative wage distribution, there is also a lower proportion of workers with a college degree and a larger proportion of less qualified occupations (especially craftsmen and unqualified workers) in comparison with the 90th quantile. Finally, it is important to note that the share of firms in the tradable sector is higher at the 90th quantile of the relative wage distribution.

This preliminary analysis supports the necessity to use an econometric model that allows us to study the conditional distribution of the ratio of permanent to fixed-term wages.

4.5 Results

We first estimate the model for the conditional mean using a fixed effects estimator[16] (Table E.1). The results show that the change in legislation increased the mean relative wage between workers on permanent and fixed-term contracts in approximately 1% and this effect is statistically significant at a 99% confidence level (column (1) of Table E.1). However, if we include year dummies, the coefficient of the legislative change becomes negative and it is still statistically significant (column (2) of Table E.1), which may be due to the fact that the dummy accounting for the change in legislation depends on the year. We, therefore, re-estimate the model with a linear trend instead, and obtain similar results to those that do not control for time effects (column (3) of Table E.1). There is also evidence, at a 90% confidence level, that the wage differential between both types of contract is larger in firms

[16] There is evidence of the existence of individual effects (the p-value of the Breusch and Pagan LM test for random effects is equal to zero), and these effects are fixed according to the Hausman test (p-value=0.0000).

Table 4.2: Quantile estimates of the elasticity of substitution between open-ended and fixed-term contracts

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	-0.0009	0.0000	-0.0006	-0.0059***	-0.0164***
	(0.0013)	(0.0005)	(0.0007)	(0.0010)	(0.0014)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	no	no	no	no	no
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs).

in the tradable sector. However, the higher wage inequality in these firms was not statistically significant when the change in legislation was in force [17].

One of the advantages of studying the distribution of the relative wage is to assess whether the estimated elasticity of substitution between open-ended and fixed-term contracts differ between firms with high and low wage inequality. The usual framework to obtain the elasticity of substitution between both types of contract is a constant elasticity of substitution production function (Dolado et al., 2002; Jimeno and Toharia, 1993; Edo, 2016). Therefore, we present the results of a quantile regression of the relative wage on the relative employment with fixed effects and without any covariates in Table 4.2.

The lack of statistical significance of the coefficients associated with the ratio of permanent to fixed-term employment for the firms in the lower and median quantiles of the distribution supports the notion that fixed-term contracts are close substitutes of open-ended contracts. This result is in line with the evidence provided by Centeno and Novo (2012). However, at the top of the distribution of the relative wages, the effect of relative employment is negative and statistically significant, which may indicate that workers on fixed-term and open-ended contracts are im-

[17] The p-value of the Wald test with the null hypothesis: $Trad_{ft} + Leg_{ft}XTrad_{ft} = 0$ is higher than 0.5 for all the regressions in Table E.1. This result holds for firms with $E_{ft} \neq 0$.

perfect substitutes in these firms. As Edo (2016) argues, if wages of workers on fixed-term and open-ended contracts are not equally flexible, as it may be the case for Portugal, the estimates of the elasticity of substitution are lower bounds of the true elasticity. This preliminary evidence shows that it is important to consider the whole distribution of relative wages when estimating the elasticity of substitution, since firms that offer higher wages to their workers on permanent contracts may not consider workers on temporary contracts as perfect substitutes of the former [18]. Additionally, the relative wage may be informative about the role that fixed-term contracts play within the firm, since this may indicate that fixed-term contracts are potentially used as buffer stocks instead of screening devices for permanent workers in firms at the higher quantiles of the relative wage distribution.

We now proceed to the analysis of the estimation results for the full model specified in Equation (4.1) (Table 4.3).

In the period in which the change in legislation was in force, the relative wage of permanent contracts increased, but the effect is only statistically significant, at a 1% and 10% significance levels, at the median and the 75th quantile of the relative wage distribution, respectively. Firms at the median of the relative wage distribution increase the ratio between permanent and fixed-term contracts' wages in approximately 1.6%, ceteris paribus. Note that the coefficient associated with the legislation dummy loses statistical significance when we include the covariates in the model[19] (Table E.2). Moreover, we only capture an indirect effect of the increase in the employment protection gap through the employment composition by type of contract in the results of the model with no covariates (except for the 90th quantile) (Table E.2). As expected, at the median and top quantiles of the distribution, the increase in the relative wage due to the change in legislation is lower for firms with a higher number of permanent to fixed-term contracts, which may give some hint that fixed-term contracts may be used as a buffer stock for open-ended contracts in

[18] This may be justified if fixed-term contracts are allocated to less qualified occupations, for example.

[19] Note, however, that the marginal effect of the change in legislation is only statistically significant, at a 5% significance level, at the 50th quantile and the 75th quantile as in the model with covariates.

Table 4.3: Quantile estimates of the effect of the change in legislation on the relative wage

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0065***	0.0060***	0.0069***	0.0029*	-0.0053**
	(0.0019)	(0.0009)	(0.0011)	(0.0016)	(0.0023)
tradable	0.0060	0.0035	0.0063	0.0118**	0.0230**
	(0.0075)	(0.0028)	(0.0040)	(0.0053)	(0.0090)
legislation	-0.0011	0.0028	0.0161***	0.0096*	0.0081
	(0.0080)	(0.0024)	(0.0037)	(0.0058)	(0.0094)
lemp_rel $\times trad$	-0.0016	-0.0002	-0.0009	-0.0002	-0.0011
	(0.0028)	(0.0014)	(0.0018)	(0.0024)	(0.0037)
lemp_rel $\times leg$	0.0027	0.0008	-0.0003	-0.0020	-0.0038*
	(0.0020)	(0.0009)	(0.0012)	(0.0015)	(0.0022)
lemp_rel $\times trad \times leg$	-0.0017	-0.0004	0.0002	0.0011	0.0035
	(0.0030)	(0.0014)	(0.0018)	(0.0023)	(0.0033)
leg $\times trad$	-0.0014	0.0002	-0.0011	-0.0038	-0.0072
	(0.0042)	(0.0015)	(0.0025)	(0.0034)	(0.0055)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

these firms (Table E.2).

We also find evidence supporting the hypothesis that firms in tradable sectors may use fixed-term contracts as a source of wage flexibility. The wage premium of open-ended contracts is larger and statistically significant, at a 95% confidence level, for firms in the tradable sector at the top of the distribution of the relative wage. For the 10% of firms paying a higher wage premium for permanent contracts, those in tradable sector pay an additional premium of 2.3% (Table 4.3). However, there is no significant evidence that the wage ratio in firms in tradable sectors is affected by changes in legislation that widen the employment protection gap between open-ended and fixed-term contracts[20].

Although the ratio of permanent to fixed-term employment has a statistically significant effect on the relative wage over the whole distribution, the economic impact is quite low (0.01% at the 10th to 50th quantiles and -0.01% at the 90th quantile).

We, therefore, conclude that firms at the median and at the top of the relative wages distribution (at least at the 75th quantile) react to changes in legislation that increase the employment protection gap between both types of contract by increasing the wage inequality between them. This may indicate that firms paying a wage premium to the workers on open-ended contracts are those that increase it even further as a response to the larger employment protection gap, but not those with the highest wage premium (at the 90th quantile of the conditional relative wage distribution) possibly due to the lower scope to increase it even more. On the other hand, those firms with lower wage inequality or that pay higher wages to workers on fixed-term contracts do not translate higher quantitative flexibility in higher wage inequality. The different roles of fixed-term contracts in the labour market- as screening devices or buffer stocks for open-ended contracts- may contribute to explain these results.

Finally, it is important to notice that the employment composition by education

[20]For the model with no covariates, at the 75th and 90th quantiles, firms in tradable sector had lower wage differentials between both contracts when the change in legislation was in force. This may be due to the fact that firms in tradable sector already have higher levels of wage inequality between open-ended and fixed-term contracts and therefore react less to the changes in legislation.

and occupation level has negligible economic effects, although they are statistically significant in most quantiles (Table E.3). The ratio between permanent and fixed-term contracts' wages seems to increase with the size of the firm in all quantiles, but especially at the bottom of the distribution. Older firms also seem to pay a higher relative wage to workers on permanent contracts (or a lower wage premium to fixed-term contracts) for those firms at the bottom of the relative wage distribution.

4.5.1 Estimation using an adaptive Markov chain Monte Carlo (MCMC) algorithm

The results discussed above were estimated using the Nelder-Mead optimisation method. However, we cannot compare the estimated coefficients over different quantiles without correcting the standard errors. Inference methods for quantile regression estimators are substantially studied and there are various resampling techniques available (Koenker and Hallock, 2001).

As Powell (2016) reports, Markov chain Monte Carlo (MCMC) is an adequate numerical optimisation technique for models with a large number of treatment variables as it is the case of the model under study. We therefore re-estimate the model using an adaptive Markov chain Monte Carlo algorithm[21] with 5,000 draws, where the first 500 were burned, and an acceptance rate of 0.4, similarly to Smith (2015)[22]. The proposal distribution is a multivariate normal density as described in Baker (2014).

In broad terms, the previous results are robust to this alternative optimisation technique, although most of the estimates of the effect of the treatment variables are now statistically significant (Tables 4.4 and E.4).

It is worth nothing that the effect of the change in legislation that increased flexibility at the margin in the Portuguese labour market is now statistically significant at the bottom of the relative wage distribution. These results may indicate that the change in legislation contributes to increase the inequalities at the bottom and at the

[21] This is not intended to be an exhaustive analysis since we plan to alternatively bootstrap the standard errors, but it is a less time consuming exercise that allows us to draw conclusions about the estimates equality along the distribution quantiles.

[22] The initial value of the random-number seed was set to 71436018. The initial parameter values and covariance matrix result from the cross-sectional quantile regression.

top of the relative wage distribution. Not only the wage disadvantage of the average permanent worker is increased in approximately 1%, at the 10th quantile, as their wage advantage is increased in firms with the largest relative wage for permanent contracts in 0.6% and 1.9%, at the 75th and 90th quantiles, respectively. This effect is therefore increasing over the firms' relative wage distribution, although firms in the middle part of the distribution appear to be unresponsive to the change in legislation. These results help to explain the negative sign of the change in legislation at the mean of the distribution when controlling for time effects (column (2) of Table E.1). Thus, mean effects of the change in legislation give us an incomplete picture on how the relative wage distribution is affected by flexibility at the margin and they are likely influenced by the lower quantiles of the relative wage distribution.

There is evidence of higher and increasing wage inequality between workers on both types of contract in firms in tradable sectors at the top of the conditional relative wage distribution. However, the marginal effect of the change in legislation is negative and statistically significant for firms in the tradable sector over the entire distribution. This indicates that in the tradable sector the wage inequalities are only magnified in firms at the bottom of the distribution where the average permanent worker may receive a lower wage than the average fixed-term contract. In the firms where fixed-term contracts face a wage disadvantage, the inequalities are reduced in approximately 0.3% at the 50th and 75th quantiles and 9% at the 90th quantile[23].

Finally, this Bayesian method allows us to conclude about the equality of the estimates over the different quantiles. The non-overlapping confidence intervals of the estimates of the dummies accounting for the change in legislation and the tradable sectors show that the effects of the treatment variables are heterogeneous over the relative wage distribution (Table E.5). A quantile regression is therefore a suitable econometric methodology to address the research questions.

[23]The p-value of these Wald tests are lower than 0.05.

Table 4.4: Quantile estimates of the effect of the change in legislation on the relative wage using an adaptive MCMC algorithm

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0058***	0.0066***	0.0145***	0.0044***	-0.0113***
	(0.0008)	(0.0018)	(0.0017)	(0.0002)	(0.0004)
tradable	0.0103***	0.0184***	0.0171***	0.0223***	0.0393***
	(0.0009)	(0.0018)	(0.0013)	(0.0003)	(0.0018)
legislation	-0.0099***	0.0004	0.0033	0.0058***	0.0186***
	(0.0024)	(0.0007)	(0.0027)	(0.0006)	(0.0011)
lemp_rel_trad	-0.0080***	-0.0038***	-0.0025	0.0007*	-0.0042***
	(0.0014)	(0.0003)	(0.0024)	(0.0004)	(0.0013)
lemp_rel_leg	0.0020	0.0017***	-0.0032	-0.0057***	-0.0149***
	(0.0022)	(0.0006)	(0.0020)	(0.0009)	(0.0008)
lemp_rel_trad_leg	0.0067**	0.0046***	0.0032*	0.0057***	0.0150***
	(0.0030)	(0.0007)	(0.0017)	(0.0015)	(0.0013)
leg_trad	-0.0133***	-0.0133***	-0.0067***	-0.0085***	-0.0273***
	(0.0009)	(0.0029)	(0.0012)	(0.0009)	(0.0016)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

4.5.2 Robustness Analysis

The positive effect of the change in legislation on wage inequality at the firm level may be the result of the behaviour of firms that already were employing workers on fixed-term and open-ended contracts before 2004 and/or the result of the behaviour of firms that start hiring fixed-term or open-ended contracts after that. In order to disentangle the source of the positive sign associated with the legislation dummy at the top quantiles of the relative wage distribution, we re-estimate the model with only those firms that already employed at least one worker on each type of contract in 2002 or 2003, before the change in legislation was implemented (Table E.6). With this sample, we still observe the statistical significant (at a 95% confidence level) and higher wage premium for open-ended contracts in firms in tradable sector at the top of the conditional relative wage distribution. However, the coefficient on the legislation dummy loses statistical and economic significance. This result may be reconciled with the evidence that wages are adjusted especially for newly hired workers rather than for insiders (e.g. Centeno and Novo, 2014; Martins and Portugal, 2014; Carneiro et al., 2012). Note that the change in legislation extended the maximum legal duration of fixed-term contracts such that firms were allowed to continue the employment relationship, and that nominal wage cuts are forbidden in the Portuguese legislation. Therefore, the effect of the change in legislation reported in Table 4.3 may be potentially explained by firms that start hiring new workers either on fixed-term contracts with lower average wages or workers on open-ended contracts with higher average wages.

Another interesting result is obtained by retrieving those firms that were excluded from the sample because the firm does not have an open-ended or a fixed-term contract in a given year, but hired them in other year(s) of the sample period. In order to perform this analysis we substitute the missings on the average wage by type of contract by the firm's time average, which more than doubles the number of firms retained in the sample. The results displayed in Table E.7 give further support to the previous conclusion that the positive effect of the change in legislation on the relative wage at the top of the distribution is undermined by lower representation of fixed-term contracts in the firm's employment composition. According to these

results, the change in legislation does not have a direct significant effect on wage inequality, but for firms with higher relative permanent employment the effect of the change in legislation is negative and statistically significant for all quantiles, except at the 10th. The higher permanent wage premium for firms in tradable sectors is still observed using this alternative dependent variable.

The results are robust to the exclusion of firms in agriculture sector[24] (Table E.8). Since firms with 11 to 20 employees experienced a larger increase in the employment protection gap between fixed-term and open-ended contracts due to the more demanding procedures to layoff workers on open-ended contracts, we interact the variables of interest with a dummy identifying these firms (d1120). The results confirm that the effect of the change in legislation on the intra-firm wage gap between both types of contract is not driven exclusively by these firms although the effect is 5% larger and statistically significant, at a 5% and 10% significance levels, in the 50th and 75th quantiles respectively (Table E.9).

4.6 Conclusion

We use exceptionally rich employer-employee data and estimate a quantile regression model with nonadditive fixed effects to study how the conditional distribution of the within-firm wage gap between open-ended and fixed-term contracts may be affected by reforms widening the employment protection asymmetries between both types of contract. We also assess whether the relative wage and the effects of employment protection reforms are heterogeneous across firms in tradable and non-tradable sectors.

Our results suggest that fixed-term and permanent workers are close substitutes in firms situated at the bottom and at the median of the relative wage distribution. However, we find that, at the top quantiles of that distribution, both types of contracts are imperfect substitutes. Our results show that, when the 2004 change in legislation was in force, wage inequality between open-ended and fixed-term con-

[24] Since the coverage of firms in agriculture sector is low in Quadros de Pessoal (some firms do not have paid workers), the sample may not be representative of this sector.

tracts significantly increased, at a 90% confidence level, in firms that pay a large wage premium to permanent workers. Notably, wage inequality increased by 1.6% in firms at the median and 1% in firms at the 75th quantile of the conditional relative wage distribution. On the other hand, wage inequality does not seem to be significantly affected by the change in legislation in firms at the bottom of the relative wage distribution. We interpret this result as possible evidence that firms attribute different uses to fixed-term contracts and this is reflected in their wage policies. Namely, reforms increasing the EPL gap only put fixed-term workers in a more disadvantaged position in firms where they face a large wage penalty. By contrast, in firms where the average fixed-term worker face a lower wage penalty or even receive a wage premium, wage inequality may not be significantly affected.

Finally, our results suggest that firms in the tradable sectors pay a higher wage premium to permanent workers at the top of the conditional relative wage distribution. This may be explained by the fact that these firms are more constrained due to the international competition and, therefore, use fixed-term contracts as a source of wage flexibility. However, they are not significantly affected by changes in legislation extending the maximum legal duration of fixed-term contracts.

This study contributes to drawing attention to the potential externalities of the promotion of flexibility at the margin. To the best of our knowledge, this study gives the first insights on how Portuguese firms with different wage policies for workers on different types of contract react to changes in EPL. Another novelty of our results is that as the relative wage of permanent contracts is higher in firms in the tradable sector, the labour demand for fixed-term contracts in these sector may be higher than that for open-ended contracts- if they are equally productive. These results are relevant for policy makers since they draw attention to the importance of encouraging investment in training for workers on fixed-term contracts and of policies promoting the contract's conversion, which may be important to sustain the necessary increase in this sector's long-run productivity.

Further research is needed to overcome some of the limitations of this study. For instance, an important exercise would be to account for possible sample selection

for firms that hire workers on fixed-term contracts[25], which may be challenging in a quantile regression with nonadditive fixed effects. Note, however, that the inclusion of fixed effects in the model reduces the concern about sample selection correction. A natural extension of this study is to assess the role of fixed-term contracts in job and worker reallocation between tradable and non-tradable sectors. Finally, it would also be interesting to analyse whether the effects of the change in legislation are due to the increase of the average wage paid to workers on open-ended contracts or to the lower average wages of fixed-term contracts and whether the probability of conversion of fixed-term into open-ended contracts plays a role in the explanation of the heterogeneous effects of changes in legislation over the relative wage distribution.

[25] We estimate the Heckman model at the mean of the distribution using a maximum likelihood estimator and did not find significant evidence of sample selection at the standard confidence levels.

Appendix E

Tables

Table E.1: Mean estimates of the effect of the change in legislation on the relative wage

	(1)	(2)	(3)
Variables	Mean	Mean	Mean
lemp_rel	-0.0062***	-0.0068***	-0.0061***
	(0.0012)	(0.0012)	(0.0012)
tradable	0.0075*	0.0075*	0.0075*
	(0.0043)	(0.0043)	(0.0043)
legislation	0.0103***	-0.0084***	0.0107***
	(0.0014)	(0.0019)	(0.0015)
lemp_rel $\times trad$	-0.0028	-0.0027	-0.0028
	(0.0019)	(0.0019)	(0.0019)
lemp_rel $\times leg$	-0.0041***	-0.0039***	-0.0042***
	(0.0012)	(0.0012)	(0.0012)
lemp_rel $\times trad \times leg$	0.0003	0.0002	0.0003
	(0.0018)	(0.0018)	(0.0018)
leg $\times trad$	-0.0053**	-0.0053**	-0.0052**
	(0.0025)	(0.0025)	(0.0025)
Constant	0.2058***	0.1775***	1.4513***
	(0.0210)	(0.0231)	(0.5133)
firm fixed effects	yes	yes	yes
time effects	no	yes	linear trend
controls	yes	yes	yes
Observations	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480
R-squared	0.0123	0.0132	0.0124

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

Table E.2: Quantile estimates of the effect of the change in legislation on the relative wage without covariates

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0006	0.0005	0.0081***	0.0034*	-0.0080***
	(0.0022)	(0.0009)	(0.0013)	(0.0018)	(0.0022)
tradable	0.0079	0.0016	0.0041	0.0090	0.0240**
	(0.0079)	(0.0026)	(0.0043)	(0.0058)	(0.0097)
legislation	0.0057	0.0005	0.0302***	0.0264***	0.0207**
	(0.0082)	(0.0018)	(0.0037)	(0.0058)	(0.0090)
lemp_rel×$trad$	0.0002	0.0005	-0.0065***	-0.0071***	-0.0025
	(0.0035)	(0.0013)	(0.0019)	(0.0024)	(0.0035)
lemp_rel×leg	0.0004	-0.0005	-0.0069***	-0.0107***	-0.0133***
	(0.0023)	(0.0009)	(0.0013)	(0.0016)	(0.0023)
lemp_rel×$trad \times leg$	-0.0063*	-0.0010	-0.0019	0.0035	0.0077**
	(0.0036)	(0.0013)	(0.0019)	(0.0023)	(0.0034)
leg×$trad$	-0.0050	0.0004	-0.0004	-0.0063*	-0.0132**
	(0.0045)	(0.0013)	(0.0027)	(0.0036)	(0.0060)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	no	no	no	no	no
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs).

Table E.3: Full quantile estimates of the effect of the change in legislation on the relative wage

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
dimension (5-9)	0.0168***	0.0034***	0.0148***	0.0165***	0.0142***
	(0.0044)	(0.0010)	(0.0022)	(0.0032)	(0.0051)
dimension (10-19)	0.0374***	0.0161***	0.0320***	0.0312***	0.0185***
	(0.0058)	(0.0016)	(0.0029)	(0.0042)	(0.0062)
dimension (20-49)	0.0527***	0.0315***	0.0490***	0.0441***	0.0257***
	(0.0073)	(0.0025)	(0.0036)	(0.0051)	(0.0074)
dimension (50-99)	0.0686***	0.0459***	0.0622***	0.0513***	0.0268***
	(0.0099)	(0.0042)	(0.0046)	(0.0066)	(0.0093)
dimension (100-249)	0.0829***	0.0586***	0.0717***	0.0544***	0.0237*
	(0.0163)	(0.0063)	(0.0065)	(0.0077)	(0.0121)
dimension (250-499)	0.1088***	0.0664***	0.0707***	0.0483***	0.0162
	(0.0251)	(0.0116)	(0.0095)	(0.0113)	(0.0187)
dimension (500-999)	0.1052***	0.0743***	0.0706***	0.0423***	0.0080
	(0.0250)	(0.0222)	(0.0137)	(0.0149)	(0.0378)
dimension (>999)	0.0707	0.0733***	0.0935***	0.0556*	0.0032
	(0.0607)	(0.0257)	(0.0174)	(0.0293)	(0.0277)
age (2-5)	0.0074*	0.0011	-0.0008	-0.0043	-0.0119**
	(0.0041)	(0.0014)	(0.0025)	(0.0038)	(0.0059)
age (6-10)	0.0174***	0.0058***	0.0024	-0.0051	-0.0167**
	(0.0050)	(0.0017)	(0.0032)	(0.0045)	(0.0074)
age (>10)	0.0285***	0.0133***	0.0130***	0.0106**	-0.0004
	(0.0060)	(0.0022)	(0.0039)	(0.0053)	(0.0086)
cspub	0.0005	0.0002	0.0002	0.0002	0.0003
	(0.0005)	(0.0002)	(0.0003)	(0.0003)	(0.0003)

(continued)

(continuation)

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
csest	0.0005***	0.0004***	0.0005***	0.0003***	0.0003***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
lemp_rel	0.0065***	0.0060***	0.0069***	0.0029*	-0.0053**
	(0.0019)	(0.0009)	(0.0011)	(0.0016)	(0.0023)
tradable	0.0060	0.0035	0.0063	0.0118**	0.0230**
	(0.0075)	(0.0028)	(0.0040)	(0.0053)	(0.0090)
legislation	-0.0011	0.0028	0.0161***	0.0096*	0.0081
	(0.0080)	(0.0024)	(0.0037)	(0.0058)	(0.0094)
lemp_rel $\times trad$	-0.0016	-0.0002	-0.0009	-0.0002	-0.0011
	(0.0028)	(0.0014)	(0.0018)	(0.0024)	(0.0037)
lemp_rel $\times leg$	0.0027	0.0008	-0.0003	-0.0020	-0.0038*
	(0.0020)	(0.0009)	(0.0012)	(0.0015)	(0.0022)
lemp_rel $\times trad \times leg$	-0.0017	-0.0004	0.0002	0.0011	0.0035
	(0.0030)	(0.0014)	(0.0018)	(0.0023)	(0.0033)
leg $\times trad$	-0.0014	0.0002	-0.0011	-0.0038	-0.0072
	(0.0042)	(0.0015)	(0.0025)	(0.0034)	(0.0055)
workers_females	0.0002*	0.0001***	-0.0001*	-0.0004***	-0.0004***
	(0.0001)	(0.0000)	(0.0000)	(0.0001)	(0.0001)
workers_immigrant	0.0005***	0.0002***	0.0004***	0.0004***	0.0001
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)
workers_highsch	-0.0000	0.0000	0.0004***	0.0005***	0.0007***
	(0.0001)	(0.0000)	(0.0000)	(0.0001)	(0.0001)
workers_bachelor	-0.0006*	0.0002**	0.0009***	0.0008***	0.0019***
	(0.0003)	(0.0001)	(0.0002)	(0.0002)	(0.0003)
workers_college	-0.0014***	-0.0002*	0.0005***	0.0006***	0.0009***
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)

(continued)

(continuation)

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
workers_managers	-0.0002	0.0008***	0.0034***	0.0054***	0.0070***
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0002)
workers_experts	-0.0011***	-0.0003***	0.0000	0.0003***	0.0010***
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
workers_intermediate	-0.0005***	-0.0001***	0.0004***	0.0006***	0.0012***
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)
workers_administrative	-0.0005***	-0.0001***	0.0003***	0.0005***	0.0010***
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)
workers_sellers	-0.0000	-0.0001***	0.0000	0.0002***	0.0002*
	(0.0001)	(0.0000)	(0.0000)	(0.0001)	(0.0001)
workers_qual_agric	-0.0001	0.0000	0.0000	0.0002	0.0003
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0003)
workers_craftsmen	-0.0001	-0.0001***	-0.0001***	-0.0001*	-0.0000
	(0.0001)	(0.0000)	(0.0000)	(0.0001)	(0.0001)
workers_machine	-0.0002	-0.0001***	-0.0003***	-0.0004***	-0.0004***
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)
age_workers	-0.0614***	-0.0091***	0.0426***	0.0262***	0.1155***
	(0.0099)	(0.0029)	(0.0061)	(0.0083)	(0.0152)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, *

p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Base cateogories are: dimension(<5 workers), age(1 year), education(share of workers with less than high school), occupation(share of unqualified workers).

Table E.4: Quantile estimates of the elasticity of substitution between open-ended and fixed-term contracts using an adaptive MCMC algorithm

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	-0.0008	0.0000	-0.0005	-0.0060***	-0.0162***
	(0.0007)	(0.0000)	(0.0004)	(0.0006)	(0.0008)
	(-0.0022 , 0.0005)	(-0.0000 , 0.0001)	(-0.0014 , 0.0003)	(-0.0071 , -0.0049)	(-0.0178 , -0.0147)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	no	no	no	no	no
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, and 95% confidence intervals in parantheses.

*** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs).

Table E.5: 95% Confidence Intervals of the quantile estimates using an adaptive MCMC algorithm

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
dimension (5-9)	(0.0361 , 0.0395)	(0.0084 , 0.0090)	(0.0234 , 0.0268)	(0.0167 , 0.0210)	(0.0055 , 0.0119)
dimension (10-19)	(0.0738 , 0.0778)	(0.0317 , 0.0344)	(0.0617 , 0.0627)	(0.0543 , 0.0566)	(0.0265 , 0.0293)
dimension (20-49)	(0.1075 , 0.1155)	(0.0643 , 0.0675)	(0.0921 , 0.0948)	(0.0779 , 0.0802)	(0.0360 , 0.0410)
dimension (50-99)	(0.1260 , 0.1365)	(0.0870 , 0.0922)	(0.1138 , 0.1167)	(0.0757 , 0.0849)	(0.0362 , 0.0445)
dimension (100-249)	(0.1507 , 0.1547)	(0.0994 , 0.1124)	(0.1438 , 0.1446)	(0.1069 , 0.1076)	(0.0374 , 0.0419)
dimension (250-499)	(0.1974 , 0.2239)	(0.1444 , 0.1495)	(0.1362 , 0.1589)	(0.0772 , 0.0958)	(0.0167 , 0.0220)
dimension (500-999)	(0.1832 , 0.2023)	(0.1332 , 0.1343)	(0.1428 , 0.1653)	(0.1053 , 0.1086)	(0.0189 , 0.0305)
dimension (>999)	(0.1463 , 0.1639)	(0.1549 , 0.1767)	(0.1614 , 0.1963)	(0.0786 , 0.1001)	(0.0388 , 0.0587)
age (2-5)	(0.0252 , 0.0355)	(0.0042 , 0.0120)	(-0.0031 , 0.0047)	(-0.0117 , -0.0078)	(-0.0308 , -0.0279)
age (6-10)	(0.0523 , 0.0593)	(0.0194 , 0.0239)	(0.0109 , 0.0193)	(0.0008 , 0.0028)	(-0.0230 , -0.0186)
age (>10)	(0.0716 , 0.0760)	(0.0279 , 0.0344)	(0.0249 , 0.0337)	(0.0174 , 0.0237)	(-0.0035 , -0.0008)
cspub	(0.0002 , 0.0008)	(0.0005 , 0.0008)	(0.0004 , 0.0006)	(0.0002 , 0.0003)	(0.0005 , 0.0005)
csest	(0.0006 , 0.0007)	(0.0005 , 0.0008)	(0.0009 , 0.0009)	(0.0010 , 0.0010)	(0.0012 , 0.0012)
lemp_rel	(0.0042 , 0.0074)	(0.0031 , 0.0100)	(0.0111 , 0.0179)	(0.0041 , 0.0047)	(-0.0122 , -0.0105)
tradable	(0.0086 , 0.0121)	(0.0149 , 0.0220)	(0.0146 , 0.0197)	(0.0217 , 0.0229)	(0.0358 , 0.0429)
legislation	(-0.0145 , -0.0053)	(-0.0010 , 0.0019)	(-0.0019 , 0.0085)	(0.0046 , 0.0070)	(0.0165 , 0.0207)
lemp_rel×$trad$	(-0.0108 , -0.0052)	(-0.0044 , -0.0032)	(-0.0071 , 0.0022)	(-0.0001 , 0.0016)	(-0.0067 , -0.0017)
lemp_rel×leg	(-0.0023 , 0.0062)	(0.0006 , 0.0028)	(-0.0072 , 0.0007)	(-0.0076 , -0.0039)	(-0.0165 , -0.0133)
lemp_rel×$trad \times leg$	(0.0009 , 0.0125)	(0.0033 , 0.0059)	(-0.0000 , 0.0065)	(0.0027 , 0.0088)	(0.0126 , 0.0175)
leg×$trad$	(-0.0150 , -0.0116)	(-0.0190 , -0.0076)	(-0.0090 , -0.0043)	(-0.0103 , -0.0067)	(-0.0304 , -0.0241)
workers_females	(-0.0000 , 0.0001)	(-0.0001 , -0.0001)	(-0.0004 , -0.0003)	(-0.0007 , -0.0007)	(-0.0010 , -0.0010)
workers_immigrant	(0.0003 , 0.0003)	(0.0002 , 0.0002)	(0.0001 , 0.0001)	(0.0001 , 0.0002)	(0.0004 , 0.0005)
workers_highsch	(-0.0000 , 0.0001)	(0.0002 , 0.0004)	(0.0006 , 0.0006)	(0.0011 , 0.0011)	(0.0014 , 0.0014)
workers_bachelor	(-0.0015 , -0.0013)	(-0.0001 , 0.0001)	(0.0015 , 0.0020)	(0.0034 , 0.0035)	(0.0037 , 0.0039)
workers_college	(-0.0016 , -0.0015)	(-0.0001 , 0.0002)	(0.0011 , 0.0012)	(0.0022 , 0.0022)	(0.0030 , 0.0030)
workers_managers	(-0.0004 , -0.0000)	(0.0005 , 0.0006)	(0.0029 , 0.0030)	(0.0053 , 0.0055)	(0.0073 , 0.0075)
workers_experts	(-0.0019 , -0.0018)	(-0.0005 , -0.0005)	(0.0001 , 0.0002)	(0.0010 , 0.0010)	(0.0019 , 0.0019)
workers_intermediate	(-0.0007 , -0.0005)	(-0.0002 , 0.0000)	(0.0008 , 0.0010)	(0.0016 , 0.0017)	(0.0024 , 0.0025)
workers_administrative	(-0.0004 , -0.0003)	(-0.0001 , 0.0001)	(0.0009 , 0.0011)	(0.0015 , 0.0016)	(0.0018 , 0.0019)
workers_sellers	(-0.0001 , 0.0001)	(-0.0001 , 0.0000)	(0.0000 , 0.0003)	(0.0002 , 0.0003)	(0.0002 , 0.0003)
workers_qual_agric	(-0.0008 , -0.0007)	(-0.0002 , -0.0001)	(-0.0002 , 0.0000)	(0.0001 , 0.0002)	(0.0004 , 0.0005)
workers_craftsmen	(-0.0004 , -0.0003)	(-0.0003 , -0.0002)	(-0.0002 , -0.0001)	(-0.0002 , -0.0002)	(-0.0003 , -0.0002)
workers_machine	(-0.0003 , -0.0002)	(-0.0005 , -0.0004)	(-0.0006 , -0.0004)	(-0.0008 , -0.0008)	(-0.0010 , -0.0009)
age_workers	(-0.1007 , -0.0846)	(-0.0221 , -0.0130)	(0.0062 , 0.0294)	(0.0753 , 0.0862)	(0.1408 , 0.1537)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: The dependent variable is the ratio of permanent wages to fixed-term wages (in logs).

Table E.6: Quantile estimates of the effect of the change in legislation on the relative wage for firms with fixed-term contracts before 2004

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0065**	0.0064***	0.0072***	0.0043**	-0.0039
	(0.0026)	(0.0012)	(0.0015)	(0.0021)	(0.0030)
tradable	0.0075	0.0062	0.0105**	0.0151**	0.0206**
	(0.0091)	(0.0040)	(0.0052)	(0.0066)	(0.0105)
legislation	0.0039	0.0019	0.0032	0.0032	0.0015
	(0.0188)	(0.0059)	(0.0077)	(0.0115)	(0.0174)
lemp_rel $\times trad$	-0.0025	-0.0006	-0.0010	-0.0011	-0.0014
	(0.0033)	(0.0018)	(0.0022)	(0.0027)	(0.0041)
lemp_rel $\times leg$	-0.0010	-0.0010	-0.0023	-0.0025	-0.0025
	(0.0025)	(0.0012)	(0.0015)	(0.0019)	(0.0030)
lemp_rel $\times trad \times leg$	-0.0021	-0.0007	0.0016	0.0022	0.0049
	(0.0033)	(0.0018)	(0.0021)	(0.0027)	(0.0038)
leg $\times trad$	-0.0007	0.0001	-0.0033	-0.0058	-0.0100
	(0.0046)	(0.0020)	(0.0030)	(0.0040)	(0.0065)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	288,807	288,807	288,807	288,807	288,807
Number of groups	60,190	60,190	60,190	60,190	60,190

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

Table E.7: Quantile estimates of the effect of the change in legislation on an alternative variable of the relative wage

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0027***	0.0031***	0.0054***	0.0043***	-0.0001
	(0.0009)	(0.0005)	(0.0005)	(0.0007)	(0.0011)
tradable	0.0028	0.0026	0.0045**	0.0074**	0.0115**
	(0.0045)	(0.0024)	(0.0022)	(0.0033)	(0.0053)
legislation	-0.0025	-0.0032	-0.0022	0.0000	0.0069
	(0.0045)	(0.0021)	(0.0025)	(0.0038)	(0.0062)
lemp_rel $\times trad$	0.0003	0.0001	-0.0005	-0.0005	-0.0022
	(0.0015)	(0.0008)	(0.0008)	(0.0011)	(0.0017)
lemp_rel $\times leg$	-0.0011	-0.0023***	-0.0032***	-0.0031***	-0.0040***
	(0.0008)	(0.0004)	(0.0004)	(0.0006)	(0.0009)
lemp_rel $\times trad \times leg$	-0.0005	0.0006	0.0012*	0.0013	0.0015
	(0.0013)	(0.0007)	(0.0007)	(0.0010)	(0.0015)
leg $\times trad$	-0.0017	-0.0004	-0.0011	-0.0017	-0.0031
	(0.0025)	(0.0012)	(0.0012)	(0.0019)	(0.0032)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	1,206,913	1,206,913	1,206,913	1,206,913	1,206,913
Number of groups	200,350	200,350	200,350	200,350	200,350

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs), where the average wage by type of contract is replaced in a given year by the respective firm's average wage over the sample period when there are no workers on fixed-term or open-ended contracts. Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

Table E.8: Quantile estimates of the effect of the change in legislation on the relative wage excluding agriculture

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0067***	0.0061***	0.0069***	0.0029*	-0.0053**
	(0.0019)	(0.0009)	(0.0011)	(0.0016)	(0.0023)
tradable	0.0065	0.0041	0.0066	0.0133**	0.0272***
	(0.0074)	(0.0029)	(0.0041)	(0.0054)	(0.0087)
legislation	0.0011	0.0024	0.0162***	0.0097*	0.0081
	(0.0078)	(0.0023)	(0.0038)	(0.0054)	(0.0084)
lemp_rel $\times trad$	-0.0017	-0.0003	-0.0008	0.0004	-0.0003
	(0.0029)	(0.0014)	(0.0018)	(0.0024)	(0.0035)
lemp_rel $\times leg$	0.0032*	0.0007	-0.0004	-0.0020	-0.0038*
	(0.0019)	(0.0009)	(0.0011)	(0.0015)	(0.0022)
lemp_rel $\times trad \times leg$	-0.0016	-0.0002	0.0004	0.0012	0.0036
	(0.0031)	(0.0014)	(0.0018)	(0.0023)	(0.0033)
leg $\times trad$	0.0028	-0.0001	-0.0012	-0.0040	-0.0078
	(0.0043)	(0.0016)	(0.0025)	(0.0035)	(0.0055)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	529,427	529,427	529,427	529,427	529,427
Number of groups	158,445	158,445	158,445	158,445	158,445

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

Table E.9: Quantile estimates of the effect of the change in legislation on the relative wage interacted with a dummy for firms with 11 to 20 employees

	(1)	(2)	(3)	(4)	(5)
Quantile	10	25	50	75	90
lemp_rel	0.0063***	0.0057***	0.0065***	0.0024	-0.0060**
	(0.0022)	(0.0010)	(0.0013)	(0.0016)	(0.0025)
tradable	0.0030	0.0025	0.0061	0.0109*	0.0282***
	(0.0076)	(0.0028)	(0.0042)	(0.0057)	(0.0095)
legislation	0.0051	0.0022	0.0155***	0.0039	0.0078
	(0.0080)	(0.0023)	(0.0036)	(0.0056)	(0.0085)
lemp_rel $\times trad$	-0.0019	-0.0005	-0.0007	0.0002	-0.0005
	(0.0033)	(0.0015)	(0.0020)	(0.0026)	(0.0040)
lemp_rel $\times leg$	0.0032	0.0010	-0.0001	-0.0015	-0.0034
	(0.0023)	(0.0010)	(0.0013)	(0.0016)	(0.0024)
lemp_rel $\times trad \times leg$	-0.0030	-0.0008	-0.0003	0.0005	0.0033
	(0.0035)	(0.0016)	(0.0020)	(0.0024)	(0.0037)
leg $\times trad$	0.0007	0.0007	-0.0003	-0.0030	-0.0067
	(0.0047)	(0.0016)	(0.0027)	(0.0037)	(0.0062)
d1120$\times lemp_rel$	0.0013	0.0013	0.0022	0.0022	0.0027
	(0.0037)	(0.0017)	(0.0019)	(0.0026)	(0.0037)
d1120$\times tradable$	0.0115	0.0047	0.0015	0.0007	-0.0021
	(0.0076)	(0.0033)	(0.0041)	(0.0062)	(0.0099)
d1120$\times legislation$	0.0058	0.0020	0.0048**	0.0052*	0.0014
	(0.0038)	(0.0018)	(0.0023)	(0.0031)	(0.0054)
d1120$\times lemp_rel \times trad$	0.0014	0.0002	-0.0006	-0.0014	-0.0016
	(0.0059)	(0.0028)	(0.0033)	(0.0045)	(0.0072)
d1120$\times lemp_rel \times leg$	-0.0033	-0.0021	-0.0023	-0.0025	-0.0011
	(0.0040)	(0.0020)	(0.0023)	(0.0030)	(0.0045)
d1120$\times lemp_rel \times trad \times leg$	0.0055	0.0025	0.0029	0.0034	-0.0003
	(0.0065)	(0.0031)	(0.0039)	(0.0050)	(0.0080)
d1120$\times leg \times trad$	-0.0086	-0.0040	-0.0058	-0.0044	-0.0019
	(0.0086)	(0.0037)	(0.0050)	(0.0072)	(0.0115)
firm fixed effects	yes	yes	yes	yes	yes
time fixed effects	yes	yes	yes	yes	yes
controls	yes	yes	yes	yes	yes
Observations	541,983	541,983	541,983	541,983	541,983
Number of groups	162,480	162,480	162,480	162,480	162,480

Source: Quadros de Pessoal, 2002-2011. Notes: Standard errors, clustered at firm level, in parantheses. *** p<0.01, ** p<0.05, * p<0.1. The dependent variable is the ratio of permanent wages to fixed-term wages (in logs). Control variables include: firm's dimension dummies and firm's age dummies, firm's capital ownership and employment composition (share of workers by education and occupation, average workers' age and proportion of females and immigrants among workforce).

Chapter 5

Conclusion

This Chapter intends to sum and discuss the main conclusions of the three essays that compose this thesis and to present a few policy implications and suggestions for further research.

We study the effects of the promotion of flexibility at the margin on the Portuguese labour market, using up to date econometric techniques and an exceptionally rich database.

During the period in which the maximum legal duration of fixed-term contracts was allowed to increase from three to six years: i) the probability of conversion of the fixed-term contract decreased; ii) fixed-term workers experienced a slower real wage growth; iii) wage inequality between permanent and fixed-term workers became more pronounced; iv) the maintenance of the fixed-term contract was promoted through a lower fixed-term job creation; and v) there was an indirect substitution of open-ended contracts for fixed-term contracts.

Our results suggest that fixed-term contracts can play different roles in the Portuguese labour market and that this should be considered when analysing the effects of the promotion of flexibility at the margin. Namely, we find evidence that fixed-term contracts are used as a screening devices, given that there is a significant correlation between the likelihood of a fixed-term worker be given an open-ended contract and experiencing an above average wage growth. Besides, workers with converted fixed-term contracts face a lower reduction in wage growth than workers that remained with fixed-term contracts in the years in which the contract was allowed to be extended for a longer period of time. Furthermore, this legislative change did not significantly impact the intra-firm wage inequality between both types of contracts in firms where the average fixed-term worker faces a lower wage disadvantage.

However, we also find evidence that fixed-term contracts are used as buffer stocks for permanent workers, given that in some firms the average fixed-term worker faces a high wage penalty relatively to the average permanent worker and in these firms the elasticity of substitution is negative and statistically significant, which supports the idea that in these cases both types of contract are imperfect substitutes.

Therefore, we show that the 2004 change in legislation that facilitated the use of fixed-term contracts had heterogeneous effects across workers and firms, which may be at least partly explained by the different roles played by fixed-term contracts. However, by facilitating the use of fixed-term contracts for a longer period of time, policy makers may shift the incentive towards the use of fixed-term contracts as buffer stocks rather than as screening devices. According to Faccini's (2014) model, if we consider that fixed-term contracts are screening devices, extending the maximum duration of the contract from two to three years is unlikely to have very significant effects since it does not greatly improve the learning about match quality.

Besides failing to foster employment growth, this legislative change generated important inefficiencies. Firstly, it not only reduced the probability that a worker on a fixed-term contract could be promoted to a more stable employment relationship but also postponed the conversion of the contract. By doing so, and according to our results, it contributed even further to increase wage inequality between workers on different contractual arrangements. The results reported in Chapter 2 and Chapter 4 show that there are mainly two sources of the increased wage inequality between both types of contract in the period in which the legislative change was in force. Not only does the average fixed-term worker face a higher relative wage penalty when compared to the average permanent worker, especially in firms at the top of the relative wage distribution, but he also experiences a lower wage growth rate when compared to the newly permanent workers that were converted from a fixed-term contract.

As documented in Chapter 3, the share of non-converted fixed-term contracts may also help to explain the job dynamics that follow an increase in flexibility at the margin. Not only did the extension of the maximum duration of fixed-term contracts increase the fixed-term job destruction in the two following periods, but it

also decreases the rate at which they are created until 3-years ahead. We, therefore, advance that the main effect of the 2004 change in legislation has been to promote the maintenance of the existing contracts, while the overall employment growth did not seem to be significantly affected. We found that it is the non-conversion of fixed-term contracts that primarily accounts for these results. In addition, the existence of a higher wage gap between both types of contracts may contribute to explaining the lack of a significantly positive effect of the change in legislation on employment growth. Namely, we find evidence that the change in legislation promoted an indirect substitution of permanent workers for fixed-term workers (Chapter 3). In particular, this could be especially relevant for the firms with lower wage differentials between the average permanent worker and the average fixed-term worker where both types of contract may be considered closer substitutes (Chapter 4).

The analysis conducted over the three essays provides new evidence on the role of fixed-term contracts in the Portuguese labour market and helps understand the main effects that the promotion of flexibility at the margin may have. We show that flexibility at the margin has a significant impact on major labour market outcomes. The main message of this thesis is that asymmetric employment protection reforms entail externalities, such as the postponement of the conversion of the fixed-term contract and higher inequality between not only permanent and fixed-term contracts but also between workers that receive an open-ended contract and those that were not able to get a stable employment relationship. Workers on fixed-term contracts are affected not only in terms of wages (in the short and long-run), but also in terms of employment opportunities and prospects. However, these costs are not spread homogeneously among them, given that the contract may serve different purposes. These effects may be attenuated if the contract is used mainly as a screening device, promoting a more efficient labour allocation. Designing effective policy measures is, therefore, very important. We formulate some suggestions of policy implications suggested by the results obtained.

Policy Implications

According to the OECD Employment Outlook "(...) policy makers are increasingly aware of the risks for efficiency and social cohesion of relying solely on tempor-

ary contracts for labour market adjustments" (OECD, 2014, p. 145). Our findings suggest that policy makers should tackle labour market segmentation through the reduction of employment protection on open-ended contracts or the increase in the restrictions on the use of fixed-term contracts. Although the former is politically more difficult to implement (Saint-Paul, 1996), the high levels of unemployment experienced during the economic and economic and financial crisis of 2008-9 combined with the increasing proportion of fixed-term contracts and their lower conversion probability may originate the necessary political support to implement it in Portugal (e.g. Saint-Paul, 1993). According to Lindbeck and Snower (2001) and Lindbeck and Snower (2002), in a labour market with high labour turnover costs for open-ended contracts, only large labour market reforms may be effective in boosting employment. Therefore, we advocate the need to promote a structural reform that reduce segmentation, which might improve labour market functioning (OECD, 2012).

According to our results, one of the most relevant policies should be to design effective incentives to promote the conversion of fixed-term contracts. This policy could have positive effects on employment dynamics, namely by increasing job creation of fixed-term contracts. Although the extension of the contract was accompanied by a penalisation in the social security contribution for firms with more than 15% of the workforce under a fixed-term contract, the results show that the measure was not effective in promoting the conversion of the contract. A potential alternative would be to implement measures encouraging the conversion of the contract in firms that penalise more fixed-term contracts, for example, in terms of wages. According to our results, these firms are those that potentially react more to changes in labour costs and that in general may use fixed-term contracts for rotation more intensively.

Another way to promote the conversion of matches started with a fixed-term contract would be to increase the mandatory hours of relevant training and to increase the monitoring of the compliance of this type of change. The mandatory training for fixed-term contracts was introduced in the Labour Code of 2003, but our results give preliminary evidence that it did not have a sufficiently positive effect to mitigate the negative effects of flexibility at the margin on wage inequality and contract's conversion.

Understanding the role played by fixed-term contracts in the Portuguese labour market and the effects of reforms that widened the employment protection asymmetries between fixed-term and regular contracts may also contribute to the ongoing discussion about the introduction of a single employment contract (e.g. Pérez and Osuna, 2014; Bentolila, Dolado and Jimeno, 2012). Though the results provided in this thesis do not allow drawing a grounded conclusion about this alternative, we believe that some light was shed on the fact that its benefits and costs have to be evaluated considering each country's institutional setting and taking into account the way in which labour market agents use fixed-term contracts.

Limitations and Further Research

The analysis reported here has some limitations that should be addressed in the future. We should be careful in extrapolating the results of this thesis to countries with different institutional frameworks and to other periods of time. For instance, the results give a good hint about the effects of the increased flexibility at the margin in Portugal in 2012 and 2013, but as Kahn (2010) argues, this type of reform is likely to have stronger effects on temporary employment during recessions than expansions. This may be partly explained by the fact that the value of a temporary contract is larger during the former phase of the cycle (Kahn, 2010). The results are, therefore, dependent on the institutional context and the economic environment of Portugal in the sample period. Nevertheless, this last concern is mitigated since the effect of the business cycle is controlled for in all regressions. We believe that the results pinpoint important hints on how to conduct the analysis of asymmetric employment protection reforms in the future, at least in countries characterised by high labour market segmentation, such as Spain or France. Other relevant institutions, such as unemployment benefits and wage setting mechanisms, may interact with flexibility at the margin and contribute to increase its effects. Therefore, future studies should also consider these interactions to better characterise those effects on wages, job flows and other labour market outcomes.

Further research on these topics should take into account that the match quality threshold at which fixed-term contracts are converted into open-ended contracts may have also been affected by the change in legislation. *A priori*, we would expect

that this threshold increases, given that firms can experiment the match longer. Besides, in Chapter 2, the sample only included continuing matches, i.e., matches whose quality was high enough such that the match survived. The analysis could be improved if we considered also those matches that did not survived. However, we consider that these limitations do not hinder the results, which are treated as lower bounds of the true effect. In the future, we intend to estimate the match quality cutoff points at which fixed-term contracts are continued and converted and how the change in legislation shifted them.

It would also be interesting to test the robustness of the results in Chapter 2 to an alternative simultaneous estimation method such as maximum likelihood or GMM instead of the two-step procedure proposed.

Furthermore, the analysis in Chapter 4 could also be extended by estimating the model with an alternative dependent variable: the wage ratio of entrants to fixed-term workers, where entrants are workers on permanent contracts with less than one year of tenure. This is especially relevant given the results previously found in the literature showing that entrants' wages may also be negatively affected by reforms increasing the EPL gap between fixed-term and open-ended contracts, namely for Portugal (e.g. Centeno and Novo, 2014) and Italy (e.g. Ordine and Rose, 2016).

Finally, the study of the effectiveness of fiscal incentives to convert fixed-term contracts in open-ended contracts implemented in some European countries has been neglected so far. Our results suggest that this type of measure is likely to be innefective when implemented with measures allowing the extension of the contract. Further research on this issue is thus called for, given that according to our conclusions, policies aiming to promote the conversion of the contract may help to attenuate the negative effects of the increasing segmentation of the Portuguese labour market.

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