

## “Making work pay” in a rationed labor market

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**Abstract** This paper empirically analyzes the labor supply effects of two “making work pay” reforms in Germany. We provide evidence in favor of policies that distinguish between low effort and low productivity by targeting individuals with *low wages* rather than those with *low earnings*. We discuss our results more generally and with comparisons to the family-based tax credits in force in the US and the UK. For the evaluation of the policies, we apply a static structural labor supply framework and explicitly account for demand-side constraints by using a double-hurdle model.

**Keywords** Tax-benefit systems · Household labor supply · Involuntary unemployment

**JEL Classification** C52 · H31 · J22

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## 1 Introduction

“Making work pay” policies are usually targeted at people who face the highest risk of unemployment. These measures have been introduced in many Organization for Economic Cooperation and Development countries in recent years. A growing body of literature discusses various issues surrounding policy design and the effectiveness of these policies in alleviating poverty and boosting employment. Most of the discussion has been conducted in the light of the earned income tax credit (EITC) in the US and the British working tax credit (WTC) (cf. Eissa and Hoynes 2004; Blundell 2000). The tax credits in the US and the UK are conditioned on joint family income and, therefore, induce negative work incentives for secondary earners in addition to positive effects on the primary earner. In contrast, transfer programs based on individual earnings do not affect the partner’s work incentives directly, and thus, they avoid the direct negative effect on the secondary earner. In this respect, individualized transfer programs could be an efficient alternative to the well established making work pay programs in the US or in the UK.

However, even when only focusing on individualized transfer programs, several questions remain concerning the optimality of the policy structure. In particular, by targeting individuals with low earnings, most making work pay policies seem to combine positive participation effects with negative effects on the population already in employment. This is indeed also a problem encountered in recent evaluations of the German mini-job reform, which is an extension of previous exemptions of social security contributions (SSC) (see, e.g., Arntz et al. 2003 and Steiner and Wrohlich 2005). The mini-job reform is conditional on *earnings* and may encourage some workers to reduce their effort to benefit from the maximum level of transfers. The negative effects on the intensive margin can be avoided when targeting only individuals with low *wage rates* instead of all workers with low earnings. This is the idea of the Belgium Employment Bonus reform, which conditions the transfers on full-time equivalized individual earnings.

It is the aim of this paper to contribute to the discussion about the optimal design of transfer programs by providing empirical evidence about labor supply and employment reactions of differently designed individualized transfer programs. More precisely, we evaluate the labor supply and employment effects of the mini-job reform and the Employment Bonus using a static structural labor supply model with demand-side rationing. Most ex-ante evaluation studies assess the potential impact of tax reforms by using jointly tax-benefit microsimulation and a structural model of labor supply. This framework relies on usual assumptions concerning household rationality, such as static joint utility maximization in a pure supply-side framework. Additionally, unemployment is assumed to be voluntarily chosen. Ignoring involuntary unemployment, however, leads to biased elasticities and wrong predictions of the employment effects of a reform. While the bias might not be so important in countries where rationing plays only a minor role, it could seriously distort the results of policy evaluations in countries with

severe demand-side constraints, as is the case in Germany. In this paper, we will quantify discrepancies from ignoring involuntary unemployment when evaluating the employment effects of the two transfer reforms. We estimate the risk of involuntary unemployment together with a structural labor supply model (double-hurdle model). The model follows previous work by Blundell et al. (1987), Bingley and Walker (1997), Duncan and MacCrae (1999), and Hogan (2004).

We characterize a triple bias implicit in unconstrained estimations: (1) misspecification, (2) erroneous freedom of choice, and (3) overstatement of the taste for leisure. Interestingly, the bias on labor supply predictions is not apparent when the effects of a policy (e.g., the mini-job reform) are small and concentrate on voluntarily unemployed workers, typically secondary earners. They become substantial with policies (e.g., the Employment Bonus) that generate large responses among primary earners and singles.

The contributions of the paper are therefore twofold. First, we cover methodological questions about the reliability of predictions based on labor supply models, and second, we derive policy conclusions about the design of in-work transfers. Our results provide empirical evidence in favor of policies that distinguish between low effort and low productivity by targeting individuals with low *wages* rather than individuals with low *earnings*. Moreover, we show that individualized policies avoid negative labor supply effects for secondary earners that have been found when introducing UK-style family-based tax credits into the German tax and transfer system (cf. Bargain and Orsini 2006 or Haan and Myck 2007).

The paper is structured as follows. In the next section, we describe the mini-job reform and a hypothetical reform inspired by the Belgian Employment Bonus. Section 3 presents data, sample selection, and the strategy to identify rationed workers. Section 4 introduces the labor supply models, whereas Section 5 presents the estimation results, focusing in particular on the concept of labor supply elasticities in rationed labor markets. Section 6 will discuss the predicted impact of both reforms on employment and Section 7 concludes.

## 2 Low earnings or low wages?

In this section, we present a brief summary of the legislation for low-paid employment in Germany before 2003 and a description of the two reforms under consideration. The main differences between the three situations are summarized in Table 1. Moreover, we discuss the two reform proposals in an international perspective.

### 2.1 The mini-job reform

Before the mini-job reform, marginal employment was defined in Germany as employment activity up to a maximum of 15 h per week and full exemption of employees' SSC below 325 Euro of monthly gross earnings. Below this

**Table 1** Pre- and postreform legislation of the German tax and benefit system

	Baseline	Mini-job reform <sup>a</sup>	Employment Bonus
Mini-job definition			
Maximum hours restriction	15 h per week	–	–
Income restriction	325 Euros	400 Euros	–
Taxes			
Income tax sets in at	326 Euros	401 Euros	1 Euro
SSC			
Full SSC set in at	326 Euros	801 Euros	1 Euro
Full time equivalent SSC rebate	0	0	140 Euros
Rebate is phased out after	–	–	1,210 Euros
Taper rate	–	–	17.80%

<sup>a</sup>Following the 2003 mini-job reform, additional earnings from a mini-job do not cumulate with earnings from other sources. This aspect is not taken up in our simulation

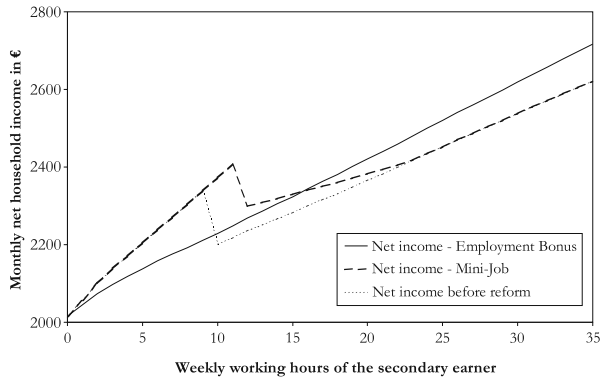
income threshold, earnings were also exempt from taxation if the employee had no other income. For those with other (nonlabor) income, the choice was given between a 20% flat-rate tax and taxation according to the progressive income tax code. Above the threshold, earnings were subject to the normal rate of SSC (about 21%) and taxation set in. Especially for secondary earners in married couples, this meant a drop in net income due to the joint taxation system (they became liable to the marginal tax rate of the primary earner), hence an incentive to remain at a low level of activity.

Following the reform, the maximum hour restriction was abolished and the range of earnings exempted from SSC was expanded up to 400 Euro. To avoid high marginal tax rates immediately above this threshold, a phasing-out of the exemption (or sliding pay-scale) was introduced: between 401 and 800 Euro, earnings are now subject to a modified SSC scheme, starting at 4% and increasing linearly up to 21%. Employees are covered by health insurance but do not acquire any pension rights unless they voluntarily add up to the normal SSC rate (Steiner and Wrohlich 2005). Income tax below the exemption earnings level is limited to a flat rate of 2%, while standard taxation sets in at 401 Euro.<sup>1</sup>

Budget lines give primary insights about the potential impact of reforms on work incentives. We depict how household net monthly income varies with the working hours of secondary earners in couples (Fig. 1) and single individuals (Fig. 2). The prereform situation displays the aforementioned drop in net income for married mini-job holders as joint taxation sets in at the 325 Euro threshold (corresponding to around 9 h/week when paid at 8 Euro/hour). The kink does not disappear with the reform but simply moves further to the right. Net household income increases in a range between 9 and 20 h.

<sup>1</sup>Another difference with the prereform situation is that income up to 400 Euro from a mini-job held as a secondary activity does not cumulate with the primary income for tax purposes, i.e., both activities are taxed independently. This may explain the apparent success of mini-jobs as a “moonlighting” activity, a feature not captured in our analysis. Modelling multiple activities is often difficult, as it requires information not covered by income surveys.

**Fig. 1** Couples: pre- and postreforms budget lines. No kids, primary earner working 40 h (median wage: 16.67 Euro/hour), secondary earner: wage: 8 Euro/hour

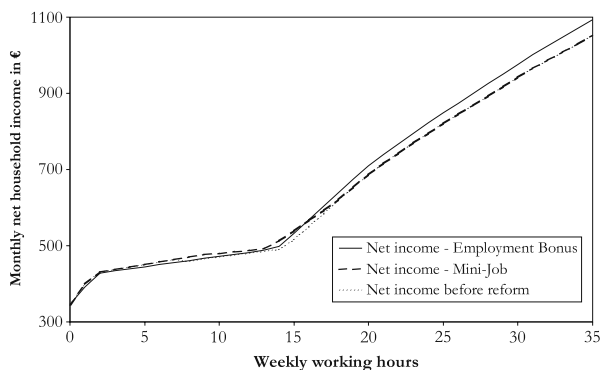


Overall, the mini-job reform seems to increase incentives to take up work for secondary earners – especially those with high fixed costs of work – and to reduce hours down to the 400 Euro threshold for those already employed. For single households, potential effects are very low. Net income increases only slightly when working less than 20 h per week. The reason for these small net gains is due to the withdrawal of means-tested social benefits as net income increases, making the budget line much flatter than in the case of secondary earners.

### 2.2 The Employment Bonus

This hypothetical reform is inspired by the Belgian Employment Bonus implemented in 2004, which consists of a substantial increase in the rebates on low-wage workers’ SSC. In Belgium, it has replaced the 2001 tax credit on low earnings, which, like the mini-job reform, rather targeted part-time employment. The Employment Bonus depends on working time so that a given worker will receive twice as many benefits if working full-time instead of working half-time. More importantly, the amount of bonus payment is conditional on the wage rate rather than on the level of earnings so that higher-wage

**Fig. 2** Singles: pre- and postreforms budget lines. Single person female, no kids, receiving social assistance, no housing benefits, wage: 10 Euro/hour



workers cannot reduce effort in order to become eligible. Full subsidy is paid up to a wage limit of 1,210 Euro/month, expressed in full-time equivalent (FTE) income. Above this threshold, it is phased out at a taper rate of 17.8% and is fully exhausted at a FTE income of 2,000 Euro (cf. Orsini 2006).<sup>2</sup>

We introduce this reform in the German system, in replacement of previous SSC exemptions, i.e., of the “old” mini-job regulation. Income is now taxed according to the progressive schedule from the first Euro earned, thus avoiding the drop in net income when secondary earners reach the 325 Euro limit. Figure 1 shows that the net gain increases with working time so that secondary earners at part-time are encouraged to increase hours or to stop working, depending on the shape of their preferences. In either case, the change in tax treatment would necessarily reduce the incentives for part-time work. The reform may encourage those with high fixed costs of working to take up a full-time activity, while the mini-job rather stimulates part-time participation. In the case of single individuals (Fig. 2), the whole budget constraint is simply shifted upwards, except for low levels of earnings, in which case, as for the mini-job, the reform is neutralized by the interaction with social assistance. The reform appears to unambiguously encourage labor supply of low-skilled workers both at the intensive and extensive margins.

### 2.3 Targeting individual or family earnings

Both the mini-job reform and the Employment Bonus are conditioned on an individualized measure of earnings. Therefore, the structure of the transfers is quite distinct from the design of the most prominent making work pay policies, namely the EITC in the US and the WTC in the UK, which are conditioned on family earnings. These Anglo-American transfer programs induce high marginal tax rates for the secondary earner in couples and, thus, create negative work incentives. Numerous empirical evaluations of the EITC and the WTC confirm this inefficiency and show that the positive employment effects for single individuals or primary earners in couple households are partly or completely offset by the negative effects on married women.

By construction, individualized transfer programs avoid these negative effects. This becomes evident in Fig. 1. At low working hours, the marginal tax rates for the secondary earner are fairly low, leading to a remarkable increase in the net household income. The high marginal tax rate induced by the mini-job reform occurs only at higher working hours. As mentioned above, this is even reinforced by the joint income taxation of married couples, which implies that the secondary earners directly face their partners' marginal tax rate on all their earnings. This is potentially responsible for a “part-time trap” further

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<sup>2</sup>Conditioning on productivity rather than earnings is a practical example of first best taxation but conveys questions about the cost and reliability of measures of wage rates (or working time) by the administration. We assume hereafter that these administrative issues can be solved and do not generate additional costs for the government (in practice, in Belgium, the information is based on the contractual hours declared by the employers to the social security institutions).

investigated below. In contrast, the Employment Bonus targets only low-wage individuals but over the entire hours distribution. Therefore, high marginal tax rates can be avoided, as shown in Fig. 1.

### 3 Data selection and identification of rationing

#### 3.1 Data selection

Our empirical assessment is based on the 2003 wave of the German Socio-Economic-Panel (GSOEP), a survey gathering socio-demographic and financial information about 11,000 representative households for the fiscal year 2002.

For the estimation, we restrict our sample and drop households where both the household head and – if present – the partner are aged between 20 and 65, or if both are self-employed, retired, disabled, on maternity leave, or in full-time education.<sup>3</sup> In couples in which one spouse falls into this group, his or her labor supply is assumed to be fixed to the observed level, while the partner can flexibly adjust his/her labor supply. In other words, labor supply of these couples is modeled according to the male or female chauvinist framework. Labor supply of single males and single females is modeled separately. We therefore distinguish between five groups in the empirical analysis: single women (1,022 observations), single men (783), couples where both spouses have a flexible labor supply (3,822), couples where the male labor supply is fixed (970), and couples where the female labor supply is fixed (562). Table 2 contains some descriptive statistics of the relevant variables.

#### 3.2 Identification of involuntary unemployment

Two questions are used to distinguish voluntary and involuntary unemployment. Each potential worker is asked (1) whether he/she has actively searched for a job within the last 4 weeks and (2) whether he/she is ready to take up a job within the next 2 weeks. We follow the International Labour Organization definition and treat those unemployed who answer both questions in the affirmative as rationed. Table 2 shows that around 6% of the individuals living in couples and around 10% of the singles are involuntarily unemployed according to this definition.<sup>4</sup>

The probability of rationing is identified by regional demand-side variables and individual characteristics. For the former, we use county information

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<sup>3</sup>As common in the labor supply literature, we do not model the behavior of the self-employed. The working behavior of the self-employed strongly depends on risk-measures, and moreover, working time and gross earnings are difficult to determine.

<sup>4</sup>Note that these rates differ from official unemployment statistics since their denominators contain some of the inactive population (precisely the voluntary unemployed) and also because of selection criteria

**Table 2** Some descriptive statistics for the estimation sample

	Singles		Couples					
	Fem.	Mal.	Both flexible		Male flexible		Female flexible	
			Fem.	Mal.	Fem.	Mal.	Fem.	Mal.
Individual information								
Age	41.87	41.77	42.15	44.71	40.89	43.65	47.33	50.64
Working hours per week	34.86	39.72	28.05	41.25	29.99	40.93	28.71	46.33
No degree	0.15	0.11	0.17	0.10	0.13	0.08	0.15	0.09
Education = 0: No degree	0.02	0.02	0.02	0.03	0.01	0.01	0.02	0.02
Education = 1: Low degree <sup>a</sup>	0.71	0.69	0.74	0.70	0.61	0.60	0.75	0.67
Education = 2: High degree <sup>b</sup>	0.27	0.28	0.23	0.28	0.36	0.39	0.23	0.30
Hourly wage (cond. on working)	14.55	17.16	13.92	19.31	13.85	20.46	13.93	16.77
Hourly wage (cond. on not working) <sup>c</sup>	9.63	11.09	9.92	13.19	10.40	12.74	10.05	11.59
Children from 0–3 years	0.04	0.00	0.13		0.23		0.08	
Children from 3–6 years	0.07	0.00	0.14		0.12		0.11	
Children over 6 years	0.45	0.07	0.86		0.74		0.64	
German	0.95	0.94	0.90		0.93		0.94	
East Germany	0.16	0.21	0.17		0.21		0.17	
Employment status <sup>d</sup>								
Involuntarily unemployed <sup>e</sup>	0.11	0.10	0.06	0.06	–	0.06	0.06	–
Voluntarily unemployed <sup>e</sup>	0.12	0.09	0.26	0.06	–	0.07	0.30	–
Employed	0.78	0.81	0.69	0.88	–	0.87	0.64	–
No. of observations	1022	783	3822	3822	562	562	970	970

We distinguish between singles, couples where both spouses have a flexible labor supply and couples where either the man or the woman has a flexible labor supply while the labor supply of the other partner is fixed

<sup>a</sup>Secondary general school (“*Hauptschule*”), intermediate school (“*Realschule*”) or other degree

<sup>b</sup>Upper secondary school (“*Gymnasium*”)

<sup>c</sup>Imputation through a standard Heckman two-step equation

<sup>d</sup>The employment status of males (females) in couples where only one spouse has a flexible labor supply is retired, self-employed, on maternity or parental leave, or disabled

<sup>e</sup>To distinguish between voluntary and involuntary unemployment, we make use of the search activity within the last 4 weeks and the readiness to take up a job within the next 2 weeks



**Table 3** Involuntary unemployment and description of the strategic types of labor office districts

Cluster	Description	No. of districts	Involuntary Unemployment (in %)	
			Females	Males
I	East German labor office districts (excluding Dresden)	38	0.124	0.117
II	Labor office districts dominated by large cities	22	0.065	0.070
III	West German labor office districts with rural elements, medium-sized industry, and average unemployment	63	0.047	0.054
IV	West German centers with good labor market prospects	10	0.049	0.049
V	West German labor office districts with the best labor market prospects	48	0.031	0.024

Source: Blien et al. (2004)

to describe the situation on the local labor market. The 181 labor office districts have been classified by Blien et al. (2004) into 12 types with similar labor market conditions that can themselves be summarized into five clusters. The classification is built upon several labor market characteristics, the most important criteria being the underemployment ratio and the corrected population density.<sup>5</sup> We assign each individual (based on his/her place of residence) to one of the five clusters and compute for each cluster the rate of involuntary unemployment as defined above.<sup>6</sup> Table 3 contains a short description of each cluster, each of which is ordered decreasingly with the level of tension on the labor market according to Blien et al.’s criteria above. Male and female unemployment rates vary consistently with the ranking of local labor markets. In particular, counties in cluster V that have the best labor market situation present the lowest rate of involuntary unemployment (3.1% for females and 2.4% for males), while cluster I (consisting of nearly all of East Germany) shows the most depressed labor market (12.4% for females and 11.7% for males). In addition to the aggregate information, we exploit the panel dimension of the GSOEP to integrate current information on past

<sup>5</sup>The underemployment ratio is defined as the relation of the number of unemployed individuals and participants in several active labor market programs to the number of all employed persons plus these programs’ participants. The corrected population density is used to improve the comparability of rural labor office districts with metropolitan and city areas. In addition, the vacancy quota, describing the relation of all reported vacancies at the labor office to the number of employed persons, and the placement quota, which contains the number of placements in relation to the number of employed persons, are used. Finally, an indicator for the tertiarization level built on the number of employed persons in agricultural occupations and an indicator for the seasonal unemployment are considered.

<sup>6</sup>Since the clusters refer to labor office districts and the individuals’ place of residence is on county level, we have to do some readjustments. We follow a rather simplistic approach and assign counties belonging to more than one labor office district to the one where the majority of inhabitants are located.

employment records. This information is valuable for identification of the individual risk of rationing.

## 4 Labor supply models

### 4.1 Unconstrained model

Discrete choice models of labor supply are based on the assumption that a household  $i$  can choose among  $J + I$  working hours (nonparticipation denoted by  $j = 0$  and  $J$  positive hours denoted by  $j = 1, \dots, J$ ). For each discrete choice  $j$ , its net income  $C_{ij}$  (equivalent to aggregate household consumption in a static framework) is computed by tax-benefit microsimulation techniques so that leisure-consumption preferences can be estimated.<sup>7</sup> The approach has become standard practice, as it provides a straightforward way to account for the nonlinear and nonconvex budget sets of complex tax and benefit systems when modeling individual and joint labor supplies of spouses. Choices  $j=0, \dots, J$  in a couple correspond simply to all combinations of the spouses' discrete hours (see van Soest 1995). Precisely, the utility  $V_{ij}$  derived by household  $i$  from making choice  $j$  is assumed to depend on a function  $U$  of spouses' leisures  $Lf_{ij}$  and  $Lm_{ij}$ , disposable income  $C_{ij}$ , and household characteristics  $Z_i$ , and on a random term  $\epsilon_{ij}$ :

$$V_{ij} = U(Lf_{ij}, Lm_{ij}, C_{ij}, Z_i) + \epsilon_{ij}. \quad (1)$$

The utility function and the choice probability of a single individual are derived in the same way as above, yet they only contain the leisure term of this individual. Couples where one spouse's labor supply is fixed are treated in the same way. For each potential worker, we allow for six discrete choices (and, hence, 36 combinations for couples where both spouses are assumed to have flexible labor supplies). The following hours classifications are used: 0, [0,12], ]12,20], ]20,34], ]34,40], >40.

### 4.2 Constrained model

Several studies have previously accounted for involuntary unemployment in labor supply estimations. Blundell et al. (1987) extend the binomial model of female participation by introducing a probability of rationing that results in a double-hurdle model. Hogan (2004) extends the approach to a panel structure, relaxing the independence of irrelevant alternatives (IIA) hypothesis through nested logit modelling. Bingley and Walker (1997) combine a latent model for the probability of involuntary unemployment with a discrete-choice multinomial probit model for the labor supply of lone mothers. Duncan and

<sup>7</sup>For this application, we use the German tax-benefit simulation model STSM. For a detailed description, see Steiner et al. (2005)

MacCrae (1999) proceed in a similar way for women in couples by using a conditional logit framework but assume unemployment for men to be completely voluntary. Laroque and Salanie (2002) model the labor supply of French women by introducing classical unemployment due to the censorship of the minimum wage; other involuntary unemployment is a residual category gathering all other explanations (frictional or business cycle unemployment). Finally, Euwals and van Soest (1999) suggest using information about desired vs actual working hours of single men and women in the Netherlands to disentangle preferences and demand-side rationing.

The *constrained* model we suggest is close to Duncan and MacCrae (1999) but differs in two aspects. First, we model involuntary unemployment for both men and women in couples. This is important, as the share of involuntary unemployed is particularly high for men (see Table 2). Second, we use information on desired working hours, being either part-time or full-time work, of unemployed workers; this way, preferences are estimated more precisely than if we simply model the probability of desired participation.

We combine the labor supply model previously described with a rationing risk model. For a single  $i$ , or spouse  $i$  in a couple, we specify a latent equation of involuntary unemployment:

$$I_i^* = \beta X_i + \eta_i \tag{2}$$

as a stochastic function of characteristics  $X_i$  thought to influence the probability of getting a job. Under the assumption of standard normality of the random term  $\eta_i$ , the risk of rationing is modelled as a standard probit. As stressed by Blundell et al. (1987), this framework allows the introduction of demand-side regional variables together with individual characteristics (mainly education and past employment history).

The model we are estimating can be seen as a double-hurdle representation. The first hurdle is the decision to be voluntarily inactive or to participate in the labor market, working either part-time, full-time, or overtime; the second hurdle describes the probability of being involuntarily unemployed for those who decide to participate in the first stage. In the technical [Appendix](#), we provide a description about the constrained choice set and derive the likelihood function.

## 5 Estimation results

### 5.1 Unemployment risk estimates

Estimates of the rationing probability are presented in Table 4.<sup>8</sup> The coefficients of the regional indicators, introduced in reference to the first cluster,

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<sup>8</sup>In the estimation, we account for the problem of matching microdata with aggregate (regional) information as described in Moulton (1990). We allow for correlation within a region and, therefore, derive consistent standard errors for the regional variables.

**Table 4** Estimation results for the unemployment probabilities

	Women <sup>a</sup>		Men <sup>a</sup>	
	Coef.	s.e.	Coef.	s.e.
Constant	<i>-1.176</i>	0.487	<b>-0.963</b>	0.304
Regional information				
Cluster 1	Reference		Reference	
Cluster 2	<b>-0.352</b>	0.016	<b>-0.451</b>	0.019
Cluster 3	<b>-0.477</b>	0.021	<b>-0.598</b>	0.013
Cluster 4	<b>-0.379</b>	0.018	<b>-0.631</b>	0.014
Cluster 5	<b>-0.578</b>	0.036	<b>-0.831</b>	0.015
Age	<i>5.374</i>	2.731	<b>5.651</b>	1.260
Age-Squared	<i>-6.521</i>	3.177	<b>-6.490</b>	1.290
Educational degree				
No degree	Reference		Reference	
Medium degree	<b>-0.519</b>	0.130	<b>-0.557</b>	0.106
High degree	<b>-0.934</b>	0.154	<b>-1.128</b>	0.152
No vocational degree	<b>0.358</b>	0.128	<b>0.368</b>	0.056
Employment status in <sup>b</sup>				
October 1998	<i>-0.099</i>	0.065	<i>-0.029</i>	0.089
October 1999	<i>-0.213</i>	0.115	<i>-0.364</i>	0.204
October 2000	<b>-0.604</b>	0.124	<b>-0.510</b>	0.100
Observations		4451		4859
Pseudolikelihood		-972.04		-1044.53

Bold letters indicate significance at the 1% level, italic letters refer to the 5% level, and underlined letters to the 10% level. Standard errors were estimated according to Moulton (1990)

<sup>a</sup>The unemployment probability has been estimated for all women (single + married) and men (single + married)

<sup>b</sup>Dummies equal 1 if regular full-time employment at indicated dates, 0 otherwise

where risk of rationing is highest, are all highly significant. They are ranked as expected, except for the third cluster in the case of women. The education variables show that higher degrees provide higher protection against unemployment. The risk of involuntary unemployment is affected by previous working history. Dummies representing employment in October of the previous 3 years show significant state dependency with respect to the last 2 years.

## 5.2 Labor supply estimates

We now turn to the estimates of the *constrained* and *unconstrained* labor supply models. Results are presented in Table 5 for single individuals and in Tables 6 and 7 for couples. In both the unconstrained and the constrained models, and for the five household types, almost all households fulfill monotonicity and concavity of the utility function with respect to the various choice variables. Most importantly, utility increases with net income for almost all households, as shown in the bottom parts of Tables 5, 6, and 7; this is the minimum requirement for the consistency of tax reform simulations hereafter. The derivatives with respect to leisure show that, for a small share of the population, positive monotonicity in leisure is not respected. As stressed by

**Table 5** Estimation results for singles: rationed and unrationed models

	Unrationed model			Rationed model		
	Women		Men	Women		Men
	Coeff.	s.e.	Coeff.	Coeff.	s.e.	s.e.
Consumption						
Age	4.181	13.528	-10.070	23.219	14.484	-18.627
Age-Squared	-7.505	14.519	12.230	-25.685	15.581	21.108
Constant	4.130	3.076	4.843	-1.802	3.294	6.866
Leisure						
Age	-0.022	0.288	-0.277	0.194	0.330	-0.599
Age-Squared	0.208	0.323	0.487	0.094	0.366	<u>0.893</u>
Child 0-3	<b>0.128</b>	0.018		<b>0.128</b>	0.015	
Child 3-6	<b>0.068</b>	0.009		<b>0.075</b>	0.009	
German	<i>0.020</i>	0.009	-0.014	<u>0.020</u>	0.010	-0.009
East Germany	0.001	0.005	<b>0.015</b>	-0.012	0.007	-0.003
Constant	<b>0.478</b>	0.074	<b>0.411</b>	<b>0.455</b>	0.082	<b>0.569</b>
Consumption	-0.471	0.113	-0.119	-0.365	0.120	-0.118
Leisure	-0.004	0.000	-0.003	-0.005	0.000	-0.004
Consumption×Leisure	-0.021	0.008	-0.028	-0.014	0.009	-0.033
Part-time category 1	-2.734	0.184	-3.855	-2.492	0.187	-3.369
Part-time category 2	-2.585	0.176	-3.875	-2.699	0.174	-3.851
Part-time category 3	-1.744	0.138	-2.966	-1.866	0.132	-3.143
Observations		1022			1022	783
Wald chi2		2.85			2.79	2.78
Log-Likelihood		-1448.25			-1394.53	-778.49
Derivatives (in percent)						
$U_c > 0$		99%			99%	100%
$U_{lf} > 0$		74%			79%	79%
$U_{lm} > 0$						89%

Bold letters indicate significance at the 1% level, italic letters refer to the 5% level, and underlined letters to the 10% level. The following hours classifications are used: 0, [0,12] (Part-time 1), [12,20] (Part-time 2), [20,34] (Part-time 3), [34,40], > 40

**Table 6** Estimation results for couples where only one spouse is flexible: rationed and unrationed models

	Unrationed model			Rationed model		
	Women Coeff.	Men Coeff.	s.e.	Women Coeff.	Men Coeff.	s.e.
Consumption						
Age	8.507	17.691	12.495	7.270	8.851	16.286
Age-Squared	-10.932	-16.938	13.496	-9.417	-7.771	17.262
Constant	0.325	-3.097	2.692	0.445	-1.016	3.620
Leisure						
Age	-0.271	-0.035	0.391	-0.428	-0.133	0.528
Age-Squared	0.537	0.204	0.438	0.745	0.373	0.580
Children 0-3	<b>0.061</b>			<b>0.070</b>		
Children 3-6	<b>0.040</b>			<b>0.039</b>		
German	-0.010	-0.009	0.010	-0.010	-0.006	0.013
East Germany	- <b>0.031</b>	0.012	0.006	- <b>0.047</b>	-0.001	0.008
Constant	<b>0.324</b>	0.088	0.086	<b>0.407</b>	0.257	0.118
Consumption	- <b>0.058</b>	0.022	0.018	- <i>0.056</i>	0.091	0.022
Leisure	- <b>0.002</b>	0.000	0.000	- <i>0.003</i>	- <b>0.002</b>	0.000
Consumption×Leisure	- <i>0.002</i>	0.001	0.003	- <i>0.002</i>	- <b>0.013</b>	0.004
Part-Time category 1	- <b>1.857</b>	0.151	0.458	- <b>1.896</b>	- <b>2.777</b>	0.466
Part-Time category 2	- <b>1.408</b>	0.167	0.426	- <b>1.557</b>	- <b>3.367</b>	0.426
Part-Time category 3	- <b>1.020</b>	0.156	0.310	- <b>1.132</b>	- <b>3.086</b>	0.308
Observations			970			562
Wald Chi <sup>2</sup>		1.3	3.97			1.07
Log-Likelihood			-1479.65			-1472.32
Derivatives						
$U_c > 0$		99%	88%			99%
$U_{lf} > 0$		70%				64%
$U_{lm} > 0$			80%			82%

Bold letters indicate significance at the 1% level, italic letters refer to the 5% level, and underlined letters to the 10% level. The following hours classifications are used: 0, [0,12] (Part-time 1), [12,20] (Part-time 2), [20,34] (Part-time 3), [34,40], > 40

**Table 7** Estimation results for couples where both spouses are flexible: rationed and unrationed models

	Unrationed model		Rationed model	
	Coef.	Std.	Coef.	Std.
Consumption				
Age – Man	–5.887	4.921	–7.588	5.389
Age-Squared – Man	4.513	5.245	6.514	5.726
Age – Woman	<b>12.136</b>	4.242	<i>11.728</i>	5.055
Age-Squared – Woman	<b>–13.203</b>	4.675	<i>–12.973</i>	5.514
Constant	1.883	1.088	2.178	1.249
Consumption-Squared	<b>–0.097</b>	0.013	<b>–0.089</b>	0.014
Leisure Man				
Age – Man	–0.191	0.154	<b>–0.590</b>	0.184
Age-Squared – Man	<u>0.321</u>	0.169	<b>0.832</b>	0.201
Constant	<b>0.334</b>	0.037	<b>0.441</b>	0.044
Leisure man-Squared	<b>–0.002</b>	0.000	<b>–0.003</b>	0.000
Leisure woman				
Age – Woman	0.266	0.136	0.231	0.153
Age-Squared – Woman	–0.169	0.154	–0.111	0.172
Child 0–3	<b>0.076</b>	0.005	<b>0.082</b>	0.005
Child 3–6	<b>0.042</b>	0.004	<b>0.041</b>	0.004
German	–0.008	0.004	<i>–0.008</i>	0.004
East Germany	<b>–0.035</b>	0.003	<b>–0.047</b>	0.003
Constant	<b>0.296</b>	0.036	<b>0.348</b>	0.040
Leisure woman-Squared	<b>–0.003</b>	0.000	<b>–0.003</b>	0.000
Consumption*Leisure man	<b>–0.014</b>	0.002	<b>–0.012</b>	0.002
Consumption*Leisure woman	<b>–0.003</b>	0.001	–0.001	0.001
Leisure Man*Leisure woman	0.000	0.000	0.000	0.000
Part-time category 1 - Women	<b>–1.659</b>	0.073	<b>–1.677</b>	0.074
Part-time category 2 - Women	<b>–1.430</b>	0.087	<b>–1.578</b>	0.086
Part-time category 3 - Women	<b>–0.995</b>	0.080	<b>–1.086</b>	0.078
Part-time category 1 - Men	<b>–3.418</b>	0.190	<b>–3.000</b>	0.193
Part-time category 2 - Men	<b>–3.911</b>	0.197	<b>–3.753</b>	0.198
Part-time category 3 - Men	<b>–3.049</b>	0.112	<b>–3.091</b>	0.112
Observations		3822		3822
Wald Chi <sup>2</sup>		16.42		12.18
Log-Likelihood		–9991.60		–9583.59
Derivatives				
$U_c > 0$		100%		100%
$U_{lf} > 0$		71%		69%
$U_{lm} > 0$		88%		92%

Bold letters indicate significance at the 1% level, italic letters refer to the 5% level, and underlined letters to the 10% level. The following hours classifications are used: 0, [0,12] (Part-time 1), [12,20] (Part-time 2), [20,34] (Part-time 3), [34,40], > 40

Euwals and van Soest (1999), there is no necessity to restrict preferences relative to the taste for leisure.

The marginal utility of income and leisure depends on individual- and household-specific variables. As expected, the presence of young children significantly increases preference for leisure of women in all groups. In line with previous studies, East German women prefer to work more. Taste shifters related to age are not always significant and do not display clear patterns. Finally,

parameters of the dummies for the part-time categories, as defined above, are negative and significant, suggesting some “disutility” stemming from this work arrangement. Coefficients are less negative for women since they are employed part-time more frequently than men. In the estimated system of indifference curves, inclusion of these dummies yields a hump in the part-time range of working hours and suggests that, other things being equal, larger participation effects are to be expected from the Bonus due to its targeting on full-time activity.

### 5.3 Predicted elasticities

In the present nonlinear model, labor supply elasticities can be obtained numerically by simulating the impact of a marginal increase in gross hourly wages on hours of work and participation.<sup>9</sup>

Table 8 presents estimated elasticities obtained with the constrained and unconstrained models. In both cases, they are computed using the whole selected population of potential workers (either constrained or unconstrained, working or not working). While the next subsections look at employment effects of the mini-job reform and the Employment Bonus, we focus here on *pure* labor supply elasticities to characterize potential working behavior. In other words, we look at changes in *desired* hours; the baseline of the *constrained* model corresponds to actual desired hours as recorded in the data while the baseline of the *unconstrained* model corresponds to observed hours, i.e., constrained workers are (mistakenly) assumed to voluntarily choose inactivity.

The elasticities from the *unconstrained* model are in line with the labor supply literature (Blundell and MaCurdy 1999) and similar in magnitude to those found in recent studies on Germany, such as Haan and Steiner (2005) or Steiner and Wrohlich (2005). For all groups, the elasticities are relatively modest. They lie in a narrow range between 0.2 and 0.3, except for women in couples (above 0.3) and men in couples where the women have fixed labor supplies (below 0.2). In this last group, a relatively high share of women are on maternity leave while the labor supply of men with small children is known to be rather inelastic. Estimates from the *constrained* model are more precise, while females in couples still display the largest elasticities.

Not considering involuntary unemployment in the *unconstrained* model leads to biased estimated elasticities for several reasons. We suggest a

<sup>9</sup>We follow a calibration method that is consistent with the probabilistic nature of the model at the individual level (Creedy and Duncan 2002). It consists in drawing for each household a set (here, 100 draws) of  $J + I$  random terms from the  $EV - I$  distribution that generates a perfect match between predicted and observed choices. The same draws are kept when predicting labor supply responses to a shock on wages or a tax reform. Averaging individual supply responses over a large number of draws provides robust transition matrices. Confidence intervals for elasticities (or labor supply responses to a reform) are obtained by repetitive random draws of the preference parameters from their estimated distributions and, for each draw, by applying the calibration procedure.



**Table 8** Labor supply elasticities in the constrained and unconstrained models

	Couples (both flexible)		Couples (only one spouse flexible)		Singles	
	Women	Men	Women	Men	Women	Men
Change in the participation rate (in percentage points)						
Unconstrained	0.150 (0.108–0.185)	0.158 (0.122–0.190)	0.224 (0.164–0.319)	0.079 (0.028–0.140)	0.197 (0.154–0.256)	0.226 (0.160–0.275)
Constrained	0.122 (0.083–0.158)	0.067 (0.052–0.078)	0.151 (0.104–0.205)	0.050 (0.022–0.080)	0.069 (0.036–0.104)	0.083 (0.048–0.119)
Change in total hours worked (in percent)						
Unconstrained	0.351 (0.303–0.425)	0.207 (0.177–0.248)	0.376 (0.263–0.504)	0.115 (0.046–0.179)	0.267 (0.210–0.344)	0.300 (0.231–0.368)
Constrained	0.316 (0.277–.362)	0.117 (0.102–0.133)	0.268 (0.189–0.339)	0.0914 (0.044–0.132)	0.104 (0.058–0.159)	0.152 (0.107–0.197)

Elasticities are computed by averaging over the whole sample the 100 simulated transitions, for each of the 100 draws of the utility parameters from the estimated asymptotic distribution of their estimator. The table reports the median values of each distribution and, in brackets, the 90% confidence interval

breakdown of these effects. First, the unconstrained model unduly allows transitions to participation for those constrained workers, assumed to voluntarily choose inactivity in the first place, whose predicted hours are positive after a wage shock. In contrast, in the *constrained* model, these workers have positive desired hours in the baseline and contribute to changes at the intensive margin. This “participation bias” leads to a clear upward bias of the labor supply elasticities from the *unconstrained* model. A second source of discrepancies is the preference bias stressed by Ham (1982), which acts in the opposite direction. In effect, ignoring involuntary unemployment necessarily leads to overstating the taste for leisure in the estimates and, therefore, to understating elasticities. Finally, a “specification bias” must affect estimates in a way that is a priori uncertain. The unconstrained model is indeed misspecified since individual characteristics are not only required to explain consumption–leisure preferences but also implicitly account for demand-side constraints. As a result, labor supply estimates reported in Tables 5, 6, and 7 are more precisely estimated with the constrained model. Moving from unconstrained to constrained estimates, the significance of some taste-shifters change in a characteristic way. In particular, the dummy for East Germany is no more significant in affecting male preference for leisure: lower employment rates of males appear to be associated with a stronger rationing effect in East Germany and not with a difference in the taste for leisure, as the unconstrained model would suggest. On the other hand, the dummy for East Germany is significant for females in both models, suggesting lower preferences for nonmarket time there.

The sign of the overall bias is not clear a priori. From results in Table 8, it turns out that the upward bias dominates. Average unconstrained elasticities of working hours are indeed larger for most groups. Differences are important in groups with a high share of the involuntarily unemployed, typically single men, while they are not significant for married women who frequently choose nonparticipation on a voluntary basis due to family constraints.<sup>10</sup> Comparing hours and participation suggests that overstatement in the former is driven by the participation bias.

In order to provide a better understanding of the differences, we distinguish the labor supply effects of three groups (voluntarily inactive, involuntarily unemployed, employed) as shown in Table 9. Instead of elasticities, we present the absolute changes in participation rates and total hours of work for each group, given a 1% uniform increase in gross wages. Results emphasize the role of the participation bias: with the *unconstrained* model, involuntarily unemployed workers markedly increase hours due to a large participation effect; this effect vanishes in the *constrained* model. The differences for the employed and the voluntarily inactive are, in general, very small and go in both directions. The overstatement of the taste for leisure, or “preference bias,” dominates for women in couples (the effect on labor supply is larger with the constrained

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<sup>10</sup>Previous findings confirm that elasticities of hours for single workers are around twice as small when accounting for demand constraints (Euwals and van Soest 1999).

**Table 9** Labor supply effects by employment status

Employment status and group	Women			Men		
	Share of obs.	Unconstr. model	Constr. model	Share of obs.	Unconstr. model	Constr. model
Couples (both flexible)						
Voluntary UE	0.25	3.60	4.50	0.05	2.73	2.61
Part.						
Work. Hours	0.25	113.84	142.30	0.05	116.76	110.84
Involuntary UE	0.05	1.11	0.00	0.06	3.08	0.00
Part.						
Work. Hours	0.05	37.70	1.25	0.06	134.49	3.38
Work. Hours	0.71	105.06	113.71	0.89	97.55	96.38
Couples (only one spouse flexible)						
Voluntary UE	0.31	0.89	1.02	0.07	0.32	0.27
Part.						
Work. Hours	0.31	27.18	31.20	0.07	13.60	11.66
Involuntary UE	0.04	0.22	0.00	0.07	0.19	0.00
Part.						
Work. Hours	0.04	7.78	0.20	0.07	8.01	0.18
Work. Hours	0.66	17.90	14.46	0.86	6.02	5.30
Singles						
Voluntary UE	0.13	0.82	0.80	0.08	0.40	0.51
Part.						
Work. Hours	0.13	29.95	28.34	0.08	0.83	21.42
Involuntary UE	0.08	0.80	0.00	0.09	0.06	0.00
Part.						
Work. Hours	0.08	30.55	0.47	0.09	19.30	0.95
Work. Hours	0.79	17.76	11.94	0.83	21.42	17.66

Numbers are absolute changes in the participation rate (in percentage points) and number of working hours (in weekly working hours) due to a 1% increase in gross hourly wage

model). Results are less clear-cut for the other groups, due to the interplay of “preference” and “specification” bias.

## 6 Employment effects of the reforms

The concept of *pure* labor supply elasticities derived from the constrained model is insightful to explain working behavior and to measure potential responsiveness to changes in financial incentives. Yet, this concept is based on desired hours and cannot provide information about the true employment effects of a reform that is the relevant information for policymakers. In the following, we account for the rationing risk and predict employment effects of the two reforms for the main labor force (about 30 million individuals).

Since our modeling of the rationing probability is a reduced form equation, we cannot assess the impact of the reform on demand-side variables, for instance, through wage rate adjustments or changes in vacancy rates simultaneous to labor supply responses. Our analysis is partial in this respect since we must assume that the individual rationing probability is not affected: constrained workers remain in their situation after the reform and do not affect the total working hours.<sup>11</sup> In other words, potential employment effects are concentrated only on those voluntarily inactive and those already working.<sup>12</sup>

For both unconstrained and constrained models, Tables 10 and 11 display total participation effects (in number of workers) and total hours effects (in FTE). The latter is broken down between hours effects of those individuals that enter the labor market (extensive margin) and that of those who had been in employment before the reform (intensive margin). We report the median, upper, and lower values of the 90% confidence interval from bootstrap simulations. Tables also indicate the proportion of new rationing.<sup>13</sup>

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<sup>11</sup>The mini-job reform has, indeed, only a small direct effect on labor cost, while the Employment Bonus has no effect whatsoever on labor demand. In the case of the mini-job, however, firms may adjust labor demand in different ways to respond to changes in legislation (e.g., by splitting previous full-time jobs into several mini-jobs). Other possible feedback effects may lead to changes in the equilibrium gross wages (here assumed to be constant). These effects are very difficult to account for without a more comprehensive framework (e.g., CGE models).

<sup>12</sup>A distinction must be made. The population already employed may freely choose to change working hours or to withdraw from the labor market (they are not rationed). The voluntarily inactive may decide to enter the labor market following the reform, but their expected labor supply is weighted by their individual rationing risk, which is deterministically predicted ex-ante.

<sup>13</sup>With the mini-job reform, for instance, almost 16% of the voluntarily inactive that are induced to enter the labor market will be rationed from the demand side. This percentage is significantly higher than the current unemployment rate of any of the groups. This is hardly surprising, given that the labor market characteristics of the inactive population are often weaker than those of the active population.

**Table 10** Labor supply effects of the mini-job reform (Figures in 1000s)

	Unconstrained model				Constrained model				Shared of newly rationed
	New participation		Working hours in FTE		New participation		Working hours in FTE		
	Total	Ext. margin	Total	Int. margin	Total	Ext. margin	Total	Int. margin	
<b>Couples (both flexible)</b>									
Women	41 (35-47)	13 (11-15)	36 (31-41)	-17 (-19--14)	36 (31-41)	11 (10-13)	-6 (-7--3)	-17 (-19--14)	16.1%
Men	3 (0-5)	9 (7-10)	0 (-1-2)	-9 (-11--6)	0 (-1-2)	4 (3-5)	-3 (-6--1)	-7 (-10--5)	15.7%
<b>Couples (only one spouse flexible)</b>									
Women	8 (6-10)	4 (2-5)	5 (3-8)	-2 (-3--2)	5 (3-8)	2 (1-2)	0 (-1-0)	-2 (-3--1)	12.8%
Men	0 (0-0)	0 (0-0)	0 (0-0)	0 (0-0)	0 (0-0)	0 (0-0)	0 (0-1)	0 (0-0)	19.6%
<b>Singles</b>									
Women	4 (2-5)	2 (1-2)	1 (0-2)	-5 (-6--3)	1 (0-2)	1 (0-1)	-2 (-4--1)	-3 (-4--2)	19.3%
Men	0 (0-1)	0 (0-0)	0 (0-0)	-1 (-1-0)	0 (0-0)	0 (0-0)	0 (0-0)	0 (-1-0)	22.5%
Total	56 (43-67)	28 (21-32)	43 (33-53)	-33 (-40--25)	43 (33-53)	18 (14-21)	-11 (-18--6)	-29 (-37--22)	15.9%

Labor supply effects are computed by averaging over the whole sample the 100 simulated transitions, for each of the 100 draws of the utility parameters from the estimated asymptotic distribution of their estimator. The table reports the median values of each distribution and, in brackets, the 90% confidence interval. Total participation is the number of individuals starting to work after the reform. Variations in working hours are decomposed into change of hours for those previously in employment (intensive margin) and change due to participation effects (extensive margin). In both cases, it is expressed in FTE computed as (change in) the total number of hours divided by 40 weekly hours. Share of newly rationed is the share of new participants who are restricted by demand-side constraints

**Table 11** Labor supply effects of the Employment Bonus (figures in 1000s)

	Unconstrained model			Constrained model			Shared of newly rationed		
	New participation		Working hours in FTE		New participation			Working hours in FTE	
	Total	Ext. margin	Int. margin	Total	Ext. margin	Int. margin		Total	Ext. margin
Couples (both flexible)									
Women	105 (93 – 114)	121 (110 – 130)	92 (84 – 99)	29 (24 – 33)	78 (71 – 84)	103 (95 – 111)	68 (62 – 74)	35 (30 – 38)	18.9%
Men	62 (54 – 71)	67 (56 – 76)	72 (64 – 80)	-5 (-8 – -2)	29 (25 – 34)	32 (25 – 38)	33 (28 – 38)	-1 (-4 – 2)	22.1%
Couples (only one spouse flexible)									
Women	23 (17 – 26)	28 (21 – 30)	20 (15 – 22)	7 (5 – 9)	13 (10 – 16)	17 (14 – 22)	11 (8 – 13)	6 (5 – 8)	15.0%
Men	5 (2 – 5)	5 (3 – 6)	5 (3 – 6)	0 (0 – 0)	2 (1 – 3)	3 (2 – 4)	2 (1 – 3)	1 (0 – 1)	21.0%
Singles									
Women	70 (52 – 77)	84 (63 – 93)	65 (48 – 73)	19 (14 – 21)	22 (16 – 29)	36 (28 – 47)	21 (15 – 27)	16 (12 – 21)	17.8%
Men	31 (25 – 37)	35 (27 – 41)	32 (26 – 38)	2 (1 – 3)	15 (11 – 21)	20 (15 – 27)	17 (12 – 22)	3 (2 – 6)	19.8%
Total	295 (243 – 330)	339 (280 – 373)	287 (240 – 318)	52 (36 – 64)	159 (134 – 187)	211 (159 – 249)	151 (116 – 177)	62 (45 – 76)	19.4%

Labor supply effects are computed by averaging over the whole sample the 100 simulated transitions, for each of the 100 draws of the utility parameters from the estimated asymptotic distribution of their estimator. The table reports the median values of each distribution and, in brackets, the 90% confidence interval. Total participation is the number of individuals starting to work after the reform. Variations in working hours are decomposed into change of hours for those previously in employment (intensive margin) and change due to participation effects (extensive margin). In both cases, it is expressed in FTE computed as (change in) the total number of hours divided by 40 weekly hours. Share of newly rationed is the share of new participants who are restricted by demand side constraints

## 6.1 The mini-job reform

On the extensive margin, the mini-job reform induces a net positive effect on participation of 56,000 (resp. 43,000) individuals according to the *constrained* (*unconstrained*) model, which is in line with the government’s goal of making work pay. Both models show that the labor supply effect is mainly borne by women living in couple households while the participation effects for singles and, in particular, for men are negligible. This result is in line with initial intuitions: the net gain of the reform concerns primarily secondary earners in couples; the budget sets of singles are hardly affected due to the fact that gains at part-time are neutralized by the social assistance scheme. Women in couples are induced to take part-time jobs so that variations at the extensive margin are smaller once translated in FTE. For instance, with the *constrained* model, the participation effect in this group represents 36,000 additional workers, which corresponds to an increase in total hours equivalent of 11,000 FTE.

The positive participation effect is counteracted by a negative effect on the intensive margin for people already at work, especially for part-timers. Thus, our results show that individualized schemes do not completely prevent the inefficiencies found in the family-based policies in force in the US and the UK. Truly, individualized transfer programs do not create high marginal tax rates for the secondary earners at the extensive margin; yet, the high withdrawal rate of the mini-job plays negatively on the intensive margin. According to the unconstrained model, women in couples have the highest negative effect on this margin: labor supply decreases by 17,000 FTE. Over the whole population, the model suggests a reduction of about 33,000 FTE, while the increased participation is about 28,000 FTE, yielding a net reduction of labor supply by around 4,000 FTE.

According to the *unconstrained* model, women in couples have the highest negative effect on this margin: labor supply decreases by 17,000 FTE. Over the whole population, the model suggests a reduction of about 33,000 FTE, while the increased participation is about 28,000 FTE, yielding a net reduction of labor supply by around 4,000 FTE. When ignoring the involuntarily unemployed, our findings are very similar to those of Arntz et al. (2003) and Steiner and Wrohlich (2005).<sup>14</sup>

The picture is only slightly different when the estimates are based on the *constrained* model. Not surprisingly, differences mainly come from (smaller)

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<sup>14</sup>Note also that our estimates seem considerably lower than the 523,000 new mini-jobs that have been created between March 2003 and March 2004 according to the Federal Employment Agency (Bundesagentur für Arbeit 2004). However, we focus here on additionally created employment, while the estimates of the FEA include persons who were already employed before the reform (241,000 with an income between 326 and 400 Euro and 196,000 with an income higher than 400 Euro) and who are now categorized as “mini-jobbers.” Official job creations then decrease down to around 86,000. Furthermore, it should be kept in mind that we concentrate on the main labor force, excluding students and pensioners from our analysis. Assuming that these groups account for about a third of the total effect, the FEA number is further reduced down to around 50,000, i.e., a number well within the estimated confidence intervals of both models.

participation effects, while effects at the intensive margin are very similar. The total net effect on hours is a reduction of 11,000 FTE, to be compared to the reduction of 4,000 FTE predicted by the unconstrained model. The difference is not statistically significant. Note that these figures hide several sources of discrepancy. First, rationed workers can contribute to the participation effect only in the unconstrained case. Second, the voluntarily unemployed who are induced to take up a job after a reform face a rationing risk with the constrained model. Third, elasticities are overstated by the unconstrained estimation. The first two effects cumulate to explain differences at the extensive margin. Interestingly, since the mini-job reform mainly targets secondary earners, i.e., the group with the highest proportion of voluntarily unemployed persons, these effects are limited.

## 6.2 The Employment Bonus

The Employment Bonus has a very strong participation effect: almost 160,000 additional individuals are estimated to enter the labor market according to the *unconstrained* model. Half of the new entrants are women in couples. Part of the success is due to the fact that the Bonus targets full-time activity and avoids the specific (dis)utility from working part-time revealed by labor supply estimates. In addition, targeting earnings at full-time allows to escape partly from the neutralizing effect of social assistance for single individuals, which explains a substantial participation effect on this group while the mini-job reform was totally ineffective. Note also that jobs created by the Employment Bonus (resp. mini-job reform) are most often full-time (resp. part-time) activities, so differences in participation effects of the two reforms are even larger when considering FTE measures.

For those in employment, the “part-time trap” is avoided by conditioning eligibility on wage rates rather than earnings. Therefore, we do not ex-ante expect negative labor supply effects at the intensive margin. A minor reduction may nevertheless be encountered amongst people working over-time since the benefit reaches a maximum at standard full-time (i.e. 40 h/week) and then decreases. The model indeed forecasts a small reduction among men in couples, the group with the largest proportion of over-time workers. In all other groups, however, hours of work increase, leading to an overall gain of 211,000 FTE.

Although the general direction of the labor supply effects is the same under both models, the size of these effects is substantially overstated by the *unconstrained* model. This result is essentially due to discrepancies in predicted participation effects, even more so in groups where the proportion of involuntary unemployment is large, i.e., men in couples and single individuals. In these groups, participation effects are two or three times larger with the unconstrained model. For married or cohabiting women, the overstatement is smaller in relative terms (“only” 50%) – this is the group with the largest voluntary unemployment – but the largest in absolute terms since this group displays large effects. Therefore, overall, our empirical results underline that



individualized transfer programs conditioned on low wages have a more efficient design in terms of employment than tax credits based on family earnings.

Finally, it is important to stress that the superiority of the Employment Bonus reform is not driven by larger budgetary costs (3.09 billion Euro per year vs 1.89 for the mini-job reform). In effect, when behavioral responses are accounted for, the net costs of the two reforms become comparable. Very clearly, the tax base of the Employment Bonus increases so that its net cost decreases substantially. A comparison of the cost efficiency of both reforms is enlightening; for this purpose, we simply divide total cost after labor supply responses by the number of new entrants after the reform (when using constrained estimations). The mini-job performs substantially worse with a unitary cost of 44,000 Euro vs 19,000 for the Employment Bonus. Naturally, efficiency concerns may need to be balanced against other social objectives.

## 7 Conclusions

In this paper, we evaluate the employment effects of two different making work pay reforms in Germany, namely, the recently implemented mini-job reform and a hypothetical scheme inspired by the Belgian Employment Bonus. Both transfer reforms are individualized and, by construction, avoid the negative effect on the participation of secondary earners in couples, as witnessed in the case of tax credits based on family income (EITC in the US, WTC in the UK). In this respect, individualized transfer programs seem to be more efficient in fostering employment. However, other features of the policy design have important implications and, in particular, the ability to distinguish between low effort and low productivity. A transfer conditioned on low wages (e.g., the Employment Bonus) avoids negative effects at the intensive margin and generates larger participation effects than a reform targeted at low earnings (e.g., the mini-job reform). The design of the mini-job reform limits its effect on labor supply, in particular, in combination with joint income taxation for married couples. This is responsible for a potential “part-time trap.” Together with incentives to reduce working hours for those already in work, this effect outweighs the positive participation effect of the reform. Overall, a move toward in-work policies conditional on wages rather than earnings seems recommendable both in terms of employment effects and cost efficiency. Naturally, conditioning on wages requires reliable information on working time and may imply additional administration costs.

For the empirical evaluation of the employment effects, we apply a double-hurdle model that accounts for unemployment risk when estimating a structural labor supply model. Most ex-ante evaluation studies assess the employment effects of tax reforms assuming unemployment to be voluntarily chosen. We suggest an original characterization of the bias affecting labor supply elasticities and the predictions of employment effects when involuntary unemployment is ignored. An unconstrained model is misspecified (specification bias) and would overstate participation effects but also taste for leisure

(preference bias). Yet, it seems that the participation bias dominates so that elasticities from unconstrained models are overstated. The overall bias is particularly apparent for groups with a high share of involuntarily unemployed amongst the nonworking, such as single individuals or men in couples.

Finally, comparing the predictions of employment effects by the standard labor supply model and by the double-hurdle model yields interesting information. Ignoring demand-side constraints essentially leads to an overstatement of effects at the extensive margin. The reasons are twofold: (1) rationed workers are unduly included in the group of those who can react to the reform and (2) the voluntarily unemployed who are induced to take a job after a reform are mistakenly not subject to the rationing risk. As illustrated by the mini-job reform, the overall bias is not critical when responses are small and driven by groups who are less affected by rationing, typically voluntarily inactive women in couples. However, when responses are substantial and when other groups are concerned (men in couple households, single individuals), the overall prediction error of employment effects becomes large. In the case of the Employment Bonus, the total employment effect is overstated by about 60%. Our findings call for a reassessment of the different policy options of make work pay schemes, accounting for the possibility of rationing. In particular, conclusions above convey that simulations of EITC and WTC for continental Europe – cf. Bargain and Orsini (2006) and Haan and Myck (2007) – should not be “too wrong” as far as the negative effect on secondary earners is concerned. Yet, positive effect at the intensive margin for all other groups may be substantially biased.

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## A Technical appendix

### A.1 Unconstrained model

We assume the error terms  $\epsilon_{ij}$  of the utility function to be i.i.d. according to an EV-I distribution. Then, the probability that alternative  $k$  is chosen by household  $i$  is given by McFadden (1974):

$$P_{ik} = \Pr(V_{ik} \geq V_{ij}, \forall j = 0, \dots, J) = \frac{\exp U(Lf_{ik}, C_{ik}, Z_i)}{\sum_{j=0}^J \exp U(Lf_{ij}, C_{ij}, Z_i)}. \quad (3)$$

Further, we model the utility function in a quadratic specification as in Blundell et al. (2000). Preferences for income and leisure coefficients vary

conditional on age, number and age of children, and region of residence. We follow van Soest (1995) and introduce dummy variables for the part-time categories in order to capture specific (dis)utility from working part-time. In the estimation, we do not consider potential effects of unobserved heterogeneity, which implies that the IIA property holds. However, Haan (2006) has shown that labor supply elasticities estimated on the same data as in the present study do not differ significantly when unobserved heterogeneity is introduced.

### A.2 Constrained model

Denoting  $d$  as the desired hours and  $p$  as a dummy representing nonrationing, we can summarize the situation of a single individual with three possible states,<sup>15</sup> to be voluntarily inactive, to be rationed, and to participate without being rationed. In the present set-up, these probabilities are written as follows:<sup>16</sup>

$$P_{i0}^{VOL} = \Pr(d_i = 0) = \frac{\exp U(L_{i0}, C_{i0}, Z_i)}{\sum_{j=0}^J \exp U(L_{ij}, C_{ij}, Z_i)}, \tag{4}$$

$$P_{ik}^{INVOL} = \Pr(d_i = k, \forall k > 0, p_i = 0) = \frac{\exp U(L_{ik}, C_{ik}, Z_i)}{\sum_{j=0}^J \exp U(L_{ij}, C_{ij}, Z_i)} [\Phi(\beta X_i)], \tag{5}$$

$$P_{ik}^{EMP} = \Pr(d_i = k, \forall k > 0, p_i = 1) = \frac{\exp U(L_{ik}, C_{ik}, Z_i)}{\sum_{j=0}^J \exp U(L_{ij}, C_{ij}, Z_i)} [1 - \Phi(\beta X_i)]. \tag{6}$$

As in Duncan and MacCrae (1999), we assume that the error terms of the labor supply model and the probability of rationing are independent, which allows us to estimate the unemployment risk separately. A more general approach could make use of a simulated maximum likelihood to introduce a correlation between these two terms. The assumption of independent error terms implies, in particular, that, conditional on observed characteristics, having positive desired hours is independent on the risk of rationing. In this sense, our specification rules out discouragement effects, which are unobservable.

<sup>15</sup>For couples, we estimate one rationing probability per spouse.

<sup>16</sup>We can reasonably assume that desired hours of employed individuals coincide with actual observed hours. This is indeed the case for over 85% of the working population, while means are, respectively, 21.9 and 21.2 h per week (when including nonparticipants). For the involuntarily unemployed, we make use of the information in the data about which type of contract they are looking for, part-time (21–34 h) or full-time (35–40 h) work. This additional information allows us to assign the involuntary unemployed to the respective group and, thus, to estimate the model more precisely. Doing so, we assume that the involuntary unemployed have only a restricted choice set when working.

Indeed, we mix those who are “voluntarily” inactive for two different reasons: either due to high fixed cost of work (due to, e.g., childcare costs) or to high search costs (e.g., due to rationing). As mentioned above, the individual unemployment risk is conditioned on the individual employment history and, hence, partly accounts for discouragement effect related to the employment status over the last 3 years (yet the initial state is assumed random). To account for discouragement, this effect could be modeled explicitly (structurally) by introducing job search cost that would increase with the risk of unemployment (cf. Duncan and MacCrae 1999). The sample log-likelihood to be maximized can be derived by summing individuals’ probabilities for their working state, using Eq. 3 for the unconstrained model or Eqs. 4–6 for the constrained model.

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