Essays on local labor markets, firm taxation and worker mobility

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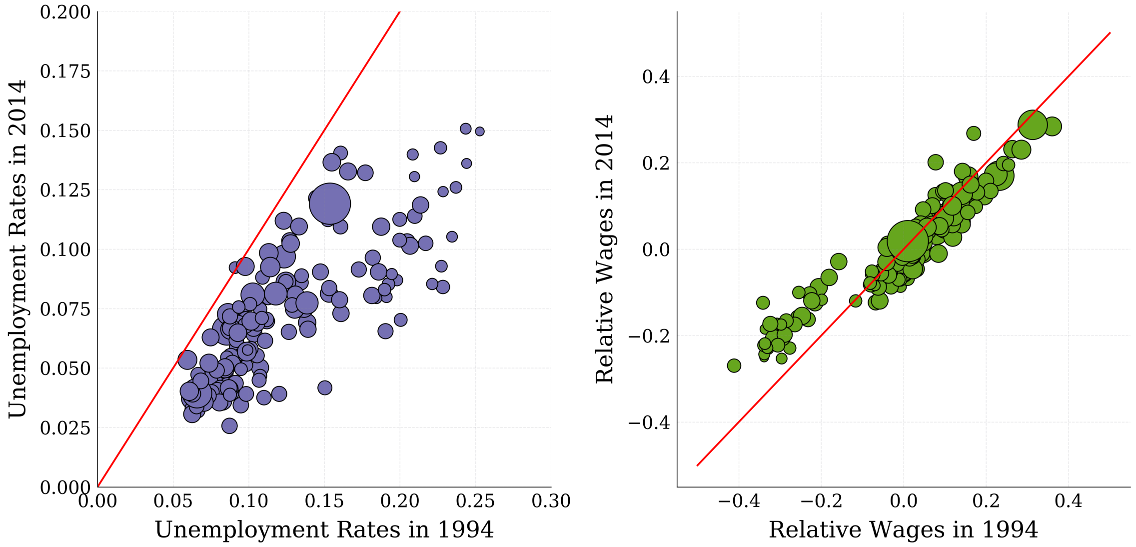
Chapter 1
Introduction

Economic opportunities are not equally distributed within a country—large geographical wage and unemployment rate differentials have been documented for a range of countries (OECD, 2013, and International Monetary Fund, 2019). These differences are rather persistent and have been linked to the deterioration of individual well being. Growing up in a distressed labor market has severe implications for happiness, health and life-expectancy (Chetty and Hendren, 2018).

This thesis investigates the causes and consequences of regional disparities in Germany and consists of three self-contained essays. Each essay utilizes spatially fine-grained microdata for Germany, whose local labor markets are far from being equal: Unemployment rates and wages exhibit stark and persistent dispersion across employment agency districts, as illustrated in Figure 1.1. The three essays differ in focus and methodology but are all inseparably related through the spatial level of analysis. The second chapter of this thesis studies the relative impact of two distinct local business tax instruments on workers' wages. While the primary contribution is the estimation of the effect of a revenue-neutral substitution between two tax instruments, the chapter provides evidence of the amplification of cross-regional wage disparities through local taxation. The third chapter relates regional unemployment differentials to worker flows, recovers underlying structural variations across regions, and investigates the impact of optimal local labor market policies that attenuate within market inefficiencies and balance the adverse effect of unemployment insurance benefits. The fourth chapter studies the characteristics of individual inter-regional migration decisions of employed workers and examines the relationship between individual earnings gains and location characteristics.

The forces behind persistent within-country disparities and their implications for policies and welfare are complex and have been widely discussed. On the one hand, long-run regional disparities can be an equilibrium outcome driven by, e.g., agglomeration forces or regional characteristics. On the other hand, they might be an outcome of market failures and possibly enhanced through government policies. Several studies have investigated the welfare implications of government policies and the results are quite mixed. When the government intervenes and supports, e.g., lagging regions, this might lead to aggregate losses as production factors relocate to or are kept in regions

Fig. 1.1: The persistence of unemployment and wage differentials



Notes: This figure displays regional unemployment rates and the (log) deviation of the regional from the average wage for 154 employment agency districts in Germany in 1994 and 2014. The size of the circles represents the size of the labor force. *Source*: Own calculations from the panel of integrated employment histories (SIAB).

where they are less efficient. In general, place-based policies must be carefully designed to ensure a beneficial adjustment. To do so, knowledge of the source of the disparities and how policies affect labor market outcomes is detrimental (Kline and Moretti, 2014).

Governments can shape the spatial distribution through policies. Local tax differences lead to variations in economic outcomes (Serrato and Zidar, 2016 and Slattery and Zidar, 2020) and fiscal autonomy of regions can lead to a misallocation of agents across space and, thus, have significant consequences for aggregate and regional welfare (Fajgelbaum et al., 2019). The second chapter of this thesis is related to this string of the literature since it demonstrates that even in a revenue-neutral scenario, the choice of the tax instrument matters for economic outcomes, in this case, workers' wages.

A specific set of place-based policies concern local labor markets where government interventions are often designed to mitigate unemployment or foster job-creation and are utilized in almost every country. Local search frictions that prevent the instant labor market clearing are a widely accepted rationale for active policy intervention. In addition, the effect of national labor market institutions, e.g., bargaining systems, unemployment insurance, or employment protection laws, can vary across local labor markets and potentially foster disparities (Boeri et al., 2019 and International Monetary Fund, 2019). The third chapter of this thesis determines structural disparities between local labor markets in Germany and evaluates the benefits of region-specific labor market policies.

As indicated before, policies designed to foster the development of specific regions also have an effect on the major factor that should contribute to the convergence of economics outcomes: worker mobility. Migration within a country is deemed as a key adjustment mechanism to restore the spatial equilibrium when local shocks hit an economy. Over the last decades, a large literature investigated the migration patterns within countries and showed that regional convergence of disparities through migration is slow (Blanchard et al., 1992 and Amior and Manning, 2018). Therefore, government intervention might be appropriate because frictions concerning worker or firm mobility could be in place and hinder a faster adjustment process. However, national policies that are in effect now may impede the convergence. For example, the current income tax system could constitute a distortion because the federal income tax is not placed based (Albouy, 2009). In general, migration decisions are known to be strongly influenced by disparities in wages and local labor market conditions (Beaudry et al., 2014). Understanding the location choice of workers is, thus, an important research subject. Therefore, chapter four investigates individual migration behavior over the life-cycle, taking into account specific destination characteristics.

Chapters two and three are concerned with how local taxation or place-based policies affect local outcomes. The former presents first evidence on the relative impact of two tax instruments on workers' wages. Chapter three is specifically concerned with local labor markets and presents new evidence on the importance of worker flow rates for unemployment rate disparities and structural variations between local markets. Also, a structural model is utilized to investigate the benefits of optimal local policies concerning welfare and employment. Chapter four, in turn, documents novel empirical evidence on workers' life-cycle migration choices and how cross-regional variations in labor market outcomes are related to the returns to migration. The remainder of this introduction explains the main contributions and approaches utilized in each chapter in more detail.

Chapter two provides evidence on the governments' ability to shape the spatial distribution of wages through local taxation. This chapter utilizes a distinct reform of the local business tax law in 1980 and causally identifies spatial wage differences that are due to regional variation in the application of two local business tax instruments: business income and payroll taxes. Existing studies provide ample empirical evidence that payroll, as well as business income taxes, are passed on to workers' wages. This chapter bridges both strings of the literature as it provides the first evidence of the average wage difference between the two tax instruments when the amount of revenue raised is similar.

I utilize an extensive establishment panel data set from the federal employment agency[1], the regional variation in the application of the payroll tax, and the repeal of payroll taxation in 1980 and provide causal evidence of the relative impact of the two instruments on wages. Before 1980 there were two local business tax regimes in place in Germany: While a small share of municipalities used a mix of payroll and business income taxation, the majority raised their revenue by taxing business income. Importantly, I provide suggestive evidence that local governments substituted the business income with the payroll tax because the per-capita revenue from local business taxation was similar across regimes. This setting establishes clear defined treatment and control regions and enables me to identify the average difference in wages between the two schemes. Furthermore, the repeal of payroll taxation in the course of a tax reform in 1980 allows me to compare firms located in treatment and control regions when they are subjected to the same tax regime.

Using event study models, I estimate that wages are 1.4 percent higher in treatment compared to control regions during the period before the reform. Thus, despite similar tax revenue, wages are higher when local governments substituted a business income with a payroll tax. Moreover, the abolition of local payroll taxation in 1980 dissolved the wage difference between regions lending support to the causal identification of the average effect. Further estimates of the average wage difference, obtained using difference-in-differences strategies, confirm the size and significance of this effect. A welfare analysis using a stylized model shows that a planner would always resort to payroll taxation as the business income tax leads to an inefficient capital choice.

In addition to the relative effect of business income and payroll taxation, this chapter provides estimates of a lower bound of the payroll tax elasticity using a difference-in-differences strategy. The estimated tax elasticity implies that a one percent decrease in labor costs, induced by a decline in payroll taxes, leads to an increase in wages of 0.72. This result provides new evidence of the importance of the tax-benefit link for the pass-through of tax burdens to workers. Since this local business tax did not contribute to any social security system, the passed-on tax burden had no offsetting effects for workers in terms of deferred compensation. While some of the existing literature emphasizes that the degree of pass-through to affected worker groups should rise with the strength of the tax-benefit link, this chapter suggests that firms shift the burden of a tax with no direct benefit link to all employees.

The following two essays are part of a joint research project with Philip Jung and Edgar Preugschat. These chapters build on social security microdata[2] and present new evidence on either local labor market disparities or individual life-cycle migration behavior.

[1]The Establishment History Panel.

[2]The panel of integrated employment histories provided by the federal employment agency.

Chapter three provides a detailed account of the evolution of local labor markets in Germany over two decades and utilizes the empirical variations in wages, unemployment rates, and vacancies to infer the underlying structural disparities. Furthermore, this chapter evaluates the welfare and employment effects of two sets of optimal labor market policies: one designed to attenuate local search externalities and externalities of national unemployment insurance and one designed to reduce aggregate unemployment by efficient relocation of the unemployed across markets.

In the first part of this study, we present new evidence on the importance of the worker flow rates from unemployment to employment (job-finding rate) and from employment to unemployment (separation rate) for the determination of regional unemployment rate differentials. Previous research highlights how labor market policies affect the unemployment rate varies conditional on which flow rate is the primary cause of its variations (Hartung et al., 2018). We utilize social-security panel data and newly digitized data on vacancies from the federal employment agency and document the development of local labor markets from 1994 to 2014. The spatial unit of analysis are employment agency districts (in the following EADs or districts). Our results show that from 1994 to 2004, separation rates explain the 70% of the variation in unemployment rates across districts and only 40% during the subsequent ten years. Since job-finding rates account for the remaining variation, this shift suggests that optimal regional policies should be able to account for both differences in separation and job-finding rates. This development is inseparably linked to the development of specific regions in Germany, which we trace out in detail in the chapter.

This finding suggests that policymakers ought to take into account both the layoff (separation) and hiring (job-finding) margin when developing policy responses to mitigate local unemployment. Therefore, in the second part of this chapter, we present a multi-island model with labor market frictions, risk avers workers, and endogenous separations and hirings. We, first, utilize the model to recover structural disparities across EADs for two time periods 1994-2002 and 2008-2014 to capture the changing importance of the worker flow rates. These parameters are identified, given the national unemployment insurance system and employment protection laws, relative to a benchmark scenario. Second, given these structural differences, we evaluate the welfare and employment effects of an optimal regional policy mix, which features unemployment insurance benefits, layoff taxes, hiring subsidies, and a production tax. The optimal mix balances the trade-off between moral hazard and consumption smoothing, given the unobservable search effort of workers, and removes imbalances from local search externalities. We find that the optimal region-specific policy mix leads to large welfare and employment gains for each district. Therefore, labor market policy that is set at the national level and does not take into account structural differences between regions implies significant welfare losses. However, we show that setting optimal policies for each region does not necessarily attenuate unemployment or welfare disparities.

We last turn to an examination of the efficiency of the actual allocation of unemployed across regions. Taking local labor market frictions and vacancies as given, we provide estimates of the potential increase in aggregate employment when the unemployed could be relocated at no cost. The efficiency measure is based on a comparison of the observed to an optimal allocation across markets. This measure provides an upper bound for the impact of any policy measure that aims to foster the reallocation of individuals. We find that the optimal allocation of unemployed workers leads to an average decrease in unemployment of 15%. Compared to the existing literature, our results are for a finer spatial level and a more extensive time horizon.

Chapter four studies the characteristics of individual inter-regional migration decisions of employed workers and the relationship between destinations and individual wage gains. This chapter complements the previous one by explicitly documenting and analyzing the individual relocation behavior across local labor markets. Building on the findings that regional wage levels are a crucial factor for individual migration decisions (Kennan and Walker, 2011), we investigate the impact of between-region wage differentials on the returns to migration. This chapter contributes to the empirical literature on regional mobility in Germany and provides a more detailed account of life-cycle migration behavior at a low spatial level. Furthermore, the chapter provides new evidence on the relationship between long-term gains from migration and wage differentials, thus, advancing the literature on returns to mobility.

We use a large German administrative panel data set which allows us to follow individual workers over relatively fine-grained work locations and the life-cycle.[3] We find, first, that most workers never leave their initial location, and only a small fraction of workers moves each year with a rate that declines with age. Conditional on moving, most workers change their work region only once or twice throughout working life and primarily to districts that are within 100 kilometers of their starting region. Second, we provide new evidence on long-term migration gains and their relationship to region-specific wage levels. We estimate these gains non-parametrically and relative to a counterfactual earnings path of a control group. Our results show that returns to mobility are persistent and substantial as movers' earnings are 12% higher than the control groups during the ten years following their relocation. When we relate these gains to wage differentials between origin and destination regions, we uncover substantial variation: A move to a district with higher wages exhibits gains that are up to four times larger compared to a transition to a region with lower wages. Although these gains increase in the average wage of the destination, a large fraction of moves (50%) is directed to regions, which have a lower wage as in the origin location. This lack of

[3]The panel of integrated employment histories provided by the federal employment agency.

aggregate directness of migration, combined with the low mobility, helps to explain the long-term stability of the interregional wage dispersion.

Chapter five provides concluding remarks.

Chapter 2
The relative impact of payroll and business income taxes on wages in Germany

2.1 Introduction

Local business taxation, which generates a significant component of municipal revenue, is recognized as an integral part of the tax system in Germany, but its design has been critically discussed for years.[1] Today, municipal revenue from local business taxation stems from a tax on business income, while until 1980, part of the revenue was raised using a payroll tax. This decision in favor of the business income tax is not straightforward. In fact, some argue for a revival of a tax on payroll because municipal revenue would not be subject to large cyclical fluctuations, unlike revenue from the business income tax (Zimmermann, 2002). Other countries have even transformed their local business tax system in the exact opposite direction. For example, the system in Austria today, which once was largely similar to the German system, is based solely on a tax on firms' payrolls.

A key policy issue in any discussion about choosing a tax instrument certainly is who bears the incidence of the tax. While the literature documents negative wage responses for payroll and business income taxes, a comparison of these two tax instruments is still missing. This article closes this gap and investigates the relative wage impact of a payroll tax compared to a tax on business income by exploiting a local business tax reform in 1980 in Germany. My results show that taxing payroll instead of business income leads to higher wages. In addition to the relative effect, this article provides an estimates of the payroll tax elasticity. Contrary to previous results, I find that firms pass on the burden of this particular payroll tax almost entirely to wages.

The design of the local business tax system in Germany before 1980 constitutes a suitable setting to causally identify the relative impact of payroll and income taxation on wages, as it enabled two local business tax regimes to be in place simultaneously: A fraction of municipalities levied a tax on firms' payroll and income, while the majority

[1] See for example Wellisch and Walz (1991), Zimmermann (2002), Fuest and Thöne (2003), Döring and Feld (2005), and Fossen and Bach (2008).

of municipalities taxed only business income.[2] This setting establishes clearly defined treatment and control regions. Hence, the identification of the relative impact on wages is based on regional variation in the application of the payroll tax. Importantly, per-capita revenues from local business taxation were similar across regimes. Therefore, differences in the burden of the local business tax do not drive the results. Moreover, the payroll tax was abolished as part of a local business tax reform in 1980, which led to a uniform local business tax system within Germany. Consequently, the regions affected by the repeal increased their business income tax to cover the foregone revenue. The post-reform period enables me to document wage differences between treatment and control regions when firms are subject to the same tax regime.

In the empirical analysis, I utilize the Establishment History Panel (BHP) combined with newly digitized municipal tax data from historical official tax reports of the Association of German Cities and the Federal Statistical Offices der Laender. The analysis is based on a subset of all municipalities due to differences in the regional level of the establishment and tax data.[3] Furthermore, due to data limitations and in line with the literature on corporate taxation, I focus on the effect of local business taxes on wages.

The first contribution of this paper is to provide a comparison of the impact of payroll and business income taxation on wages. To validate the identification, I, first, estimate an event study model to establish flat after-treatment trends. The estimates show no significant differences in wages between treatment and control regions after the reform was enacted and all firms were subject to the same tax regime. Wages before the reform differed significantly across regimes: Establishments located in municipalities that levied the payroll tax paid higher wages compared to firms located in regions that taxed only business income. Thus, the repeal of the local payroll tax in 1980 caused wages to decline in previous payroll tax collecting regions relative to non-collecting regions. Second, I obtain estimates of the average wage difference between the two tax regimes using difference-in-differences (DiD) models. The estimates show that wages are higher when payroll instead of business income is taxed. Overall, the repeal of the payroll tax led to relative wage reductions of 1.4% in municipalities that levied the tax. Moreover, I provide evidence on the heterogeneity of the impact by firm size and industry sector. The estimates are significant for all industries and small and medium sized firms. This finding complements the recent research of Fuest, Peichl and Siegloch (2018) (FPS), who provide compelling evidence that the German business income tax is passed on to workers, as my result suggests that substituting the business income

[2]In addition to the tax on payroll and business income, municipalities did tax the firms' capital stock. The tax rate was proportional to the business income tax rate, as will be discussed further below, and will be treated as part of the income tax.

[3]The expression establishment is used because the observation unit is a regionally and economically delimited individual plant, which may belong to a larger company. Another term often used for such economic units is *permanent establishment*. In the remainder of this paper, I use the terms firm and establishment interchangeably.

tax with a revenue equivalent payroll tax would increase wages.

The observed wage effect can be rationalized within a partial competitive equilibrium model with immobile labor. Within this framework, both tax instruments lower the wage. While the payroll tax is passed on entirely onto wages, the income tax lowers wages through capital demand (i.e., through the marginal product of labor). Wages are higher when payroll is taxed because the business income tax leads to an inefficient investment decision, whereas the payroll tax decreases wages without any impediment to labor.

As a second contribution, I isolate the wage effect of the payroll tax and estimate the corresponding elasticity. Here, I additionally utilize the substantial variation in payroll and income tax rates for identification. Both tax rates exhibited substantial variation across regions as local governments set them autonomously. The point estimate, obtained using my preferred specification, states that a 1% increase in labor costs, due to a rise in the payroll tax rate, causes wages to decline by at least 0.72%. This estimate is, by itself, of interest as the revenues from the payroll tax contributed solely to the municipal budget. The majority of prior payroll tax research does not distinguish between contributions to social security and payroll taxes, as they are both levied on the sum of wages. However, a key difference is that social security contributions come with a direct tax-benefit link (Gruber, 1997), which is often used to rationalize the pass-through of the burden onto wages. The almost complete pass-through of payroll taxes in Germany indicates that firms can shift the burden, despite no tax-benefit link.

Literature This article is related to the two strings of the literature that investigate the wage response to payroll or business income taxation. The consensus of the empirical literature is that both taxes are at least partially passed on to workers' wages. Studies on the incidence of the payroll tax mostly focuses on developing countries in Latin America (Gruber, 1997; Kugler and Kugler, 2009; Cruces et al., 2010), northern European countries (Bennmarker et al., 2009; Korkeamäki and Uusitalo, 2009; Saez et al., 2019) or Canada (Deslauriers et al., 2018) and find a partial to complete pass-through of the tax onto workers wages.While the literature concerning the payroll tax is scarce, there are several studies that investigate the impact of business income taxation.[4] Three articles confirm a partial pass-through of the income tax to wages in Germany: While Dwenger et al. (2019) focuses on the impact of the federal cor-

[4]Earlier studies concerning (business) income taxation(e.g. Felix (2007), Desai et al. (2007) and Hassett and Mathur (2015)) use cross-country variation in tax rates and find a negative impact of corporate income taxes on wages. In more recent contributions, studies turned to within country analysis using administrative firm or worker panel data (e.g., Liu and Altshuler (2013), McKenzie and Ferede (2017) and Serrato and Zidar (2016)). Another string of the literature on corporate income taxation introduce wage bargaining as a possibility for firms to shift the corporate tax burden to worker (e.g., Arulampalam et al. (2012), Felix and Hines (2009) and aus dem Moore (2014)). For a thorough review of the empirical literature on corporate taxation the reader is referred to Clausing (2012) or Gravelle (2013).

porate income tax, Bauer et al. (2012) and FPS analyze the incidence of the local business income tax. FPS is certainly the most influential study, credibly identifying the wage elasticity with respect to business income taxation. They use matched employer-employee panel data, covering the years 1994 to 2010, and estimate a net of tax elasticity of 0.38%. The present study combines both strings of the literature, as my results suggest that a revenue-equivalent payroll tax has a less severe impact on wages than a business income tax.

Concerning the literature on payroll taxation, this article is closest to Deslauriers et al. (2018), who emphasize the importance of investigating a payroll tax that does not contribute to social security systems. They use administrative firm and worker data from Canada and find no impact of the payroll tax on firms' employment or profits. Their results with respect to wages are in line with the evidence presented in the present article: the payroll tax is almost entirely passed on to workers. In addition, I estimate a substantially higher elasticity compared to Neumann (2017), who finds that the social security burden is equally split between employers and employees.

Moreover, the paper adds to the literature investigating the importance of the tax-benefit link for the pass-through of social security contributions (SSCs) or payroll taxes. The importance of this link is documented by Bozio et al. (2019). They show that the impact of SSCs on wages increases with the strength of the tax-benefit link. The almost full pass-through of the payroll tax in Germany, which did not contribute to a social security system and, hence, had no tax-benefit link stand in contrast to their results. The crucial difference between Bozio et al. (2019) and the present article, is that they analyze how firms pass on SSCs to the affected worker groups. Similar, Saez et al. (2019) show that firms in Sweden are not able to pass the tax burden with a weak benefit link onto a specific worker group but rather shift the entire burden onto all workers. The results presented here support the firm-level shift argument of Saez et al. (2019) as the flat payroll tax in Germany with no benefit link is passed on almost completely to workers.

This article is organized as follows. Section 2.2 describes the data and discusses the local business tax system and its reform in 1980. In section 2.3, the implications of the reform are quantified in a partial equilibrium framework. Section 2.4 lays out the empirical strategy and presents the results. Section 2.5 concludes.

2.2 Institutional setting and data

This section describes the institutional setting, tax and establishment data, the sample selection and the reform of local business taxes in 1980. Government tasks in Germany are distributed across territorial entities and the local business tax (Gewerbesteuer) falls into the jurisdiction of municipalities (Gemeinden). In the 1970s, there were approximately 8500 municipalities in West-Germany and revenues from local business taxation were a significant part of the municipal budget. Until 1980, municipalities set three local business tax rates: income, capital and payroll. The last had to be approved by the respective federal state, which led to regional variations in its application. As a consequence, only 7.4% of local governments levied the tax. In Bavaria and Saarland, for example, no municipality collected payroll taxes and in Baden-Württemberg, only one city (Mannheim) levied a tax on firms' payrolls. States with the highest share of municipalities that raised payroll taxes are North-Rhine Westphalia (36%) and Hesse (26%). The tax base and a federal rate for each local business tax are set at the national level. Municipalities decide on a local multiplier and, therefore, have autonomy over the tax rate, which led to substantial variation in the latter.

2.2.1 Local business taxation

Local governments' revenue from local business taxation from 1977 to 1979 was raised using income, capital and, in some municipalities, payroll taxes.[5] Tax rates consist of two components: a basic federal rate t (Steuermeßzahl), which is valid for all municipalities, and a multiplicand λ (Hebesatz) set by each local government. Corporate and non-corporate firms are subject to local business taxation, while most firms in the public, agricultural and fishing sector or liberal professions are not liable (§§2, 3 GewStG, 1977).

The payroll tax (Lohnsummensteuer) has three noteworthy characteristics. First, the tax was levied on the sum of pre-tax wages paid by firms and the tax rate was flat. Second, revenue raised from payroll taxation did not contribute to any social security system but exclusively to the municipality budget, i.e., a worker would only benefit indirectly through amenities provided by the local government (Andreae, 1958). Third, a municipality could only levy a tax on payroll if the respective federal state approved payroll taxation.

Payroll tax rates $\tau_w = t^w \lambda_w$ consisted of a local multiplier λ_w and a federal rate t^w. The amount of payroll taxes payable for a firm i residing in county j thus reads

$$T^w_{i,j} = \tau_{w,j} \bar{w}_i N_i,$$

where $\bar{w}_i$ is the average wage at firm i and N_i is the number of employees. Throughout the observation period, the basic federal rate was 0.2%.

The local business income and capital tax rates (Gewerbesteuer auf Ertrag und Kapital) are constructed in the same way as the payroll tax rate. A basic firm income tax rate t^y is set at the federal level, and every municipality decides on a local multiplier

[5] Due to the application of capital taxation the tax regime with business income and capital taxation is not comparable to the comprehensive business income tax (U.S. Treasury (1992)).

λ_y, leading to the income tax rate $\tau_y = t^y \lambda_y$.[6] During the observation period, the local income tax multiplier was applied in local capital taxation as well, and only the federal multiplier differed: $t^k = \theta t^y$. Hence, the capital tax was proportional to the income tax: $\tau_k = t^k \lambda_y = \theta \tau_y$. The capital tax base was the modified assessed value of the capital stock, which until 1984 basically was the current period capital stock K_t (§12 GewStG, 1977).[7] A firm's capital tax liability thus is $T^k_{i,j} = \theta \tau_{p,j} K_i$.

Capital, payroll and income tax liabilities were deductible from the income tax base while capital costs were not, leading to an income tax liability $T^y_{i,j}$ of[8]

$$T^y_{i,j} = \tau_{y,j}[F(K_i, N_i) - w_i N_i - T^k_{i,j} - T^w_{i,j} - T^y_{i,j}].$$

Federal rates were set at 5% for income and 0.2% for capital taxation.

2.2.2 Data and sample selection

This section describes the firm panel data, the local business tax data and the sample selection. The administrative establishment data is from the Establishment History Panel (Schmucker et al., 2016) provided by the Institute for Employment Research.[9] The annual panel covers 50% of all establishments in Germany from 1975 to 2014, employing at least one worker subject to social security contributions. This panel contains information on the establishments' industry sector, location and the number of full- and part-time workers. Furthermore, different wage measures are available (e.g., mean and median daily wages for full-time employees). Wages are recorded in June in every year. As the data stems from social security notifications, wages are only reported up to a contribution limit and approximately 10% of full-time employees' wages are affected by a censoring limit each year. To circumvent the censoring problem, I follow FPS and use the median wage of full-time employees as the outcome variable in my analysis.[10] Industries are classified in 269 classes based on the industrial classification of economics activities from 1973, which is valid for my observation period. The location information indicates the county (*Kreis*) in which the company is located. County borders are based on the territorial classification of 31.12.2014.

The establishment data is combined with data on local business taxes digitized from

[6] Note, that business income is taxed at the federal as well. This description focuses on local business taxation.

[7] I assume, that the capital stock consists of debt and abstract from tax exempt amounts and reductions following Fecher (1980) and Richter and Wiegard (1991).

[8] With the income tax being deductible itself the effective tax rate is $\hat{\tau}_y = \frac{\tau_y}{1+\tau_y}$.

[9] This study uses the weakly anonymous Establishment History Panel 1975-2014 (BHP 7514). Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB) and/or remote data access.

[10] The results do not change if wages are adjusted for inflation. "Preisindex für die Lebenshaltung für alle privaten Haushalte im früheren Bundesgebiet (1995=100)". Verbraucherpreisindizes für Deutschland. Lange Reihen ab 1948. Statistisches Bundesamt (Destatis), 2018.

historical publications. The Statistical Yearbook of German Municipalities published by the Association of German Cities (*Statistisches Jahrbuch Deutscher Gemeinden, Deutscher Städtetag*) contains information on multiplier and tax revenue for the 503 largest municipalities in Germany. The data includes information on all cities. Aggregate statistics, as well as data on the weighted tax multiplier for each county, are digitized from the *Realsteuervergleich* published by the Federal Statistical Offices der Laender. The digitized data is supplemented by data on aggregate municipal revenue by different local tax sources covering 1966 to 1981 compiled by Köster (1984). Overall, I obtain data on local business tax rates for all cities and weighted tax rates for all counties for 1977 to 1983.

The combination of both data sets is not straightforward, as tax rates vary across municipalities, and the establishment data is on the county level. However, merging the data for cities (*kreisfreie Städte, Stadkreise*) is not a problem, as cities belong to both regional classifications − they are counties and municipalities. Furthermore, for most of the main analysis, assigning establishments to a tax regime is sufficient. An establishment can be assigned to a regime with certainty if none or all county municipalities levy a payroll tax. I use the weighted tax rate information to identify such counties and add them to the sample in parts of the analysis.

Moreover, large territorial reform in the 1970s, which led to many municipality mergers and border adjustments, complicates the assignment of municipalities to counties because the tax data is for non-fictional territory each year. Therefore, my observation period starts in 1977, when the majority of reforms had already been completed. Counties subjected to a territorial reform after 1977 are excluded from the sample if the inclusion could bias the results. A detailed description provided in the appendix (Section 2.A.1).

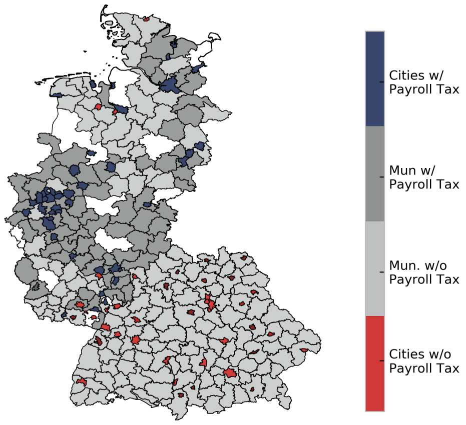
Figure 2.1 presents a map of all West German counties based on the territorial classification of 2014. The map serves merely as a visualization of the regional dispersion of payroll taxation in 1979. Counties displayed in white are excluded from the sample due to prior abolishment of payroll taxation or regional mergers. Observations of establishments based in Berlin are disregarded.

Focusing on 1977 to 1983, I construct two samples.[11] First, whenever I use the information on statutory tax rates, the results are presented for establishments located in cities (city sample). Second, I extend the city sample and include all counties in which no municipality levied a payroll tax (baseline sample). The expansion of the city sample with noncollecting counties adds 36 out of 125 counties located in the five states where some municipalities levied payroll taxes[12] and 109 counties located in the

[11] I do not use any data after 1984 since there was a change in the notification procedure now including one-time payments in the earnings measure that led to an increase in average daily wages.

[12] Schleswig-Holstein, Lower-Saxony, North Rhine-Westphalia, Rhineland-Palatinate, and Hesse

Fig. 2.1: Regional variation in payroll taxation in 1979



Notes: The figure displays a map of West German counties based on jurisdictional borders as of 2014. The counties are sorted into five categories: Cities with payroll taxation; cities without payroll taxation; counties where no municipality levied a payroll tax; counties where some municipalities levied a payroll tax; and counties and cities that abolished payroll taxation before 1980 or were subject to a merger (white areas).

other federal states to the sample. Due to regional mergers or prior abolishment of payroll taxation, 8 out of 92 cities are excluded from both samples.[13] The city sample covers 84 of 92 cities and approximately 55,000 establishments. The baseline sample contains information on above 100,000 establishments in 229 counties.

2.2.3 The tax reform of 1980

To reduce the firms' tax burden, the federal government enacted a tax reform (Steueränderungsgesetz 1979), announced on November 30, 1978, which came into effect on January 1, 1980.[14] The decree consisted mainly of three parts. First, municipalities' share of federal income taxes was increased by one percentage point to 15%. Second, the Gewerbesteuerumlage (i.e., the share of municipal local business tax revenue transferred to the federal government) was reduced by one-third. These first two modifications are federal changes valid for all municipalities and designed to allow local governments to reduce local business tax rates. Third, the annulment of payroll taxation affected only a minor share (7%) of municipalities. These municipalities lost, on average, 39% of their yearly revenue from local business taxation and were asked to modify their income and capital tax multiplier to balance their budget, given that the additional revenue from federal income or the decreased payments to the federal government might not be enough to cover the foregone income from payroll taxation.[15]

[13]Note, there were 92 cities in West-Germany in 1977. Three were merged with other counties and do not have the city status today. They are part of the eight cities that are left out of the samples.

[14]A reform in 1984 again modified the laws of local business taxation. The resulting changes applied to all liable firms.

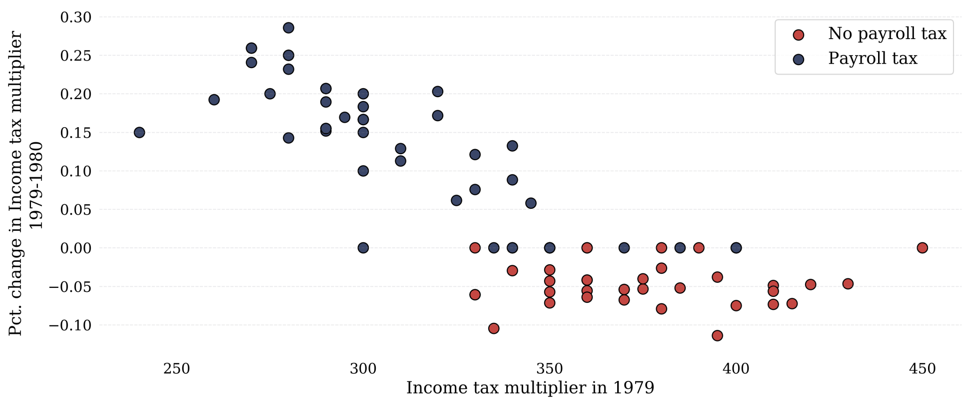
[15]Calculated using the data on the 503 largest municipalities.

Municipalities that did not levy a payroll tax should have used the additional revenue to relieve firms from a share of the local income and capital tax burden.[16] In the following paragraphs, I provide evidence that the financing needs were similar across local business tax regimes both prior to and after the reform.

The reform of local business taxes in 1980 severely affected the revenue of municipalities that levied a payroll tax. In 1979, aggregate revenue from income and capital taxation was above DM 15 billion, while revenue from payroll taxation amounted to DM 3.3 billion.[17] Although municipalities vote on their local multipliers at the end of each year, the reform announcement in late 1978 did not lead to local business tax rate adjustments in 1979. Hence, there was no pre-reform adjustment: 39 of 40 cities not collecting payroll taxes and 41 of 44 cities raising payroll taxes did not adjust their multiplier between 1978 and 1979.

Figure 2.2 displays the percentage change in the income tax multiplier from 1979 to 1980, as well as its distribution in 1979. In 1979, municipalities that collected a payroll tax clearly imposed a lower income tax multiplier than the other municipalities. In the course of the reform, most nonpayroll tax collecting municipalities adhered – to some degree – to the balanced budget policy and lowered their income and capital tax multiplier. The mean non-zero change was -5.7%, while 11 cities did not decrease the tax rate. Furthermore, nine former payroll tax collecting municipalities did not change their income and capital tax multiplier while the others increase the multiplier by 17.1%, on average.

Fig. 2.2: Changes in the business income tax multiplier (1979-1980)



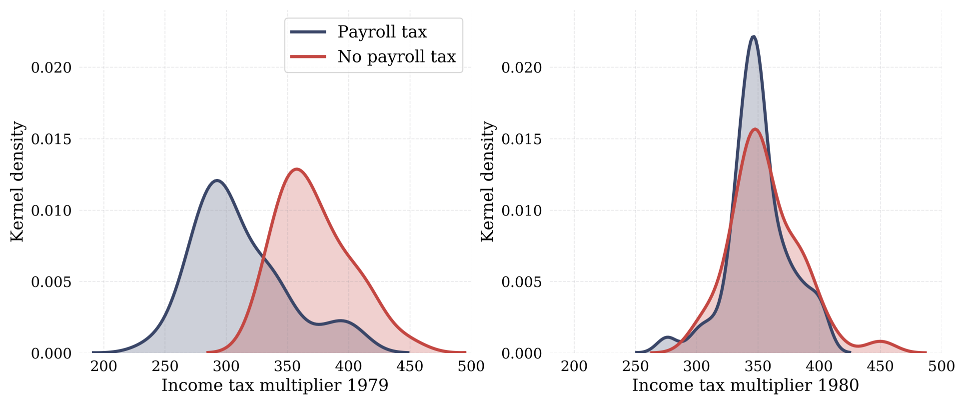
Notes: The figure displays the percentage change of the income tax multiplier, which is set at the municipal level, for the city sample.

[16]Some of the federal states provided oversight of adherence to the balanced budget. Whether the decree legally bound local governments to pass on their increased federal revenue to local firms was critically discussed (Ifst, 1979a; Ifst, 1979b).

[17]The average payroll tax rate τ_w in cities was 1.7% during the observation period. The average multiplier was 864% with a minimum of 200% and a maximum of 1160%. Average values are slightly lower for the 503 largest municipalities ($\lambda_w = 830\%$).

Figure 2.2 also illustrates that the changes in municipal revenue through the higher share of federal income taxes and lower Gewerbesteuerumlage affected all nonpayroll tax collecting cities almost equally as they all decreased the multiplier by approximately the same percentage. Moreover, the additional financing needs of former payroll tax collecting cities fall in the level of the income tax multiplier: a lower income tax multiplier in 1979 indicates a larger increase between 1979 and 1980. This is due to the inverse relationship of the payroll and income tax rates in cities that raise both taxes (Figure 2.7). This finding suggests that cities that levied the payroll tax did not use it to collect additional revenue but instead substituted business income with payroll taxation. Concurrently, the per-capita revenue raised from local business taxation was similar between business tax regimes in 1979 (Figure 2.9).

Fig. 2.3: The distribution of the business income tax multiplier (1979, 1980)



Notes: The figure displays the distribution of the local business income tax multiplier for the city sample in 1979 and 1980.

Kernel density estimates of the income and capital tax multiplier of cities prior to and after reforming the local business tax system are displayed in Figure 2.3. As the figure displays, the application of payroll corresponds to lower income and capital tax rates (left panel). In the aftermath of the reform (right panel), all municipalities seem to have the same financing need from local business taxation as both groups' median and mean multiplier are similar. The distribution of the income tax multiplier for the 503 largest municipalities provides the same results (Figure 2.8).

2.3 The model

I use a stylized partial equilibrium model to derive the implications of business income and payroll taxation for wages and to highlight the transmission mechanisms. I treat municipalities as small open economies since local governments have autonomy over tax rates. Different from cross-country studies, capital and goods are more mobile across regions and changes in good or capital prices are likely of minor importance for local wages (Fuest et al. (2018)).

Firms produce with a strictly concave production function $F(K,N)$ with decreasing returns to scale.[18] Each firm is subject to income (τ_y), capital (τ_k) and payroll (τ_w) taxation, based on the local multiplier chosen by the municipality j, in which the firm operates. While capital costs r are not deductible from profits, labor costs and capital tax payments are fully deductible. Each county is treated as a small open economy with perfectly mobile output and capital. Furthermore, I assume that the output price is not affected by municipal taxation and set the price equal to one. A firm maximizes its after-tax profit by choosing capital K and labor N:[19]

$$\max_{K,N} \Pi = (1-\tau_y)[F(K,N) - (1+\tau_w)wN - \tau_k K] - rK \tag{2.1}$$

The first-order conditions for labor (2.2) and capital (2.3) describe the factor demand functions of the firm. Labor demand is set based on the wage and a payroll tax premium

$$w(1+\tau_w) \quad = \quad F_N, \tag{2.2}$$

where $F_N = \frac{\partial F(K,N)}{\partial N}$ is the marginal product of labor. Capital demand is affected by income and capital taxation since the capital stock is taxed directly and capital costs are not deductible.[20]

$$\frac{r}{1-\tau_y} + \theta\tau_y \quad = \quad F_K \tag{2.3}$$

As discussed in previous sections, the capital tax is proportional to the income tax as the national rates differ but the local multiplier is the same ($\theta = \frac{\tau_k}{\tau_y}$). I express the capital tax as a proportion of the income tax, since both taxes work through capital demand. Hence, overall capital costs are $R = \theta\tau_y + \frac{r}{1-\tau_y}$.

Totally differentiating both demand equations allows me to study the adjustment of capital, wages, and labor in response to a tax change. I assume, that the interest rate is determined on the international market and does not respond to tax changes ($dr = 0$). The total differential for labor demand (2.2) is

$$F_{NN}dN + F_{NK}dK \quad = \quad (1+\tau_w)dw + wd\tau_w \tag{2.4}$$

[18] $\frac{\partial F}{\partial i} > 0$, $\frac{\partial^2 F}{\partial i^2} < 0$ and $\frac{\partial^2 F}{\partial ij} > 0$ for $i \neq j$ and $i, j \in \{N, K\}$

[19] I abstract from tax-exempt amounts. An additional assumption is that firms' capital stock consists only of debt. Firm (i) and municipality (j) indices are dropped for notational convenience.

[20] Dividend payments are not deductible and capital costs are not deductible up to 1984, from which point onward half of the costs $\alpha = 0.5$ can be subtracted (cf., Richter and Wiegard, 1991).

and for capital demand (2.3) is

$$F_{KK}dK + F_{KN}dN = \left[\theta + \frac{r}{(1-\tau_y)^2}\right] d\tau_y. \tag{2.5}$$

In the following paragraphs, I analyze the reaction of capital and wages to changes in either income or payroll taxation under inelastic labor supply. The assumption of immobile workers and inelastic labor supply, concerning the payroll tax, is supported by some of the recent studies (e.g., Deslauriers et al., 2018; Cruces et al., 2010). These authors find no change in employment in reaction to a payroll tax change and argue that this result is in line with models in which labor demand is more elastic than supply and payroll taxes are shifted completely onto wages. The analysis of the other polar case is provided in the appendix (Section 2.B.1).

A payroll tax is passed on to wages completely Assuming labor remains constant ($dN = 0$) and the income tax rates do not change in reaction to a payroll tax change ($d\tau_y = 0$) equation (2.5) becomes $F_{KK}dK = 0$. Hence, an increase in payroll taxation does not affect a firm's investment decision. Inserting the result into equation (2.4)

$$\frac{dw}{d\tau_w} = -\frac{w}{1+\tau_w} < 0$$

shows that the additional costs are levied onto wages.

Income taxes affect wages through the capital stock Given a change in income taxation while labor is immobile and the payroll tax does not react ($d\tau_w = 0$) equation (2.5) simplifies to

$$\frac{dK}{d\tau_y} = \frac{1}{F_{KK}}\left[\theta + \frac{r}{(1-\tau_y)^2}\right] < 0.$$

Inserting the result into (2.4) and rearranging yields

$$\frac{dw}{d\tau_y} = \frac{F_{NK}}{1+\tau_w}\frac{dK}{d\tau_y} < 0.$$

Hence, an increase in income taxation lowers the capital stock, which in turn reduces the wage.

To summarize, in this stylized model − with fixed labor supply − payroll taxation lowers the wage, while capital is not affected. Business income taxation reduces the capital stock and, since labor cannot adjust, decreases the wage. The strong assumption of labor immobility is not easy to uphold, especially at this low geographical level, although Neumann (2017) finds no evidence of aggregate labor supply changes in reaction to a change in payroll taxes (social security contributions). However, the transmission of business income taxes onto wages goes through the capital-labor ratio even if one allows for a degree labor mobility across counties or imperfect competition in product markets, which will affect the size of the reaction (Gravelle, 2013, and McKenzie and Ferede, 2017). This study does not attempt to analyze the general equilibrium effects, and the model serves as a tool to illustrate the underlying mechanisms that potentially govern the pass-through of payroll and business income taxes onto wages.

2.3.1 Quantitative implications

To gain insights into the prospective size of wage differences, I utilize the model presented above and compute wage differences based on the observed average differences in tax rates. I assume strictly concave production function with decreasing returns to scale $F(K, N) = K^{\beta} N^{\gamma}$, where $\beta + \gamma < 1$. Labor is immobile and inelastically supplied at $N = 1$.

Using the first-order conditions, the analytical solution for the capital stock is

$$K^{*} = \left[\left(\frac{\beta}{\frac{r}{1-\tau_y} + \theta\tau_y} \right) N^{\gamma} \right]^{\frac{1}{1-\beta}}$$

and for the wage

$$w = \frac{\gamma}{1+\tau_w}(1-\beta) K^{*\beta} N^{(\gamma-1)}.$$

Deriving the elasticity of wages with respect to income taxation (τ_y) yields

$$\epsilon_{w,\tau_y} = -\frac{\beta}{1-\beta}\psi_0,$$

where $\psi_0 = \frac{\tau_y}{1-\tau_y} \frac{r+(1-\tau_y)^2\theta}{r+(1-\tau_y)\theta\tau_y} > 0$. Hence, ϵ_{w,τ_y} depends on the interest rate r and increases in the capital share β. Similar to the results discussed using the differentials, income taxation lowers the wage via a lower capital stock and the size of the effect depends on the capital share and the interest rate.

The elasticity of the wage with respect to payroll taxation is

$$\epsilon_{w,1+\tau_w} = -1.$$

Payroll taxes are passed onto wages while not affecting a firm's capital choice. In general, the wage change of moving from payroll and income to just income taxation is contingent – in this simplistic model – on the real interest rate and the capital share of the production function. The wage difference between the two tax regimes negatively depends on the interest rate because the difference in capital stocks decreases as interest rates increase. The latter effect is the result from decreasing differences in capital costs $R = \frac{r}{1-\tau_y} + \theta\tau_y$, as r increases.

I test the model's predictions for relative wage effects using a simple counterfactual analysis: A city j raises its income using payroll and income taxation in period t and moves to sole business income taxation in period $t+1$. Tax rates in t are calibrated to observed average tax rates in payroll tax collecting cities and rates in $t+1$ are calibrated to observed average rates in all other cities in 1979. Assuming labor remains constant, I calculate $\Delta w = \ln(w_{j,t+1}) - \ln(w_{j,t})$. As a starting point, I assume standard values for the production function ($\beta = \frac{1}{3}$) and an interest rate of $r = 0.045$. Note that wage differences do not depend on the labor share γ. In this benchmark scenario, the model predicts a wage decrease of 1.1%. This decrease can be decomposed as a 1.7% increase, due to the abolishment of payroll taxation, and a 2.8% decrease, due to the increase in business income taxes. The adjustment of the capital stock causes the latter. This adjustment depends negatively on the capital share and the interest rate. Hence, the wage differences between regimes decrease in β and r and might become negative. A more extensive analysis, accounting for different values for r and β is provided in the appendix 2.B.2. Overall, the wage difference between the two tax regimes is expected to be positive based on the implications of the model calibrated to standard values.

Introducing firm-union bargaining into the model, as first demonstrated by Arulampalam et al. (2012), provides a different channel for taxes to influence wages. I extend their framework, including a payroll tax and illustrate in appendix 2.B.3 that income and payroll taxes lead to lower wages via the bargaining process. Still, the overall effect on wages depends on the employment adjustment. The predicted wage differences using the bargaining model and assuming constant labor are only slightly higher and decrease in the capital share and the interest rate.

2.3.2 Welfare evaluation

This section analyzes the welfare implications of business income and payroll taxes. The representative household owns the firm, receives the profits Π and maximizes the following utility function

$$U(C, N) = C - V(N) \quad s.t. \quad C = wN + \Pi, \tag{2.6}$$

where $V(N)$ represents disutility from labor.

The planner then maximizes the following welfare function

$$\begin{aligned}
\max_{\tau_w,\tau_y,w,K} W &= wN + \Pi - V(N) \\
s.t. \quad G &= \tau_y[F(K,N) - (1+\tau_w)wN - \theta\tau_y K] + \tau_w wN + \theta\tau_y K \\
w &= V'(N) \\
F_N &= w(1+\tau_w) \\
F_K &= \frac{r}{1-\tau_y} + \theta\tau_y
\end{aligned}$$

accounting for the exogenous level of local government spending G and the optimality conditions of the firm and the household.

The following welfare analysis is based on a polar case where labor is immobile (i.e., N is fixed). I abstract from capital taxation ($\tau_k = 0$) and derive closed-form welfare solutions for two local business tax regimes: for sole business income, and sole payroll taxation.

Let variables X^w denote the realized value under sole payroll and X^y under sole business income taxation. Furthermore, assume that the government revenue needs are given by $\bar{G}$, which is the same in both regimes

Proposition: Welfare is always higher when payroll (W^w) rather than business income (W^y) is taxed to raise the same amount of revenue ($\bar{G}$).

Welfare differences are given by

$$W^w - W^y = \left[\frac{\beta}{r}\right]^{\frac{\beta}{1-\beta}} \left[1 - (1-\tau_y)^{\frac{\beta}{1-\beta}}\right] - r\left[\frac{\beta}{r}\right]^{\frac{1}{1-\beta}} \left[1 - (1-\tau_y)^{\frac{1}{1-\beta}}\right] > 0, \tag{2.7}$$

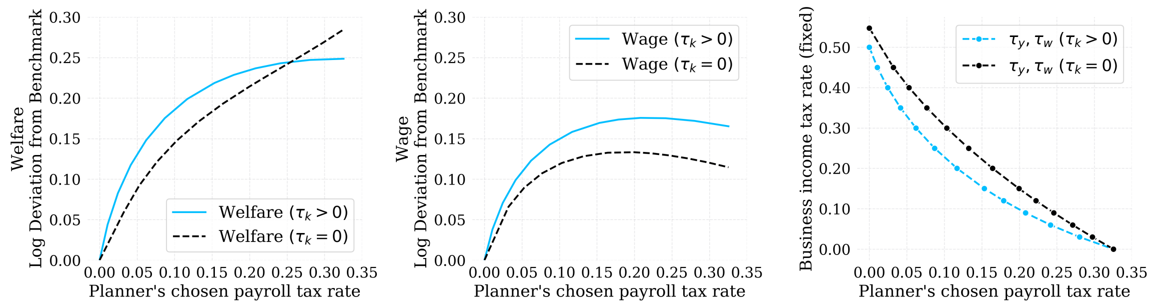
for all $\beta, \tau_y \in (0,1)$. For the proof, see Appendix 2.B.5.

The intuition behind the result is simple: business income taxation prevents the firm from efficiently choosing capital, while payroll taxes are passed on to the worker without any impediment to labor. Moreover, welfare is similar when no tax is raised.

To illustrate welfare under capital taxation, I solve the planner's problem for different tax rate combinations and compare the resulting welfare level to a benchmark. The model is calibrated using the same parameter values as in the previous section ($\beta = \frac{1}{3}$, $r = 0.045$). In the benchmark-setting, government revenue is raised using the income tax τ_y (and capital taxation, as their are mutually dependent). Specifically, the level of government spending G is determined for $\tau_w = 0$ and an arbitrary value for the business income tax set to $\tau_y = 0.5$. This spending level is kept constant in the following quantitative exercise. The welfare level achieved in this situation serves as the benchmark. Moreover, given the level of G, I additionally solve for the optimal business income tax rate and the corresponding welfare level for the model without capital taxation ($\tau_k = 0$).

Subsequently, the planner chooses optimal values for τ_w given a decreasing, exogenous sequence of τ_y to fulfill the budgetary needs, i.e., G is kept constant. This exercise is carried out for both models, with ($\tau_k > 0$) and without ($\tau_k = 0$) capital taxation. Figure 2.4 presents the resulting sequences for welfare (left panel), wages (center) and optimal payroll tax rates (right panel). Welfare rises unambiguously whenever the

Fig. 2.4: Welfare and wage differences between payroll and business income taxation



Notes: This figure displays the (log) deviation of welfare (wage) for fixed government spending and an exogenous business income tax rate from a benchmark. The benchmark welfare level is based on the optimal allocation of capital for ($\tau_w = 0, \tau_y = 0.5, \tau_k = \theta\tau_y$).

income tax is lowered, and an increase in payroll taxes compensates for the forgone government income. Allowing for capital taxation does not change the result, as the tax also solely impacts a firm's capital choice. Wages are always higher, compared to sole business income taxation. However, a compensated decrease in payroll taxes does not necessarily lead to a decrease in wages (center panel). The higher optimal payroll tax rate in the case without capital taxation (right panel) is the result of the lack of revenue from capital taxation. The optimal increase in payroll tax rates is lower than the decrease in business income tax rates: A payroll tax rate of 11% compensates for a decrease in τ_y from 50% to 20%. A similar pattern is observed for the actual tax rates: while payroll tax rates differ by 1.7%, the average difference for business income tax rates across tax regimes is 3.2%.

2.4 Empirical strategy and results

The empirical strategy, as well as the results, are presented in this section.[21] The outcome variable of interest is the log median wage of full-time employees of firms that are observed for the whole period and which employed at least three workers. For identification, I exploit the regional variation in the application of payroll taxation and the abolishment of the tax in 1980. To obtain the average wage difference between the two local business tax regimes, I compare wages of firms located in regions that raise

[21] The estimation strategy follows Fuest et al. (2018).

payroll and income taxes to wages of firms located in counties that levy only an income tax, before and after the local business tax reform. Moreover, to estimate the payroll tax elasticity, I additionally utilize variations in payroll and business income tax rates in cities.

2.4.1 Estimation strategy

In a first step, the average difference in wages between the two tax regimes is estimated for each year during $1977-1983$ using an event study model (2.8). This model is primarily used to establish common trends in outcomes between firms located in treatment and control regions after the abolishment of payroll taxation. Flat after treatment trends are essential for the causal identification of the average wage difference.

$$\ln(w_{i,j,t}) = \sum_{n=-3}^{3} \gamma_n D_{j,t}^n + \mu_i + \eta_j + \psi_{r,t} + \epsilon_{i,j,t}. \tag{2.8}$$

Dummies $D_{j,t}^n$ indicate an event in region j that occurs n periods away from time t. The coefficients γ_n capture the difference between wages paid in treatment and control counties. Treatment - payroll taxation - ended on January 1st, 1980. The baseline event window runs from three years before to three years after the event. As the firm panel is balanced in event time, the coefficients of interest capture the wage difference in each year from 1977 to 1983. I include firm (μ_i) and county (η_j) fixed effects. "State $\times$ year" or "region $\times$ year" fixed effects ($\psi_{r,t}$) are added to account for regional shocks. In the "state $\times$ year" fixed effect specification, the states Hamburg and Bremen (and Bremerhaven) are assigned to the states Schleswig-Holstein and Lower-Saxony, respectively. For the higher aggregated specification, I group the federal states into three regions: North, which contains the states Bremen, Hamburg, Lower-Saxony and Schleswig-Holstein; West, which contains the states North-Rhine Westphalia, Hesse, Rhineland Palatinate and Saarland; and South, which contains the states Bavaria and Baden-Wuerttemberg. The error term is $\epsilon_{i,j,t}$. Standard errors are clustered at the county level, which is the level of the identifying variation.

The generalized DiD model (2.9) is then used to estimate the average wage difference between the two local business tax regimes.

$$\ln\left(w_{i,j,t}\right) = \chi D_{j,t}^{\tau_w} + \mu_i + \eta_j + \psi_{r,t} + \epsilon_{i,j,t} \tag{2.9}$$

The dummy variable $D_{j,t}^{\tau_w}$ indicates active payroll taxation in year t and county j. The coefficient χ captures the average difference in wages paid by firms located in treatment, compared to firms located in control regions before and after the reform. I also include a firm μ_i and a county η_j fixed effect as well as "state $\times$ year" or "region $\times$ year" ($\psi_{r,t}$) fixed effects. Standard errors are clustered at the county level.

Estimating the payroll tax elasticity The joint identification of the elasticity of the wage with respect to payroll (α), income (δ) and capital (ζ) taxation estimating model (2.10) is not feasible because income and capital tax are mutually dependent ($\tau_y = \theta\tau_k$). All taxes lower the wage, and all rates were adjusted at the same time in close to all regions.

$$\ln(w_{i,j,t}) = \alpha \ln(1 + \tau_{w,j,t}) + \delta \ln(1 - \tau_{y,j,t}) + \zeta \ln(\tau_{k,j,t}) + \mu_i + \eta_j + \psi_{r,t} + \epsilon_{i,t} \quad (2.10)$$

To provide a lower bound estimate of the payroll tax elasticity, I adjust each median wage observation for the effect of business income taxation using the corresponding elasticity estimate of Fuest et al. (2018) and abstract from capital taxation ($\zeta = 0$). Hence, to estimate the elasticity I also use variation in the intensive margin – i.e., tax rates – of income and payroll taxes. The adjusted wages are constructed using the business income tax elasticity estimate of $\hat{\delta} = 0.388$ and deriving $\widehat{w}_{i,j,t} = \ln(w_{i,j,t}) - \hat{\delta}\ln(1 - \tau_{y,j,t})$.

$$\widehat{w}_{i,j,t} = \alpha \ln(1 + \tau_{w,j,t}) + \mu_i + \eta_j + \psi_{r,t} + \epsilon_{i,j,t} \quad (2.11)$$

The estimate of α using the model (2.11) will be biased because, in contrast to the period analyzed by FPS, capital is taxed directly and capital costs are not partially deductible. The tax on capital will likely lead to a higher elasticity δ.[22] Hence, differences in adjusted wages between firms based in treatment and control regions using the estimate stated above will be lower as if corrected by the true estimate. As a consequence, the elasticity with respect to payroll taxation will be underestimated.

Adjusting for anticipatory reactions The event study estimates will demonstrate a clear adjustment of the outcome variable before the reform, but after its announcement. Based on the date and credibility of the reform announcement, I treat the adjustment as an anticipatory reaction of firms.The reform decree was passed in November 1978, including detailed information on how the local multiplier should be set after January 1, 1980. The first wage observation after the reform announcement is from June 1979. Hence, firms had excellent information about the new local multiplier and they had approximately one year to adjust wages. Following Malani and Reif (2015), I account for the anticipatory effects of the reform and include leading indicator variables for whether the reform is adopted in models (2.9) and (2.11). Specifically, the payroll

[22] In a model with partial deductibility and no capital tax, as in Fuest et al. (2018), the derivative of capital costs w.r.t the income tax is $\frac{\partial \bar{R}}{\partial \tau_y} = (1-\alpha)\frac{r}{(1-\tau_y)^2} > 0$. The respective result for the model utilized here is $\frac{\partial R}{\partial \tau_y} = \theta + \frac{r}{(1-\tau_y)^2} > 0$. Hence, the response to an increase in income taxation is larger, because interest payments are not deductible ($\alpha = 0$) and an additional tax, proportional to the income tax, is levied directly on capital ($\theta = \frac{\tau_k}{\tau_y}$).

tax elasticity is estimated using model (2.12)

$$\ln \widehat{w}_{i,j,t} = \alpha \ln(1 + \tau_{w,j,t}) + \sum_{n=1}^{S} \gamma_a D_{j,t+n} + \mu_i + \eta_j + \psi_{r,t} + \epsilon_{i,t}, \tag{2.12}$$

where $D_{j,t+n} = 1$ if the reform occurs in n periods from today and γ_a captures the anticipatory effects. The method assumes perfect foresight of agents and presumes that expectations longer than the horizon S do not influence outcomes. However, this strategy leads to inconsistent estimates if any of these assumptions do not hold. Arguably, perfect foresight is given in 1979 since the reform was announced one year prior and the decree contained clear instructions on how the new local tax multiplier should be set. Based on the announcement a little more than a year before the reform, a leading indicator of one period should also suffice. μ_i is a firm and η_j a county fixed effects. $\psi_{r,t}$ represents the "region×year" fixed effect, as before. The error term is $\epsilon_{i,j,t}$. Standard errors are clustered at the municipality level, which is the level of the identifying variation.

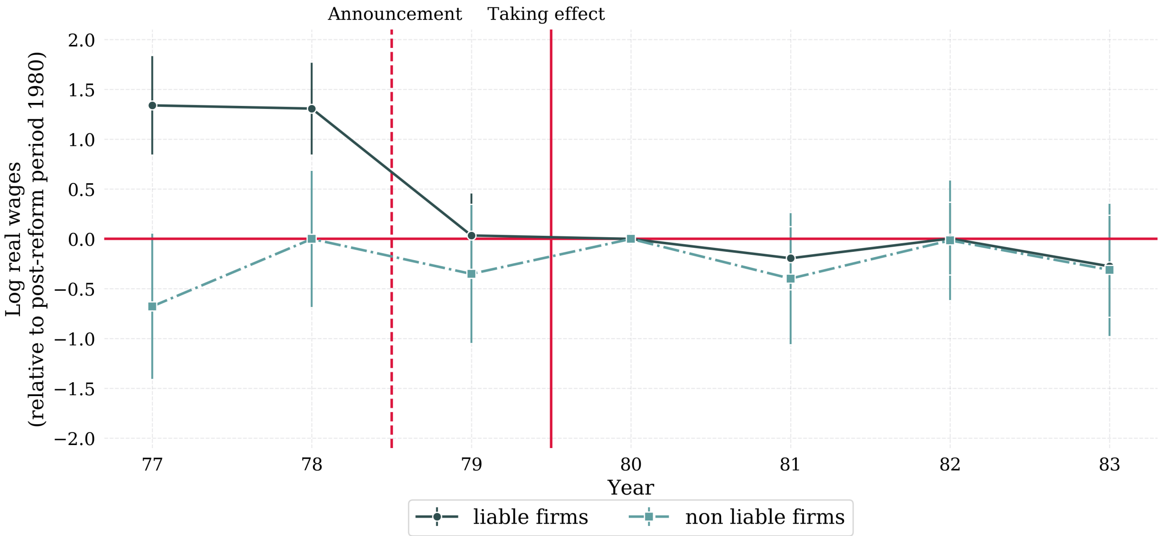
2.4.2 The average wage difference between tax regimes

The event study estimates (2.8) are presented in Figure 2.5. The estimated coefficients are displayed for the group of firms liable and not liable to local business taxation. These coefficients capture the differences in wages between firms in treatment and control regions in each year relative to the first after-treatment period (1980). The treatment period runs from 1977 to 1979. During this time, revenue from local business taxation in treatment regions was raised using payroll and business income taxation, while control regions levied only a business income tax. Hence, the presented coefficients reflect the wage difference that is due to the tax on payroll rather than on business income. These wage differences are striking: Before the reform liable firms located in payroll tax collecting municipalities pay wages 1.4% higher than firms only subject to business income taxation.

However, in the direct succession of the reform announcement in 1978, the wage difference dissolved, which suggests an anticipatory adjustment of wages. This adjustment is not due to a change in tax rates as only a fraction of municipalities changed their rates between 1978 and 1979. Furthermore, wages of firms that are not subject to local business taxation (non-liable firms) do not exhibit the same relative decrease. One would expect the wages paid from these firms show the corresponding adjustment if the pre-reform reaction observed for liable firms was due to, for example, local shocks not captured by the fixed effects. Given that the reform was announced in November 1978 and the reform decree contained detailed information on how the local multiplier should be set in 1980, I treat this adjustment as an anticipatory reaction of firms.

After 1980, firms in all regions were subjected to the same local business tax regime. Reassuringly, there is no significant wage difference in treatment and control regions after the reform was enacted and regions were subjected. Moreover, the estimate for non-liable firms (Column 6) shows no significant differences in wages at any point in time, supporting the identification of the average wage effect.

Fig. 2.5: Event study estimates



Notes: This figure presents coefficients γ_n, estimated using model (2.8). Regressions include county, firm, and "state × year" fixed effects. Not liable firms are neither subject to the local income nor the payroll tax. Estimation is on the sample of firms based in 229 counties: All cities and all counties where payroll tax was not collected. The time of observation is 1977 to 1983. Standard errors are clustered at the county level.

The results using the city sample are comparable (Figure 2.10), although the drop in wages in 1979 is not as strong. Furthermore, the results are robust to using "region × year" rather than "state × year" fixed effects (Figure 2.11). As an additional sensitivity check, the observation period is extended to cover the years 1975 to 1983. I obtain similar results using the extended sample period (Figure 2.12).

Turning to the average effect, I estimate the wage difference between tax regimes using the generalized DiD model (2.9). The estimated coefficient represents the average difference in median wages of firms subjected to payroll and income taxation relative to wages of businesses subjected only to an income tax. A positive coefficient indicates that wages were higher before the reform. Table 2.1 presents the results for the baseline sample. The point estimates show, depending on the specification used, a significant difference in wages in treatment and control regions of at least 1%. In other words, wages were higher when part of the municipal revenue was raised using payroll rather than business income taxes. My preferred result is presented in Column (2): The reform lowered wages in treatment regions by 1.43%, when state-specific annual shocks, as well as the anticipation of the reform, are accounted for. Not accounting for firms' anticipatory reaction leads to lower estimates (Columns 1, 3, and 4) as the pre-reform adjustment is attributed to the average wage difference before the reform.

Table 2.1: Difference-in-differences: Average wage difference between tax regimes

	Average wage difference					
	Liable firms					Non liable firms
	(1)	(2)	(3)	(4)	(5)	(6)
$\hat{\chi}$	1.01	1.44	1.18	1.48	1.62	−0.16
	(0.15)	(0.19)	(0.17)	(0.17)	(0.20)	(0.29)
State × year FE	√	√				
Region × year FE			√		√	√
Year FE				√		
Expectation		√			√	
Observations			104,934			29,832
Number of regions			229			229

Notes: The tables displays the estimates of the coefficient χ using model (2.9). Regressions include county and firm fixed effects. Non-liable firms are neither subject to the local income nor the payroll tax. Estimation is on the sample of firms based in 229 counties: All cities and all counties where payroll tax was not collected. The time of observation is 1977 to 1983. Standard errors are clustered at the county level.

The last column displays the average wage differences between non-liable firms in treatment regions and control regions. I find a slightly negative coefficient that is not significantly different from zero. Hence, when firms are not subjected to the local business tax, wages do not differ. Moreover, restricting the control group to cities does not alter the results (Table 2.5), while the number of firms decreases to just over 50,000. The point estimates fall from 1.44 to 0.92 and 1.62 to 1.49 when controlling for state or region × year fixed effects and reform anticipation, respectively.

The estimated average wage difference between tax regimes is in line with the prediction of the model as wages were higher when governments raise part of their revenue using payroll taxes instead of business income taxes. However, the estimates are higher, on average, than the wage difference of 1% predicted by the model, suggesting that other margins should be considered to explain the wage differences between the tax regimes. One possibility is a higher firm entry rate in municipalities with low local business income taxes, which leads to higher labor demand.[23]

The estimates presented above rely on wage differences between firms based in cities and subject to a payroll tax and firms based in cities or counties that did not levy the tax. To ensure the result is generalizable to firms subject to a payroll tax and not

[23] A more formal argument is provided in appendix Section 2.B.4.

based in a city, I extend the sample. Specifically, I include a set of counties where some municipalities levy the payroll tax. The selected counties are sorted into the control and treatment groups based on the relative intensity of payroll taxation. The apportionment is discussed in section 2.6. In total, the number of counties and firms increases to 292 and 140,000, respectively. The estimates are similar to the results of the city sample (Table 2.7), although standard errors reduce by more than half. Note that the results are (downward) biased as some firms are attributed to be subject to the incorrect local business tax regime.

2.4.3 The average wage difference between tax regimes by firm type

To foster comparison to FPS and to ensure that differences in industry composition or firm size do not drive the results, I estimate model (2.9) stratified by firm size or industry sector. Table 2.2 presents the DiD average reform results for the baseline sample, which covers 229 districts. Concurrent to the finding of FPS, small and medium establishments react strongest to the change of the tax system. The insignificant estimates for medium and large size firms might be the result of a higher wage-setting power of larger firms. FPS provide evidence of such a mechanism. Moreover, they show that large firms tend to operate in more than one county and are more often foreign-owned, leading to more opportunities to shift profits to another establishment located in another county or state.

Table 2.2: Difference-in-differences: Average wage difference between tax regimes by firm type

	Average wage difference			
Firm size	Below 10	10 to 99	100 to 499	Above 500
	1.82	1.25	0.44	0.49
	(0.24)	(0.18)	(0.25)	(0.35)
Sector	Manufacturing	Construction	Trade	Services
	1.03	2.71	0.92	0.75
	(0.21)	(0.28)	(0.24)	(0.28)

Notes: The table presents the estimates of the coefficient χ using model (2.9) stratified by either firm size or sector. Regressions include county, firm, "state $\times$ year" and "firm type $\times$ year" fixed effects. Estimation is on the sample of firms based in 229 counties: All cities and all counties where payroll tax was not collected. The time of observation is 1977 to 1983.

Tuning to industry sectors, I find significant point estimates for all four industries.

Manufacturing and construction exhibiting the most substantial effects: Wages in manufacturing and construction decreased by 1% and 2.71%, respectively. The significant wage effects for trade and service sectors are lower compared to the other two sectors. A possible explanation is that firms in these less tradable sectors might be able to shift a share of the tax burden to consumers (cf., Fuest et al. (2018))

2.4.4 Payroll tax elasticity estimates

This section presents lower bound estimates of the uncompensated payroll tax elasticity. Wages are adjusted using the estimate of FPS and abstracting from capital taxation, as discussed. The wage adjustment ensures that the estimate captures the uncompensated payroll tax elasticity. This estimate is of particular interest because an elasticity estimate for Germany where the tax does not contribute to any social security systems is still missing from the literature. The underlying sample is the city sample.[24]

The elasticity estimates are presented in Table 2.3. My preferred result is stated in Column (1). Controlling for annual state-specific shocks and for the anticipation of the reform, I estimate an elasticity of -0.72%. The sign of the elasticity is in line with the theory: An increase in payroll taxes lowers the wage. Moreover, the pass-through of the payroll tax onto wages holds when higher level region-year fixed effects are used (Columns 3, 4 and 5), although the elasticity estimate decreases up to -0.44%. Not controlling for anticipatory effects leads to a higher estimate of -0.73% (Column 2).

The presented estimates represent a lower bound because I do not control for the effect of capital taxation. As the capital tax is a linear transformation of the business income tax, which is higher in municipalities that do not raise the payroll tax, an accurate adjustment for capital taxation would increase the wage difference between treatment and control regions before the reform. This increase would lead to a more substantial decrease afterward. Hence, the correct payroll tax elasticity is larger. Moreover, the estimate used to adjust median wages for the impact of the business income tax is valid for a different local business tax regime, where capital costs are partially deductible. The partial deductibility attenuates the effect of business income taxes on wages.

The elasticity of wages concerning a change in payroll taxation is in line with the results of Deslauriers et al. (2018). They find that the payroll tax that has no direct tax-benefit link in Canada is almost entirely shifted onto wages. Moreover, the lower bound elasticity estimate of -0.72 indicates that firms are able to shift the tax burden to workers when payroll taxes do not contribute to a social security system, i.e., when

[24] Expanding the sample with counties in which no municipality levied a payroll tax would bias the results at two margins: The use of weighted tax rates would lead to biased adjusted wages and to biased elasticity estimates.

Table 2.3: Difference-in-differences: Payroll tax elasticity

	Payroll tax elasticity				
	Dependent variable: $\widehat{w}$				
	(1)	(2)	(3)	(4)	(5)
$\ln(1+\tau_{ls})$	-0.72	-0.73	-0.61	-0.44	-0.56
	(0.27)	(0.25)	(0.20)	(0.17)	(0.22)
State $\times$ year FE	$\checkmark$	$\checkmark$			
Region $\times$ year FE			$\checkmark$		$\checkmark$
year FE				$\checkmark$	
Expectation	$\checkmark$				$\checkmark$
Observations			50,008		
Number of regions			84		

Notes: The table presents estimates of the coefficient α using model (2.11) or (2.12). Regressions include county and firm fixed effects. The underlying sample is the city sample. The time of observation is 1977 to 1983. Standard errors are clustered at the city (i.e., county) level.

the tax burden is not associated with a direct benefit for workers. This result gives support to the firm-level tax shifting argument provided by Saez et al. (2019): The burden of a tax with a weak tax-benefit link cannot be passed on to affected worker groups but is rather shifted to all employees.

To assess the incidence of the payroll tax, I calculate the effect of a DM 1 increase of payroll tax revenue on wages, following Liu and Altshuler (2013). In 1979, the aggregate revenue of payroll taxation was DM 3.3 billion, and the average payroll tax rate was 1.7%. The tax base, which is the sum of labor income, can be recovered directly and amounts to 194.1 billion. Note that the incidence only differs from the elasticity as the latter is a response to a change in labor costs $(1+\tau_w)$. An increase of the tax rate by one percentage point would increase aggregate revenue by 1.94 billion, assuming the tax base does not change. The rise of τ_w implies an increase in labor costs of 0.98% and a decrease in wages of 0.71% or 1.38 billion. Hence, workers' share of the tax burden is 71%.

2.5 Conclusion

This article provides the first evidence of the compensated wage effect of a local payroll compared to a business income tax. I utilize the repeal of the payroll tax, as well as the regional variation in the application before the abolishment, and provide empirical evidence that wages are higher when local business tax revenue is raised using payroll rather than business income taxation. While prior research has focused on the isolated effect of each tax instrument, this study shows that payroll taxation has a lesser impact on wages compared to business income taxation when the amount of revenue raised is similar. More accurately, I find that a payroll tax reduction that is compensated by an increase in business income taxes led to wage decreases.

The observed wage difference can be rationalized in a stylized competitive model with fixed labor. Payroll taxes are completely shifted onto wages, and business income taxes affect wages indirectly through the capital choice. In a revenue-neutral scenario, a tax on payroll will lead to higher wages compared to a tax on business income due to an inefficient capital choice of the firm. A welfare analysis unambiguously favors payroll taxation.

An important limitation of this study is the abstraction from any labor adjustment or labor mobility. A change in employment could potentially overturn the welfare results and immobility is a strong assumption, especially at this local level. In addition, the empirical results depend on the specific design of the local business tax during 1977 to 1983 which raises the issue of external validity. However, while the theoretical results derived in this article are contingent on the model, there are additional margins that advocate for a payroll tax. The tax has, concerning government income, certain advantages: The inclusion of the payroll broadens the tax base and leads to a more reliable stream of revenue, compared to business income taxation. Moreover, the addition of immobile factors (labor income) into the tax base ensures, that amenities, provided by the local governments, are paid for by the firms that use them (Andreae (1958), Zimmermann (2002)). For example, Fossen and Bach (2008) show that a revenue-neutral reform, where wage expenses are included in the tax base, can lead to a substantial reduction in tax rates and a more equal dispersion in revenue from local business taxation across German municipalities. This study provides an additional argument in favor of a tax on payroll: higher wages.

Moreover, I isolate the effect of the payroll tax on wages and provide a lower bound estimate of the corresponding elasticity. The estimate implies that firms shift payroll taxes almost entirely onto workers' wages. This finding suggests, complementary to Saez et al. (2019) and Deslauriers et al. (2018), that firm pass on the payroll burden, despite the absence of direct worker benefits through social security contributions.

Appendix

2.A Empirical results

2.A.1 Sample selection

This section gives an overview of the sample of counties during the observational period. Payroll tax rates were set on the municipal level, but the establishment data is on the county level and represents the 2014 allocation of territory. The main result, the average wage difference between the two tax regimes, is unbiased if firms are assigned to the correct regime. Therefore, counties are included in the sample if either all or no municipalities have levied a payroll tax. Changes in county territory make the assignment more difficult. In states where no municipality levied a tax on payroll, territorial changes do not affect the results (Bavaria, Baden-Württemberg, Saarland), and all counties are included in the baseline sample. For all other states, a merger of municipalities with different tax regimes might lead to a false assignment. Table 2.4 displays the number of counties, which are included in the baseline sample, by state.

Table 2.4: Regional application of payroll taxation

State	Number of counties	
	No Payroll Tax	Total
Schleswig-Holstein	3	11
Lower-Saxony	21	38
North Rhine-Westphalia	5	31
Rhineland-Palatinate	4	24
Hesse	0	21

Changes in territory The major territorial reform of the German counties in the 1970s was completed in most states before 1977. The territories of the federal states Hamburg and Bremen were not subject to adjustments. The reform commenced in 1974 in Saarland, Schleswig-Holstein, and Rhineland-Palatinate and there were no territorial reforms of cities in the time since. However, in 2012 and 2014, two municipalities had minor territorial adjustments Rhineland-Palatinate. The reform was completed in Hesse in 1977 with one major territorial change afterward: In 1979, Lahn was split up and merged into Lahn-Dill-Kreis and Gießen. Lahn-Dill-Kreis and Gießen are dropped from the sample because Gießen did not levy a tax on firms' payroll.

After 1976 there was only one major adjustment of territorial borders of counties in North Rhine-Westphalia: The city Aachen was merged with the county Aachen into'Städteregion Aachen in 2009. I do not include Aachen into the sample because the city Aachen did not raise payroll taxes while the county Aachen did.

In Lower-Saxony, the majority of adjustments concluded in 1977, and there was one territorial reform involving a city afterward: Hannover was merged with the county Hannover into Region Hannover in 2001. The region Hannover is not included in the sample because not all municipalities belonging to the county Hannover did levy a payroll tax.

In Baden-Württemberg, the majority of territorial reforms were completed in 1975. Freiburg received a municipality in 1978 (57 inhabitants today), and there was an additional minor territorial change in 2011. In Bavaria, there were only two small territorial changes between 1990 and today. The majority of regional modifications completed in 1978.

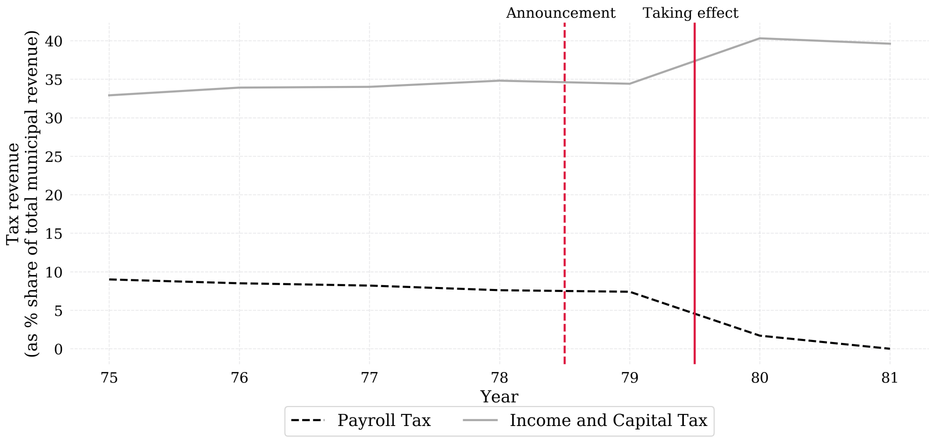
The selection criteria for the city sample, as well as the baseline sample, differ. I disregard minor territorial changes after the reform. The city sample is restricted to cities where the unweighted multiplier can be identified. In total, three West-German cities have undergone substantial changes in territory. Furthermore, cities that abolished payroll taxation between 1977 and 1979 are not considered in the analysis. In total, 8 of 92 cities are left out of the sample. Aachen, Lahn, and Hannover because of regional mergers. Berlin, because of its unique status. Mannheim, Trier, Frankenthal, and Speyer because they raised payroll taxes in 1977.

For the baseline sample, the application of payroll taxation has to be identified. The restrictions of the city sample apply. Moreover, mergers in Bavaria, Baden-Württemberg, and Saarland do not lead to any exclusions of counties as the application of the tax can be credibly assigned. Counties, where any municipality raised a payroll tax from 1977 to 1979, are disregarded. The minor territorial reforms after 1980 in Rhineland-Palatinate are neglected.

2.A.2 Local business taxes

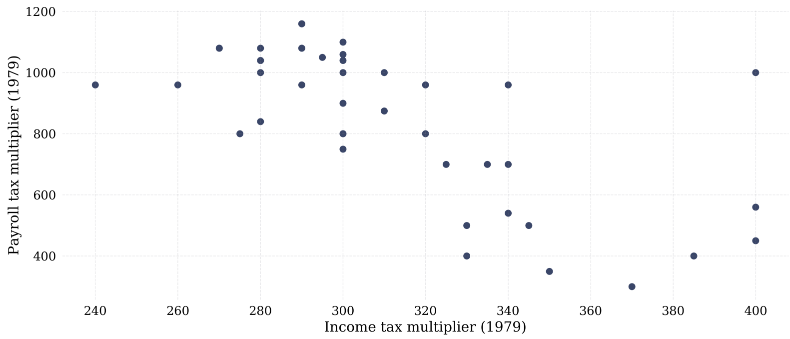
Figure 2.6 displays the share of income and capital as well as payroll tax revenue on total municipal for the years 1975 to 1981. Both percentages are almost constant at 34 and 8% up to 1980, respectively. In the aftermath of the reform, the revenue share of income and capital tax increased by 6 percentage points to above 40%. The revenue from payroll taxation in 1980 is still positive due to deferred payments from 1979. Firms were not subject to payroll taxation from January 1, 1980, onward.

Fig. 2.6: Revenues from local business taxes over time



Notes: The measure includes the share of local business taxes payable to the federal state. This share decreases by $\frac{1}{3}$ on 1.1.1980. *Source*: Köster, Thomas, (1984 [2011]) Die Entwicklung kommunaler Finanzsysteme am Beispiel Großbritanniens, Frankreichs und Deutschlands 1790 bis 1980. GESIS Köln, Deutschland ZA8458 Datenfile Version 1.0.0..

Fig. 2.7: Payroll and income tax multipliers in cities

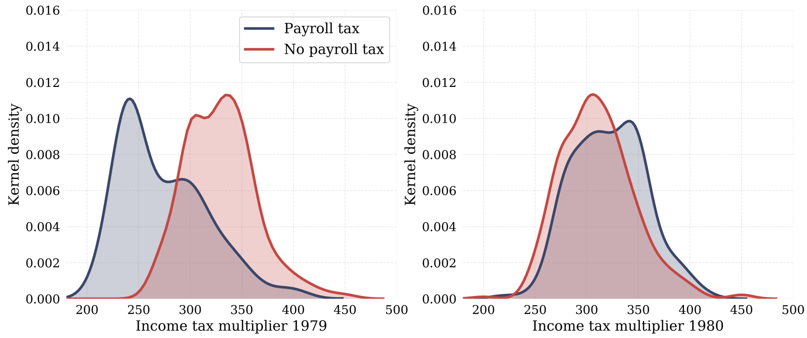


Notes: This figure displays payroll and income tax multiplier in 1979 for payroll tax collecting cities.

Figure 2.8 displays the distribution of the income tax multiplier for the 503 larges municipalities.

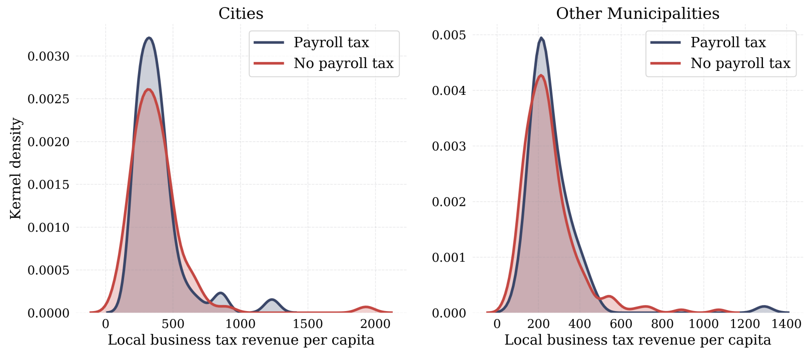
Figure 2.9 displays the distribution of the total per capita revenue from local business taxation for the 503 largest municipalities. The application of payroll taxation groups local governments. The left panel shows revenues raised in cities while the right panel shows revenues of municipalities belonging to a county. The mean difference of revenues from cities (municipalities) that levy a payroll tax compared to cities (municipalities), which only collect an income tax, is 0.01% (1.4%).

Fig. 2.8: Distribution of the income tax multiplier (503 municipalities)



Notes: This figure displays the income tax multiplier for the 503 largest municipalities in 1979.

Fig. 2.9: Per-capita revenue from local business taxation



Notes: This figure displays the per-capita revenue raised from local business taxation in 92 Cities and the largest 411 other municipalities in 1979.

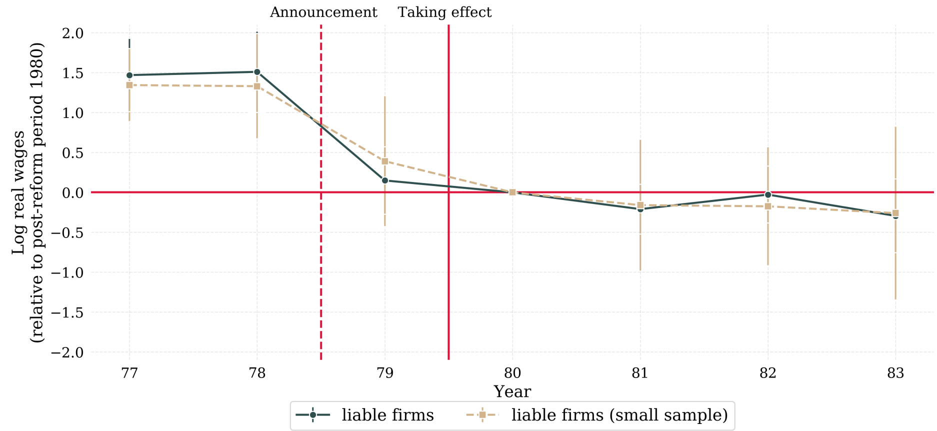
2.A.3 Estimation results: Sensitivity

Figure 2.10 displays the event study estimates for the city sample and the estimates for the extended sample, which were presented in the main text. The estimated coefficients are similar across samples.

Figure 2.11 displays event study results for the baseline sample under "region $\times$ year" fixed effects. The estimated coefficients are similar across both samples.

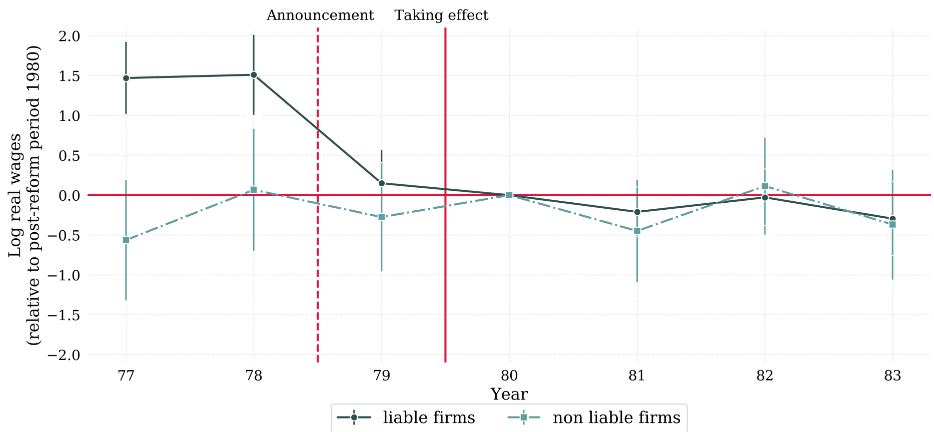
Figure 2.12 displays the event study estimates for the time period 1975 to 1983. I do not control for any territorial changes before 1977. Falsely attributing firms based in payroll tax collecting municipalities to the control group should decrease wage differences as average wages in the control group increase; Falsely attributing firms based in non-collecting municipalities to the treatment groups should, on average, lower observed differences in wages. Importantly, the wage difference between collecting and non-collecting municipalities before the reform is significantly different from zero for the years 1975 and 1976.

Fig. 2.10: Event study estimates: A comparison between samples



Notes: This figure displays the event study estimates for the baseline and the city sample. Regressions include county, firm, and "state × year" fixed effects. Non-liable firms are neither subject to the local profit nor the payroll tax. The time of observation is 1977 to 1983.

Fig. 2.11: Event study estimates: Higher level fixed effects

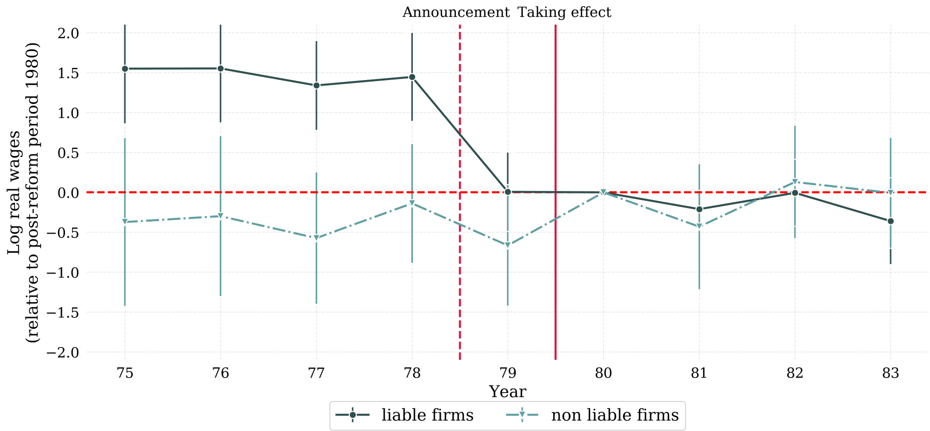


Notes: This figure displays the event study estimates for the baseline sample. Regressions include county, firm, and "region × year" fixed effects. Non-liable firms are neither subject to the local income nor the payroll tax. The time of observation is 1977 to 1983.

The average wage effect: City Sample Table 2.5 displays DiD estimates for the city sample only including self-governing municipalities.

Average wage effect: Robustness To ensure that the effect is not entirely driven by the large cities that levy a tax on payroll, I extend the sample. I include most counties in which some municipalities applied payroll taxation. Specifically, I account for the different intensity of payroll taxation across these counties and sort them into the control or treatment group based on the implied average wages by payroll tax revenue in 1979. Average wages are approximated using data on tax revenue, tax rates, and employment. The revenue from payroll taxation is given by $r_{ls} = \tau_{ls}(\bar{w}\ E)$, where E is the total number of employees and $\bar{w}$ the average wage. The aggregate county

Fig. 2.12: Event study estimates: Extended sample period (1975-1983)



Notes: This figure displays the event study estimates for the baseline sample. Regressions include county, firm, and "region × year" fixed effects. Non-liable firms are neither subject to the local income nor the payroll tax. The time of observation is 1975 to 1983.

employment data is obtained from the federal employment agency. I use monthly average employment in 1984 to approximate the number of employed in 1979.[25]

$$\bar{w} = \frac{r_{ls}}{\tau_{ls} E}$$

I sort counties with an average annual wage below (above) DM 2000 into the control (treatment) group. Table 2.6 shows the additional number of counties added to the treatment (payroll tax) and the control (no payroll tax) group. The other counties are not considered in the robustness exercise due to territorial mergers or the abolishment of payroll taxation before 1980. I expect the overall reform effect to be lower because municipalities that do not raise a payroll tax are treated as payroll tax collecting and vice versa.

Table 2.7 displays DiD estimates for the extended baseline sample, which includes all counties that can be assigned to a tax regime based on the criteria discussed above.

[25]The employment variable is recovered using the unemployment rate as well as the number of unemployed. "Arbeitslose und Arbeitslosenquote (abh. EP) sowie ausgewählte Strukturen (Frauen, Ausländer, unter 25 Jahre, Langzeitarbeitslose, Schwerbehinderte) nach Kreisen ab Dezember 1984", Bundesagentur für Arbeit.

Table 2.5: Difference-in-differences: Average wage difference between tax regimes (City sample)

	Average wage effect					
	Liable firms					Non liable firms
	(1)	(2)	(3)	(4)	(5)	(6)
Reform Effect	0.71	0.92	1.18	1.04	1.49	0.29
	(0.32)	(0.39)	(0.28)	(0.25)	(0.35)	(0.51)
State × year FE	√	√				√
Region × year FE			√		√	
Year FE				√		
Expectation		√			√	
Observations			50,008			14,642
Number of regions			84			84

Notes: The table presents estimates of the coefficient χ using model (2.9) for the city sample. All estimations include county and firm fixed effects. Non-liable firms are neither subject to the local profit nor the payroll tax. The time of observation is 1977 to 1983.

Table 2.7: Difference-in-differences: Average wage difference between tax regimes (Extended baseline sample)

	Average wage effect					
	Liable firms					Non-liable firms
	(1)	(2)	(3)	(4)	(5)	(6)
Reform Effect	0.68	0.93	0.78	1.06	1.05	0.24
	(0.14)	(0.17)	(0.18)	(0.15)	(0.21)	(0.23)
State × year FE	√	√				
Region × year FE			√		√	√
year FE				√		
Expectation		√			√	
Observations			139,944			40,027
Number of regions			292			292

Notes: The table displays estimates of the coefficient χ using model (2.9). All estimations include county and firm fixed effects. Non-liable firms are neither subject to the local profit nor the payroll tax. The model is estimated on an extended sample of firms which are in all counties not subject to territorial changes or prior abolishment of payroll taxation. The time of observation is 1977 to 1983.

Table 2.6: Regional application of payroll taxation: The extended sample

State	Number of counties		
	Payroll Tax	No Payroll Tax	Total
Schleswig-Holstein	2	8 (+5)	11
Lower-Saxony	11	22 (+1)	38
North Rhine-Westphalia	24	6 (+1)	31
Rhineland-Palatinate	12	12 (+8)	24
Hesse	15	2 (+2)	21

Average wage effect: Effect heterogeneity Table 2.8 displays DiD estimates by firm type for the baseline sample with "region×year" fixed effects.

Table 2.8: Difference-in-differences: Average wage difference between tax regimes by firm type (Higher level fixed effects)

	Average wage effect			
Firm Size	Below 10	10 to 99	100 to 499	Above 500
	2.01	1.44	0.54	0.64
	(0.23)	(0.20)	(0.24)	(0.37)
Sector	Manuf.	Const.	Trade	Serv.
	1.15	2.86	1.11	0.95
	(0.24)	(0.30)	(0.24)	(0.26)

Notes: The table displays estimates of the coefficient χ using model (2.9) stratified by either firms size or sector. All estimations include county, firm and "region × year" fixed effects. Estimation is on the sample of firms based in cities as well as counties which do not levy a payroll tax. The time of observation is 1977 to 1983.

2.B Theory appendix

2.B.1 Mobile labor

For completeness, I analyze the other polar case under the assumption of perfect labor mobility. In this setting labor demand will adjust in reaction to a tax change. The wage is not affected ($dw = 0$), as it is determined on the national market.

Mobile labor: Payroll taxation. Again, I start with a change in payroll taxes and assume that income tax rates are not adjusted ($d\tau_y = 0$). Equation (2.4) becomes $dN = \frac{-F_{NK}dK + wd\tau_w}{F_{NN}}$. Inserting into (2.5) leads to

$$\frac{dK}{d\tau_w} = \frac{-wF_{KN}}{F_{KK}F_{NN} - F_{NK}^2} < 0$$

and reinserting yields

$$\frac{dN}{d\tau_w} = \underbrace{-\frac{F_{NK}}{F_{NN}}\frac{dK}{d\tau_w}}_{<0} + \underbrace{\frac{w}{F_{NN}}}_{<0} < 0.$$

With perfectly mobile worker the capital stock and labor decrease in reaction to an increase in payroll taxation.

Mobile labor: Income taxation Given a change in income taxation while labor is immobile and the payroll tax does not react equation (2.4) becomes $dN = -\frac{F_{NK}}{F_{NN}}dK$. Inserting into the total differential for capital and rearranging yields

$$\frac{dK}{d\tau_y} = \frac{F_{NN}}{F_{KK}F_{NN} - F_{KN}^2}\left[\theta + \frac{r}{(1-\tau_y)^2}\right] < 0.$$

An increase in income tax rates leads to a lower capital stock. Hence, labor also declines

$$\frac{dN}{d\tau_y} = -\frac{F_{NK}}{F_{NN}}\frac{dK}{d\tau_y} < 0.$$

Result In a simple model with labor mobility an increase in payroll as well as income taxation decrease capital and labor demand.

2.B.2 Quantitative implications: Sensitivity

I assume standard values for the capital share of the production function: $\beta = \frac{1}{3}$. Note that wage differences do not depend on the labor share γ. Log wage differences (Δw) are computed for two interest rates, 4.5 and 7.5%, because interest rates in Germany in the late 1970s were quite high and subject to large fluctuations due to various crises. Table 2.9 presents the results. I differentiate between the effect of the decrease of the payroll tax to zero (Column 1 and 4) and the increase in income tax rates (Column 2 and 5). The overall reform effect is stated in Columns 3 and 6. The model predicts an average wage difference between tax regimes of 1.1 and 0.8% for the low and high-interest rates, respectively. Decreasing the capital share by 15% decreases wage differences by 55 and 75% to 0.5 and 0.2%, respectively. Average wage effects are driven by the change of capital stock as a reaction to business income tax rate changes. The capital stock adjustment depends negatively on the capital share and the interest rate. Hence, the wage differences between regimes decrease in β and r and might well become negative.

Table 2.9: Quantitative implications: Sensitivity

	(1)	(2)	(3)	(4)	(5)	(6)
	$r = 4.5\%$			$r = 7.5\%$		
	$\Delta w(\tau_w)$	$\Delta w(\tau_y)$	Δw	$\Delta w(\tau_w)$	$\Delta w(\tau_y)$	Δw
$\beta = 0.33$	1.70	−2.76	−1.06	1.70	−2.45	−0.75
$\beta = 0.28$	1.70	−2.18	−0.48	1.70	−1.94	−0.24

Notes: The calculations are based on observed average tax rates of cities in 1979. $\Delta w(\tau_w)$ indicates the resulting wage change from the observed average change in payroll tax rates. $\Delta w(\tau_y)$ indicates the resulting wage change from the observed average change in business income tax rates. Δw indicates the overall wage effect of the tax regime change.

2.B.3 Wage bargaining framework

Following FPS, I use the efficient bargaining model developed in McDonald and Solow (1981) and extend the model by including payroll and capital taxation. Firms bargain with a union over employment N and wages w. Unions and firms operate only in one county. Indices are dropped for notational convenience.

Wage and employment levels are determined as outcomes of the bargaining problem

$$w*, N^* \;=\; \arg\max\;(w-\bar{w})N^{1-\mu}\;\mathit{\Pi}^{\mu} \tag{2.13}$$

where $1-\mu$ is the bargaining power of the union and $\bar{w}$ the outside option of worker.[26] If the bargaining fails firms will earn zero profits and workers will earn their reservation wage $\bar{w}$.

[26]The problem can also be interpreted as bargaining over a wage premium $s = w - \bar{w}$.

The bargaining problem can be rearranged to

$$\arg\max_{w,N} \ (1-\mu)\ln[(w-\bar{w})N] + \mu\ln[\Pi]. \tag{2.14}$$

Rearranging first order conditions yields the bargaining wage (2.15)

$$w = \mu\bar{w} + (1-\mu)\frac{F(K,N) - (\tau_k + \frac{r}{(1-\tau_y)})K}{(1+\tau_w)N} \tag{2.15}$$

and the optimal employment level (2.16)

$$F_N = (1+\tau_w)w - \frac{1-\mu}{\mu}\frac{\Pi}{N(1-\tau_y)}. \tag{2.16}$$

After the bargaining process has concluded, firms set capital to maximize profits under the optimality condition

$$F_K = \frac{R_t}{1-\tau_y} + \tau_k \tag{2.17}$$

An increase in payroll as well as income tax leads to lower wages but the overall effect of course depends on the adjustment in employment.

$$\begin{aligned}
\frac{\partial w}{\partial \tau_w} &= -(1-\mu)\frac{F(K,N) - \frac{(1-\tau_y)\theta\tau_y + r}{(1-\tau_y)}K}{(1+\tau_w)^2} \\
\frac{\partial w}{\partial \tau_y} &= -(1-\mu)(\theta - \frac{r}{(1-\tau_y)^2})\frac{K}{(1+\tau_w)N}
\end{aligned}$$

Keeping capital and labor fixed I can analyze the direct effect of each tax. The bargaining power is set to $\mu = 0.5$ and the reservation wage is set to $\bar{w} = 0.7w$. Both values do not influence the relative wage results. All other values are similar those introduced in the main text.

Table 2.10: Quantitative implications: Bargaining model

	(1)	(2)	(3)	(4)	(5)	(6)
	$r = 4.5\%$			$r = 7.5\%$		
	$\Delta w(\tau_w)$	$\Delta w(\tau_y)$	Δw	$\Delta w(\tau_w)$	$\Delta w(\tau_y)$	Δw
$\beta = 0.33$	1.72	−2.84	−1.17	1.72	−2.51	−0.84
$\beta = 0.28$	1.72	−2.24	−0.57	1.72	−1.98	−0.31

Notes: The calculations are based on observed average tax rates of cities in 1979. $\Delta w(\tau_w)$ indicates the resulting wage change from the observed average change in payroll tax rates. $\Delta w(\tau_y)$ indicates the resulting wage change from the observed average change in business income tax rates. Δw indicates the overall wage effect of the tax regime change.

Bargaining and competitive solution lead to the same result if capital would adjust before bargaining. The direct incidence of payroll relative to income taxation − i.e., in the situation where the firm cannot adjust the input factors − leads to a higher wage decrease when the model is calibrated to the same values as before.

2.B.4 Firm entry

The location decision of a firm is influenced by region-specific tax rates (Serrato and Zidar, 2016). Given the framework presented in the main text as well as immobile and inelastically supplied labor, a new firm will choose a location based on expected profit, i.e., the firm will locate in municipality j if $\Pi_j > \Pi_k$ for all k. As payroll taxes are passed on completely onto workers' wages, the location decision of the firm is driven by the differences in business income tax rates. Hence, municipalities with lower tax rates will exhibit a higher number of firm foundations. The increased labor demand through the larger number of firms, combined with inelastic labor supply, will lead to even larger wage differences between local business tax regimes.

2.B.5 Welfare

The household maximizes utility $U(C, N) = C - V(N)$ s.t. $C = wN + \Pi$. Labor N is held immobile and constant. The production function is $F(K, N) = N^\gamma K^\beta$.

$$\max_N wN + (1 - \tau_y)\Pi - V(N). \tag{2.18}$$

The planner maximizes the following welfare function

$$\max_{\tau_w,\tau_y,w,K,N} W = wN + (1-\tau_y)\left[F(K,N) - (1+\tau_w)wN - \tau_k K\right] - rK - V(N). \quad (2.19)$$

subject to:

$$\begin{aligned}
F_K &= \tau_k + \frac{r}{1-\tau_y} \\
F_N &= w(1+\tau_w) \\
G &= \tau_w wN + \tau_k K + \tau_y[F(K,N) - (1+\tau_w)wN - \tau_k K].
\end{aligned}$$

The following analysis is based on immobile and constant labor (e.g., N is fixed). Substituting in the first order condition with respect to the wage $(w = \frac{1}{1+\tau_w}F_N = \frac{\gamma}{1+\tau_w}N^{\gamma-1}K^{\beta})$, the welfare function can be rearranged to

$$\begin{aligned}
W &= wN + (1-\tau_y)\left[F - (1+\tau_w)wN - \tau_k K\right] - rK \\
&= \frac{\gamma}{1+\tau_w}N^{\gamma-1}K^{\beta}N + (1-\tau_y)\left[F - (1+\tau_w)\frac{\gamma}{1+\tau_w}N^{\gamma-1}K^{\beta}N\right] \\
&\quad -(1-\tau_y)\tau_k K - rK \\
&= \frac{\gamma}{1+\tau_w}F + (1-\tau_y)\left[F - \gamma F\right] - rK \\
&\quad -(1-\tau_y)\tau_k K - rK \\
&= \left[\frac{\gamma}{1+\tau_w} + (1-\gamma)(1-\tau_y)\right]F \\
&\quad -(1-\tau_y)\tau_k K - rK.
\end{aligned}$$

In the following, I compare welfare for households for two local business tax regimes: business income or payroll taxation. The capital tax is set to zero ($\tau_k = 0$). Let X^w denote outcomes where $\tau_y = 0$ and X^y where $\tau_w = 0$. The business income tax rate needed to achieve the same amount of government revenue compared to payroll taxation (i.e., $G^y = G^w$) is

$$G^y = \tau_y(F^y - w^y N) \overset{!}{=} G^w = \tau_w w^w N \quad (2.20)$$

$$\tau_w = \frac{\tau_y(1-\gamma)(1-\tau_y)^{\frac{\beta}{1-\beta}}}{\gamma - \tau_y(1-\gamma)(1-\tau_y)^{\frac{\beta}{1-\beta}}}. \quad (2.21)$$

Welfare under sole payroll taxation reduces to

$$W^w = \left[\frac{\gamma}{1+\tau_w} + (1-\gamma)\right]F^w - rK^w. \quad (2.22)$$

Welfare under sole business income taxation is given by

$$W^y = [\gamma + (1-\tau_y)(1-\gamma)] F^y - rK^y \tag{2.23}$$

For the evaluation of welfare differences between both tax regimes, welfare achieved under sole business income taxation (W^y) is compared to welfare under sole payroll taxation W^w. The latter can be reformulated to

$$W^w = [1 - \tau_y(1-\gamma)(1-\tau_y)^{\frac{\beta}{1-\beta}}]F^w - rK^w \tag{2.24}$$

when substitution for the revenue equivalent business income tax (equation 2.20).

The welfare difference then is

$$\begin{aligned} W^w - W^y &= \left[1 - \tau_y(1-\gamma)(1-\tau_y)^{\frac{\beta}{1-\beta}}\right] F^w - [\gamma + (1-\tau_y)(1-\gamma)] F^y && (2.25) \\ &= -r[K^w - K^y] && (2.26) \\ &= \left[\frac{\beta}{r}\right]^{\frac{\beta}{1-\beta}} \left[1 - (1-\tau_y)^{\frac{\beta}{1-\beta}}\right] && (2.27) \\ &\quad -r\left[\frac{\beta}{r}\right]^{\frac{1}{1-\beta}} \left[1 - (1-\tau_y)^{\frac{1}{1-\beta}}\right] > 0 \quad \forall\, \beta, \tau_y \in (0,1) && (2.28) \end{aligned}$$

where I utilize that $F^y = (1-\tau_y)^{\frac{\beta}{1-\beta}} F^w$. Hence, the planner always taxes payroll rather than business income. When no tax is raised ($\tau_w = \tau_y = 0$) the welfare difference is zero.

Chapter 3
Granular labor market flows and local unemployment

joint with Philip Jung and Edgar Preugschat

3.1 Introduction

Stark spatial dispersion of unemployment is a common phenomenon across countries, and Germany certainly is no exception: Unemployment rates in eastern Germany were more than twice as high compared to southern regions for the better part of the last two decades. These disparities call for region-specific policies if local markets are inefficient (Glaeser and Gottlieb, 2008 and Kline and Moretti, 2013). The design of these policies will depend crucially on the determinants of local unemployment: A measure designed to increase hirings in eastern Germany might not be well suited when unemployment is high because employment relationships are terminated more often. Yet, there is little research on the role that worker flows—job losses and hirings—play in determining unemployment rate differentials between local labor markets. In this paper, we investigate the importance of these flows for regional variation in unemployment rates and evaluate how they, in combination with vacancies and wages, map into structural variations across regions. Moreover, we study how the attenuation of inefficiencies within local labor markets affects welfare and unemployment rate disparities.

This paper makes three contributions. First, we document that regional unemployment differentials are driven, to a large degree, by differences in flows into unemployment (separation rates). The importance of separation rates for the determination of unemployment differentials, however, varies substantially, suggesting that the hiring (job-finding) and layoff margin should be taken into account for a holistic view of optimal local labor market policy. Second, we present a multi-region model featuring three labor market policy instruments and a tax on production that endogenously accounts for job loss and hiring variations. We quantitatively evaluate the optimal local policy mix and examine the corresponding welfare and employment effects as well as the impact on the unemployment rate disparities in Germany. Third, we investigate the efficiency of the actual distribution of unemployed across local labor markets. Here, we provide an upper bound estimate for policy measures designed to reduce aggregate unemployment by breaking down mobility frictions that prevent an efficient allocation of the unemployed across sectors.

Our empirical analysis is based on a new data set on monthly vacancy stocks as well as workers' employment histories. We construct the latter using social security microdata from the employment panel of integrated employment biographies. The data on vacancy stocks is obtained from the federal employment agency for the period 2000-2014. Earlier vacancy data is newly digitized from official reports of the employment agency. At the smallest level of disaggregation, the data includes observations for every employment agency district (in the following EAD or district) in Germany and covers the years 1994-2014.[1]

We document that, over time, regional unemployment rates have almost halved, which mirrors the development of aggregate unemployment as Germany developed into an "economic superstar" (Dustmann et al., 2014). Our evaluation at the local level, however, shows that this sharp reduction has not led to a more similar unemployment distribution within Germany. In fact, the dispersion of unemployment rates is larger in 2014 compared to 1994. An examination of the evolution of the underlying worker flows reveals two opposing movements: On the one hand, regional separation rate differences have almost vanished, which coincides with a sharp decrease in these rates. On the other hand, differences in job-finding rates have increased, while the average rise in job-finding rates is small. The substantial decreases in separation rates, in combination with the minor increases in job-finding rates, suggests that the development of the former causes the decline in unemployment rates, which is documented for the aggregate level by Hartung et al. (2018). Concerning the importance of worker flows, we determine that separation rates account for 70% of the regional variation in unemployment rates during 1994-2004, but their importance has decreased to 41% during 2005-2014, which mirrors the decreasing dispersion of separation rates. This development, in turn, increased the importance of policy measures that target job-finding rates to attenuate local unemployment since they account for the remaining variation. These new results highlight that job losses and hirings are essential components of regional unemployment differentials.[2]

The story we tell is a tale of two regions: East Germany and North Rhine-Westphalia (NRW). Districts located in these two regions exhibited the highest unemployment rates already in 1994 but due to different reasons. In NRW, unemployment was high because job-finding rates were low, while East German districts exhibited above-average

[1] Our primary focus is on the development across EADs in reunited Germany. However, in some sections, we provide additional results for West Germany during 1980-2014.

[2] In addition to regional labor markets, we provide new insights on occupational labor markets: In the appendix, we show that higher occupational unemployment in East compared to West Germany is driven by differences in separation rates. Moreover, within East and West Germany separations explain the bulk of the unemployment rate variation across occupations. We observe no change in the importance of separation rates over time.

separation rates. Throughout our observation period, unemployment rates in East Germany approached West German levels and this catch up was driven by the most substantial decreases in separation rates and happened despite declining job-finding rates. Districts in NRW, however, fell behind as job-finding rates did not rise, and the declines in separation rates were the lowest in the country. Therefore, during the first ten years (1994-2004), regional unemployment disparities were driven by East German districts, which exhibited high separation rates, while during the subsequent years (2005-2014), the variations were primarily caused by EADs in NRW, which boosted the lowest job-finding rates.

Building on the finding that the layoffs and the hires are important for the determination of regional unemployment, the second part develops a multi-island model that endogenously accounts for both margins and is based on Jung and Kuester (2015). Workers are immobile, risk-averse, and unable to save. Additionally, they choose search effort that is unobservable to a planner. The local government sets unemployment benefits, hiring subsidies, layoff taxes, and a production tax. There are no transfers between regions. Each local labor market is subject to matching frictions following Mortensen and Pissarides (1994). Within this environment, a planner has two incentives to intervene. First, deviations from the Hosios (1990) condition give rise to search externalities and may lead to inefficient levels of unemployment. Second, the planner cannot observe the search effort of the workers. Therefore, she is not able to fully insure unemployed workers as she balances the trade-off between moral hazard and consumption smoothing.

To evaluate the impact of region-specific policies on welfare and unemployment differentials, we first have to determine structural differences between local labor markets that explain the empirical variations. These underlying structural disparities will determine how the local policies affect each labor market. As we investigate regional labor markets within Germany, we do not face institutional differences in unemployment insurance systems or employment protection laws but have to take into account that local labor markets might differ in, e.g., matching technology, productivity, or the degree of unionization. To do so we utilize the model, in combination with the regional variation in wages, workers flows, and vacancies to recover four region-specific parameters. First, we calibrate the model to a fictional, average employment agency district during 1994-2002 to determine region-invariant preferences and technology parameters. Subsequently, we establish how empirical variations in separations, hirings, vacancies, and wages relate to structural differences between regions—i.e., we identify key model parameters for each local labor market relative to the benchmark scenario. The targeted parameters are productivity, matching efficiency, job-match uncertainty, and the firms' bargaining power. To account for the documented varying importance of the worker flow rates, we recover these structural parameters within the model for two periods: 1994-2002 and 2008-2014.

While our primary concern is the policy evaluation, the development of the structural parameters over time provides suggestive evidence concerning two explanations that were brought forward to account for the favorable development of the German labor market. We document a substantial increase in bargaining power across districts, which supports the view that the decline in union coverage that commenced in the early 1990s could be an important factor for the decrease in unemployment. Furthermore, we find a substantial decrease in job-match uncertainty, which we argue to be closely related to the unemployment insurance benefit cut in the course of the labor market reforms in the mid-2000s (the *Hartz* reforms). While the model yields ambiguous implications for the increase in bargaining power concerning its effect on unemployment and separation rates, the decrease in unemployment insurance benefits implies a decrease in these outcomes, which is in line with the empirical development.

We then evaluate how the optimal policy mix varies across geographical locations in the steady-state. The previously recovered structural variations determine the shape of the policy mix across regions and the strength of their impact on local labor market outcomes. With the optimal policy mix, the planner implements the constraint-efficient allocation, given the unobservable search effort of workers. We find almost no role for variations of the unemployment insurance benefit level. The planner's optimal hiring subsidies and layoff taxes, however, vary substantially across districts. In both periods, East Germany is subject to rather average policies, while the layoff taxes and hiring subsidies are particularly high in NRW and lowest in Bavaria. This result is due to coexisting high separation rates and high matching efficiency, where the latter implies that the average unemployment duration in East Germany is lower, which in total, leads to moderate policy responses.

In total, the optimal policy mix leads to gains in welfare and employment of around 5% and 10%, respectively. These improvements are, however, heterogeneous across districts: While labor market inefficiencies within regions are removed, the policy mix does not reduce welfare or unemployment differentials between EADs in Germany. This result is not generic but the outcome of the structural disparities. For example, we document a decrease in unemployment rate dispersion for West Germany in 2008-2014. Optimal labor market policies that increase welfare and attenuate unemployment can, thus, lead to a rise in unemployment disparities between regions.

In the final part, we extend our analysis to policies designed to promote regional mobility. With the growing importance of the job-finding rate in the determination of geographical unemployment rates, measures that bring jobseekers and vacancies closer together ought to become more critical. We draw from the newly compiled vacancy data and analyze the importance of mismatch between labor demand (posted vacancies) and supply (jobseekers) in segmented markets for aggregate unemployment. To evaluate the efficiency of the actual distribution of unemployed, we employ a measure developed by Sahin, Song, Topa and Violante (2014). This measure captures the impact of misallocation between markets on unemployment. It is based on a comparison of an optimal and the actual distribution of the unemployed. The optimal allocation is derived from a dynamic search and matching model, where a planner redistributes unemployed at no cost to maximize employment.

We find that geographical mismatch in Germany in $1994 - 2014$ at the district level is non-negligible as eliminating mismatch can increase aggregate hires by $10 - 15\%$. The increase in hires corresponds to a decrease in aggregate unemployment of approximately the same magnitude. We indeed find an increasing importance of mismatch unemployment at the end of the sample, which is closely related to the decrease of separation rates. Furthermore, a large share of the unemployed has to be reallocated to achieve the reduction ($30 - 38\%$). In the first decade, the primary relocation flow goes from East Germany to predominantly Bavaria. Afterward, unemployed workers from primarily NRW are relocated, again, to Bavaria. The changes in the reallocation pattern mirror the changes in the distribution of the unemployment rate: While the eastern districts have caught up substantially, in terms of distance to the average unemployment rate, unemployment in West German regions is stagnating.

Literature Our research is related to the literature on the ins and outs of labor markets (e.g., Elsby et al., 2011). Although differences in regional unemployment rates in Germany are well documented (Patuelli et al., 2012), an analysis of the underlying labor market flows is still missing from the literature. The new evidence we provide on the importance of separation rates for unemployment differentials echoes the findings of the research on the importance of worker flows for the development of the aggregate unemployment rate in Germany. For example, Hertweck and Sigrist (2015) and Jung and Kuhn (2014) show that the separation rate explains most of the variation in the unemployment rate over time. Moreover, complementary to Hartung et al. (2018), who document that the decrease in separations accounts for the major part of the decline in aggregate unemployment, we find that this decrease was heterogeneous across regions and led to more similar local separation rates. Concerning other countries, our results are in line with those of Bilal (2019), who finds that job-losses explain the majority of the variations in regional unemployment rates in the U.S. and France. We add that the explanatory power of separation rates can vary over time.

We add to the literature on the efficiency of regional labor markets. Our contribution is twofold. First, our results concerning mismatch unemployment contribute to the studies that focus on the efficiency of the spatial unemployment distribution (Sahin et al., 2014, and Marinescu and Rathelot, 2014). Our contribution, relative to the existing research on mismatch in Germany, is the extension of the mismatch measure by more than 20 years and the analysis at a much more disaggregated regional level.[3]

Second, we provide new results concerning the efficiency rationale for local labor market policies. While Kline and Moretti (2013) abstract from the layoff margin and study the relations ship between job-finding rates and hirings subsidies, there is little work that accounts for the importance of both layoffs and hirings. The present article is closest to the study of Bilal (2019), who rationalizes regional variations in job losses and hirings in a model with worker and firm mobility. In his model, spatial unemployment rate differentials are due to pooling externalities that drive the location choice of firms: High productive firms co-locate in slack labor markets, while low productive firms locate in tight labor markets. The co-location of the latter leads to high unemployment that is driven by high separation rates in tight labor markets. Our empirical results point to a more complex mechanism, as we do not observe that higher separation rates coincide with tighter labor markets. While we abstract from worker mobility, we provide a more extensive policy analysis where we account for the interaction of local matching frictions and the design of unemployment benefits.

Our research is related to the literature on optimal unemployment insurance as we build on the work of Jung and Kuester (2015). We show that, similar to unemployment insurance over the business cycle, in the presence of hiring subsidies and layoff taxes, the need for region-specific unemployment insurance benefits is small.

As the development of the identified region-specific structural parameters encapsulates the structural changes that underly the favorable development of the German labor market during 1994-2014, this research is related to the literature investigating this drop in unemployment. We provide evidence that it was rather the reform of the unemployment insurance benefits in the course of the labor market reforms in the mid-2000s, as argued by Hartung et al. (2018), instead of the liberalization of the bargaining agreements (Dustmann et al., 2014) that is responsible for lower unemployment.[4]

This paper is organized as follows. Section 3.2, describes the data, discusses the sample selection and provides a detailed overview on the development of local labor markets in Germany. In section 3.3, we present the model. Section 3.4 calibrates the model to a benchmark economy. Moreover, we relate regional variations to structural model parameters and quantitatively evaluate the optimal policy mix in the steady state. Section 3.5 introduces the framework to measure mismatch and the discusses

[3] While Bauer (2013) analyses mismatch during $2000-2010$ in West Germany with a special focus on the effect of the Hartz reforms; we describe the development of regional mismatch in Germany over three decades. Hutter and Weber (2017) explore the role of mismatch for accuracy of matching function forecasts. They use official vacancy and unemployment data from the federal employment agency and compute a simplistic mismatch measure for 21 occupations and 50 regions. Other studies are, for example, Erken et al. (2015) (Netherlands); Patterson et al. (2016) and Turrell et al. (2018) (both U.K.), who analyze mismatch and quantify the associated output loss.

[4] The literature investigating the Hartz reforms is extensive and additionally investigates the effect of the reform on the matching efficiency. See, for exmaple, Hertweck and Sigrist (2015), Launov and Wälde (2016), Klinger and Rothe (2012), Krause and Uhlig (2012), Klinger and Weber (2016).

the aggregate implications. Section 3.6 concludes.

3.2 Data and labor market development

This section introduces the microdata we utilize to construct local unemployment rates and workers flows. We provide a comprehensive overview of the heterogeneous developments of local labor markets in Germany and document the importance of worker flows for unemployment differentials.

3.2.1 Data and sample selection

Our primary data source is the employment panel of integrated employment biographies (SIAB) provided by the Institute for Employment Research (IAB).[5] The individual employment histories are used to construct monthly unemployment, job-finding, and separation rates. This data is merged with monthly vacancy data. We obtain data on the vacancy stock for $2000-2014$ from the federal employment agency and combine them with newly digitized data from their official historical reports. The segmentation of the historic data imposes the lowest possible level of disaggregation for our analysis.

Individual employment histories The SIAB is a two percent representative sample off all individuals who are unemployed or subject to social security contributions and covers the years $1975-2014$. The panel covers approximately 80% of the labor force, as self-employed and civil servants are excluded. We focus our analysis on the period $1980-2014$ and include East German workers from 1994 onward.[6]

The construction of monthly worker histories follows Jung and Kuhn (2014). In particular, we aggregate the daily employment histories to a monthly frequency using predefined reference weeks. Labor market states are assigned to a worker based on a hierarchical ordering. In general, parallel notifications are treated as follows: an employment notification replaces an unemployment notification, which in turn replaces non-participation. We exclude individuals with no information on their employment status or geographic location. Inactive employment relationships (e.g., maternity leave) are excluded as are marginally employed workers if they do not have a parallel unemployment spell. An individual is defined as employed if we observe a full-, part-time, or apprentice employment notification. A worker is marked as unemployed if she registered as unemployed at the federal employment agency. The registration-based

[5]This study uses the weakly anonymous Sample of Integrated Labour Market Biographies (Years 1975 - 2014). Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB) and subsequently remote data access.

[6]Information on workers in East Germany in 1992 and 1993 is used to impute the location variable if applicable.

measure is not consistently recorded for the years before 2000. Therefore, we use information on the unemployment benefit recipient status to construct prior histories. The flow data based on the benefit recipient definition is adjusted following Hartung et al. (2018). We remove the differences in levels between registered unemployed and benefit recipients and extend the registration-based flow rates backward using the growth rates of the benefit-recipient rates.[7] We control for inflows from nonemployment to unemployment in early 2005 that are due to regulatory changes. We follow Hartung et al. (2018) and exclude individuals that entered unemployment from nonemployment in the first six months of 2005 and did not transition to employment up to 2007.

Fig. 3.1: Aggregate unemployment rates (East and West Germany)

Notes: The figure displays the average monthly unemployment rates in East and West Germany. The left panel displays the official unemployment rates (FEA) as well as the unadjusted unemployment rates from SIAB. The right panel displays the SIAB unemployment rate with and without the inflow correction (SIAB adj). The official unemployment rates are measured in terms of the dependent civilian labor force. The constructed unemployment rates from SIAB data take into account the average annual number of civil servants in East and West Germany.

Figure 3.1 shows the resulting aggregate unemployment rates for East and West Germany as well as the corresponding official rates. The left panel shows the unadjusted time series, while the right panel displays the effect of the inflow adjustment. The constructed series closely resembles the official rate in level and cyclicality. The modification of the inflows from nonemployment mainly affects the unemployment rates after 2005.[8]

[7]The adjustment preserves the cyclical variation of each rate but adjusts for the level difference. As worker flow rates are highly volatile, we set the level difference to the average monthly difference in the first six months of 2000. Note that the mismatch index is invariant to aggregate changes in unemployment or vacancies. The relative distribution determines the size of mismatch.

[8]The official unemployment rates are measured in terms of the dependent civilian labor force, which we obtain from "Arbeitslosigkeit im Zeitverlauf: Entwicklung der Arbeitslosenquote (Strukturmerkmale)", 2019, Statistik der Bundesagentur für Arbeit. The constructed unemployment rates from SIAB data take into account the average annual number of civil servants in East and West Ger-

The regional level of analysis is the employment agency district. The territory of each EAD is based on the geographical division of 2014. There are 154 districts. We treat the three EADs covering Berlin as a single one because the territorial borders were subject to many changes during the observation period. Each district includes approximately three municipalities (*Kreise*, NUTS3). The SIAB data contains information on the municipality in which the workplace is located for every employed worker. Information on the place of residence for employed and unemployed is only available from 1999 onwards. To obtain consistent data on regional transitions covering the whole observation period, we assign the district of the last employment spell as the current place of residence when an individual becomes unemployed. We impute the location information of unemployed entering from non-employment by setting it to the location of their next employer if the employment is consecutive.

Vacancies We obtain data on registered vacancies from the federal employment agency for the years 2000 to 2014. Earlier data is digitized using the official monthly as well as annual reports (*Amtliche Nachrichten der Bundesagentur für Arbeit*) of the employment agency. We observe the number of vacancies that are reported to the FEA, which cover approximately $40 - 50\%$ of all vacancies (Brenzel et al. (2016)). However, other sources are not representative at this highly disaggregated regional level. Moreover, the reported vacancies before 2000 include job offers for seasonal workers or promoted vacancies, while the vacancies received from the FEA are based on a modified concept, where seasonal jobs and promoted jobs are excluded. We re-scale vacancies before 2000 by a constant factor to account for the level shift.[9]

The geographical area of responsibility of the EADs was not subject to larges changes up to 2012 when a significant reorganization of the territory took place. As the regional level of the SIAB data is based on the division of territory in 2014, we adjust the vacancy data correspondingly. Specifically, we use overlapping data for old and new territorial classifications to modify the vacancy counts to be representative of the 2014 allocation of territory. The modifications of the vacancy data are discussed in detail in appendix 3.A.3. Overall, we assemble a regional data set spanning the years $1980 - 2014$ for West and $1994 - 2014$ for West and East Germany.

many. We obtain the number of civil servants from the Statistical Offices of the Federation and the Länder ("74111-01-04-4-B" and "74111-01-05-4", Statistical Offices of the Federation and the Länder, Germany, 2019.).

[9]We re-weight vacancies before 2000 by a factor of 0.7 to account for job offers for seasonal workers or promoted vacancies. The adjustment factor is based on overlapping aggregate vacancy data from Hartmann and Reimer (2010) for the period $2001 - 2009$.

3.2.2 Regional labor market disparities

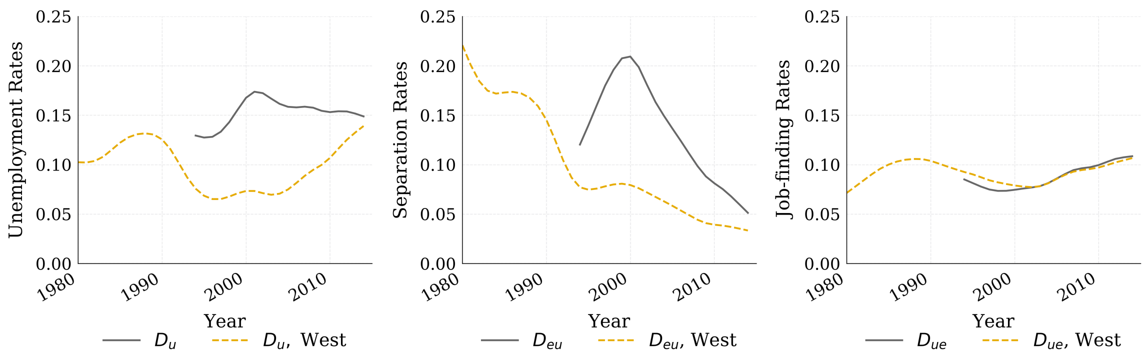
This section provides a comprehensive overview of the development of disparities in regional labor markets. We first focus on the dispersion of three key labor market variables across time and districts: unemployment rates, separation and job-finding rates. These three indicators give a good overview of how equal local labor markets in Germany are and which flow rate is most likely the dominant source of the evolution of the unemployment rate dispersion. The dispersion (D_t) is defined as the squared deviation of the respective variable from its average across districts

$$D_{m,t} = \frac{1}{I}\sum_{i=1}^{I}\left(\frac{m_{it} - \bar{m}_t}{\bar{m}_t}\right)^2 \quad \text{for m} \in \{\xi, f, u\}, \tag{3.1}$$

where $\bar{m}_t$ represents the average at time t. ξ is the separation rate, f the job finding rate and u the unemployment rate. An increase in dispersion is associated with a more unequal distribution across markets.

Figure 3.2 displays the results for West and united Germany (East and West Germany). Germany and West Germany show a large level difference in the dispersion of the unemployment rate, which indicates that the rates across districts located East and West differ substantially (left panel). We find that, despite the decrease in aggregate unemployment (Figure 1), the regional dispersion for both regions is higher in 2014, compared to 1994. Hence, the changes in unemployment rates were heterogeneous. Furthermore, the declining dispersion in united Germany during the 2000s, in combination with the increasing dispersion in West Germany starting in 2005, suggests that East German labor markets converge toward the average unemployment rate while West German EADs fall behind.

Fig. 3.2: Regional dispersion of labor market variables



Notes: The figure displays the annual dispersion of unemployment rates, separation and job finding rates across employment agency districts. All series display the HP-filtered trend ($\lambda = 6.25$).

Concerning separation rates (center panel), we find that the addition of East Germany leads to a substantial increase in the dispersion, compared to West Germany, which suggests that eastern EADs exhibit considerably different separation rates. However, separation rates have become much more similar throughout the sample period. In West Germany, we observe a substantial decrease in dispersion from 25% to below 5%. The average squared deviation from the mean separation rate in Germany increases up to the year 2000. Afterward, we observe a decrease from 20% to 5%. The dispersion of job finding rates is similar in level and trend between West and united Germany (right panel). Hence, job-finding rates in eastern districts are similar to those in western districts. Concerning the development over time, West Germany exhibits a decline from 1990 to 2004, after which the dispersion increases. Across all districts, the increase in dispersion commences already in 1997.

To summarize, the dispersion of local unemployment rates, in West and united Germany, has increased over the last two decades—despite a decrease in aggregate unemployment. Variations in separation rates did not cause this development since their dispersion across districts reduces substantially. It will most likely be related to the development of job-finding rates in West Germany because western districts are the primary driver of the dispersion of the job-finding rate over time.

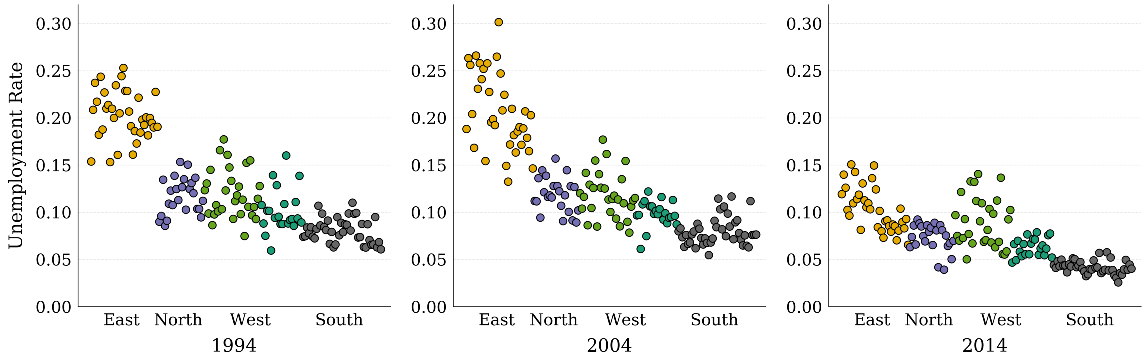
We now turn to a more detailed analysis and document the level differences across regions and their development over time. To visualize the spatial variations, we group the districts into four regions based on the geographical location of the respective federal state within Germany.[10] Although we regard North Rhine-Westphalia as a "West" region, the corresponding districts are shaded in a different color to ease the visual identification. Figure 3.3 shows the development of unemployment rates during 1994-2014. In 1994, eastern EADs were decoupled from the others as they exhibit unemployment rates almost twice as high. While the average difference to West German EADs had declined up to 2004, unemployment had increased to even higher levels across all districts. During the subsequent ten years, unemployment rates decreased significantly and have become more similar, which was determined by the dispersion measure presented before.

The development also provides suggestive evidence about which districts are responsible for the increase in the dispersion of the unemployment rates in West Germany and the stagnating dispersion in Germany. We find that it is predominantly districts located in North-Rhine Westphalia that exhibit high unemployment rates in 2014, while eastern EADs unemployment rates have almost converged to West German levels. Overall, while the dispersion across EADs increased, regional unemployment rates decreased by -54.0%, on average, from 1994 to 2014.

In the subsequent paragraphs, we link the observed differences in unemployment

[10] South: Bavaria and Baden-Wuerttemberg; West: North Rhine-Westphalia, Hesse, Saarland, and Rhineland Palatinate; North: Schleswig-Holstein, Hamburg, Bremen, and Lower-Saxony; East: Berlin, Saxony, Saxony-Anhalt, Thuringia, and Mecklenburg-West Pomerania.

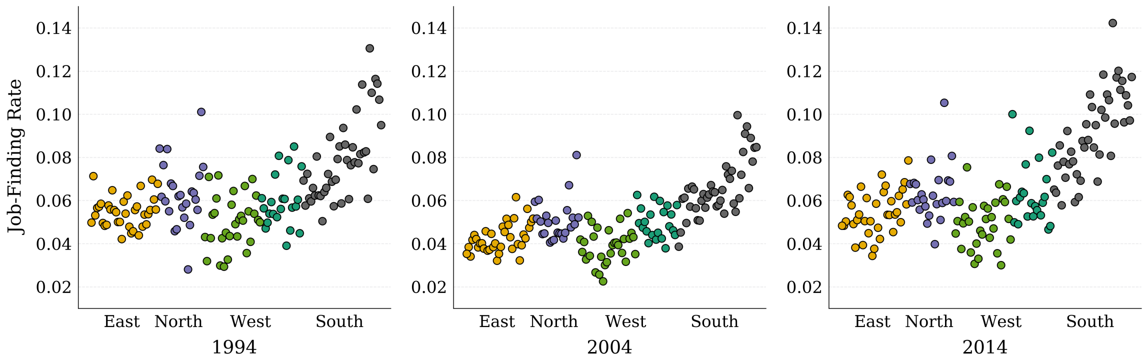
Fig. 3.3: Unemployment rates by employment agency district



Notes: The figure displays the average monthly unemployment rate in 154 employment agency districts for the years 1994, 2004 and 2014.

rates to the underlying variations in worker flows. As can be seen from Figure 3.4, the high unemployment rates in East Germany are not the result of lower transitions out of unemployment. The job-finding rates are similar across all EADs but the southern, where we observe approximately 50% above-average rates for all points in time. The rates are stable over the observation period. In fact, the average change between 1994 and 2014 was 5.0%. This small change indicates that job-finding rates are not the major driver of the between district or over time variation in unemployment rates. The distribution of separation rates, which is shown in Figure 3.5, paints a different

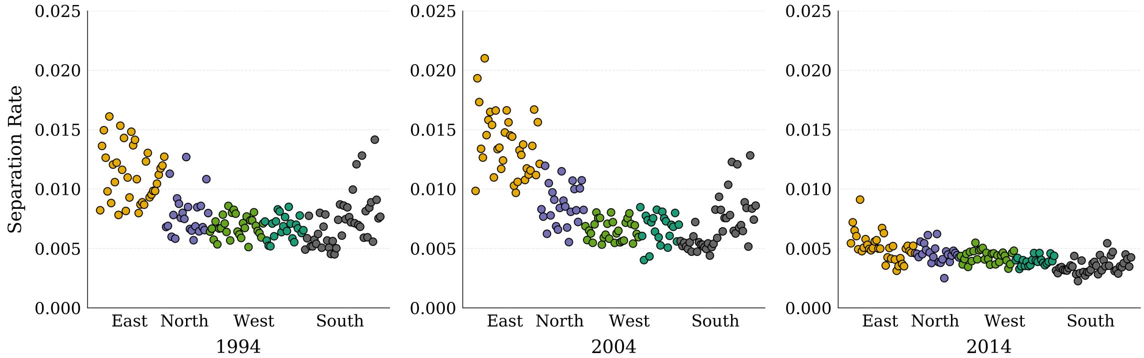
Fig. 3.4: Job-finding rates by employment agency district



Notes: The figure displays the average monthly unemployment inflow rate in 154 employment agency districts for the years 1994, 2004 and 2014.

picture of regional disparities. Foremost, it is remarkably similar to the distribution of unemployment rates: Separation rates in East German districts are more than twice as high in 1994 and 2004, compared to all other regions. In 2014, the differences had almost completely vanished, which is reminiscent of the low dispersion. Overall, we observe an average decrease in the separation rates of −63.0% throughout the sample period.

Fig. 3.5: Separation rates by employment agency district



Notes: The figure displays the average monthly unemployment inflow rate in 154 employment agency districts for the years 1994, 2004 and 2014.

Unemployment rate decomposition We employ a two-state stock-flow model to quantify the importance of the transitions in and out of unemployment for the variation in unemployment rates across districts. Specifically, we exploit the steady state relationship of the unemployment rate where $u_j^* = \frac{\xi_j}{\xi_j + f_j}$ approximately holds (Fujita and Ramey, 2009), where f_j denotes the job-finding and ξ_j the separation rate in district j. We decompose the volatility around the average unemployment rate in each year

$$ln(u_j^*/\bar{u}) = (1-\bar{u})\ln(\xi_j/\bar{\xi}) - (1-\bar{u})\ln(f_j/\bar{f}) + \epsilon_j, \tag{3.2}$$

or, in short, $du = d\xi + df + \epsilon$. A bar denotes the average rate across districts and ϵ denotes the error term. Fujita and Ramey (2009) show that $ln(u_j/\bar{u})$ can be decomposed to

$$1 = \frac{Cov(du, d\xi)}{Var(du)} + \frac{Cov(du, df)}{Var(du)} + \frac{Cov(du, \epsilon)}{Var(du)}, \tag{3.3}$$

where, e.g., $\frac{Cov(du,d\xi)}{Var(du)}$ measures the contribution of the separation rate to the variation of the unemployment rate across districts. Table 3.1 presents the results. They mirror the observed changes of the distribution of job-finding and separation rates, as displayed in Figures 3.4 and 3.5. Over the sample period, separation rates account for, on average, 59% of the unemployment rate variation across EADs. Their contribution decreases from 70% in 1994-2004 to 41% in 2005-2014. Variations in job-finding rates account for, on average, 42% and their importance increases from 31 in 1994-2004 to 58% in 2005-2014.

Determining which flow rate drives unemployment differences is important because worker flows are informative about how business cycle shocks or unemployment insurance benefits changes affect these differences, which in turn will affect local labor market policies (Jung and Kuhn, 2014, and Hartung et al., 2018). Our descriptive evidence shows that differences in separation rates are the main reason for the variation in

Table 3.1: Decomposition of regional unemployment rate disparities

	1994 - 2014	1994 - 2004	2005 - 2014
Job-finding	0.42	0.31	0.58
Separation	0.59	0.70	0.43

Notes: The table presents the results of a variance decomposition of the unemployment rates across districts for different time intervals. "Job-finding" ("Separation") indicates the contribution of the rate to the variation of unemployment rates across districts.

unemployment rates during 1994-2004. However, separation rates have become much more similar across EADs, while the differences in job-finding rates remained near to constant. As a result, the latter have become more critical in the determination of regional differences in unemployment rates.

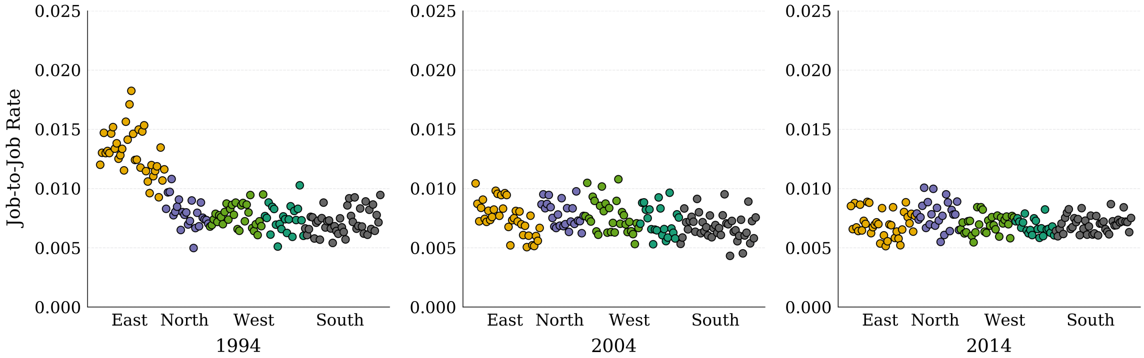
Our findings for Germany complement the results of Jung and Kuhn (2014) who show that variations of the separation rate largely explain aggregate unemployment volatility over time. Furthermore, the fall of the aggregate separation rate, which is documented by Hartung et al. (2018), is mirrored by regional separation rates. These decreases, however, were heterogeneous across districts, led to more equal separation rates, and increased the importance of the job-finding rate for regional unemployment differentials. Bilal (2019) also finds that variations in job losses explain a significant part of regional unemployment rate differentials in the U.S. and France. He does not provide any evidence on the evolution over time.

In the appendix, we provide additional evidence on differences in occupational unemployment rates in West and East Germany (Section 3.B.1). Our findings show that East Germany exhibits a similar distribution of separation and job-finding rates across occupations compared to West Germany. However, we show that occupational unemployment rates are higher in East Germany due to higher separation rates. Furthermore, disparities in occupational unemployment rates are driven by differences in separation rates for East and West Germany (Table 3.7, and 3.8). Separation rates explain the majority (above 80%) of variations in the unemployment rate throughout 1994-2014. We observe a slight increase in the importance of the job-finding rate up to 18%.

Wages, vacancies and job-to-job transitions Figure 3.6 shows that differences in job-to-job transition rates cannot explain regional unemployment rate differentials. In 1994, the districts with the highest unemployment even exhibit the highest employer to employer transitions. For all other points in time, we find no substantial variation of job-to-job transitions rates across EADs.

While unemployment rates are one of the most important labor market indicators,

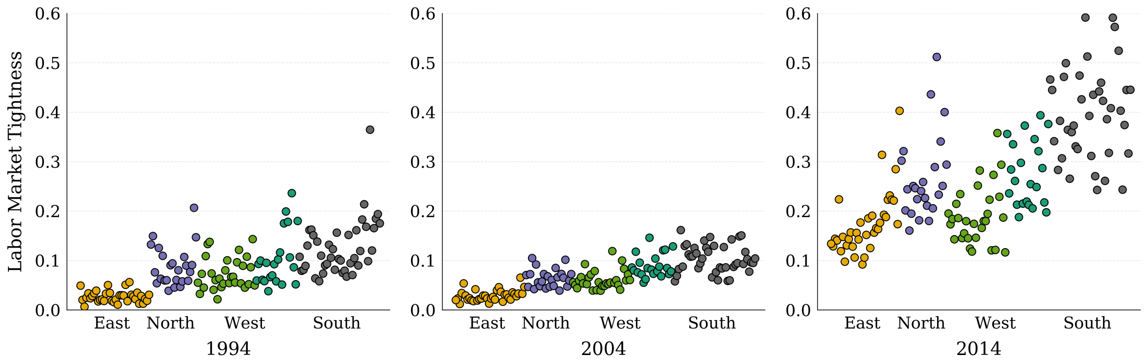
Fig. 3.6: Job-to-job rates by employment agency district



Notes: The figure displays the average job-to-job transition rates in 154 employment agency districts for the years 1994, 2004 and 2014.

labor market tightness is also informative with respect to the state of the economy. Figure 3.7 displays labor market tightness −the ratio of vacancies over unemployed− across EADs. From 1994 to 2004, tightness has decreased across all districts mirroring the increase in unemployment rates. In the next 10 years, markets have become substantially more tight, i.e., tightness increases. This is generally favorable for job-seekers because they compete for relatively more open vacancies.

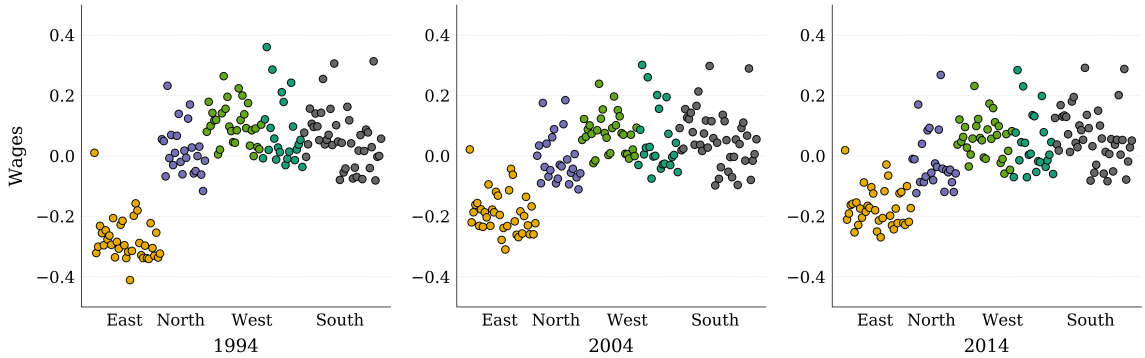
Fig. 3.7: Labor market tightness by employment agency district



Notes: The figure displays the average labor market tightness (the ratio of vacancies over unemployed) in 154 employment agency districts for the years 1994, 2004 and 2014.

Figure 3.8 displays the deviations of the wage from the average wage across districts in each year. Wages exhibit a substantial dispersion across EADs and this dispersions is very persistent as the correlation between 1994 and 2014 is above 0.9. They are lowest in East German EADs. However, eastern districts catch up to the average wage as the deviation reduces from −30% in 1994 to −20% in 2014.

Fig. 3.8: Wages by employment agency district



Notes: The figure displays the average (log) deviation of the wage from the average wage in each year in 154 employment agency districts for the years 1994, 2004 and 2014.

3.3 The model

There is a small open economy that consists of i islands/regions populated by workers of mass ϖ_i. A fraction e_i of these workers is employed and the remainder is unemployed $u_i = 1 - e_i$ at the beginning of the period. Workers can not move from one region to another. Regions take the unemployment insurance (UI) benefit system, layoff taxes, hiring subsidies, and production taxes as given. There is no self-insurance by workers. A federal UI scheme makes unemployment-dependent transfers to the member-state government. Each island produces a freely tradable homogeneous good with a price normalized to 1.

3.3.1 Workers

Time is discrete. A worker living in region i and discounting the future with $\beta \in (0, 1)$ has lifetime utility of

$$E_0 \left\{ \sum_{j=0}^{\infty} \beta^t [U(c^i_{n,t}) + \overline{h} \cdot I(\text{not working}_t) - \iota \cdot I(\text{search}^i_{n,t})] \right\}. \tag{3.4}$$

E_0 denotes the expectation operator. The worker enjoys utility from consumption, $c^i_{n,t}$ with a standard felicity function $U(c) : \mathcal{R} \to \mathcal{R}$ which is twice continuously differentiable, strictly increasing and concave. n indicates whether the worker is employed (e) or unemployed (u). I is the indicator function. If not employed, the worker makes a 0-1 decision to search for a job or not. Workers differ by a utility cost of search, ι, incurred only if the worker searchers for a new job. Both across workers and time $\iota \sim F_\iota(0, \sigma^2_\iota)$ is *iid*, where $F_\iota(\cdot, \cdot)$ marks the logistic distribution with mean 0 and variance $\sigma^2_\iota = \pi \frac{\psi^2_\iota}{3}$, with $\psi_\iota > 0$ and π being the mathematical constant. We denote with $\overline{h}$ the average value of leisure net of mean search cost, which can be negative if search is very costly.

Workers cannot self-insure against income fluctuations through saving or borrowing but might receive dividend income Π^i (described below) from owning the firms in their region. Ownership rights are distributed equally across inhabitants of the region.

Let w_t^i be the wage that an employed worker earns, consumption of the worker is given by

$$\begin{aligned} c_{u,t}^i &:= b_t^i + \Pi^i \quad \text{if unemployed at the beginning of } t, \\ c_{e,t}^i &:= w_t^i + \Pi^i \quad \text{if employed at the beginning of } t. \end{aligned} \tag{3.5}$$

If the worker enters the period unemployed, the worker receives an amount b_t^i of unemployment benefits. The government, by assumption, conditions payment only on the worker's current employment status. A worker who enters the period employed receives wage income or, if separated, a severance payment equal to the period's wage. Given that we abstract from aggregate risk, we will describe the value functions and profit equations recursively and denote with primes future values.

Value of an employed worker Let ξ^i be the separation rate of existing matches in region i. Before separations occur, the value of an employed worker is

$$V_e^i = \quad U(c_e^i) + [1-\xi^i]\beta V_e^{i\prime} + \xi^i[V_u^i - \mathtt{U}(c_u^i)]. \tag{3.6}$$

A worker who is employed at the beginning of the period consumes c_e^i irrespective of the separation decision (due to severance payments). With probability $1-\xi^i$ the match continues. With probability ξ^i, instead, the match separates and the worker immediately start to searching for new employment. V_u^i is the value of a worker who starts the period unemployed.

Value of an unemployed worker and search An unemployed worker chooses a cut-off strategy balancing her search costs ι with the expected gain from search:

$$\iota^{s,i} = f^i \, \beta \, \Delta'^i \tag{3.7}$$

Here $\Delta^i \equiv V_{e,t}^i - V_{u,t}^i$ denotes the gain from employment relative to unemployment and f^i marks the job-finding rate. Using the properties of the logistic distribution, the share of unemployed workers who search is given by

$$s^i = Prob(\iota \leq \iota^{s,i}) = 1/[1+\exp\{-\iota^{s,i}/\psi_s\}]. \tag{3.8}$$

and the conditional expectation can be shown to be

$$\int_{-\infty}^{\iota_t^{s,i}} \iota^{s,i} dF_\iota(\iota) \equiv \Psi(s^i) = -\psi_s[(1-s^i)\log(1-s^i) + s^i \log s^i] \tag{3.9}$$

which can be interpreted as the option value of having a choice to search.

With this, the value of an unemployed worker at the beginning of the period, before the realization of the search preference shock, is given by

$$\begin{aligned} V_u^i &= U(c_u^i) + \overline{h} + \Psi(s^i) + s^i f^i \beta \Delta^{i\prime} + \beta V_u^{i\prime} && (3.10) \\ &= U(c_u^i) + \overline{h} - \psi_s \log(1-s^i) + \beta V_u^{i\prime} && (3.11) \end{aligned}$$

The worker consumes c_u^i unemployment insurance income, enjoys utility of leisure $\overline{h}$ and has an option value of searching. With probability s^i the worker decides to search and obtains a job offer with probability f^i, so the product is the transition probability of moving to employment if unemployed today. The second line follows by substituting in the optimal choice s^i into the option value $\Psi(s^i)$. If the worker does not search or does not receive an offer, she remains unemployed receiving a discounted value of unemployed $\beta V_u^{i\prime}$. For later reference we can derive

$$\Delta^i = U(c_e^i) - U(c_u^i) - [1-\xi^i][\overline{h} - \psi_s \log(1-s^i)] + [1-\xi^i]\beta E_0 \Delta^{i\prime} \tag{3.12}$$

3.3.2 Firms

Firms are owned in an equal amount by the inhabitants of the island. Firms discount future profits with a given discount factor R^i that we allow to be island specific (derived below). Firms build a match with one worker to produce output. A firm that enters the period matched to a worker can either produce or separate from the worker. Production entails a firm-specific resource cost, ϵ_j. This fixed cost is independently and identically distributed across firms and time with distribution function $F_\epsilon(0, \sigma_\epsilon^2)$. $F_\epsilon(\cdot, \cdot)$ is the logistic distribution with mean zero and variance $\sigma_\epsilon^2 = \pi \frac{\psi_\epsilon^2}{3}$, with $\psi_\epsilon > 0$. The firm separates from the worker (first line) whenever the idiosyncratic cost shock, ϵ_j, is larger than a state-dependent threshold $\epsilon_t^{\xi,i}$. Using again the properties of the logistic distribution, conditional on the threshold, the separation rate can be expressed as

$$\xi^i = Prob(\epsilon_j \geq \epsilon^{\xi,i}) = 1/[1 + \exp\{(\epsilon^{\xi,i})/(\exp\{a^i\}\psi_\epsilon)\}]. \tag{3.13}$$

and the option value of having a choice can be denoted by $\Psi(\xi^i) = -\exp\{a^i\}\psi_\epsilon[(1-\xi^i)\log(1-\xi^i) + \xi^i \log \xi^i]$.

Ex ante, namely, before the idiosyncratic cost shock ϵ_j is realized, the value of a firm that has a worker is given by

$$J^i = \begin{array}{l}(1-\xi^i)\left[\exp\{a^i\} - \tau_J^i - w^i + E_0 \frac{J^{i\prime}}{R^i}\right] \\ -\xi^i\left[\tau_\xi^i + w^i\right] + \Psi(\xi^i)\end{array} \tag{3.14}$$

Upon separation, the firm has to pay a layoff tax τ_ξ^i and severance payments equal to the current wage. If the match survives it will produce output with the regional specific productivity a_t^i, pays the worker a wage w^i and faces a production tax tax τ_J^i proportional to output (or wages in equilibrium). A match that produces this period continues into the next. Firms on the main island pay a cost $\exp\{a_t^i\}\kappa_v^i$ to create a vacancy in island i again assumed to be proportional to productivity (i.e., expressed in opportunity cost of time of the local manager of the firm). If the firm finds a worker, the worker is hired this period and can start producing from the next period onward. If the firm hires a worker the firm might receive a hiring subsidy τ_q^i payed out at the end of this period after matching took place. Firms post vacancies on each region as long as the cost of a vacancy equals its gains

$$\exp\{a^i\}\kappa_v^i = q^i E_0 \frac{J^{i\prime}}{R^i} + q^i \tau_q^i, \tag{3.15}$$

where q^i is the probability of filling a vacancy. Let v^i be the number of vacancies posted. The number of matches m^i and the separation rate determine the evolution of employment

$$e^{i\prime} = [1-\xi^i]\cdot e^i + m^i.$$

and we impose a constant-returns matching function:

$$m^i = \chi_i \cdot \left[v^i\right]^\gamma \cdot \left[[\xi^i e^i + 1 - e^i]s^i\right]^{1-\gamma}, \; \gamma \in (0,1). \tag{3.16}$$

Here, $\chi_i > 0$ is matching efficiency,market tightness $\theta^i := v^i/([\xi^i e^i + 1 - e^i]s^i)$, the job-finding rate as $f^i := m^i/([\xi^i e^i + 1 - e^i]s^i) = \chi_i[\theta^i]^\gamma$, and the job-filling rate as $q^i := m^i/v^i = \chi_i[\theta^i]^{\gamma-1}$. The mass of workers who potentially search is $\xi^i e^i + 1 - e^i$, with $\xi^i e^i$ being workers separated at the beginning of the period. s^i is the share of those who do actually search.

Total production of output in region i is given by

$$y_t^i = e^i(1-\xi^i)\exp\{a^i\} \tag{3.17}$$

where $e^i(1-\xi^i)$ is the mass of existing matches that are not separated at the beginning of the period.

Dividends Dividends per inhabitant accruing from region i arise from firm profits, namely,

$$\begin{aligned} \Pi^i \;=\; & e^i\Psi(\xi^i) + e^i(1-\xi^i)\left[\exp\{a^i\} - \tau_J^i - w^i\right] - e^i\xi^i\left[w^i + \tau_\xi^i\right] \\ & - \exp\{a^i\}\kappa_v v^i + v^i q^i \tau_q^i. \end{aligned} \tag{3.18}$$

and are distributed back lump sum.

3.3.3 Bargaining between firm and worker

At the beginning of the period, workers and firm bargain over the wage and the severance payment as well as over a state-contingent plan for separation (as a function of the realization of the cost shock $\epsilon^{\xi,i}$)[11] using a standard Nash-bargaining protocol

$$(w^i, \epsilon^{\xi,i}) = \arg\max_{w^i,\epsilon^{\xi,i}} (\Delta^i)^{1-\eta}(J^i)^{\eta}, \tag{3.19}$$

where η measures the bargaining power of the firm. The first-order condition for the wage is as follows

$$(1-\eta^i)J^i = \eta^i \frac{\Delta^i}{U'(c_e^i)}. \tag{3.20}$$

The first-order condition for the separation cutoff yields

$$\epsilon^{\xi,i} = \left[\exp\{a^i\} - \tau_J^i + \tau_\xi^i + \frac{J^{i\prime}}{R^i}\right] + \frac{\beta\Delta_{u,}^{i\prime} + \psi_s\log(1-s^i) - \overline{h}}{U'(c_e^i)}.$$

3.3.4 Government

The government of the region finances its spending by imposing a tax on firm τ_J^i and possibly on separation τ_ξ^i. It runs an UI benefit scheme to finance the unemployment payments $u^i b^i$ and possibly offers hiring subsidies per region.

$$[e^i(1-\xi^i)\tau_J^i + e^i\xi^i\,\tau_\xi^i] + \Delta B = u^i b^i + q^i v^i \tau_q^i \tag{3.21}$$

and ΔB is a financing item the region receives/pays from the general country-wide government that mechanically balance the local budget for a given policy instruments.

3.3.5 Planner

To determine the optimal labor market policy instruments, we will focus on a constrained-efficient allocation where the local government respects the constraint that all unemployed and all employed workers have to be treated equally, i.e., we do not allow the

[11]The firm will insure the risk-averse worker against the idiosyncratic risk associated with ϵ_j so that the wage, w, is independent of the realization of ϵ_j, hence the severance payment equals the wage

government to condition on the duration of unemployment. Moreover we look for a solution without interregional transfers, i.e., we impose the constraint that the regional governmental budget has to balance. Let Δ^i be the promised utility difference between the employment state and unemployment. The government starting with a particular promise and employment level e^i then solves the following maximization problem

$$\begin{aligned}
W(\Delta^i, e^i) &= \max_{\xi^i,\theta^i,c^i_e,c^i_u,\Delta^{i\prime}} e_i U(c_{e,i}) + (1-e^{i\prime})U(c_{u,i}) \\
&\quad +(e^i\xi^i + (1-e^{i\prime}))(\Psi_s(s_i) + \overline{h}) + \beta W(\Delta^{i\prime}, e^{i\prime}) \\
&\quad s.t. \\
e^{i\prime} &= e^i(1-\xi^i) + (e^i\xi^i + 1 - e^{i\prime})s^i\chi_i(\theta^i)^\gamma \\
s^i &= 1/[1+\exp\{-s^i\chi_i(\theta^i)^\gamma\beta\Delta^{i\prime}/\psi_s\}] \\
\Delta^i &= U(c^i_e) - U(c^i_u) - (1-\xi^i)[\overline{h} - \psi_s\log(1-s^i)] + [1-\xi^i]\beta E_0\Delta^{\prime i} \\
e^i(1-\xi^i)\exp\{a^i\} &= e^i c^i_e + u^i c^i_u - e^i\Psi(\xi^i) + (e^i\xi^i + 1 - e^{i\prime})s^i\theta^i\exp\{a^i\}\kappa^i_v
\end{aligned}$$

The objective function of the local government is utilitarian welfare maximization with equal weighting of employed and unemployed workers. It takes promised utility differences as given, i.e., respects the moral hazard constraint that idiosyncratic search cost are unobservable to the government, so the government respects the privately optimal search decision. The final constraint imposes a local balanced budget rule, i.e., it rules out any redistribution across regions. This assumption avoids the political economy involved in regional redistribution and focuses instead on welfare improvements within a balanced budget world.

Let us state for later reference the implied discount kernel of the social planner $Q = \beta\frac{\lambda^{i\prime}}{\lambda^i}$ where

$$\lambda^i = (\frac{e^i}{u'(c^i_e)} + \frac{u^i}{u'(c^i_u)})^{-1}$$

is given as weighted marginal utility of consumption. We will impose that the discount kernel of the firms in a decentralized equilibrium is given by the planners kernel $R = Q$. This assumption ensures that differences in allocations do not arise due to differences in discounting.

3.3.6 The optimal policy mix

With these assumptions in place, it follows from arguments in Jung and Kuester (2015) that an optimal set of instruments that decentralizes the constrained-efficient allocation of the above planner's solution can be characterized as summarized by the following proposition.

Proposition 1. *Consider the economy described in the prior sections. Let* $\Omega := \frac{\eta}{\gamma}\frac{1-\gamma}{1-\eta}$

be the Hosios measure of search externalities and $\zeta = \frac{\psi_s}{f(1-s)} \frac{1-e}{[\xi e+(1-e)]} \frac{\mathtt{U}'(c_u)-\mathtt{U}'(c_e)}{\mathtt{U}'(c_u)\mathtt{U}'(c_e)}$ *be a measure of tension between moral hazard and insurance of the unemployed. The following tax rules then implement the constrained-efficient allocation in steady state:*

$$
\begin{aligned}
\tau_q^i &= \exp\{a^i\}\kappa_v^i \frac{\theta^i}{f^i}\left[1-\Omega^i\right] + \frac{\eta^i}{1-\eta^i}\zeta^i \\
\tau_\xi^i &= \tau_J^i + \tau_q^i + (1-s^i f^i)\zeta^i \\
b^i &= \tau_q^i \frac{(1-\beta)}{\beta} e + \frac{e^i}{\beta}[1-\beta(1-s^i f^i)(1-\xi^i)]\zeta^i \\
\tau_J &= \frac{1-e}{e} b^i - \xi^i(1-s^i f^i)\zeta^i
\end{aligned}
$$

In our model, the government would intervene in the labor market for two reasons. The proposition captures these by the terms Ω and ζ which reflect, respectively, that the optimal policies are affected by search externalities arising due to a violation of the Hosios (1990) condition and tensions between moral hazard and insurance.

To see how the planner counteracts the search externalities, it is useful to consider a case where workers are risk-neutral and, thus, the moral hazard insurance motive is absent ($\zeta = 0$). Then, hiring subsidies are the primary means to counteract the search externalities. When the bargaining power of firms is below the elasticity of the matching function with respect to vacancies ($\gamma > \eta$), the planner subsidizes hirings to foster vacancy creation. In this case, wages are too high compared to productivity, firms post too few vacancies, and unemployed jobseekers crowd each other out. On the other hand, when the firms' bargaining power is above the vacancy share, the planner will levy a hiring tax to lower the number of vacancies posted. Moreover, layoff taxes would balance the budget, while unemployment insurance benefits and the production tax would be close to zero.

In our model, the planner additionally insures the unemployed as they cannot save. However, by providing this insurance, he changes the workers' outside option, which, in turn, affects their unobservable search effort and all of their labor market decisions. The moral hazard insurance motive implies positive hiring subsidies as more vacancies increase the workers' search incentives. However, in contrast to the search externalities, the importance of this motive increases with the firms' bargaining power. Layoff taxes act as a means to finance the hiring subsidies and ensure that firms take into account the prospective social costs of an additional unemployed worker. To see the main trade-offs more easily it is helpful to focus on a particular case with logarithmic utility evaluated at the Hosios-condition where we allow the discount factor to go to $\beta \to 1$.

Corollary 2. *Under the same conditions as in Proposition 1, assume furthermore that* $\beta \to 1$*, the Hosios condition holds,* $\gamma = \eta$ *(that is,* $\Omega = 1$*). Define* $D_2 = D - 1$ *as the adjusted duration* $D = \frac{1}{sf}$ *over which the government on average pays unemployment benefits to an unemployed worker, let* $\epsilon_{D_2} := \frac{D}{D_2}\frac{f}{\psi_s}(1-s)c_u U'(c_u)$ *be the elasticity of* *duration D_2 with respect to an increase in consumption of an unemployed worker in the next period. Assume logarithmic utility U(c)=log(c),then $\zeta = \frac{(c_e - c_u)}{\epsilon_{D_2}}$ and we get*

$$
\begin{aligned}
\frac{b^i}{w^i} &= \frac{1}{1 + D\epsilon_{D_2}} && (3.22)\\
\tau_q^i &= \frac{\eta^i}{1-\eta^i}\frac{D}{1 + D\epsilon_{D_2}}w && (3.23)\\
\tau_\xi^i &= \tau_q^i + \frac{D_2}{1 + D\epsilon_{D_2}}w && (3.24)\\
\tau_J &= 0 && (3.25)
\end{aligned}
$$

The first equation tells us how to set the net-replacement rate optimally, namely, according to the standard Bailey-Chetty formula (Baily, 1978, and Chetty, 2006); the lower the micro-elasticity of search is the larger the government can grant consumption insurance. The second equation shows that even at the Hosios-condition, the government grants hiring subsidies. When making a hiring decision of an unemployed, firms do not take into account the relaxation of the financing constraint of the social security system that arises from their private action. To align private and social incentives, it is optimal to subsidize hirings proportional to the duration weighted net benefit payments. The factor of proportionality depends on the bargaining share of the firm as a lower share implies higher search incentives for workers through higher wages. Layoffs are taxed to ensure that firms internalize the additional costs for the social system incurred through the separation. First, firms bear the hiring subsidies that are utilized to reemploy the worker. Second, the planner ensures that the firms take the costs of the UI benefit payments for the duration of the unemployment spell following the first month into account, as firms are only obliged to pay a severance payment equal to the monthly wage upon separation.

To summarize, the vacancy elasticity of the matching function and the bargaining power will determine whether the planner mitigates search externalities by subsidizing or taxing hirings. Concerning the moral hazard insurance trade-off, the duration of the average unemployment spell is the primary driver of the optimal policy mix.

3.4 Quantitative evaluation

The purpose of this section is twofold: First, we utilize the model to translate regional labor market disparities into structural variations. Based on a benchmark calibration, region-specific wages, vacancies, and job-finding and separation rates, we identify four structural parameters for each district and two time intervals: 1994 to 2002 and 2008 to 2014. We then relate the recovered structural changes to existing theories brought forward to explain the labor market development in Germany.

In the second part of this section, we assume that the structural parameters are invariant to labor market policy instruments and evaluate the welfare and employment effects of the optimal policy mix for each region. This exercise allows us to provide an important insight on local labor market disparities as we determine to what degree these disparities are driven by local inefficiencies.

3.4.1 Calibration

We initially calibrate the model to resemble a fictional unemployment agency district. This calibration will serve as our benchmark for the subsequent quantitative exercise. The fictional economy is formed based on average worker flows, vacancies, and wages in all EADs in Germany between 1994 and 2002. Our measure of the wage is the ratio of the average wage of all employees to the employment-weighted average. Annual averages of monthly job-finding and separation rates are taken directly from the data. Labor market tightness is the weighted average vacancy to unemployment ratio.[12]

We assume log utility and also evaluate the model for log utility in the subsequent sections. While we set some parameters outside the model, others are determined within the model by targeting the average job-finding and separation rate, labor market tightness, and wages. One period represents a month.

Table 3.2 summarizes the calibrated parameters. The time-discount factor is calibrated to $\beta = 0.996$ to match a monthly interest rate of 0.04. We set the utility from leisure to $h = 0.476$ to target an unemployment rate of 0.105. In the model, not all of the unemployed search, this calibration ensures that in the steady-state, the share of non-searching unemployed amounts to 81 percent. This share matches the proportion of unemployed workers that do not search actively but are available on short notice in 2014.[13] We set $\psi_s = 0.569$ to match an elasticity of the average duration of unemployment concerning benefits of 0.5. This value is well in the range of the estimates of the literature (Schmieder and Von Wachter, 2016).

We calibrate vacancy posting costs to $\kappa_v = 2.5$. This value is based on the estimated vacancy posting costs of Muehlemann and Pfeifer (2016) that amount to 4,700 Euro and an average net wage of approximately 1,800 Euro. The elasticity of the matching function with respect to vacancies is calibrated to $\gamma = 0.25$, which is estimated directly from the data and discussed in a later section (3.5.2). To determine the bargaining parameter, we target the average monthly labor market tightness in the fictional economy. This leads to $\eta = 0.156$. The matching efficiency parameter in the benchmark economy is determined within the model and set to $\eta = 0.142$. This value leads to an

[12]We re-weight vacancies before 2000 by a factor of 0.7 to account for job offers for seasonal workers or promoted vacancies. The adjustment factor is based on overlapping aggregate vacancy data from Hartmann and Reimer (2010) for the period $2001 - 2009$.

[13]We obtain the number of individuals participating in employment agency measures on unemployment in 2014 from Fuchs et al. (2014).

Table 3.2: Calibrated parameters

Parameter		Value
Preferences		
β	discount factor	0.996
$\bar{h}$	disutility of work	0.476
ψ_s	scaling parameter utility of search costs	0.569
Vacancies, matching and bargaining		
κ_v	vacancy posting costs	2.500
γ	match. func. elasticity w.r.t vacancies	0.250
χ	matching efficiency	0.142
η	bargaining power	0.156
Production and layoffs		
A	Productivity	0.893
ψ_ϵ	disp. of idiosyncratic costs	4.051
Labor Market Policy		
b	Replacement rate	0.670
τ_q	Vacancy posting subsidy	0
τ_ξ	Layoff tax	3

Notes: The table presents the calibrated parameter values. We calibrate the model to resemble the average unemployment agency district during $1994 - 2002$.

average job-finding rate of 0.06, as in the data.

Concerning productivity, we target the average wage and set productivity to $A_i = 0.893$. The dispersion parameter for the cost shock is calibrated to $\psi_\epsilon = 4.051$, to match the average separation rate.

Last, we set the policy variables. These variables are scaled in terms of wages. We set the replacement rate, $b = c_u/c_e$, such that in the steady-state benefit payments replace 67% of wages. This value corresponds to the official replacement rate for a worker with one or more children in the first year of unemployment. We set vacancy subsidies to $\tau_q = 0$. Layoff taxes are set to a constant value, in terms of wages, of $\tau_\xi = 3$. This assumption is a compromise, as there are no layoff taxes, per se, but the government protects workers through a statutory notice period of three months. The government balances its budget through the production tax τ_J.

3.4.2 Structural disparities

We utilize the calibrated benchmark model and the regional variation of wages, labor market tightness, job-finding and separation rates to calibrate four region-specific structural parameters. Moreover, we capture the documented varying importance of the worker flow rates, as we determine these structural parameters for two time periods from 1994 to 2002 and 2008 to 2014.[14] Our approach is as follows, we assume that the workers' preferences do not differ across districts and set them to the respective parameters of the benchmark calibration. Additionally, each local labor market exhibits a similar vacancy elasticity of the matching function, and the policy instruments are valid for all regions. The four remaining parameters are productivity, matching efficiency, bargaining power, and the dispersion of the cost shock. We interpret the latter as the uncertainty of a job-match, as it governs separation decisions. These parameters are re-calibrated for each district and time period such that the model matches our empirical observations of wages, labor market tightness, job-finding and separation rates. Hence, the four parameters are determined relative to the benchmark scenario, as the preference and some technology parameters are held constant.

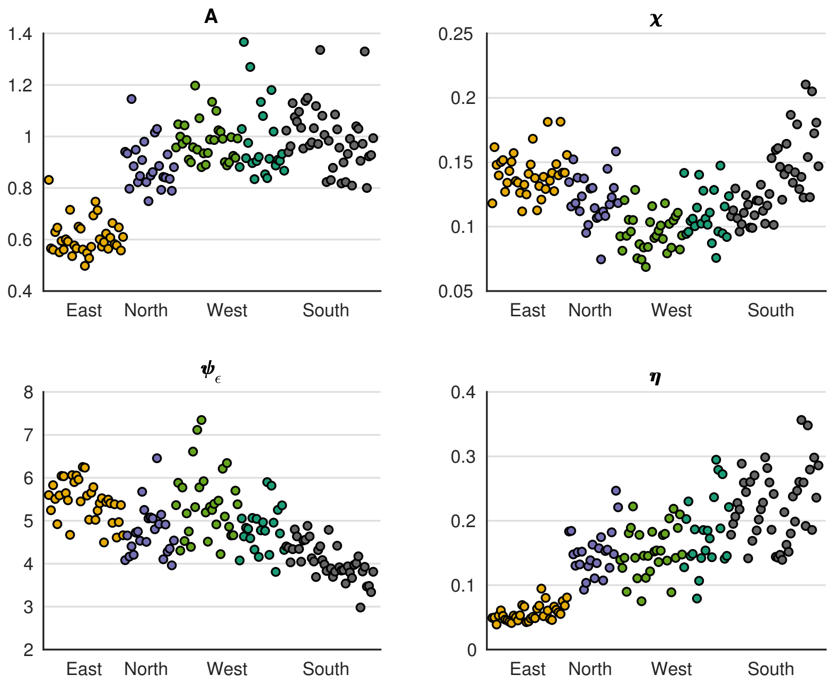
Given the model, the four parameters are jointly identified based on observed variations of wages, labor market tightness as well as job-finding and separation rates. We provide an intuition for the explanatory power of each variable in the appendix (Section 3.C). In brief, wages solely affect the productivity parameter. Separation rates shape the distribution of the uncertainty of a job-match, while they have a small effect on regional productivity. Job-finding rates, in combination with labor market tightness, determine the matching efficiency, which is identified with almost no uncertainty. Given the constant vacancy share, tightness, and job-finding rates, the matching function determines the efficiency parameter. Moreover, tightness and job-finding rates influence job-match uncertainty through the surplus of a match: ceteris paribus, higher job-finding rates lead to a larger value of unemployment, thus leading to a smaller surplus and less match uncertainty. The bargaining power is almost exclusively determined by labor market tightness, as the number of vacancies posted and the bargaining power are closely related.

Figure 3.9 displays the identified structural disparities in 1994-2002. The productivity parameter A is identified using the empirical wage observations. We find that productivity in eastern districts is well below the West German levels. Besides this vast difference, the variations across EADs in the North, West, or South are substantial. The seven districts with the highest productivity in their respective region are Berlin (East), Hamburg (North), Dusseldorf, Frankfurt, Bad Homburg (West), Ingolstadt, and Munich (South). They are all major centers of economic activity. Overall, the

[14]Note, that we abstract from any transitional dynamics due to the labor market reforms during the mid-2000s and exclude the years 2003 to 2007.

productivity distribution closely resembles the wage distribution (Figure 3.8).

Fig. 3.9: Structural labor market disparities (1994 – 2002)



Notes: The figure displays the identified values of the four region-specific structural parameters: Productivity (A), matching efficiency (χ), match uncertainty (ψ_ϵ), and bargaining power (η). We determine the parameters within the model given the observed variation in wages, labor market tightness, separation, and job-finding rates. We target the averages of these four variables within 1994 – 2002.

Matching efficiency shows a U-shaped pattern, where eastern and southern districts exhibit the highest values (upper right). This parameter is jointly identified by labor market tightness and the job-finding rate. The fact that labor markets are less tight in East Germany, but the job-finding rate is similar to the other districts' rates is captured by the higher matching efficiency. Hence, high unemployment in East Germany is not attributable to low efficiency in the matching market. EADs located in NRW exhibit the lowest efficiency, which mirrors the low job-finding rates. Similarly, the high frequency of flows into employment, in combination with a tight labor market, in the South identify an above-average efficiency of the matching market.

The uncertainty of a match (ψ_ϵ) captures the variations of the separation rates (lower left). Consequently, we find high uncertainty in East Germany. As shown in the empirical section, separation rates do not vary much across the other districts. The structural parameter, however, does exhibit significant variation. This variation is due to differences in the surplus of a match, which additionally shapes the identified distribution. The match surplus depends on the job-finding rate: A higher job-finding rate increases the value of unemployment and, thus, leads to a smaller surplus and lower match uncertainty, given constant separation rates. This second effect determines the higher and lower parameter values of ψ_ϵ in NRW and the South, respectively.

Concerning the bargaining power of firms, we uncover an increasing pattern from East to South (lower right). The shape of the identified parameter across districts matches the distribution of labor market tightness closely. The bargaining power governs the relationship between wages and productivity. Hence, while eastern workers are paid wages that are high, most of the southern workers receive wages that are lower relative to productivity in the respective region. This finding is consistent with the explanation that the wages in the East that were bargained by West German labor unions after the reunification were too high relative to productivity (e.g., Burda and Hunt, 2001).

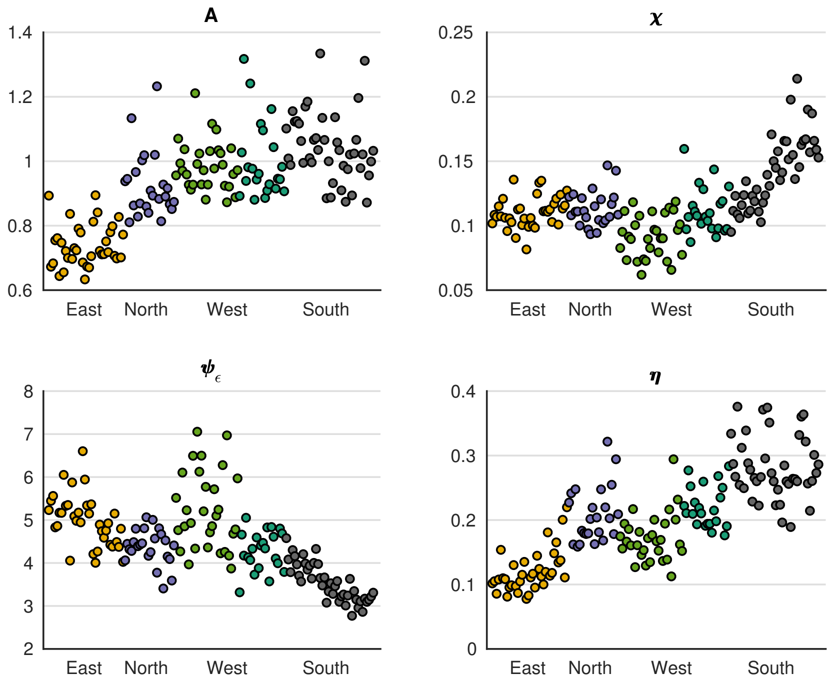
Figure 3.10 displays the structural disparities in 2008-2014. The qualitative profile of the parameters changes only slightly. Productivity and the bargaining power of firms are still lowest in the East and highest in parts of the South. Note that we measure wages in relative terms and do not capture wage growth, but rather the distance to the average wage.[15] While the matching market is still most efficient in the South, matching efficiency has decreased in the East. Also, districts located in East and West exhibit the highest job-match uncertainty, similar to the previous period. While uncertainty is high in the East due to higher separation rates, it is high in the West because of the low efficiency of the matching market. We discuss the changes in more detail next.

Table 3.3 summarizes the structural and empirical (log) changes between 1994-2002 and 2008-2014 for the four regions. There are three particularly pronounced structural changes in East Germany: Productivity rises while matching efficiency decreases, and the bargaining power of firms almost doubles. Tighter labor markets identify the increase in the firms' bargaining power. However, the matching efficiency decreases due to the average fall in job-finding rates. The rise in wages, combined with higher bargaining power, determines a substantial increase in productivity. Furthermore, the large decrease in separation rates leads to a decrease in match uncertainty.

The bargaining power of firms in the West increases due to tighter labor markets. Higher tightness and increasing job-finding rates lead to a small decrease in matching efficiency. Hence, the increase in tightness was relatively large compared to the rise in job-finding rates. Also, lower separation rates lead to lower match uncertainty, on average. A closer analysis of the West reveals heterogeneous developments (lower panel). Districts in North Rhine-Westphalia exhibit the most moderate decreases in separation rates, no changes in job-finding rates, and the smallest increases in tightness, while job-finding rates and tightness improve substantially in the other districts. Thus, districts in North Rhine-Westphalia display the least favorable development in all dimensions captured.

[15] Wages are adjusted to $w_i = \frac{w_i \sum_j e_j}{\sum_j w_j e_j}$.

Fig. 3.10: Structural labor market disparities (2008 – 2014)



Notes: The figure displays the identified values of the four region-specific structural parameters: Productivity (A), matching efficiency (χ), match uncertainty (ψ_ϵ), and bargaining power (η). We determine the parameters within the model given the observed variation in wages, labor market tightness, separation, and job-finding rates. We target the averages of these four variables within 2008 – 2014.

Concerning changes in the northern and southern districts, we observe an increase in firms' bargaining power and productivity. The higher bargaining power can be attributed to an above 50% increase in tightness. The rise in tightness, combined with a surge in job-finding rates, identifies the change in matching efficiency, which is positive or negative depending on the relative increase in both variables. Also, the reduction in separation rates explains less uncertain job-matches. Note that the rise in job-finding rates will also be partially captured in the decrease of job-match uncertainty following from the arguments provided before.

The structural parameters provide a complex picture of the causes of regional unemployment differentials and provide suggestive evidence concerning the origins of their evolution over time. In the empirical section, we documented that primarily districts in East Germany or North Rhine-Westphalia suffer from high unemployment during both periods. The model provides two distinctive explanations for these two regions: While the job-match uncertainty is high in the East and North Rhine-Westphalia, the efficiency of the matching market is high in the former and low in the latter. Hence, unemployment in the East is driven by a high frequency of job losses, while unemployed workers in NRW find jobs less often, due to a low matching efficiency. Overall, East German local labor markets have become much more similar to western and northern districts, while NRW's districts are falling behind.

Table 3.3: Structural and empirical changes from 1994 – 2002 to 2008 – 2014

	Structural changes			
	Prod. (A)	Matching Eff. (χ)	Match uncert. (ψ_ϵ)	Bargaining (η)
East	0.20	−0.24	−0.09	0.76
North	0.04	−0.06	−0.08	0.33
West	0.01	−0.02	−0.07	0.18
South	0.05	0.06	−0.14	0.24
	Empirical changes			
	Wages	Job-find. rates (f)	Separation rates (ξ)	Tightness (θ)
East	0.07	−0.04	−0.67	1.08
North	−0.02	0.03	−0.45	0.63
West	−0.03	0.05	−0.31	0.43
South	0.00	0.13	−0.44	0.53
	Empirical changes: The West			
	Wages	Job-find. rates (f)	Separation rates (ξ)	Tightness (θ)
NRW	−0.03	−0.00	−0.28	0.38
Other	−0.02	0.10	−0.35	0.47

Notes: The table presents the average (log) changes of the identified structural parameters and the four target variables between 1994-2002 and 2008-2014. We group districts into four regions: South: Bavaria and Baden-Wuerttemberg; West: North Rhine-Westphalia, Hesse, Saarland, and Rhineland Palatinate; North: Schleswig-Holstein, Hamburg, Bremen, and Lower-Saxony; East: Berlin, Saxony, Saxony-Anhalt, Thuringia, and Mecklenburg-West Pomerania.

In general, the joint development across EADs is the substantial increase in bargaining power and the decline of job-match uncertainty. Explanations for both changes have been previously brought forward as a reason for the favorable development of the German labor market. We discuss them one at a time next.

First, the substantial decrease in match uncertainty can be attributed, at least partially, to the Hartz labor market reforms, which we did not model this exercise. In the course of these reforms, benefits for the long-term unemployed were cut substantially. In the model, match uncertainty is closely related to the replacement rate b: As the workers' outside option decreases, the surplus of a match rises. This increase of the surplus will either be captured by lower separation rates or by lower values of ψ_ϵ. As we keep the replacement rate constant between periods, the factual decrease of b, in combination with the lower separation rates, will be captured, at least partially, by the evolution of match uncertainty. Thus, a different way to think about the observed changes in job-match uncertainty is that they merely capture the lower UI benefits. The importance of the link between unemployment benefits and separation rates has been investigated by Hartung et al. (2018). They show that the reform induced decrease of the replacement rate for long-term unemployed is responsible for a large portion of the decline of the aggregate separation rate.

To ensure that our indirect approach, where we do not account for the reforms, associates the UI benefits decrease with ψ_ϵ and does not affect the other identified parameters, we provide a sensitivity analysis where we keep the mean of ψ_ϵ constant between observation periods. We back out the implied change in benefits needed to offset the decrease in match uncertainty and redo the exercise using the corresponding replacement rate. We find that, while job-match uncertainty is kept constant and the replacement rate decreases by approximately 10%, the changes of the other structural parameters between the two periods are comparable to the changes under varying uncertainty (Table 3.16). Foremost, the increase in firms' bargaining power is only moderately affected when we account for the decrease in benefits.

Second, the increase in firms' bargaining power might be the result of more flexible collective bargaining agreements. This reasoning was put forward by Dustmann et al. (2014) who find that the share of workers covered by an industry-wide bargaining agreement decreased by almost 20 percentage points to 56% from 1995 to 2008. They argue that more flexible agreements led to an increase in competitiveness through lower wages and increased labor demand in Germany. Boeri et al. (2019) provide evidence that this decline was stronger in East than in West Germany within 1996 and 2013: While the share of workers subject to an industry-wide agreement decreased by 38% in East Germany, the decrease in West Germany amounts to 25%.[16] The more substantial decrease in East Germany mirrors the evolution of the bargaining power uncovered here.

Another prominent view is that the Hartz reforms substantially improved the efficiency of the matching market, which subsequently led to the favorable development in Germany.[17] The heterogeneous developments of regional matching efficiencies determined in the section partially contrast this view: While we document increases in job-finding rates, matching efficiency decreases in some districts as labor market tightness exhibits substantial increases.[18] We confirm the heterogeneous changes at the federal-state level using linear-least-squares regression models, to mitigate concerns

[16]cf., Boeri et al. (2019), Table 3.

[17]For example, see Klinger and Weber (2016) and Launov and Wälde (2016).

[18]We find no increase in aggregate matching efficiency after the mid-2000s, while Klinger and Weber (2016) document an increasing efficiency. Compared to Klinger and Weber (2016), the development of our aggregate job-finding rate differs substantially. These differences might be due to a different definition of worker states, most likely due to our exclusion of marginal employment. We provide further arguments in the appendix (Section 3.B.3). Our estimated changes of region-specific matching efficiency are comparable to Bauer (2013) for West Germany.

that not accounting for local spillovers might bias our results (Stops and Fedorets, 2019). The estimates are provided in the appendix (Table 3.11).

Overall, the decreases in matching efficiency imply that gains in efficiency were not responsible for a significant share of the unemployment decreases up to the year 2014. In our model, an increase in bargaining power alone has an ambiguous implication for separation and unemployment rates: The higher bargaining power leads to two counteracting effects. First, lower wages that decrease the workers' search incentives. Second, more vacancies that lead to a higher job-finding rate. Overall, the impact on the value of unemployment is ambiguous and, given constant match uncertainty, separation rates may increase or decline. The reduction of the replacement rate, however, implies substantially lower separation rates since the lower value of unemployment will lead to a larger match surplus, which in line with the empirical development.

Furthermore, our results provide a slightly different picture compared to the mechanism proposed by Bilal (2019) to account for regional differences in separation rates. In his model, more productive firms sort into locations where the contact rate ($\chi\theta^{-\alpha}$) is higher, which is inversely related to labor market tightness. He argues that more productive firms forego relatively higher profits the longer they do not employ a worker. Therefore, they are willing to pay higher wages in less tight labor markets. He infers that productive (or unproductive) firms co-locate. This co-location leads to an equilibrium where unemployment is high in low productive regions, which also exhibit tight labor markets. In these regions, unemployment is high due to more frequent job-losses. High productive regions exhibit slack labor markets, high wages, and low rates of job loss. Empirically, our findings concerning the relationship between worker flows and productivity are similar: separation rates decrease in local productivity; the job-finding rate is slightly negatively correlated with the separation rate; and the contact rate is positively correlated with the separation rate. However, we find almost no supporting evidence that productivity is higher in slack labor markets.[19] In contrast, productivity slightly increases with labor market tightness.

3.4.3 The optimal policy mix

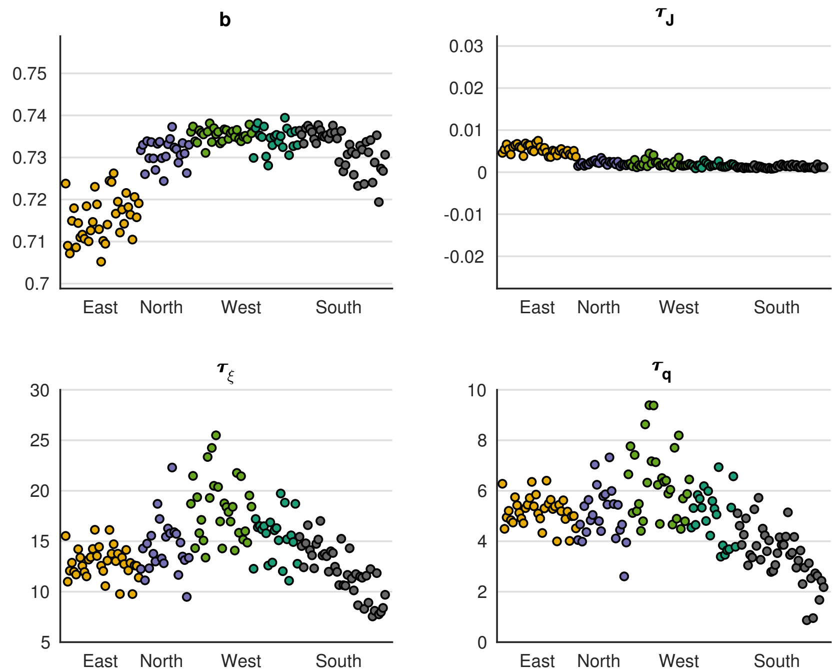
In this section, we quantitatively evaluate the optimal labor market policy mix that achieves the constraint-efficient solution in each employment agency district. We take the region-specific values of A, χ, ψ_ϵ and η as given for both time intervals, which we assume to be invariant to policy changes. These parameters shape the policy variations.

Figure 3.11 displays the regional variation of the policy instruments for the period 1994-2002. Note that we scale the instruments in terms of regional wages. We find that the planner does not fully insure the unemployed (upper left), as their search effort

[19]Note that for some districts in Bavaria where separation rates are relatively high, productivities (wages) are relatively low and these regions exhibit high tightness.

is unobservable. However, unemployment benefits increase substantially compared to the steady-state value of 0.67. Without the layoff taxes and hiring subsidies (lower left and lower right), the higher benefits would lead to a lower search effort. The hiring subsidies counteract this effect since they reward firms if they relieve the social system by employing a worker. With higher subsidies, firms are more likely to post a vacancy, which increases the job-finding rate and workers' search incentives. Similarly, the layoff taxes ensure that firms internalize the social costs of their separation decision, i.e., the benefit payments for the duration of the unemployment spell and the subsidies paid to reemploy the worker.

Fig. 3.11: The optimal policy mix (1994 – 2002)



Notes: The figure displays the optimal policy mix in the planner's economy (1994 – 2002). b represents the replacement rate ($\frac{c_u}{c_e}$) and τ_J the optimal production tax. τ_ξ and τ_q are the layoff taxes and hiring subsidies, respectively.

The observed benefit structure is in line with the intuition provided in Corollary 2. In East Germany and parts of the South, an increase of unemployment insurance benefits has a more substantial impact on the unemployeds' search effort due to higher matching efficiency. Moreover, b decreases in separation rates. In general, our results highlight that unemployment benefits do not need to vary a lot across districts when the other tax instruments are utilized. Also, the production tax is small and does not show significant variations across regions.

The planner instead resorts to hiring and layoff taxes to achieve the constraint-efficient allocation. She subsidizes hirings in Bavaria despite that, according to the Hosios condition, unemployment is below its social optimum, as the firms' bargaining power ($\eta > \alpha$) is too high. Due to relatively low wages, there is excessive vacancy creation, and the posted vacancies crowd each other out. In the absence of the moral-hazard insurance motive, the planner would levy a hiring tax. However, she grants subsidies because of the positive effect of hiring a worker on the social security system. Concerning the other districts, unemployment is above the social optimum, and the planner grants hiring subsidies to foster vacancy creation and restore efficiency. Here, the moral hazard insurance motive of the planner is small as the workers' high bargaining power ensures that wages are relatively high and, thus, the search incentives for workers are high.[20] Hires in eastern districts are not excessively subsidized as the matching efficiency is quite high, which leads to substantial increases in hires with lower subsidies. Layoffs are taxed to cover the costs of the hirings subsidies and to make firms internalize the additional costs for the social system incurred by separating from a worker. These two motives for layoff taxes depend on the duration of the average unemployment spell. Thus, it is more beneficial to keep the marginal worker in NRW in employment because her average unemployment duration is longer, compared to workers in eastern or southern districts.

Table 3.4 displays the policy-induced changes in employment, job-finding rates, search effort, separation rates, and welfare. We compare the allocation in the steady-state to the planner's constraint-efficient allocation. Employment rises in all regions. These employment increases are achieved through fewer separations and more hires, as firms internalize the respective social costs or benefits through the layoff tax and hiring subsidy. Moreover, the workers' search effort increases, despite the higher benefit level. The decreases of separations go through changes in the surplus of a match. The surplus increases in layoff taxes but also depends on all labor market decisions of the workers, which in turn are affected by the replacement rate. For example, layoffs in eastern districts fall the least compared to the other regions, despite high match uncertainty. This outcome indicates that the change of the match-surplus was small. Overall, the optimal policy mix induces a welfare increase of 4% to 5%.

Table 3.4: Steady-state compared to the planner economy (1994 – 2002)

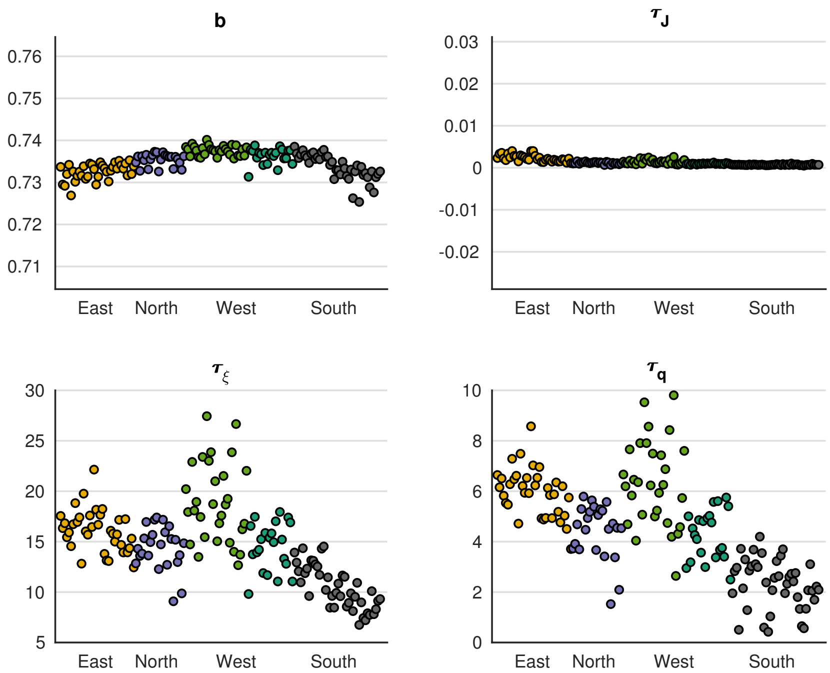
	Employment	ξ	s	f	Welfare
East	0.18	-0.45	0.03	0.62	0.04
North	0.12	-1.27	0.01	0.32	0.04
West	0.12	-1.56	0.01	0.30	0.05
South	0.08	-1.39	0.00	0.20	0.05

Notes: The table presents the (log) change in employment, separation rates (ξ), search effort (s), job-finding rates (f) and welfare under the optimal regional labor market policy mix. The change is measured relative to the individual steady state.

[20]The importance of this motive increases in the bargaining power (Corollary 2).

The regional variation of the optimal policy mix in 2008-2014, which is displayed in Figure 3.12, closely resembles the one during the earlier period. The planner utilizes all tax instruments: benefits increase, layoffs are taxed and hires subsidized. Moreover, the production tax and the replacement rate show little regional variation. We observe small changes for districts located in the North or South. The most pronounced difference can be observed for the EADs in eastern Germany and NRW. Eastern districts now receive an increase in benefits that is comparable to the other EADs, which is due to the decrease in matching efficiency. When efficiency declines, UI benefits have a smaller impact on the unemployeds' search effort. Also, the decrease in the efficiency of the matching market leads to higher layoff taxes and hiring subsidies as the average duration of an unemployment spell is prolonged. In NRW, the planner grants even more substantial hiring subsidies compared to the previous period due to lower matching efficiency. In general, layoff taxes and hiring subsidies decrease with the firms' bargaining power, i.e., decrease as the districts' unemployment level approaches the socially optimal level. When unemployment is above its efficient level, as in some northern and southern regions, the planner grants hiring subsidies because of the moral hazard insurance motive.

Fig. 3.12: The optimal policy mix $(2008-2014)$



Notes: The figure displays the optimal policy mix in the planner's economy $(2008-2014)$. b represents the replacement rate $(\frac{c_u}{c_e})$ and τ_J the optimal production tax. τ_ξ and τ_q are the layoff taxes and hiring subsidies, respectively.

Table 3.5 displays the policy-induced changes for the period 2008-2014. Compared to the steady-state, employment rises in all regions; however, by a more modest amount of 4% to 13%. Again, these increases are the result of the lower separation and higher job-finding rates. Workers' search effort shows only minor changes compared to the steady-state. Reductions in layoffs are more substantial as the uncertainty of a match is lower compared to the earlier period, which strengthens the reaction of separation rates to match-surplus changes. Welfare increases in all regions by 5%, on average.

Table 3.5: Steady-state compared to the planner economy ($2008 - 2014$)

	Employment	ξ	s	f	Welfare
East	0.13	-1.27	0.01	0.39	0.05
North	0.08	-1.63	0.00	0.22	0.05
West	0.09	-1.78	0.00	0.25	0.05
South	0.04	-1.48	-0.01	0.12	0.06

Notes: The table presents the (log) change in employment, separation rates (ξ), search effort (s), job-finding rates (f) and welfare under the optimal regional labor market policy mix. The change is measured relative to the individual steady state.

In general, the planner's allocation leads to heterogeneous changes across districts, while welfare and employment increase substantially. We assess the impact of the optimal policy mix on the welfare and unemployment distribution in the subsequent section.

Sensitivity analysis As discussed in the prior section, the labor market reforms in the mid-2000s led to benefit cuts for the long-term unemployed. Within the model, ψ_ϵ is closely linked to the benefit level. In the appendix, we back out the implied change in b that is necessary to keep the mean of ψ_ϵ constant between our two observation periods. The optimal policy mix shows minor variations to the mix presented in the present section (Figure 3.37). However, the recalibration attenuates the differences of separations, hirings, and employment between the steady-state and the planner's economy in 2008-2014, as the higher value of match uncertainty leads to a lower elasticity of layoffs with respect to benefit changes (Table 3.17).

Furthermore, the identified parameters depend on the initial policy values. We provide a sensitivity analysis in the appendix where we increase the layoff tax by introducing additional severance payments. The results do not change qualitatively, and the quantitative changes between 1994-2002 and 2008-2014 are similar (Table 3.13). However, a larger initial value of τ_ξ increases the identified ψ_ϵ and, thus, attenuates the responses of separation rates in response to policy changes. Hence, the optimal policy mix leads to smaller changes in employment, separations and hirings (Tables 3.14, and 3.15).

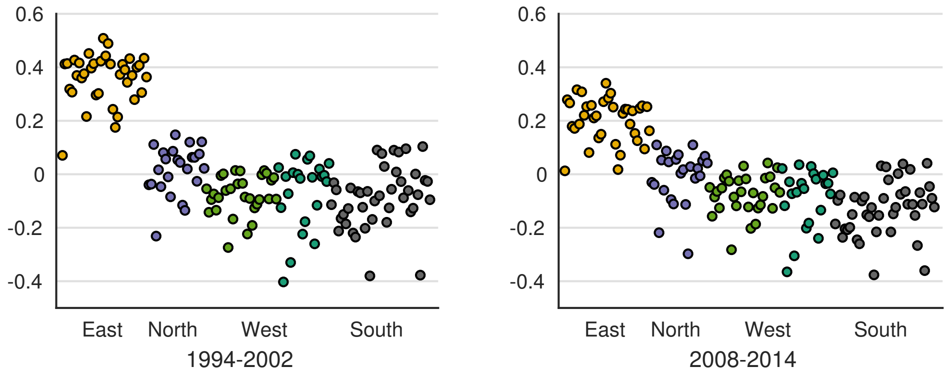
In the previous analysis, vacancy posting costs depend on local productivity as we treat them as region-specific. When we set fixed posting costs, however, the lack of regional variation is primarily reflected in the bargaining power (Figure 3.35). Consequently, when firms face the same costs in every district, the shape of the identified parameters is qualitatively similar, but the bargaining power is rescaled in terms of productivity (Figure 3.36).

3.4.4 Local inefficiencies, welfare, and unemployment differentials

To illustrate the welfare differences across regions, we compare the utility of individuals in the steady-state—absent the optimal policy mix. Furthermore, we assess the impact of the optimal policy mix on the welfare and unemployment rate distribution.

We visualize the welfare distribution by comparing each district to the benchmark region. Figure 3.13 shows the (log) utility difference, which we interpret as a worker's willingness to pay to relocate to the benchmark district. In 1994-2002 (left panel), the dispersion of these welfare gains is substantial: Most workers from West German districts would need to be compensated to relocate to the benchmark district, while workers from eastern EADs would gain at least 20%. We find a more equal distribution for the period 2008-2014 (right panel), but the utility differences are still substantial.

Fig. 3.13: Comparison of steady-state welfare disparities



Notes: The figure displays the (log) welfare difference between the individual steady-states and the benchmark economy. The benchmark district represents the average employment agency district during $1994-2002$. A value larger than zero indicates a welfare gain.

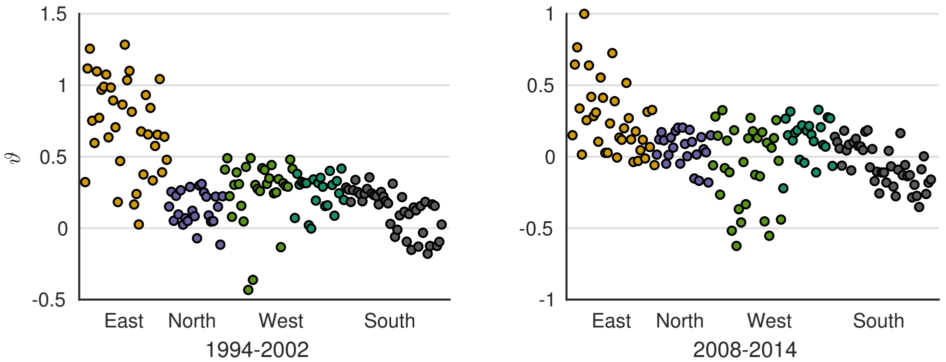
Concerning the impact of the optimal policy mix and unemployment and welfare differentials between districts, we find that the optimal policy mix increases the differences. We observe an increase in the coefficient of variation of welfare from 1.1 to 1.5 during the earlier and 1.2 to 2.0 in the later period. The unemployment rates' coefficient of variation rises from 0.4 (0.4) in the steady-state to 0.8 (0.5) in the planner's economy in 1994-2002 (2008-2014). Thus, the optimal policy mix does not mitigate welfare or unemployment rate differentials across EADs in Germany.

This result, however, is not generic but contingent on the structural disparities between regions. Figure displays the absolute deviation of the local unemployment rates from the average unemployment rate in the steady-state relative to the planner's economy

$$\vartheta = \left| \left(\frac{u_j^p}{\bar{u}^p} - 1 \right) \right| - \left| \left(\frac{u_j^{ss}}{\bar{u}^{ss}} - 1 \right) \right|, \tag{3.26}$$

where "p" indicates the planner's allocation and a bar indicates the average across districts, i.e., we measure if the distance to the average unemployment rate has increased or decreased. A value larger than zero indicates a percentage point increase in the distance to the average unemployment rate. Note that an increase in distance captures both relatively lower or higher local unemployment rates. The results show that in the constraint-efficient allocation for 1994-2002, almost all districts are farther away from the average unemployment rate compared to the steady-state.

Fig. 3.14: Deviations of the distance to avg. unemployment rate



Notes: The figure displays the percentage point deviation welfare difference between the individual steady-states and the benchmark economy. A value larger than zero indicates the percentage point increase in the distance to the average unemployment rate.

In 2008-2014, the change in distance is substantially lower compared to the previous period, which is also indicated by the smaller increase in the coefficient of variation. Furthermore, the coefficient of variation in West Germany even reduces from 0.36 to 0.26. Therefore, the optimal policy mix does not *necessarily* mitigate welfare or unemployment rate differentials across EADs.

3.5 Mismatch unemployment

The prior analysis abstracted from mobility between regions and focused on the local effect of optimal labor market policies. In this section, we provide an upper bound estimate for the potential increase in aggregate employment when mobility frictions are absent. We measure the efficiency gain—the reduction of aggregate unemployment—of eliminating the mismatch across distinct regional labor markets. Mismatch might arise if many unemployed seek jobs in markets with low vacancies and frictions to mobility impede an adjustment. This misallocation then leads to an inefficient level of aggregate unemployment. While we do not analyze the possible sources of mismatch, the measure can be interpreted as an upper bound for the importance of frictional worker mobility across sectors for aggregate unemployment.[21]

We document mismatch during two periods: 1980-2014 as well as 1994-2014. The analysis of the more extended period focuses on the development in West Germany. For 1994-2014, we investigate mismatch in East and West Germany. The subsequent sections introduce and discuss the measurement approach and present the results.

3.5.1 Measuring mismatch

We utilize the framework of Sahin et al. (2014) to measure the degree of mismatch across regional labor markets in Germany. The strategy is as follows. We measure the effect of mismatch across local labor markets on aggregate unemployment comparing an optimal to the actual distribution of the unemployed. The optimal allocation is derived within a dynamic search and matching model, where a planner redistributes unemployed at no cost to maximize employment. The economy consists of distinct labor markets, and each market is subject to frictions captured by a matching function. Based on the planner's optimal allocation rule, we then estimate the fraction of hires lost due to mismatch and recover counterfactual unemployment rates.

Model The economy consists of a number of, non-overlapping, sectors $i \in I$ of whom each is subject to frictions. We assume that these frictions are captured by a Cobb-Douglas matching function of the form: $h_{it} = \Phi_t \chi_{it} v_{it}^{\gamma} u_{it}^{1-\gamma}$. Hires, unemployment and vacancies at time t in sector i are given by h_{it}, u_{it} and v_{it}, respectively. Unemployed workers can only search in one sector. The vacancy elasticity of the matching function γ is assumed to be constant across time and sectors. Φ_t represents aggregate and χ_{it} sector specific matching efficiency. Individuals are risk neutral and are either employed or unemployed ($\sum_{i=1}^{I} e_{it}+u_{it} = 1$). There is no on the job search, matches are destroyed at the exogenously give rate ξ_t and vacancies arise exogenously.

In the benchmark environment, the planner minimizes unemployment, given sector-specific matching efficiencies and vacancy shares. Moreover, there are no reallocation costs. A more detailed model description, as well as the derivation of the allocation rule, are relegated to the appendix (Section 3.D).[22] The planner's optimal allocation rule across sectors can be written as

$$\chi_{1t} \left(\frac{v_{1t}}{u_{1t}^{*}} \right)^{\gamma} = \cdots = \chi_{It} \left(\frac{v_{It}}{u_{It}^{*}} \right)^{\gamma}, \tag{3.27}$$

[21] Herz and van Rens (2020) estimate the underlying frictions of mismatch. Their results point to wage frictions as the main cause of mismatch. Marinescu and Rathelot (2018) present evidence on mismatch for the U.S. taking into account the geography of job search. They find that the departure from the assumption of distinct markets does not lead to a large increase in mismatch.

[22] An extension of the model, where the planner takes into to account differences in separation rates and productivity, is presented in appendix (Section 3.D).

where u_i^* represents the optimal number of unemployed that should be searching in sector i. The allocation rule implies that (3.28) has to hold for each sector pair j and k.

$$\frac{v_j}{u_j^*} = \left(\frac{\chi_k}{\chi_j}\right)^{\frac{1}{\gamma}} \frac{v_k}{u_k^*} \tag{3.28}$$

In short, the optimal allocation equalizes matching efficiency weighted vacancy to unemployment ratios across sectors. The planner re-allocates unemployed to sectors with higher matching efficiencies and labor market tightness.

Mismatch Index We utilize this optimality condition to construct a measure for hires lost due to mismatch. Specifically, we measure the impact of the misallocation of unemployed on hires by comparing the planner's optimal hires to the number of actual hires. The mismatch index (3.29) then captures the fraction of hires lost due to misallocation.

$$M_{\chi,t} = 1 - \frac{h_t}{h_t^*}, \tag{3.29}$$

The aggregate number of hirings h_t in period t can be recovered by summing h_{it} across all sectors. The corresponding expression is

$$h_t = \Phi_t v_t^{\gamma} u_t^{1-\gamma} \left[\sum_{i=1}^{I} \chi_{it} \left(\frac{v_{it}}{v_t}\right)^{\gamma} \left(\frac{u_{it}}{u_t}\right)^{1-\gamma}\right].$$

Optimal hirings h_t^* are based on the planner's allocation of unemployed across sectors (u_{it}^*), taking vacancies and sectoral matching efficiencies as given

$$h_t^* = \Phi_t v_t^{\gamma} u_t^{1-\gamma} \left[\sum_{i=1}^{I} \chi_{it} \left(\frac{v_{it}}{v_t}\right)^{\gamma} \left(\frac{u_{it}^*}{u_t}\right)^{1-\gamma}\right]$$

Substituting for the planner's allocation rule for each district pair (3.28) the equation simplifies to $h_t^* = \bar{\chi}_t \Phi_t v_t^{\gamma} u_t^{1-\gamma}$. Here, sector specific matching efficiencies are weighted by the respective vacancy shares and aggregated to a scaling factor $\bar{\chi}_t = [\sum_i^I \chi_{it}^{\frac{1}{\gamma}} \left(\frac{v_{it}}{v_t}\right)]^{\gamma}$, which is constant across sectors. The mismatch index then simplifies to

$$M_{\chi,t} = 1 - \sum_{i=1}^{I} \left(\frac{\chi_{it}}{\bar{\chi}_t}\right) \left(\frac{v_{it}}{v_t}\right)^{\gamma} \left(\frac{u_{it}}{u_t}\right)^{1-\gamma}. \tag{3.30}$$

As demonstrated by Sahin et al. (2014), the index is invariant to aggregate shocks and lies within zero and one. In addition, it is increasing in the level of disaggregation – i.e., indices with a different disaggregation level are not directly comparable. Moreover, the index can be calculated solely on the basis of the empirical distribution of unemployed, vacancies and matching efficiencies.

Counterfactual unemployment The fraction of hires lost has direct implications for aggregate unemployment. We derive the planner's optimal job-finding rate and subsequently construct a counterfactual unemployment rate − absent any mismatch. The actual job-finding (f_t) is given by:

$$f_t = \frac{h_t}{u_t} \quad = \quad (1 - \mathcal{M}_{\chi,t})\bar{\chi}_t \Phi_t \left(\frac{v_t}{u_t}\right)^{\gamma},$$

where we exploit that equation (3.30) is equivalent to $h_t = (1 - M_{\chi,t})h_t^*$. Mismatch lowers the job-finding rate as the planner's distribution is more efficient. Based on the planner's allocation of unemployment across sectors, the optimal job-finding rate is

$$f_t^* = \chi_t \Phi_t \left(\frac{v_t}{u_t^*}\right)^{\gamma} = f_t \frac{1}{1 - \mathcal{M}_{\chi,t}} \left(\frac{u_t}{u_t^*}\right)^{\gamma}.$$

Hence, we can recover the optimal rate using the actual rate and the mismatch index. Note that the optimal job-finding rate is higher than the actual rate because of the efficient allocation of unemployed and because the optimal aggregate unemployment rate is most likely lower ($u_t^* < u_t$). We then derive the sequence for the counterfactual unemployment rate using the flow equation (3.31), f_t^* and an initial value u_0^*.

$$u_{t+1}^* = \xi_t + (1 - \xi_t - f_t^*)u_t^* \tag{3.31}$$

The separation rate is taken as exogenous, similar to the model. We select $u_0^* = \frac{\xi_0}{\xi_0 + f_0^*}$, where $f_0^* = f_t \frac{1}{1 - \mathcal{M}_{\chi,t}}$.

The share of unemployment that is caused by mismatch is measured as the difference between actual and optimal unemployment rate. Note that mismatch not only directly affects unemployment through lower hires but also through lower unemployment u_t^*, which leads to an even higher job-finding rate.

3.5.2 Estimates of the matching parameters

To derive the mismatch indices, we estimate the vacancy share of the matching function (γ) as well as the sectoral matching efficiency parameter (χ_i). Geographical labor markets are treated as non-overlapping. Hence, we exclude inter-district job-transitions. The matching process in each market can be described by a reduced-form Cobb-Douglas matching function with constant returns to scale. We estimate model (3.32)[23] using linear least squares at the aggregate level to recover the vacancy share of the matching

[23] We chose the following functional form $m_t = \Phi_t V_t^{\gamma} U_t^{1-\gamma}$, where Φ_t represents aggregate matching efficiency (Barnichon and Figura, 2015). The function relates total hires (m_t) to the stock of vacancies (V_t) and unemployed (U_t). Model (3.32) is a transformation where divide by the stock of unemployed, take logs, and model aggregate matching efficiency using a time trend.

function (γ).

$$\ln(f_t) = \ln(\Phi_0) + D_t + D_m + \gamma \ln(\theta_t) + \epsilon_t \tag{3.32}$$

Labor market tightness is denoted by $\theta_t = \frac{v_t}{u_t}$ and f_t is the job-finding rate. To account for the endogeneity problem of the matching function, we model the time-varying component of the aggregate matching efficiency using time trends and structural break dummies similar to Borowczyk-Martins et al. (2013) and Sedláček (2016). In the baseline specification, D_t includes a quartic time trend. In a sensitivity analysis, we include dummies for the implementation of the Hartz reforms, the dot-com bubble, the great recession as well as the reunification of Germany in 1991. Moreover, we include monthly time dummy variables (D_m) to account for seasonal fluctuations.

We estimate model (3.32) for various time periods. The results are presented in Table 3.9 of the appendix. Our estimates are in line with prior results (e.g., Burda and Wyplosz, 1994 and Hertweck and Sigrist, 2015). Our preferred estimate is $\gamma = 0.25$, which we use to calculate our mismatch indices in the subsequent sections.[24]

On more disaggregated sectoral levels, we estimate panel regressions of the form

$$\ln(f_{it}) = D_t + D_m + \ln(\chi_i) + \gamma \ln(\theta_{it}) + \epsilon_{it}, \tag{3.33}$$

where χ_i is the time-invariant district-specific matching efficiency. While we treat matching efficiency as time-invariant in our benchmark analysis, we provide evidence on the time variation of χ_i in the appendix (Table 3.11).

3.5.3 Regional mismatch in Germany

We estimate regional mismatch at the EAD level and present two mismatch measures: The M index, which is driven by the distribution of unemployment and vacancies, and the M_χ index, which additionally accounts for the constant heterogeneity in matching efficiency.[25]

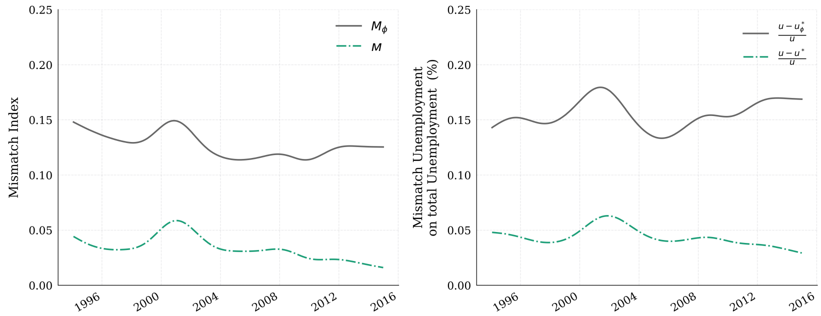
The mismatch indices and the contribution of mismatch to aggregate unemployment in 1994-2014 are presented in Figure 3.15. The M_t index shows a declining trend from 2001 onward. Moreover, the estimated degree of mismatch is low as the fraction of hires lost due to misallocation is below 6%. The $M_{\chi,t}$ index indicates that local labor market frictions − captured by matching efficiency − are an important source of heterogeneity. Aggregate hires can be increased by, on average, 12% more, when the planner takes differences in matching efficiency into account. The contribution to aggregate unemployment (right panel) shows a u-shaped pattern over the period

[24] Note that, based on the construction of the data, hires in period t are workers who transition from unemployment to employment in between the months t and $t + 1$.

[25] Without heterogeneity in matching efficiency the index reduces to $M = 1 - \sum_{i=1}^{I} \left(\frac{v_{it}}{v_t}\right)^{\gamma} \left(\frac{u_{it}}{u_t}\right)^{1-\gamma}$ (c.f. Sahin et al., 2014).

2000-2005 but stabilizes at around 15% from 2012 onward. This development suggests no role for geographical mismatch in the increase in unemployment during the mid-2000s. The increase of the contribution to aggregate unemployment from 2005 onward

Fig. 3.15: Regional mismatch across employment agency districts



Notes: The figure presents the mismatch indices M and M_χ as well as the corresponding contribution of mismatch to aggregate unemployment. The unit of analysis are 154 employment agency districts. All series display the HP-filtered trend (λ =100,000).

is attributable to the increase of mismatch (left panel) and to the overall decrease of separations, which was documented in a previous section. More formally, the flow equation of unemployment can be rearranged to

$$\Delta u_{t+1} - \Delta u^*_{t+1} = -s_t(u_t - u^*_t) + f^*_t u^*_t - f_t u_t. \tag{3.34}$$

Hence, the change in differences between actual and counterfactual unemployment rate depends negatively on the separation rate if $u_t > u^*_t$. Therefore, the decrease in separation rates has increased the importance of mismatch.

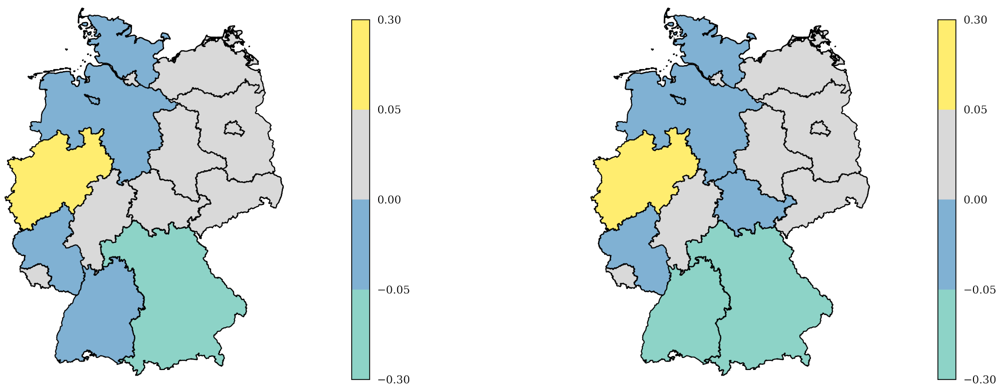
Visualizing the share of unemployed, which has to be reallocated to eliminate mismatch unemployment, is instructive to understand the direction and size of the changes the planner undertakes to eliminate mismatch. We chose federal states to illustrate regional differences. Mismatch at the federal state level is lower but exhibits similar movements over time compared to mismatch across employment agency districts (Figure 3.28). Figure 3.16 displays the fraction of misallocated unemployed by state

$$\frac{u^M_{jt}}{u_t} = \frac{u_{it}}{u_t} - \left(\frac{\chi_i}{\bar{\chi}}\right)^{\frac{1}{\gamma}} \frac{v_{it}}{v_t}. \tag{3.35}$$

A negative (positive) value indicates that actual unemployment is below (above) its optimal value. In 1994 (left panel), the planner reallocates a total of 37% of the unemployed to a different state. One-third of this fraction lives in NRW, while two-thirds come from the six East German states. Almost all (81%) are relocated to Bavaria. In 2014 (right panel), the planner moves 30% of the unemployed to a different state. Of this 30%, 60% (25%) are from NRW (eastern states) and 80% are moved to

Bavaria. These results are congruent with the change in the unemployment distribution presented in prior sections. During the first years of our sample, unemployment rates are highest in East Germany, while during the later years, districts in North-Rhine Westphalia (NRW) exhibit similar unemployment but substantially lower job-finding rates compared to the eastern states.

Fig. 3.16: The deviation of actual from optimal unemployment at the federal state level



(a) 1994 (b) 2014

Notes: This figure displays the redistribution of unemployed across federal states based on the allocation rule (3.35). The planner takes the heterogeneity in matching efficiency into account.

Bavaria is the central receiving region due to its above-average job-finding rates. Hence, to equalize matching efficiency weighted unemployment-vacancy ratios, the number of unemployed residing in the south has to rise.

Mismatch across East German EADs is low: The optimal allocation of the planner would increase hires by below 5%. Here, mismatch is low as unemployment rates are high in all districts and decrease homogeneously up to 2014. Moreover, labor market tightness does not vary a lot. Similar unemployment rates and labor market tightness leave no room for the planner to minimize aggregate unemployment in East Germany (Figure 3.27).

For West Germany, we extend the index backward to 1980. The index unadjusted for heterogeneity (M_t) decreases throughout the period and lies below 5%. Accounting for differences in matching efficiency ($M_{\chi,t}$) leads to a considerable rise in the level of mismatch: Redistributing unemployed based on the planner's rule leads to an average increase in hires of 15%. Moreover, mismatch in West Germany is similar in level and cyclicality during 1994-2014 compared to Germany (Figure 3.30).

We provide a sensitivity analysis in the appendix. We compare versions of the mismatch indices estimated using SIAB data to indices based on official unemployment data (Section 3.B.4). We conclude that the SIAB data provides a good approximation of the actual distribution of unemployment across regions. Section 3.B.5 of the appendix offers a supplementary analysis where differences in separation rates and productivity, as well as time variation in matching efficiency, are taken into account. Time variation in matching efficiency leads to minor level changes, while heterogeneity in separations and productivity decreases mismatch (Figure 3.29).

Note that the results should be interpreted as an upper bound, as the planner reallocates the unemployed at no cost. The presence of regional preferences or moving costs, for example, might substantially change the implied reallocation pattern as both are essential frictions of inter-regional migration (e.g., Heise and Porzio (2019) and Schmutz and Sidibé (2019)). Since at the lowest spatial level mismatch accounts for only 15% of aggregate unemployment, which implies that removing mismatch leads to a two percentage point decrease of the unemployment rate, and the relocation costs are potentially extensive, reducing unemployment through local policies seems to be the more promising approach.

3.6 Conclusion

In this paper, we analyze regional disparities in unemployment rates and the underlying labor market flows in Germany. Our analysis is based on administrative social-security records and newly compiled vacancy data on the employment agency district level. We present new evidence on the importance of worker flows for regional unemployment differentials. A decomposition of the unemployment rate variation across districts shows that separation rates explain 70% of the regional differences during 1994-2004. Over time, however, their importance in the determination of unemployment rate differentials diminished, leaving job-finding and separation rates responsible for 54% and 41% of the variation in 2005-2014, respectively. This finding suggests that when designing policies to mitigate local unemployment, the government should take into account both margins because the transmission of policies goes through the flow rates (Jung and Kuhn, 2014, and Hartung et al., 2018).

We then utilize a multi-island model with search and matching frictions to identify structural differences between regions. Given these structural disparities, we evaluate how an optimal policy mix that jointly balances local search externalities and externalities caused by the provision of unemployment insurance benefits affects welfare and unemployment disparities. The optimal region-specific policy mix leads to significant increases in welfare and employment compared to the steady-state. This result suggests that local labor market policies are an effective means to increase overall welfare.

Last, we evaluate the effect of mismatch between vacancies and unemployed in local labor markets on aggregate unemployment. Quantitatively, we find regional mismatch to be responsible for, on average, 15% of unemployment. Therefore, eliminating mismatch leads to lower decreases in unemployment compared to the average decline under the optimal region-specific policy mix. Taking into account potential relocation costs provides further arguments in favor of local policies compared to reallocating workers.

Appendix

3.A Data

3.A.1 Individual employment histories

A contribution limit to the social security system leads to top coding of wages. Following Gartner (2005), we impute all censored wages at the social security threshold in each year.

We provide additional evidence on the regional development across occupations in the appendix. Our data covers 25 occupations for the years $1980 - 2013$. Information on the occupation in which a person is working is available for every employment spell. Occupations are classified based on the "Klassifikation der Berufe 1988", which was not subject to changes during the observation period.[26] If a worker becomes unemployed, we assume that she is searching for a job in the occupation of her last employment. To estimate sector-specific matching efficiencies, we need detailed data on monthly unemployment in and outflows for each sector. Therefore, we drop occupations where we either have to many missing data points or observe only a low number of monthly transitions. In total, we drop 8 of 33 occupations from the sample. In the empirical analysis, we drop occupations with less than 100 monthly transitions out of unemployment or more than five missing observations of unemployment flow or stock variables. Specifically, we exclude the following occupations: Agriculture and fishery (1-6) miners (7-9), ceramists (12), paper manufacturer (16), printer (17), wood conditioner (18) and leather manufacturers (37). The occupation unskilled labour (53) is disregarded due to missing observations for East Germany.

3.A.2 Vacancies

This section describes the frequencies of the vacancy data and the regional adjustment procedure. We digitize monthly or quarterly data on vacancy stocks for employment agency districts and occupational classifications from official (archived) publications of the federal employment agency. From 2000 onward, we obtain the vacancy data di-

[26]The first version was developed in the 1960s. It was modified in 1975, as well as in 1988. Importantly, the 2-digit level was not subject to any change during the observation period; only 4-digit and lower categories were modified. The KldB 88 was replaced in 2011 by the KldB 2010, adjusting the classification scheme to international standards. Employment notifications after 30.11.2011 follow the new classification of 2010 but are recoded, and inaccuracies may occur.

rectly from the FEA. Table 3.6 shows the frequencies and sources of the vacancy data. In addition to registered vacancies, we data on unemployment (rates and stocks) when possible. The data on unemployment is utilized to demonstrate the comparability of SIAB to official data when computing the mismatch index. In general, we interpolate quarterly data to a monthly frequency. Moreover, we linearly interpolate missing values.

Vacancy data for employment agency districts are digitized from monthly ANBA publications for the years 1976 – 2000. Data for East Germany is available from 1991 onward. To obtain regional data comparable to the SIAB data, we correct the vacancy stocks for territorial changes. Specifically, we adjust for changes in district territory within the years 2012 and 2013. We provide a discussion of the adjustment procedure discussed in a subsequent section.

Table 3.6: Vacancy data: Sources and frequencies

Sector	Time	Frequency	Source	Other data
	West Germany			
Occupation	1972 – 1981	m	ANBA_a	U
	1982 – 1997	q	ANBA_a	U
	1998 – 2000	m	ANBA_m	
	2000 – 2013	m	FEA	
EAD	1976 – 1999	m	ANBA_m	U, u
	2000 – 2014	m	FEA	
	East Germany			
Occupation	1992 – 1997	q	ANBA_a	
	1998 – 2000	m	ANBA_m	
	2000 – 2013	m	FEA	
EAD	1991 – 1999	m	ANBA_m	U, u
	2000 – 2014	m	FEA	

Notes: ANBA_a (ANBA_m) indicates that the report was published annually (monthly). We list the total time period but use data from 1980 onward. "Other data indicates the other time series digitized: U number of unemployed; u unemployment rate.

For occupations, we obtain a monthly vacancy series from annual ANBA publications for the years $1972 - 1981$. Within $1982 - 1996$, we interpolate the quarterly data collected from annual ANBA publications to a monthly frequency. For the period from 1997 to 2000, we digitize registered vacancies from monthly ANBA publications. East German vacancy data by occupation are available from 1992 onward. For occupations, the vacancies include seasonal jobs up to 2005. These seasonal jobs are predominantly in agricultural occupations. Here, we observe a drop in the average number of vacancies from almost 50000 in 2004 to 4000 in 2005. Therefore, we exclude the agricultural sector from our sample. The occupational classification scheme corresponds to one of the SIAB data. The vacancy data digitized from the ANBA is reported for 33, 2-digit occupational codes, which were not subject to changes during the observational period. During some years, the data is only available at a quarterly frequency. Vacancy data received from the FEA covers the years $2000 - 2013$ and is aggregated in accordance to the classification of the ANBA data.

3.A.3 Re-confinement of the employment agency districts

The SIAB data utilized in this article is based on territorial borders of 2014. Between 2012 and 2013, the territory of the employment agency districts was partially re-confined. In particular, the total number of EADs reduces from 176 to 156.[27]

Our vacancy data before 2000 comes from historical annual reports and employment agency districts are based on non-fictional territorial borders in each year. Thus, we cannot recalculate the districts' vacancy stocks. Also, vacancy data received from the FEA for the years $2000 - 2007$ is based on territorial borders before the re-confinement.

To match the regional unit of the individual employment histories, we approximate the vacancy distribution for all years before 2007. The approximation is based on overlapping vacancy data for old and new territorial classifications. Specifically, we obtain vacancy data for the old and new territorial borders for each employment agency district. This data covers the period from January 2007 to June 2012. Hence, we can infer the number of vacancies gained (lost) due to re-confinement.

Our adjustment procedure is as follows: Assume there are two districts X and Y, where X received parts of the territory of Y during the re-confinement in 2012. Let the districts based on the new territory be denoted by X' and Y'. The FEA provides us with non-fictional (X, Y) and fictional (X', Y') EAD vacancy data for the period 2007 to 2012, where fictional data is based on the territorial borders after 2012. We use this overlapping period to identify the average share of vacancies lost in Y ($\bar{\alpha} = \frac{1}{T}\sum_t^T \alpha_t$)

[27] The reassignment happened in three steps on 1.7.2012, 1.10.2012, and 1.1.2013.

solving the system of equations for each t:

$$\begin{aligned} V_{X',t} &= V_{X,t} + \alpha_t V_{Y,t} \\ V_{Y',t} &= V_{Y,t} - \alpha_t V_{Y,t}. \end{aligned}$$

The average shares are then used to adjust the number of vacancies in both districts for all months before 2007 (e.g., $V_{Y',t} = V_{Y,t} - \bar{\alpha} V_{Y,t}$). If an exchange of territory happened between only two districts, we can infer the exact number of vacancies lost (gained) for the giving (receiving) district. However, in a situation where multiple districts exchange territory, we estimate the mean share.

Concerning East Germany, we digitize ANBA publications from January 1991 up to December 1999 and append them with data from the FEA. Again, we have to correct for territorial changes using data for the territorial borders in 2014 available for the period 2007 to 2014 and vacancy data for the old territorial boundaries spanning 2000 to 2012. Three employment agency districts are newly formed: Greifswald, Freiberg, and Bernburg. Two EADs are discontinued because of a merger (Wittenberg and Gera). Overall, all average shares of vacancies lost (gained) can be calculated. Nine districts were not subject to any territorial changes.

The number of vacancies of 68 out of the 156 districts does not need to be approximated because the district's borders do not change, or a new district is an exact combination of two or more old EADs. The following subsection lists the EADs where the number of vacancies was approximated and describes the identification of the average shares. We compare our corrected time series to the official vacancy series during the overlapping period 2007 to 2012 in Section 3.A.3. The mean percentage derivation from the correct number of vacancies is 3% in West and 6% in East Germany.

Territorial changes

North Rhine-Westphalia In North Rhine-Westphalia, all shares of vacancies lost (gained) can be calculated. The following districts were subject to territorial changes: On the 1.7.12, Hamm receives territory from Dortmund, and Recklinghausen receives territory from Gelsenkirchen. On the 1.10.12, Aachen-Düren is newly formed out of Aachen and Düren. Meschede-Soest is formed out of Meschede and Soest, and Ahlen-Münster is formed out of Ahlen and Münster. On the 1.1.13, Mettmann is newly formed out of parts of the districts Düsseldorf and Wuppertal. Solingen-Wuppertal is formed out of Solingen as well as parts of Wuppertal.

Baden-Wuerttemberg In Baden-Wuerttemberg, all shares of vacancies lost (gained) can be calculated. The following districts were subject to territorial changes: On the 1.7.12, Heidelberg receives territory from Mannheim. On the 1.10.12 Ulm receives territory from Ravensburg. Karlsruhe-Rastatt is formed out of Karlsruhe and Rastatt. Konstanz-Ravensburg is formed out of Konstanz and parts of Ravensburg. Nagold-Porfzheim is formed out of Nagold and Pforzheim. Schwäbisch Hall – Tauberbischofsheim is created out of Schwäbisch Hall and Tauberbischofsheim. Rottweil – Villingen-Schwenningen is formed out of Rottweil and Villingen-Schwenningen.

Hesse In Hesse, all shares of vacancies lost (gained) can be calculated. The following districts were subject to territorial changes: On the 1.10.12, Bad Hersfeld-Fulda is newly formed out of the complete territory of Fulda as well as parts of Bad-Hersfeld. Kassel receives territory from Bad-Hersfeld. Korbach receives territory from Marburg and Kassel. Marburg receives territory from Wetzlar. Limburg-Wetzlar is newly formed out of Wetzlar and parts of Marburg. On the 1.1.13 Offenbach receives territory from Frankfurt. Gießen receives territory from Frankfurt. Bad Homburg is newly formed out of parts of Frankfurt and Darmstadt.

Rhineland-Palatinate and Saarland In Rhineland-Palatinate and Saarland, all shares of vacancies lost (gained) can be calculated. The following districts were subject to territorial changes: On the 1.10.12, Saarland is newly formed out of Saarlouis, Saarbrücken, and Neunkirchen. Bad-Kreuznach receives territory from Koblenz. Trier receives territory from Mayen. Koblenz-Mayen is newly formed out of parts of Koblenz and Mayen. Landau receives territory from Ludwigshafen.

Schleswig-Holstein and Hamburg On the 1.10.12, Bad Oldesloe receives parts from Elmshorn and Lübeck. Elmshorn receives parts from Lübeck and Neumünster. Flensburg receives parts from Heide and Neumünster. Heide receives parts from Elmshorn. Kiel receives parts from Neumünster. Neumünster receives parts from Kiel and Flensburg. Hamburg does not change.

Due to multiple territorial trades, we estimate the mean share. Since Lübeck does not receive any additional territory and, on average, 99% of vacancies stay in Lübeck we assume, that flows to Elmshorn and Bad-Oldesloe are zero. This assumption enables us to identify the flow from Elmshorn to Bad-Oldesloe.

We estimate the following system of equations for districts: Kiel (K,K'), Flensburg (F,F'), Neumünster (N,N'), Heide (H,H') and Elmshorn (E,E').

$$\begin{aligned}
V_{K'} &= (1-\alpha)V_K + \beta V_N \\
V_{E'} &= (1-\tau)V_E + \psi V_N \\
V_{H'} &= (1-\phi)V_H + \zeta V_N + \tau V_E \\
V_{F'} &= (1-\delta)V_F + \gamma V_N + \phi V_H \\
V_{N'} &= (1-\beta-\gamma-\zeta-\psi)V_N + \alpha V_K + \delta V_F
\end{aligned}$$

The coefficients are estimated using a seemingly unrelated regression with the following restrictions: $\tau = 0, \delta = 0, \gamma = 0, \beta = 0$.

Lower Saxony and Bremen

1. Nordhorn and Emden-Leer
   Nordhorn only receives territory. Emden is merged with parts of Leer to form Emden-Leer. Leer is discontinued, and the area is divided between the districts of Nordhorn and Emden-Leer. Hence, the inflow from Leer to Nordhorn as well es Emden-Leer can be calculated.

2. Lüneburg-Uelzen and Celle
   Lüneburg is merged with parts of Uelzen to form Lüneburg-Uelzen. Uelzen is discontinued, and the area is divided between the districts of Lüneburg-Uelzen and Celle. Celle receives territory from Uelzen and loses territory to Hannover. The shares of territory lost (gained) can be calculated

3. Bremen-Bremerhaven, Stade and Nienburg-Verden
   Bremen is merged with parts of Bremerhaven to form Bremen-Bremerhaven. Hence, the share of vacancies lost from Bremerhaven to Bremen-Bremerhaven can be calculated. As Bremerhafen is split up and the area is divided between Bremen-Bremerhaven and Stade, the latter share can be calculated. Stade receives parts of Bremerhaven and Verden. The share of Verden to Stade can be calculated using prior results. Verden is discontinued and merged with parts of Nienburg to Nienburg-Verden. The share of vacancies lost from Nienburg to Nienburg-Verden and Nienburg to Hannover can be calculated.

4. Helmstedt, Hannover, and Hameln
   Helmstedt receives parts of Braunschweig and to Helmstedt. Hence, the share of vacancies transferred from Braunschweig to Helmstedt can be calculated. Hannover receives territory from Celle, Nienburg, and Hameln. As the shares of Celle and Nienburg to Hannover have been calculated, the share of Hameln to Hannover follows as residual. Using the last calculated share number of vacancies transferred from Hildesheim to Hameln follows as a residual.

5. Göttingen, Braunschweig-Goslar and Hildesheim

   (a) Braunschweig-Goslar (BSG) is newly formed out of parts of Braunschweig (BS), without territory given to Helmstedt and Goslar (G).

   (b) Göttingen (GOE') consists of Göttingen and parts of Goslar and Hildesheim (H), without territory given to Hameln and Göttingen.

(c) Hildesheim (H') consists of parts of Braunschweig and parts of Hildesheim.

The shares of vacancies lost (gained) cannot be calculated directly. Hence, we need to solve the following system of equations:

$$\begin{aligned} V_{H'} &= (1-\alpha)V_{BS} + (1-\gamma)V_H \\ V_{BSG} &= \alpha V_{BS} + \beta V_G \\ V_{GOE'} &= V_{GOE} + (1-\beta)V_G + \gamma V_H \end{aligned}$$

We apply the restriction $\gamma = 0$ because Hildesheim loses only parts of the small municipality Northeim (29000 inhabitants) to Göttingen. This restriction enables us to calculate the other shares each month.

Bavaria On the 1.10.12, Bamberg-Coburg is newly formed out of Bamberg and Coburg. Bayreuth-Hof is newly formed out of Bayreuth and parts of Hof. Regensburg receives parts of Landshut. Weiden receives parts of Hof. Donauwörth receives parts of Memmingen. Freising receives parts of München. Kempten-Memmingen is newly formed out of Kempten and parts of Memmingen. Landshut-Pfarrkichen is newly formed out of parts of Landshut and Pfarrkirchen. München loses parts of its territory to Freising and Weilheim. Traunstein receives parts of Pfarrkirchen. Weilheim receives parts of München.

On the 1.1.13, Ansbach-Weißenburg is newly formed out of Weißenburg and parts of Ansbach and Nürnberg. Fürth is newly formed out of parts of Nürnberg and Ansbach.

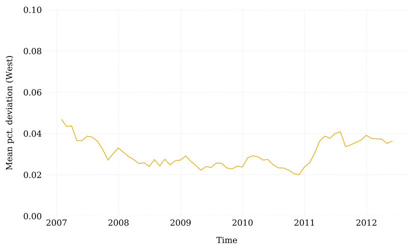
All shares from reassignment on 1.10.12 can be calculated. Due to multiple territorial trades on 1.1.2013, we apply a restriction to identify the flows. Nürnberg gives only parts of the municipality Roth (24000 inhabitants) to Ansbach-Weißenburg. Hence, we assume this flow to be zero.

Brandenburg On the 1.10.12, Cottbus receives territory from Potsdam. The mean share of lost territory, measured in percent of vacancies lost to Cottbus, can be calculated.

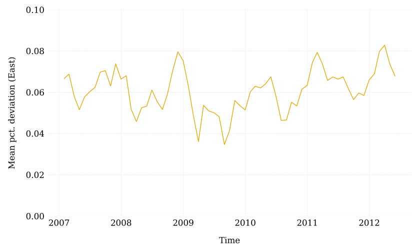
Mecklenburg-West Pomerania On the 1.1.13, Greifswald is formed out of parts of Neubrandenburg and Stralsund. Rostock loses parts to Stralsund. The mean share of vacancies lost can be calculated.

Saxony On the 1.7.12, Bautzen and Riesa receive parts of Dresden. On the 1.1.13, Freiberg is newly formed out of parts of Chemnitz, Leipzig, and Ostschatz. On an unknown date, Annaberg-Buchholz receives parts of Zwickau. Vacancy flows can be identified, and the mean share of vacancies lost can be calculated for every district.

Fig. 3.17: Impact of the adjustment for territorial changes on the vacancy data



(a) West EADs



(b) East EADs

Notes: This figure shows the average deviation of the adjusted from the official number of vacancies in all UADs subject to territorial changes.

Saxony-Anhalt On the 1.7.12, Dessau-Roßlau-Wittenberg is formed out of Wittenberg and parts of Halle and Dessau-Roßlau. Halle receives parts of Weißenberg. On the 1.1.13, Bernburg is newly formed out of parts of Madgeburg, Sangerhausen, and Dessau-Roßlau. Vacancy flows can be identified, and the mean share of vacancies lost can be calculated for every district.

Thuringia On the 1.7.12, Altenburg-Gera is formed out of Altenburg, Gera, and parts of Jena. Erfurt receives parts of Suhl. Gotha loses parts to Suhl. Vacancy flows can be identified, and the mean share of vacancies lost (gained) can be calculated for every district.

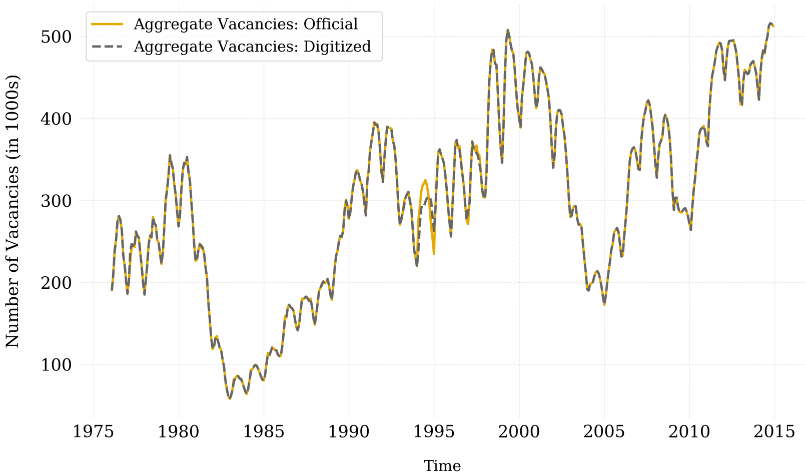
Berlin We treat Berlin as one employment agency district and do not correct for territorial changes between the single districts.

The adjusted vacancy data

After the adjustment, we compare the adjusted time series to the true distribution of vacancies across districts. In particular, we compute the average percentage deviation from the official series. Figure 3.17 shows the average - non zero - deviation during the overlapping period. The average deviation of our adjusted series is around 3% in West and 6% in East Germany.

Figure 3.18 shows the official as well as our constructed series for vacancies in Germany. The digitized data is aggregated from the adjusted district level. Our territorial adjustment procedure does not affect the aggregate number of vacancies. The deviations of the digitized from the official series in the mid-1990s are the result of differences in printed and official data.

Fig. 3.18: Comparison of official and digitized vacancy data



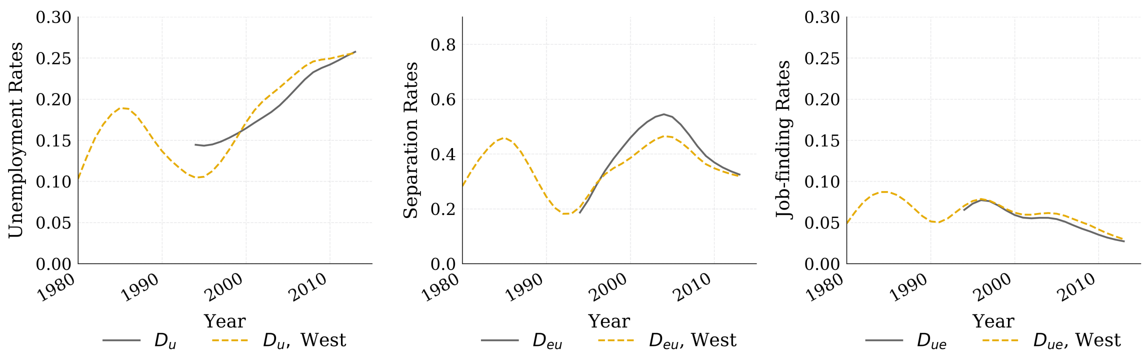
Notes: This figures displays the official as well as the adjusted series for registered vacancies in Germany (West Germany until 1992). The adjusted series is based on (1) adjusted digitized data up to 2000, (2) the FEA vacancy data adjusted for territorial changes and (3) the FEA data for new territorial borders.

3.B Descriptive evidence

3.B.1 Occupational labor markets

The occupational dispersion of unemployment and separation rates as well as labor market tightness – displayed in Figure 3.19 – is more substantial, compared to EADs but homogeneous in West and reunited Germany. Hence, the addition of East German vacancies or unemployed does not change the distribution of the respective variable. What stands out is that the development of the dispersion over time is entirely different than in the districts. Unemployment rates (left panel) are much more unequally dispersed in 2013 compared to 1994. The decrease in the dispersion in separation rates (center panel) starts only in 2005 and the dispersion of job-finding rates (right panel) shows a decline from 1996 onward.

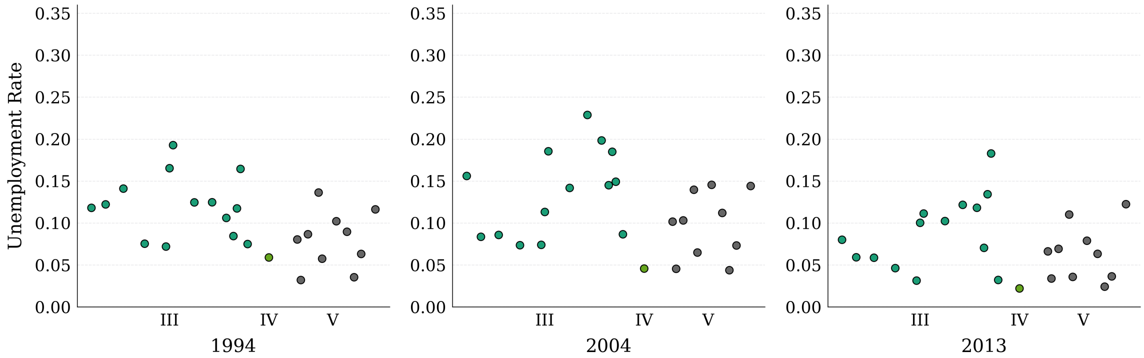
Fig. 3.19: Occupational dispersion of key labor market variables



Notes: The figure displays the annual dispersion of unemployment rates, separation rates and labor market tightness across 25 occupations. All series display the HP-filtered trend ($\lambda = 6.25$).

Figure 3.20 shows the unemployment rate across occupations for the years 1994, 2004, and 2013. The average decrease in occupational unemployment rates from 1994 – 2013 is substantial (−38.0%), albeit with an increasing dispersion.

Fig. 3.20: Unemployment rates by occupation

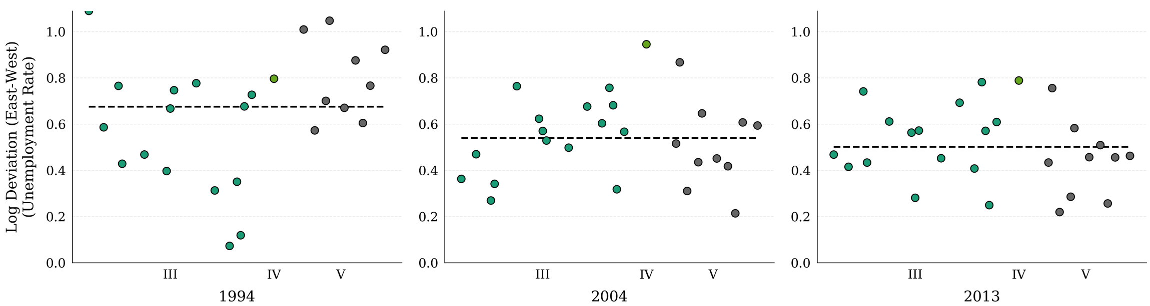


Notes: The figure presents the average monthly unemployment rates of occupations in Germany. Occupations in agriculture, fishing and mining, which correspond to sections I and II are dropped from the sample.

Within the same period job-finding rates increased on average by 8.0% while separation rates decreased by −30.0%. The relative changes in inflow rate suggest that separations are most important for the decrease in unemployment over time.

Occupational unemployment in West and East Germany Differences between East and West occupational unemployment rates are large, as displayed in Figure 3.21. Unemployment in eastern occupation was above 70% (50%) higher compared to western occupations in 1994 (2004). The difference decreased to 46% in 2013.

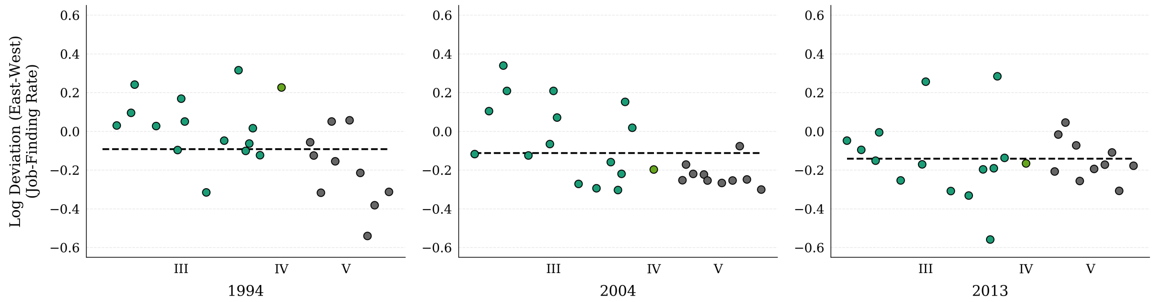
Fig. 3.21: Differences in occupational unemployment rates (West/East)



Notes: The figure presents the log deviation between East and West unemployment rates as well as the mean difference. A positive value indicates higher unemployment in East Germany. Occupations in agriculture, fishing and mining, which correspond to sections I and II are dropped from the sample.

Job-finding rates (Figure 3.22) show a negative difference between East and West occupations for the years 1994, 2004 and 2013.

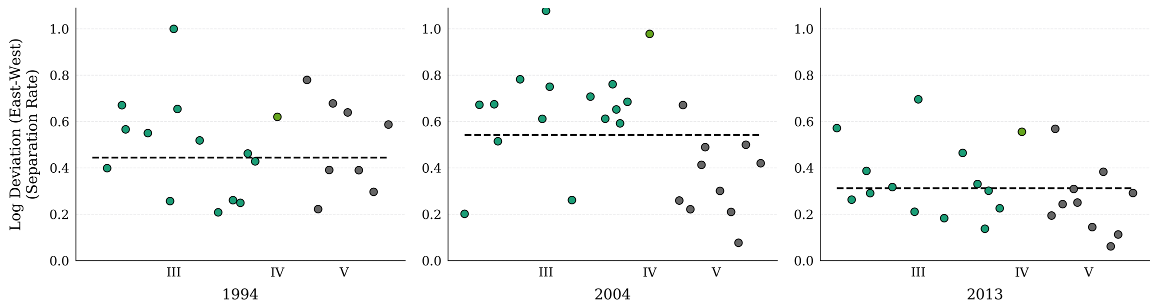
Fig. 3.22: Differences in occupational job-finding rates (West/East)



Notes: The figure presents the log deviation between East and West job-finding rates as well as the mean difference. A positive value indicates higher job-finding rates in East Germany. Occupations in agriculture, fishing and mining, which correspond to sections *I* and *II* are dropped from the sample.

Disparities between the two regions are, to a major degree, determined by separations rates. Figure 3.23 shows that the average rate was 35 to above 50% higher in eastern occupations.

Fig. 3.23: Differences in occupational separation rates (West/East)



Notes: The figure presents the log deviation between East and West separation rates as well as the mean difference. A positive value indicates higher separations in East Germany. Occupations in agriculture, fishing and mining, which correspond to sections *I* and *II* are dropped from the sample.

Occupational unemployment the contribution of job-finding and separation rates We employ the decomposition method that was presented in the main text.

Table 3.7: Decomposition of occupational unemployment rate disparities (West Germany)

	1994 - 2014	1994 - 2004	2005 - 2014
Job-finding	0.09	0.09	0.18
Separation	0.90	0.89	0.81

Notes: The table presents the results of a variance decomposition of the unemployment rates across districts for different time intervals. "Job-finding" ("Separation") indicates that the job-finding (separation) rate is held constant.

Table 3.8: Decomposition of occupational unemployment rate disparities (East Germany)

	1994 - 2014	1994 - 2004	2005 - 2014
Job-finding	0.10	0.12	0.15
Separation	0.88	0.86	0.84

Notes: The table presents the results of a variance decomposition of the unemployment rates across districts for different time intervals. "Job-finding" ("Separation") indicates that the contribution of the respective rate to the unemployment rate variations across occupations.

3.B.2 Estimates of the matching efficiencies

This subsection presents the estimated vacancy shares and sector-specific matching efficiencies. Table 3.9 presents the estimates of model (3.32) for different time intervals and specifications. Our preferred estimate is shown in Column (1).

Table 3.9: Estimates of the vacancy elasticity in Germany

	Time Frame			
	(1)	(2)	(3)	(4)
Parameter	1994 to 2014	2000 to 2014	1994 to 2014	2000 to 2014
α	0.25***	0.26***	0.24***	0.29***
D_{05}		√	√	√
$D_{00,09}$			√	√
D_m	√	√	√	√
$R2$	0.78	0.78	0.78	0.79

Notes: The table presents estimates of model 3.32. Regressions include a qEadratic time trend as well as structural break dummies. "***" indicates significance at the 99 percent level. Values are rounded.

Individual matching efficiency parameter ϕ_i are estimated using model (3.33). The estimated vacancy share, the standard deviation as well as minimum and maximum matching efficiency are presented in Table 3.10, i.e., χ_{min} (χ_{max}) is the minimum (maximum) matching efficiencies across sectors.

Table 3.10: Vacancy shares and sectoral matching efficiencies

	Sector	α	$std(\chi)$	$\chi_{min},\ \chi_{max}$	Time Frame
(1)	States (16)	0.19***	0.13	$0.81, 1.34$	1994 – 2014
(2)	UA Districts (156)	0.16***	0.24	$0.55, 1.80$	1994 – 2014
(3)	UA Districts (36, E)	0.13***	0.11	$0.81, 1.31$	1994 – 2014
(4)	UA Districts (118, W)	0.19***	0.27	$0.57, 1.80$	1980 – 2014

Notes: Regressions include a quartic time trend, monthly dummy variables to control for seasonal fluctuations and structural break dummies. "***" indicates significance at the 99 percent level.

To determine time-varying matching efficiency we estimate

$$\ln(f_{it}) = D_t + D_m + I_{t<2005} \ln(\chi_i) + I_{t>2004} \ln(\chi_i) + \alpha \ln(\theta_{it}) + \epsilon_{it}, \tag{3.36}$$

where $I_{t<2005}$ is an indicator for the months prior to 2005.

Table 3.11 presents estimates of the occupation (groups) and state specific matching efficiencies. Estimates across occupational groups are comparable to the estimates of Bauer (2013). The relative ranking of state-specific matching efficiencies is as well similar, with Bremen (Bavaria) exhibiting the lowest (highest) efficiency. Furthermore, she documents comparable changes in state-specific matching efficiencies in West Germany. However, she finds increasing efficiency in East German states. Our results might be the result from the longer time period, as the analyses the period 2000 to 2009.

Table 3.11: Estimates of occupation- and state-specific matching efficiencies (χ_i)

Sector	1994 − 2014	1994 − 2004	2005 − 2014
		Occupations	
III Manufacturing	1.50	1.51	1.50
IV Technical professions	0.58	0.55	0.63
V Service professions	1.14	1.09	1.21
		States	
Schleswig-Holstein	1.07	1.13	1.00
Hamburg	0.94	0.98	0.89
Lower Saxony	1.01	1.02	0.99
Bremen	0.84	0.87	0.81
North Rhine-Westphalia	0.82	0.84	0.78
Hesse	0.95	0.95	0.94
Rhineland Palatinate	1.01	0.97	1.02
Baden-Wuerttemberg	1.06	1.01	1.07
Bavaria	1.34	1.26	1.39
Saarland	0.84	0.82	0.86
Berlin	0.92	1.01	0.86
Brandenburg	1.05	1.16	0.97
Mecklenburg-West Pomerania	1.08	1.22	0.97
Saxony	1.06	1.13	1.00
Saxony-Anhalt	1.00	1.10	0.94
Thuringia	1.15	1.25	1.07

Notes: The table presents estimates obtained using model (3.33). Estimations include a qEadratic time trend, monthly dummy variables to control for seasonal fluctuations as well as structural break dummies (when applicable). Efficiencies for 1994 − 2004 and 2005 − 2014 (Column 3, and 4, respectively) are estimated using model (3.36).

3.B.3 Aggregate changes in job-finding rates and the relation to changes in matching efficiencies

The development of local labor markets suggests a decline in matching efficiency in some states between $1994 - 2002$ and $2008 - 2014$. We argue that the differences from our results compared to Klinger and Weber (2016), who observe a large increase in the trend of matching efficiency from 2006 to 2011, are due to a different treatment of marginal employment and mini-jobs. While we treat marginally employed as unemployed if they have a parallel unemployment spell and as out of the labor force otherwise, Klinger and Weber (2016) treat them as employment in the case of no parallel unemployment notification. The inclusion of marginal employment leads to a higher job-finding rate and potentially induces a break in 1999, as marginal jobs are fully recorded from this point onward (cf., Hartung et al. (2018), Figure 17). Table 3.12 presents our observed increases of the aggregate job-finding rates between $1994-2002$ and $2008-2014$. While Klinger and Weber (2016) document an increase of the aggregate job-finding rate from (approximately) 7% to 9% (i.e., of 29%) we find that the rates have increased by 8% to 11%.[28]

Table 3.12: Aggregate changes in job-finding and separation rates

	1994 - 2002	2008 - 2014	2010 - 2014	$\Delta 2008 - 2014$	$\Delta 2010 - 2014$
Job-find.	5.36%	5.80%	6.00%	8%	11%
Separat.	0.79%	0.52%	0.49%	−34%	−38%

Notes: The table presents the average monthly aggregate job-finding and separation rate for different time intervals. $\Delta 2008 - 2014$ denotes the percentage change between $1994 - 2002$ and $2008 - 2014$. $\Delta 2010 - 2014$ denotes the percentage change between $1994 - 2002$ and $2010 - 2014$

A difference in treatment of the vacancy series will most likely not cause the different results. While we adjust our vacancy series (downwards) before 2000 to account for the change in composition as seasonal jobs were excluded from the measure from this point onwards, Klinger and Weber (2016) level up the vacancies series after 2000.[29] This adjustment should not affect the differences in results as the market tightness exhibits only a level-shift.

Moreover, the decrease in matching efficiency in some regions cannot be the result of our inflow-correction. This adjustment lowers the number of unemployed after 2005 while leaving the number of hires mostly unaffected as the excluded unemployed are detached from the labor market. Thus, the job-finding rate increases.

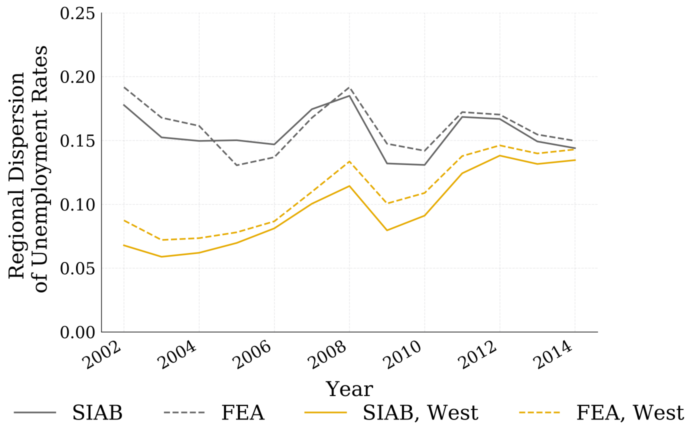
[28] Our results deviate from Hartung et al. (2018) as we consider both, West and East Germany

[29] " *The official definition of vacancies changed in 2000 and does no longer include vacancies for so-called minijobs and other marginal jobs. As we consider all types of dependent employment, however, an analogue definition is preferable – we corrected the break by additional statistics and the multiplier gained from the overlapping year.*", Klinger and Weber (2016) (Online appendix, page 2).

3.B.4 Fit of the unemployment distribution across regions

This section provides evidence that the regional variations of SIAB unemployment rates closely resemble the actual distribution of unemployment rates across EADs.[30] Figure 3.24 shows the annual dispersion of unemployment rates across EADs for unemployment data obtained from our sample (SIAB) as well as from the federal employment agency (FEA). The annual dispersion of the SIAB data matches the dispersion of the FEA data in level and trend for West and united Germany.

Fig. 3.24: Dispersion of unemployment rates (Robustness)



Notes: The figure displays the annual dispersion of unemployment rates for unemployment data obtained from our sample (SIAB) as well as from the federal employment agency (FEA). Data on employment is taken from SIAB.

To compute the mismatch indices presented in section 3.5, sectoral matching efficiencies need to be estimated using outflow rates for highly disaggregated sectors. A mismatch index abstracting from heterogeneity in matching-efficiency relies solely on the distribution of unemployment and vacancies

$$M_t = 1 - \sum_{i=1}^{I} \left(\frac{v_{it}}{\bar{v}_t}\right)^{\alpha} \left(\frac{u_{it}}{\bar{u}_t}\right)^{1-\alpha}.$$

We utilize this specific index to facilitate a comparison between data sources. Concerning geographical mismatch, we highlight the comparability of the unemployment data sourced from SIAB and official unemployment data. The vacancy share is set to 0.25.[31] We conclude that the SIAB data is a good representation for the true distribution of unemployment across districts and occupations.

[30]We use different data sources for official regional unemployment rates. Rates from 1994 to 1999 are taken from the ANBA publications; Rates from 2001 onward are obtained from "Arbeitsmarktstatistik der Bundesagentur für Arbeit" from Statistische Ämter des Bundes und der Länder, Deutschland, 2017 (13211-02-05-4).

[31]This index is similar to the one developed by Jackman and Roper (1987) if one sets $\alpha = 0.5$.

Figure 3.25 shows the annual average benchmark mismatch index constructed using SIAB data (M, SIAB) and using different data sources for West Germany. We additionally compute the M index using official unemployment data from the FEA (M, BA) and using monthly vacancy and unemployment data from ANBA (M, ANBA). For the latter, we do not correct for territorial changes between employment agency districts. The comparison of the M index across different data sources shows almost no differences in cyclicality. Indices computed using the official data for unemployment are only slightly higher, compared to the index based on SIAB data. At the same time, the absolute difference is quite stable. Hence, the sample from SIAB used in this paper is a good approximation of the actual distribution of unemployment across districts. Furthermore, the territorial correction of the vacancy series does not lead to differences in the cyclicality of the indices.

Fig. 3.25: Mismatch index, Districts (West Germany)

Notes: The figure presents baseline mismatch indices computed for West German employment agency districts using three different data sources. The mismatch index M, SIAB is the yearly average of the mismatch index without district specific matching-efficiency computed using SIAB micro-data as well as the adjusted vacancy series. M, ANBA is computed using unadjusted unemployment as well as vacancy data sourced from the ANBA publications. M, BA is computed using the adjusted vacancy data as well as official unemployment data from the FEA.

Figure 3.26 plots the index for East Germany. We compute the M index using official unemployment data from the FEA (M, BA, and M, ANBA) and using annual average unemployment data based on SIAB data (M, SIAB). The indices are comparable in level and cyclicality. Overall, mismatch driven by heterogeneity in the distribution of unemployment and vacancies is low. The unemployment data sourced from SIAB is a good representation of the actual distribution of unemployed across EADs.

Fig. 3.26: Mismatch index, Districts (East Germany)

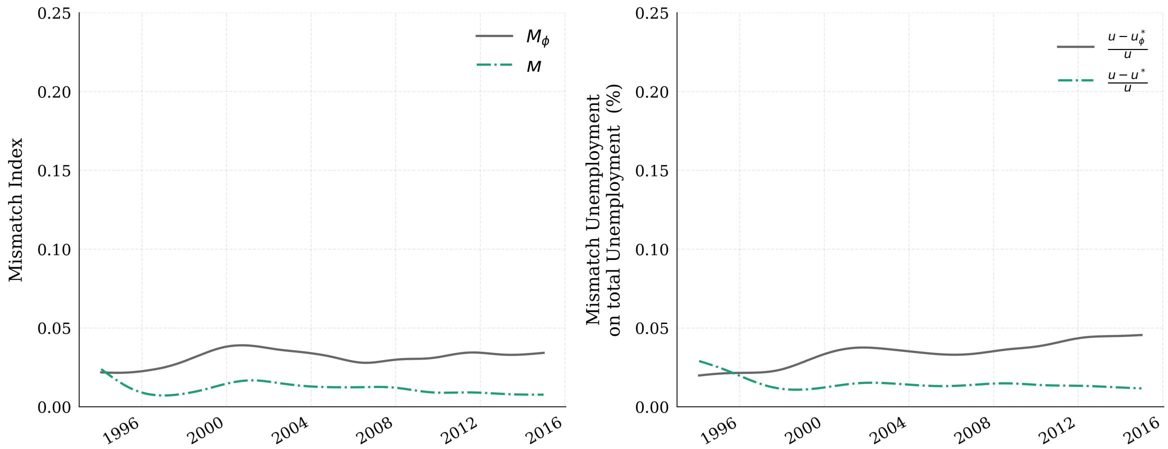
Notes: The figure presents baseline mismatch indices computed for East German employment agency districts using three different data sources. The mismatch index M, SIAB is the yearly average of the mismatch index without district specific matching-efficiency computed using SIAB micro-data as well as the adjusted vacancy series. M, ANBA is computed using unadjusted unemployment as well as vacancy data sourced from the ANBA publications. M, BA is computed using the adjusted vacancy data as well as official unemployment data from the FEA.

3.B.5 Regional mismatch: Additional evidence

In this subsection, we provide evidence on mismatch across employment agency districts in West and East Germany and across federal states. Furthermore, we account for heterogeneity in separation rates, productivity, and allow for time-varying matching-efficiency.

Mismatch across East German EADs is low (Figure 3.27, left panel): The optimal allocation of the planner would increase hires by below 5% for both measures of mismatch.

Fig. 3.27: Regional mismatch across East German EADs

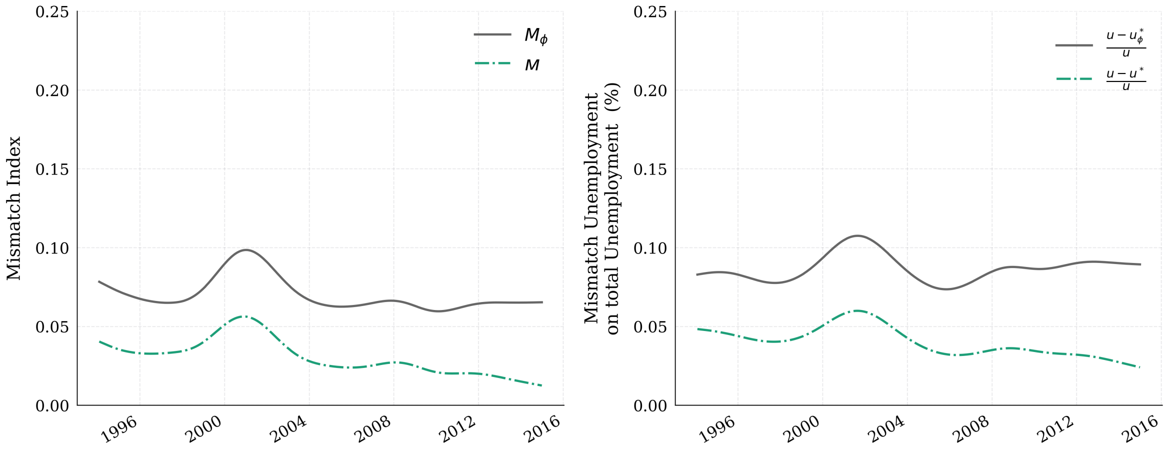


Notes: The figure presents the mismatch indices M and M_ϕ as well as the corresponding mismatch unemployment rates. The indices are HP-filtered with a smoothing parameter $\lambda = 100{,}000$.

The optimal allocation of the planner would increase aggregate hires by below 10% for both measures of mismatch by eliminating mismatch across federal states (Figure 3.28).

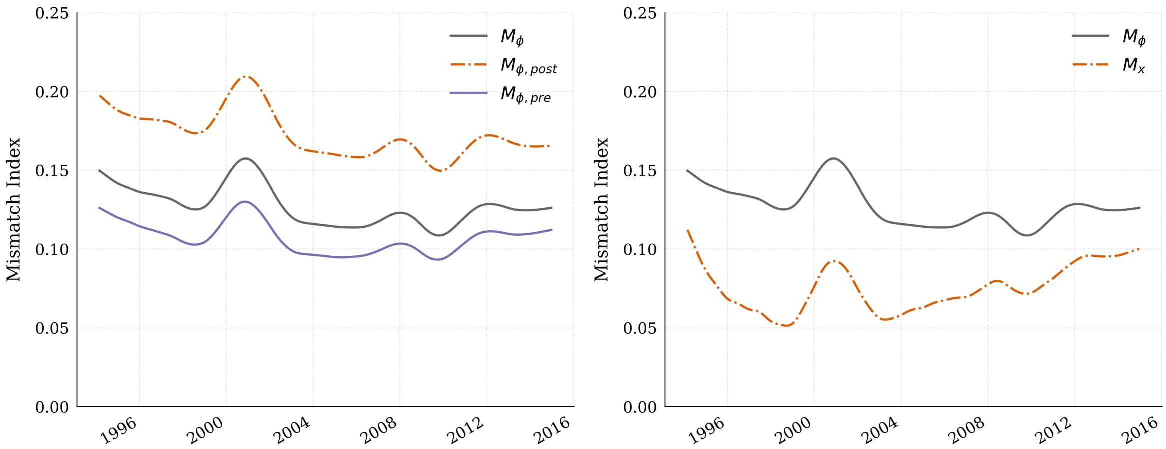
Time variation in matching efficiency leads a level change in mismatch, while heterogeneity in separations and productivity (wages) decreases mismatch (Figure 3.29; left and right panel, respectively). The index M_{x_t} is derived in section 3.D. We approximate productivity by the wage.

Fig. 3.28: Regional mismatch across federal states



Notes: The figure presents the mismatch indices M and M_ϕ as well as the corresponding mismatch unemployment rates. The indices are HP-filtered with a smoothing parameter $\lambda = 100{,}000$.

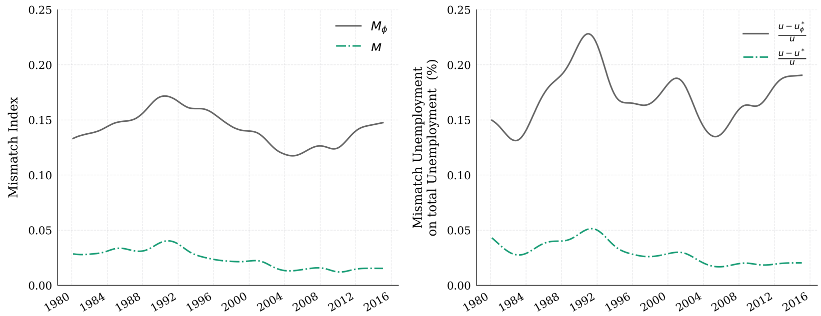
Fig. 3.29: Regional mismatch across employment agency districts: Sensitivity



Notes: The figure presents the mismatch indices M_{θ_t}, M_{x_t} and M_{ϕ_t}. The indices are HP-filtered with a smoothing parameter $\lambda = 100{,}000$. M_{θ_t} represents the estimates presented in the main text. M_{ϕ_t} is estimated using the region-specific matching efficiencies provided in Table 3.11. Post indicates the use of the matching efficiencies valid for the period after 2005 and pre for the period before 2006. M_{x_t} is estimated based on the measure derived in Section 3.D.

Regional mismatch in West Germany For West Germany, we extend the index backward to 1980. The fraction of hires lost due to misallocation of unemployment across West German EADs and the corresponding contribution to the aggregate unemployment rate during 1980−2014 are displayed in Figure 3.30. The index unadjusted for heterogeneity (M_t) decreases throughout the period and lies below 5%. Accounting for differences in matching efficiency ($M_{\chi,t}$) leads to a considerable rise in the level of mismatch: Redistributing unemployed based on the planner's rule leads to an average increase in hires of 15%. In general, mismatch in West Germany is similar in level and cyclicality during 1994−2014 compared to Germany.[32] Moreover, the right panel shows

Fig. 3.30: Regional mismatch across West German EADs



Notes: The figure presents the mismatch indices M and M_χ as well as the corresponding contribution of mismatch to aggregate unemployment. The unit of analysis are 118 employment agency districts. All series display the HP-filtered trend (λ =100,000).

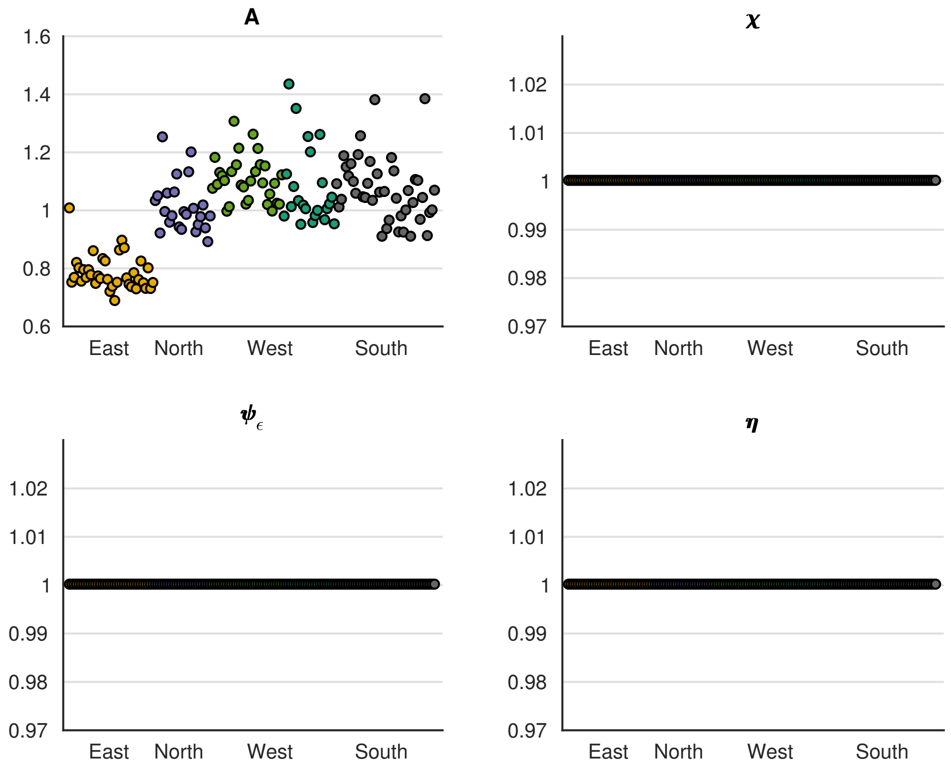
that regional mismatch is not the leading cause of high unemployment during the 1990s and early 2000s. The declining contribution of mismatch in 1992 and 2001 coincides with a high and long-lasting increase in separations (Hartung et al. (2016)). A change in separations generally leads to a more substantial change in the counterfactual rate. This is due to lower (higher) optimal compared to actual unemployment (employment). Therefore, the contribution to aggregate unemployment behaves countercyclical to the actual unemployment rate.

[32] The size of the indices is not comparable across different these regional classifications as the number of segments, vacancies and unemployed differ (cf., Sahin et al., 2014). However, the mismatch unemployment rate ($u_t - u_t^*$) can be used compare the severity of mismatch: When accounting for heterogeneity in matching efficiency, the mismatch unemployment rate is 16% higher in Germany than in West Germany.

3.C Quantitative evaluation: Additional results

Identification: Wages Figure 3.31 displays the difference of the identified parameters when wages are fixed at the benchmark districts' value, e.g., $\frac{A}{\bar{A}}$, where $\bar{A}_i$ represents the parameter value with constant wages. For example, productivity is approximately 20% lower in eastern districts when we account for the empirical variation in wages. Moreover, wages solely affect the productivity parameter. Hence, the empirical variation in wages does not affect the other three parameters.

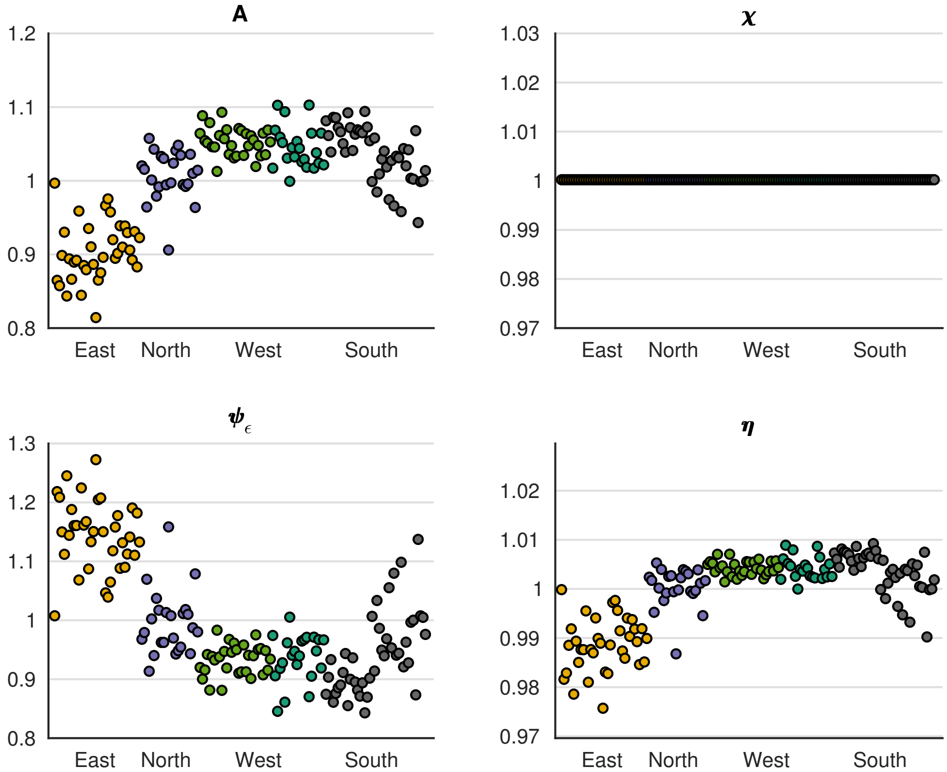
Fig. 3.31: Relative structural disparities with constant wages (1994 – 2002)



Notes: The figure displays the relative values of the four region-specific structural parameters for the steady-state in 1994 – 2002 and a counterfactual steady-state. We determine the parameters in the steady-state within the model given the observed variation in wages, tightness, job-finding, and separation rates. We determine the parameters in the counterfactual steady-state given the observed variation in tightness, job-finding, and separation rates and set wages to the benchmark districts' counterpart.

Identification: Separation rates Figure 3.32 displays the difference of the identified parameters between the steady-state and a counterfactual steady-state where separation rates are fixed at the benchmark districts' value. For example, the lower left panel displays $\frac{\psi_\epsilon}{\bar{\psi}_\epsilon}$, where $\bar{\psi}_\epsilon$ represents the parameter value with constant separations rates. Constant separations rates primarily affect the match uncertainty. This parameter is rescaled proportionally to the variations in separations. We observe minor variations of productivity and the bargaining power. Separation rates do not affect the matching efficiency.

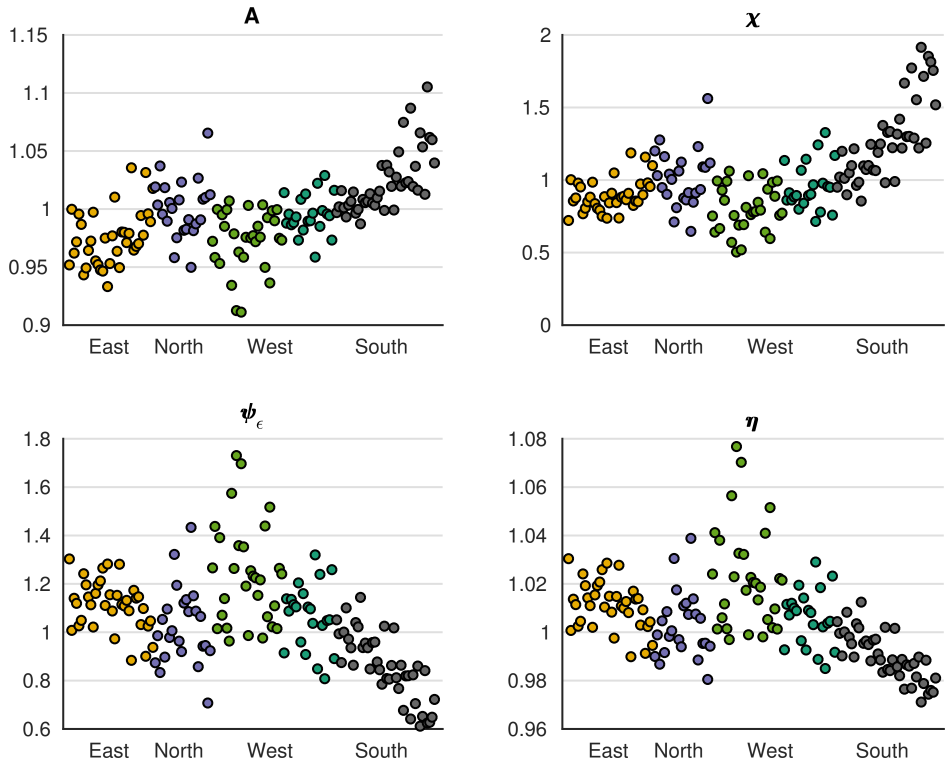
Fig. 3.32: Relative structural disparities with constant separation rates (1994 – 2002)



Notes: The figure displays the relative values of the four region-specific structural parameters for the steady-state in 1994 – 2002 and a counterfactual steady-state. We determine the parameters in the steady-state within the model given the observed variation in wages, tightness, job-finding, and separation rates. We determine the parameters in the counterfactual steady-state given the observed variation in wages, tightness, and job-finding rates and set separation rates to the benchmark districts' counterpart.

Identification: Job-finding rates Figure 3.33 displays the difference of the identified parameters between the steady-state and a counterfactual steady-state where job-finding rates are fixed at the benchmark districts' value. For example, the lower-left panel displays $\frac{\psi_\epsilon}{\bar{\psi}_\epsilon}$, where $\bar{\psi}_\epsilon$ represents the parameter value with constant job-finding rates. Constant job-finding rates primarily affect the match uncertainty and matching efficiency. The latter parameter is rescaled proportionally to the variations in the job-finding rate. Match uncertainty is rescaled inversely: a higher job-finding rate, ceteris paribus, increases the value of unemployment and, thus, lowers the match surplus. Match uncertainty decreases when the surplus decreases, and separation rates do not change. Hence, when we account for the higher job-finding rates in Bavaria, the job-match uncertainty is lower, as displayed in the lower-left panel.

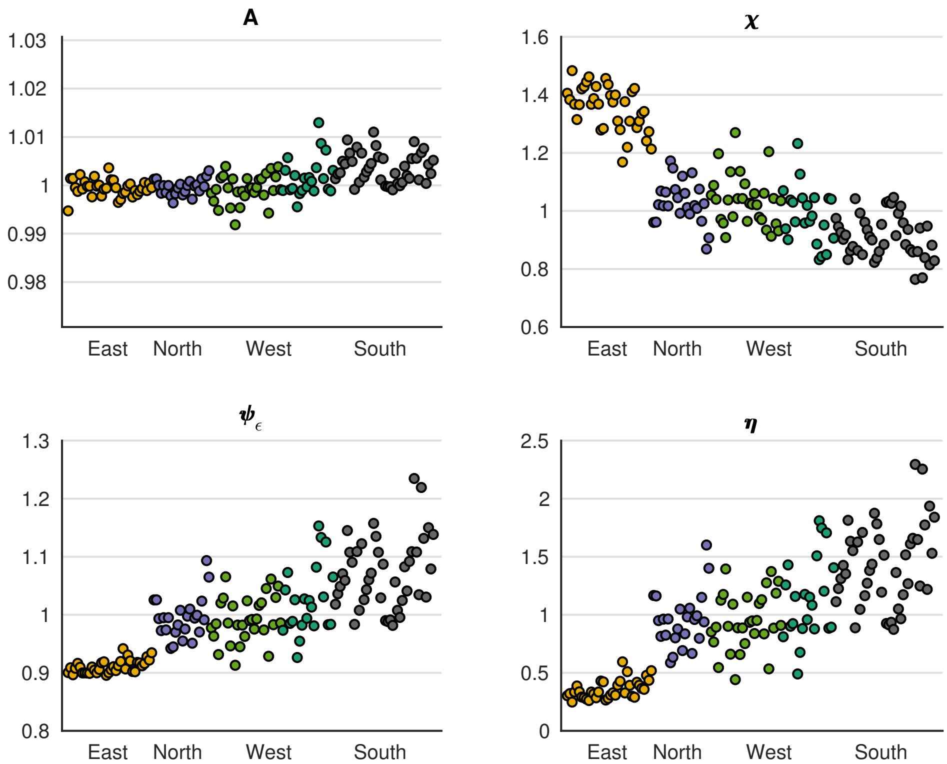
Fig. 3.33: Relative structural disparities with constant job-finding rates (1994 – 2002)



Notes: The figure displays the relative values of the four region-specific structural parameters for the steady-state in 1994 – 2002 and a counterfactual steady-state. We determine the parameters in the steady-state within the model given the observed variation in wages, tightness, job-finding, and separation rates. We determine the parameters in the counterfactual steady-state given the observed variation in wages, tightness, and separation rates and set job-finding rates to the benchmark districts' counterpart.

Identification: Tightness Figure 3.34 displays the deviation from the individual steady-state parameters when labor market tightness is set to the benchmark districts' value. The parameters which are primarily affected are the matching efficiency, the bargaining power, and match uncertainty —— the changes in the bargaining power mirror the distribution of labor market tightness. When tightness is kept constant, the matching efficiency increases in East Germany by 40%, which reflects the lower tightness in these local markets. Tightness also determines matching efficiency through the matching function. Furthermore, higher bargaining power implies a higher match-surplus, which translates, given constant separation rates, into higher job-match uncertainty.

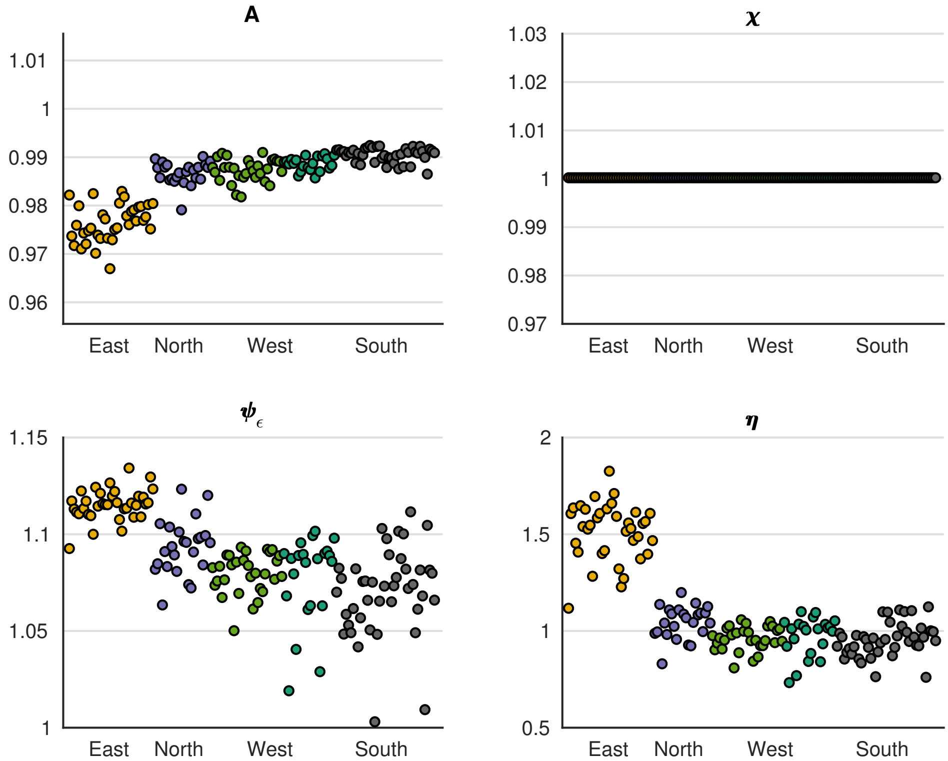
Fig. 3.34: Relative structural disparities with constant tightness (1994 – 2002)



Notes: The figure displays the relative values of the four region-specific structural parameters for the steady-state in 1994 – 2002 and a counterfactual steady-state. We determine the parameters in the steady-state within the model given the observed variation in wages, tightness, job-finding, and separation rates. We determine the parameters in the counterfactual steady-state given the observed variation in wages, job-finding, and separation rates and set tightness to the benchmark districts' counterpart.

Region-specific and constant vacancy posting costs Figure 3.35 shows that if vacancy posting costs do not depend on regional productivity, the bargaining power rescales accordingly. We display the parameter under constant posting costs relative to the parameter under region-specific costs.[33] Hence, bargaining power in the East (lower right) with fixed vacancy posting costs is 50% higher than with variable posting costs. Region-specific costs have a minor impact on productivity and do not influence matching efficiency. Moreover, the variability has minor impacts on match-stability.

Fig. 3.35: Relative structural disparities: fixed vs. region-specific vacancy posting costs

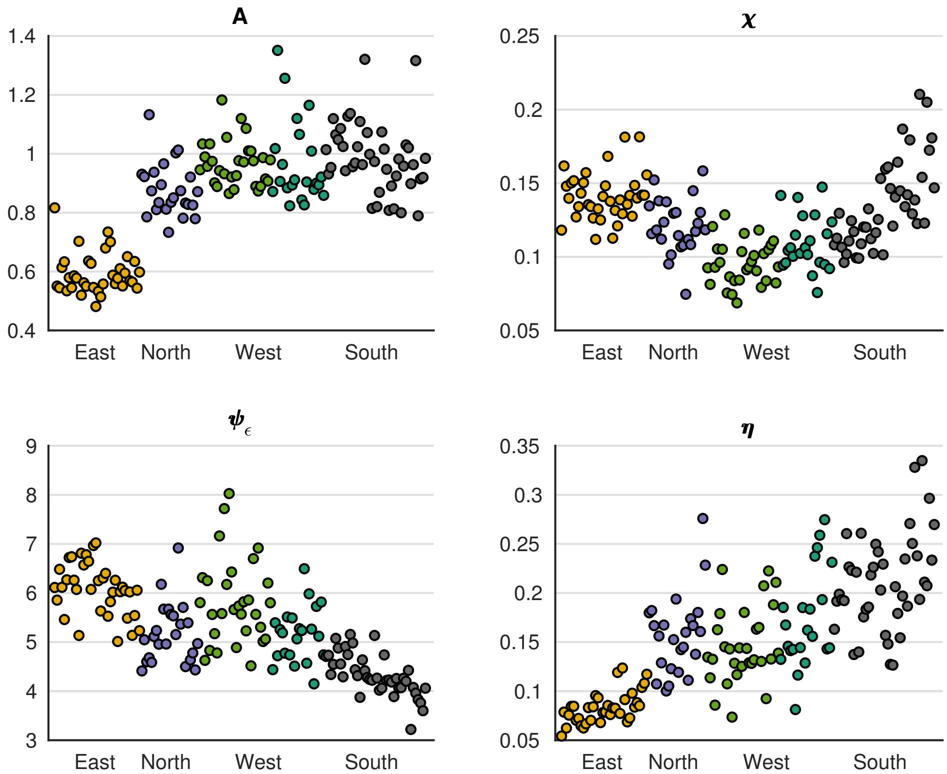


Notes: The figure displays the identified values of the four region-specific structural parameters (A, χ, ψ_ϵ and η) under constant relative to region-specific vacancy posting costs. We determine the parameters within the model given the observed variation in wages, labor market tightness, separation, and job-finding rates. We target the averages of these four variables within 1994 − 2002.

Figure 3.36 illustrates that the relative structure of the identified parameters is preserved.

[33]For example, relative productivity in the upper left panel is A(fix vacancy posting costs) / A(region-specific vacancy posting costs)

Fig. 3.36: Regional differences in 1994 – 2002 with fixed vacancy posting costs



Notes: The figure displays the identified values of the four region-specific structural parameters (A, χ, ψ_ϵ and η) under constant vacancy posting costs. We determine the parameters within the model given the observed variation in wages, labor market tightness, separation, and job-finding rates. We target the averages of these four variables within 1994 – 2002.

Severance payments We recalibrate the initial point and assume that each firm has to pay an additional amount upon separating from the worker. These severance payments amount to three months worth of wages − i.e., the layoff tax increases to 7. The government redistributes the payments by an increase in benefits. With a prior replacement rate of 67% and an average unemployment duration of 20 months ($\frac{1}{sf}$), the replacement rate increases to 87%.

Table 3.13: Structural changes with redistributed severance payments (1994 − 2002 to 2008 − 2014)

	Productivity (A)	Matching χ	Shock Dispersion ψ_ϵ	Bargaining η
East	0.19	−0.24	−0.13	0.74
North	0.03	−0.06	−0.09	0.31
West	0.01	−0.02	−0.07	0.17
South	0.05	0.06	−0.13	0.23

Notes: The table presents the average (log) changes of the identified structural parameters between 1994 − 2002 and 2008 − 2014. We group districts into four regions: South: Bavaria and Baden-Wuerttemberg; West: North Rhine-Westphalia, Hesse, Saarland, and Rhineland Palatinate; North: Schleswig-Holstein, Hamburg, Bremen, and Lower-Saxony; East: Berlin, Saxony, Saxony-Anhalt, Thuringia, and Mecklenburg-West Pomerania.

Table 3.14: Steady-state compared to the planner economy with redistributed severance payments (1994 − 2002)

	Employment	ξ	s	f	Welfare
East	0.13	0.07	0.09	0.64	0.04
North	0.09	−0.48	0.06	0.33	0.02
West	0.10	−0.85	0.06	0.30	0.02
South	0.04	−0.39	0.05	0.21	0.03

Notes: The table presents the (log) change in employment, separation rates (ξ), search effort (s), job finding rates (f) and welfare under the optimal regional labor market policy mix. The change is measured relative to the individual steady-state.

Table 3.15: Steady-state compared to the planner economy with redistributed severance payments $(2008-2014)$

	Employment	ξ	s	f	Welfare
East	0.10	-0.59	0.07	0.40	0.02
North	0.06	-0.71	0.05	0.23	0.02
West	0.08	-0.99	0.06	0.25	0.02
South	0.02	-0.27	0.04	0.12	0.03

Notes: The table presents the (log) change in employment, separation rates (ξ), search effort (s), job finding rates (f) and welfare under the optimal regional labor market policy mix. The change is measured relative to the individual steady-state.

Hartz reforms In the course of the Hartz reforms, the replacement rate for long-term unemployed was decreased by 25%. Within the model, we cannot disentangle a change in b from variations in ψ_ϵ. Therefore, we back out the implied change in b by matching the average value of ψ_ϵ during the earlier period. Hence, we allow for changes in variation across the mean but attribute all level shifts to a change in the benefit level. We find that a 10.7% decrease in b suffices to keep the unweighted mean constant.[34] This is a large drop in the benefit level compared to the small fraction of long-term unemployed. We view this choice as the extreme case and show that the identified parameters and policy implications do not vary qualitatively compared to not accounting for the policy change. We adjust the replacement rate (b) to 89.3% percent of its prior value to account for the changes in benefits, recover the structural parameters in $2008 - 2014$ and evaluate the optimal policy mix. Table 3.16 shows that the changes between $1994 - 2002$ are $2008 - 2014$ similar to before and only ψ_ϵ varies.

Table 3.16: Structural changes with an adjusted replacement rate from $1994 - 2002$ to $2008 - 2014$

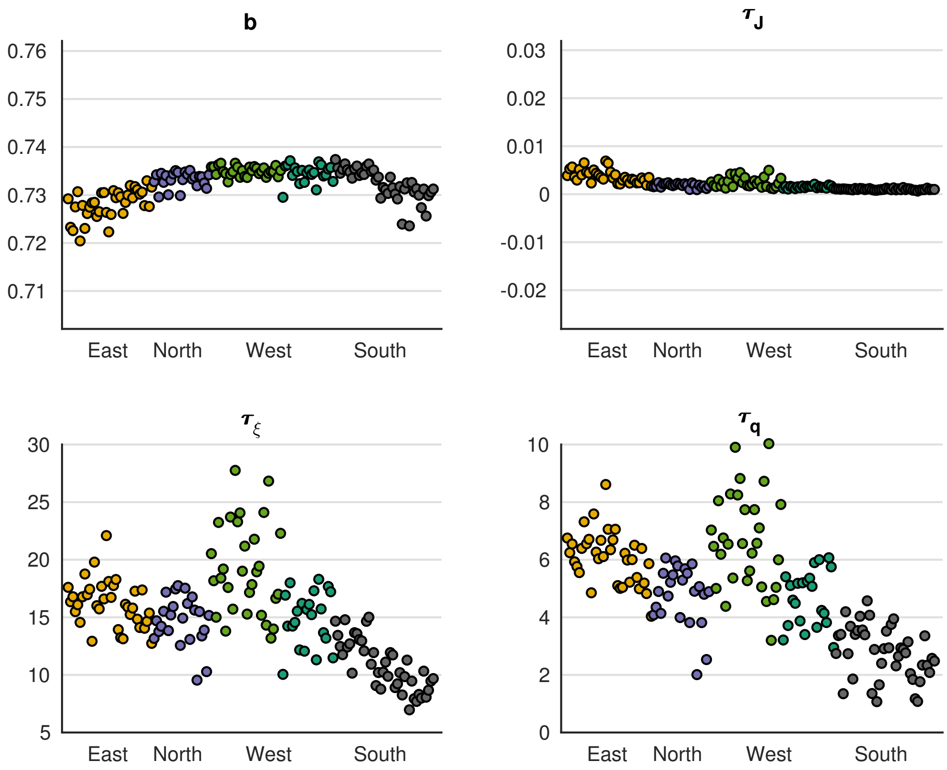
	Productivity (A)	Matching (χ)	Shock Dispersion (ψ_ϵ)	Bargaining (η)
East	0.18	-0.24	0.01	0.64
North	0.03	-0.06	0.01	0.22
West	0.0	-0.02	0.02	0.07
South	0.04	0.06	-0.07	0.15

Notes: The table presents the average (log) changes of the identified structural parameters and the four target variables between $1994-2002$ and $2008-2014$ with an adjusted replacement rate. We group districts into four regions: South: Bavaria and Baden-Wuerttemberg; West: North Rhine-Westphalia, Hesse, Saarland, and Rhineland Palatinate; North: Schleswig-Holstein, Hamburg, Bremen, and Lower-Saxony; East: Berlin, Saxony, Saxony-Anhalt, Thuringia, and Mecklenburg-West Pomerania.

Concerning the optimal policy mix, the deviation of the planner's solution from the steady-state is quantitatively smaller. The respective results are displayed in Figure 3.37 and Table 3.17.

[34] We do not use a weighting scheme to determine the change. However, a larger drop in benefits is needed to keep the mean constant if we weigh the change by employment.

Fig. 3.37: The optimal policy mix with an adjusted replacement rate (2008 – 2014)



Notes: The figure displays the optimal policy mix in the planner's economy (2008 – 2014). b represents the replacement rate ($\frac{c_u}{c_e}$) and τ_J the optimal production tax. τ_ξ and τ_q are the layoff taxes and hiring subsidies, respectively.

Table 3.17: Steady-state compared to the planner's economy with an adjusted replacement rate (2008 – 2014)

	Employment	ξ	s	f	Welfare
East	0.1	-0.73	-0.02	0.40	0.05
North	0.07	-1.11	-0.04	0.22	0.05
West	0.08	-1.20	-0.03	0.25	0.05
South	0.04	-1.04	-0.04	0.12	0.06

Notes: The table presents the (log) change in employment, separation rates (ξ), search effort (s), job finding rates (f) and welfare under the optimal regional labor market policy mix. The change is measured relative to the individual steady-state.

3.D The model (Mismatch)

This section briefly describes the framework developed by Sahin et al. (2014) and derives the optimal allocation rule.

Benchmark The economy is populated by risk-neutral individuals who can either be employed or unemployed. Unemployed workers search for a job. Labor markets are distinct: an individual can only search in one market (district) i. Hence, $\sum_i^I (e_{it} + u_{it}) = 1$ Vacancies arise exogenously. There is no on-the-job search. Let next period's values be denoted by " $'$ ".

New matches in each district are determined by a Cobb-Douglas matching function $m_{it} = \Phi_t \chi_{it} v_{it}^{\gamma} u_{it}^{1-\gamma}$, where $\Phi_t \chi_{it}$ is the matching efficiency and γ the vacancy share.

Each employed worker produces Z units of output. All existing matches are destroyed at the exogenous rate ξ_t. Aggregate shocks, Z_t, ξ_t, Φ_t are drawn from the conditional distribution $\Lambda_{Z,\xi,\Phi}(Z', \xi', \Phi'; Z, \xi, \Phi)$, i.e., the realization of next period's parameters depend only on their current state. The vector of vacancies is drawn from the conditional distribution function $\Lambda_v(v'; v, Z', \xi', \Phi')$, i.e., they depend on their current state and next period's realization of aggregate productivity, separation rate, and matching efficiency. The sector-specific matching efficiencies are independent across sectors and are drawn from the independent distribution function $\Lambda_\chi(\chi'; \chi)$, where χ is the vector of matching efficiencies.

Each period consists of three stages. In the first stage, the distribution of employment and the vector of vacancies, matching efficiencies, and the aggregate shocks ($\{v, \phi, \xi, Z, \Phi\}$) are observed. The planner then decides how to allocate unemployment across districts. In the second stage, new hires are formed based on the matching-function, and production takes place. At the last stage, a fraction of job-matches separates exogenously, and the employment distribution in the next period is determined. The planner's problem is the following

$$V(e; v; \phi, Z, \xi, \Phi) = \max_{0 \le u_i} \sum_i^I Z(e_i + h_i) + \beta \mathbb{E}[V(e'; v', \phi', Z'\xi', \Phi')] \tag{3.37}$$

$$s.t. \quad : \tag{3.38}$$

$$\sum_i^I u_i \le u \tag{3.39}$$

$$\sum_i^I (e_i + u_i) = 1 \tag{3.40}$$

$$h_i = \Phi \chi_i v_i^{\gamma} u_i^{1-\gamma} \tag{3.41}$$

$$e_i' = (1 - \xi)(e_i + h_i) \tag{3.42}$$

$$\Lambda_{Z,\xi,\Phi}(Z', \xi', \Phi'; Z, \xi, \Phi), \Lambda_v(v'; v, Z', \xi', \Phi'), \Lambda_\chi(\chi'; \chi) \tag{3.43}$$

Th Lagrangian for the planner's problem is

$$\max_{u_i} \mathcal{L} = \sum_i^I Z(e_i + h_i) + \beta \mathbb{E}[V(e'; v', \chi', Z', \xi', \Phi')] - \mu(u_i - u)$$

Taking the first derivative with respect to u_i and rearranging yields

$$Z(1-\gamma)\Phi\chi_i v_i^\gamma u_i^{-\gamma} + \beta \mathbb{E}\left[\frac{\partial V'}{\partial e_i'}\right](1-\xi)(1-\gamma)\Phi\chi_i v_i^\gamma u_i^{-\gamma} = \mu,$$

where we use $\frac{\partial V'}{\partial u_i} = \frac{\partial V'}{\partial e_i'}\frac{\partial e_i'}{\partial u_i}$. Using the envelope theorem gives

$$\frac{\partial V}{\partial e_i} = \frac{\partial \mathcal{L}}{\partial e_i} = Z + \beta(1-\xi)\mathbb{E}\left[\frac{\partial V'}{\partial e_i'}\right].$$

Iterating this equation forwards yields: $\mathbb{E}\left[\frac{\partial V'}{\partial e_i'}\right] = \frac{Z}{1-\beta(1-\xi)}$.

Substituting the latter result into the first order condition and rearranging gives

$$(1-\gamma)\chi_i\left(\frac{v_i}{u_i}\right)^\gamma = \frac{\mu}{\Phi(Z + \beta\frac{Z}{1-\beta(1-\xi)}(1-\xi))}.$$

Note that all right hand side variables are independent of the sector. Hence, the planner's optimal allocation rule then follows directly, denoting the planner's allocation with "*":

$$\chi_1\left(\frac{v_1}{u_1^*}\right)^\gamma = ... = \chi_I\left(\frac{v_I}{u_I^*}\right)^\gamma. \tag{3.44}$$

Heterogeneity in productivity and separation rates Allowing productivity and separation rates to differ between districts leads to a modified optimality condition. We approximate productivity by wages. The planner now re-allocates the unemployed taking into account regional differences in matching-efficiency, number of vacancies, separation rates as well as productivity.[35]

The benchmark index is driven by two sources of heterogeneity: the distribution of vacancies and differences in matching efficiencies. Two other relevant sources of heterogeneity are productivity and separation rates. The model can be extended to account for these additional factors. A key difference is that the planner now maximizes employment. Allowing productivity (z_{it}) and separation rates (ξ_{it}) to differ between districts leads to a modified optimality condition. The modified planner's allocation rule implies that

$$\frac{v_{jt}}{u_{jt}^*} = \left(\frac{x_{kt}}{x_{jt}}\right)^{\frac{1}{\gamma}}\frac{v_{kt}}{u_{kt}^*} \tag{3.45}$$

must hold in equilibrium for each district pair j and k. To ease notation z_i denotes

[35]For a detailed derivation the reader is referred to Sahin et al. (2014).

district specific matching efficiency weighted by productivity and separation rate $x_{it} = \frac{z_{it}\chi_{it}}{1-\beta(1-\xi_{it})}$. The planner now reassigns the unemployed taking into account sectoral differences in matching efficiency, vacancies, separation rates and productivity. Going through the same steps as for the benchmark index, the fraction of hires lost due to mismatch is capture by

$$M_{x,t} = 1 - \sum_{i=1}^{I} \left(\frac{\chi_{it}}{\bar{\chi}_{xt}}\right) \left(\frac{v_{it}}{v_t}\right)^{\gamma} \left(\frac{u_{it}}{u_t}\right)^{1-\gamma}, \tag{3.46}$$

where

$$\bar{\chi}_{xt} = \sum_{i=1}^{I} \chi_{it} \left(\frac{x_{it}}{\bar{x}_t}\right)^{\frac{1-\gamma}{\gamma}} \left(\frac{v_{it}}{v_t}\right) \quad \text{and} \quad \bar{x}_t = \left[\sum_{i=1}^{I} x_{it}^{\frac{1}{\gamma}} \left(\frac{v_{it}}{v_t}\right)\right]^{\gamma}.$$

The counterfactual unemployment rate can be constructed in the same way as for the M_{χ} index.

3.E Classification

Table 3.18: District Classification

Code	District	Code	District	Code	District
030	AA Greifswald	373	AA Paderborn	729	AA Fürth
031	AA Neubrandenburg	375	AA Recklinghausen	735	AA Nürnberg
032	AA Rostock	377	AA Rheine	739	AA Regensburg
033	AA Schwerin	381	AA Siegen	743	AA Schwandorf
034	AA Stralsund	383	AA Meschede – Soest	747	AA Schweinfurt
111	AA Bad Oldesloe	387	AA Wesel	751	AA Weiden
115	AA Elmshorn	391	AA Solingen – Wuppertal	759	AA Würzburg
119	AA Flensburg	411	AA Bad Hersfeld – Fulda	811	AA Augsburg
123	AA Hamburg	415	AA Darmstadt	815	AA Deggendorf
127	AA Heide	419	AA Frankfurt	819	AA Donauwörth
131	AA Kiel	427	AA Gießen	823	AA Freising
135	AA Lübeck	431	AA Hanau	827	AA Ingolstadt
139	AA Neumünster	433	AA Bad Homburg	831	AA Kempten – Memmingen
211	AA Braunschweig – Goslar	435	AA Kassel	835	AA Landshut – Pfarrkirchen
214	AA Bremen – Bremerhaven	439	AA Korbach	843	AA München
221	AA Celle	443	AA Limburg – Wetzlar	847	AA Passau
224	AA Emden – Leer	447	AA Marburg	855	AA Rosenheim
231	AA Göttingen	451	AA Offenbach	859	AA Traunstein
234	AA Hameln	459	AA Wiesbaden	863	AA Weilheim
237	AA Hannover	511	AA Bad Kreuznach	035	AA Cottbus
241	AA Helmstedt	515	AA Kaiserslautern – Pirmasens	036	AA Eberswalde
244	AA Hildesheim	519	AA Koblenz – Mayen	037	AA Frankfurt (Oder)
251	AA Lüneburg – Uelzen	523	AA Ludwigshafen	038	AA Neuruppin
257	AA Nordhorn	527	AA Mainz	039	AA Potsdam
261	AA Oldenburg – Wilhelmshaven	535	AA Montabaur		AA Berlin Süd
264	AA Osnabrück	543	AA Landau	955	AA Berlin Nord
267	AA Stade	547	AA Neuwied		AA Berlin Mitte
274	AA Vechta	555	AA Saarland	041	AA Bernburg
277	AA Nienburg – Verden	563	AA Trier	042	AA Dessau-Roßlau – Wittenberg
311	AA Aachen – Düren	611	AA Aalen	043	AA Halberstadt
315	AA Bergisch Gladbach	614	AA Balingen	044	AA Halle
317	AA Bielefeld	617	AA Freiburg	045	AA Magdeburg
321	AA Bochum	621	AA Göppingen	046	AA Weißenfels
323	AA Bonn	624	AA Heidelberg	047	AA Sangerhausen
325	AA Brühl	627	AA Heilbronn	048	AA Stendal
327	AA Coesfeld	631	AA Karlsruhe – Rastatt	093	AA Erfurt
331	AA Detmold	634	AA Konstanz – Ravensburg	094	AA Altenburg – Gera
333	AA Dortmund	637	AA Lörrach	095	AA Gotha
337	AA Düsseldorf	641	AA Ludwigsburg	096	AA Jena
341	AA Duisburg	644	AA Mannheim	097	AA Nordhausen
343	AA Essen	647	AA Nagold – Pforzheim	098	AA Suhl
345	AA Gelsenkirchen	651	AA Offenburg	071	AA Annaberg-Buchholz
347	AA Hagen	664	AA Reutlingen	072	AA Bautzen
351	AA Hamm	671	AA Waiblingen	073	AA Chemnitz
353	AA Herford	674	AA Schwäb. Hall – Tauberb.	074	AA Dresden
355	AA Iserlohn	677	AA Stuttgart	075	AA Leipzig
357	AA Köln	684	AA Ulm	076	AA Oschatz
361	AA Krefeld	687	AA Rottweil – Villingen-Schw.	077	AA Pirna
364	AA Mettmann	711	AA Ansbach – Weißenburg	078	AA Plauen
365	AA Mönchengladbach	715	AA Aschaffenburg	079	AA Riesa
367	AA Ahlen – Münster	723	AA Bayreuth – Hof	080	AA Freiberg
371	AA Oberhausen	727	AA Bamberg – Coburg	092	AA Zwickau

Employment agency districts according to territorial status in 2014.

Chapter 4

Regional job-mobility, wage differentials and long-term earnings gains

joint with Philip Jung and Edgar Preugschat

4.1 Introduction

Wages in Germany exhibit large and highly persistent spatial dispersion with an average coefficient of variation of about 11% for the period from 1994-2014. Figure 4.1 shows the distribution of regional average wages relative to the mean wage in Germany for the years 1994 and 2014. The correlation of wages between these years is strong (0.94), and dispersion has been relatively stable.[1]

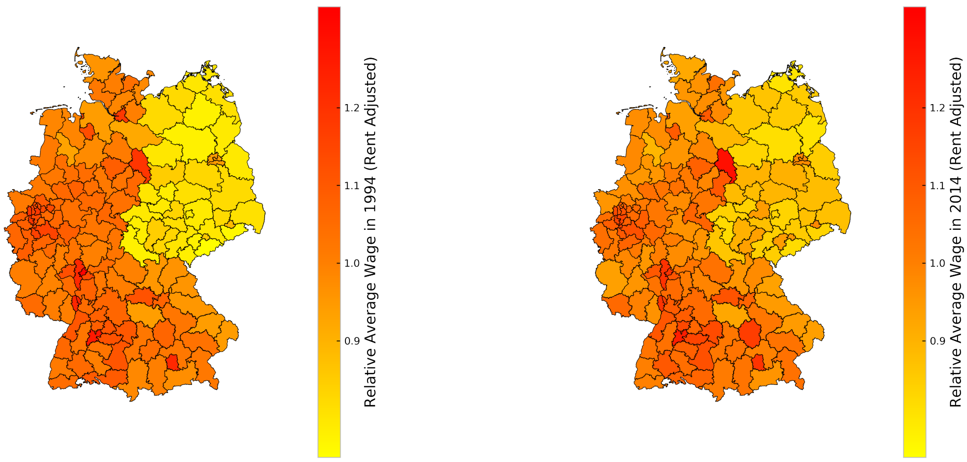
Given these substantial, persistent wage differences across regions, we investigate the characteristics of individual inter-regional migration decisions and the associated wage gains of employed workers. While there is evidence that differences in regional wage levels drive aggregate migration flows (e.g., Hunt, 2006, and Bauer et al., 2019 for the case of Germany) and cause significant variations in the lifetime earnings of movers (Kennan and Walker, 2011), the interactions between individual migration decisions, and aggregate migration outcomes, in particular, the relationship between individual wage gains and regional average wages have not been studied much.

To study migration incentives and outcomes, we use a large German administrative panel data set, which allows us to follow individual workers across relatively fine-grained work locations and over the life-cycle. We find, first, that most workers never leave their initial location, and only a small fraction of workers moves each month. The moving rate sharply declines with age, and most workers, conditional on relocating at all, change their work region only once or twice throughout working life. Moreover, most moves are to regions that are within 100 kilometers of the initial location. Second, workers do realize permanent wage gains relative to stayers of 10%. These gains increase in the average wage of the destination, hence the regional wage level is an essential

[1]The dispersion has slightly decreased over time, which can be attributed to a catch up of eastern regions. The average wage difference between East and West Germany decreased within this period by 10 percentage points. The coefficient of variation has decreased from 0.13 in 1994 to 0.10 in 2014.

determinant of the size of earnings gains from mobility. However, at the same time, a majority of moves is directed to regions that have a lower wage as the origin location. This lack of aggregate directness of migration helps to explain the long-term stability of the interregional wage dispersion.

Fig. 4.1: The relative wage distribution in 1994 and 2014



(a) 1994 (b) 2014

Notes: This figure displays rent deflated wages relative to the average wage in employment agency districts in Germany for the cross-section of workers in 1994 and 2014. Wages are deflated using regional rent data with a cost-share of 24% (see Section 4.C.3 for a description of the data and adjustment procedure). The three employment agency districts of Berlin are treated as one.

As individual reallocation is often related to a new job opportunity (Kaplan and Schulhofer-Wohl, 2017, and Schmutz and Sidibé, 2019) we focus on inter-regional job-to-job transitions. We document and analyze job-related regional mobility and the corresponding earnings gains using German social-security micro-data: the employment panel of integrated employment biographies. The spatial unit of analysis is the employment agency district (in the following EAD, district, or region). We follow two cohort groups over their working life—first, workers aged 25 in 1980 – 1984. Second, workers aged 25 in 1994 – 1996. The first group of workers allows us to document mobility over the whole working life. The second group captures mobility patterns when wages are much more dispersed – due to the reunification of Germany.

In the first part of the empirical analysis, we provide a comprehensive overview of the job-related migration behavior in Germany. Our new findings show that the individual intensity of reallocation is highest during the first ten years of working life, as almost half all of inter-regional transitions take place between ages 25 and 34. However, the individual migration frequency is low: The average migrant relocates one or two times. Also, the majority of workers do not relocate to another district. Concerning differences between our cohort groups, we observe an increase in the intensity of regional mobility: While 40% of workers of the 1980s cohorts have left their initial region at age 54, the same share of the 1990s cohorts has migrated at least once at age 44. In addition to the individual intensity of migration, we document the life-cycle profile of the individual moving probability and the allocation of movers by distance to the destination. Our results show that mobility decreases by age: The average probability of migrating to another region is more than twice as high for younger as older workers. In terms of distance moved, we find that most destinations are within a radius of less than 100 km of the home district. The distribution of moves is not directly related to wages (based on a logit model), and the distance distribution of relocations stays stable across age. Furthermore, we observe no tendency to move to a district that exhibits higher average wages: half of the inter-regional transitions are to regions with a lower average wage. Our results suggest that regional wages have little informative value concerning the distance moved or destination chosen. These findings raise the question of whether workers take region-specific wage differentials into account in their moving decision. Are regional wage levels at all an indicator of prospective earnings gains?

We answer this question in the second part of the paper. Specifically, we complement the prior literature by estimating the relocation and earnings gain pattern across regions conditional on average wage differences between destination and origin districts. When studies analyze gains conditional on the destinations of migrants, they, for the most part, focus on regional characteristics such as urbanization (Lehmer and Ludsteck, 2011, and Huttunen et al., 2018). The average wage level correlates with the degree of urbanization, but this correlation is only moderate (0.58) across employment agency districts in 2014.[2]

Our non-parametric, long-term gain measure is based on the earnings loss literature (Jacobson et al., 1993, and Couch and Placzek, 2010). We match inter-regional movers to a control group to account for their counterfactual earnings path, given no inter-regional transition would take place. This control group consists of individuals that never leave their initial district ("stayers"). One specific advantage of our cohort approach is that we can sort individuals into groups of stayers or movers based on their migration decisions over their whole working life. The gain from mobility is the difference in accumulated income between movers and stayers over the ten years that follow the moving event. To account for potential selection bias, we control for between-group earnings differences before the moving event.

The estimated long-term earnings gains of inter-regional job-changers, relative to the earnings of stayers, are substantial and amount to 12%, on average. To measure the impact of wage levels on mobility gains, we assign districts to quartiles based on a districts' rank in the average wage distribution. While gains do not vary conditional on the origin districts' rank, we find that they increase in the rank of the destination

[2] We use rent deflated average wages. Employment agency districts are sorted into three categories of urbanization based on the classification scheme "Regionstyp" of the Federal Institute for Research on Building, Urban Affairs and Spatial Development (BBSR). The correlation in 1994 is 0.60.

district: they range from 4% to 16%. Moreover, permanent gains differ considerably depending on the direction of the transition within the distribution of wages across spatial units: they increase substantially with the wage differential between origin and destination district. Here, the estimated gains range from 6% to 18%. These results suggest that the wage-level is a crucial determinant of the returns to mobility. Our findings are robust to an adjustment for regional rent differentials, are not driven by the gains from job-to-job mobility and are not education specific.

We further refine our analysis and estimate gains and moving probabilities in dependency of the quartile of origin and destination district in the average wage distribution. We find that a mover can increase his earnings more, the lower the rank of the district of origin, and the higher the rank of the destination district. Permanent earnings gains differ by almost 30% when we compare a move from the lowest to the highest quartile with a transition in the opposite direction. We also uncover substantial and persistent earnings losses for movers who relocate to districts with low average wages. Most informative are the relative moving probabilities as the majority of transitions are upward or to a district ranked in the same wage quartile. Hence, the average wage differential of a transition to a district with a lower average wage is smaller than the differential of a transition to a district with a higher wage. Thus, it is crucial where a worker comes from and migrates to. However, the fact that a considerable proportion of inter-regional job transitions are to a district with a similar or lower rank, which entails a significantly lower long-term income gain, suggests that frictions play an important role in relocation decisions, an issue which we analyze in a companion paper using a structural model.[3]

Literature Internal migration has been studied extensively but evidence on individual life-cycle mobility at a fine spatial level is scarce.[4] For Germany, a substantial share of studies focused on the migration flows between East and West. During the consecutive years after the reunification, a large number of East Germans, most of them young, moved to West Germany—by 2001 the net number of migrants was 1.3 million. While their motives were manifold, the higher possible income in western regions is seen as the main motivation to leave (Krueger and Pischke, 1995, Burda and Hunt, 2001, Hunt, 2006, and Brücker and Trübswetter, 2007).

Most of the subsequent literature analyzes migration pattern of adults across federal states. They document migration propensities by age brackets or education categories (Smolny and Peukert, 2012, Arntz, 2010, and Arntz et al., 2014), or analyze the motives of repeated movers (Hunt, 2004). At a more disaggregated level Bauer et al. (2019) show that young workers constitute the largest group of migrants and highlight age-varying

[3] See also Desmet and Rossi-Hansberg, 2013, and Schmutz and Sidibé, 2019 for related approaches.
[4] See Greenwood (1997) for an overview of the literature.

incentives of migration. Bonin et al., 2008 use retrospective survey data and document that 77% of individuals changed their residence at least once after leaving the parental home. However, they cannot distinguish between inter-regional moves and relocations within one city. Haas (2000) analyses cross-sectional job-related inter-regional mobility throughout 1980 and 1995. She documents the importance of employer-to-employer relocation and finds an increase in aggregate mobility within her sample period. We add to this string of the literature as we provide a detailed life-cycle profile of the individual intensity of job-related migration.[5]

We contribute to the literature on gains from internal migration. In general, inter-regional migration is found to go hand in hand with considerable wage gains, especially for job-to-job transitions (Glaeser and Maré, 2001, and Yankow, 2003). Johnson and Schulhofer-Wohl (2019) find no realized return to mobility across all types of migrants, while they do not consider the reason for migration. For Germany, substantial differences in wage growth for transitions between rural and urban areas or East and West Germany are well documented (Lehmer and Ludsteck, 2011 and Heise and Porzio, 2019). This paper is closest to the work of Lehmer and Ludsteck (2011), who show that job-to-job transitions across regions lead to larger wage growth compared to within region transitions. They analyze differences in wage growth of inter-regional job-to-job movers by four region types, i.e., the degree of urbanization. Besides the different focus, our gain measure differs in several aspects. First, we control for the unobservable heterogeneity between movers and stayers before the moving event. Second, their control group consists of individuals who stay within their region-type for one year, while we sort individuals into stayer and mover groups based on their long-term relocation behavior and focus on the returns to the first inter-regional transition. Last, we explicitly adjust wage gains for differences in local housing costs.

A share of articles identifies wages as one of the main regional factors that explain aggregate migratory movements in Germany (Mitze and Reinkowski, 2011, and Bauer et al., 2019). We connect to this work by documenting how individual migration decisions are related to regional wage differences.

Another strand focuses on the aggregate effect of migration on regional disparities in Germany − predominantly on differences in wages and unemployment rates. Or on the determinants of the spatial wage gap in Germany.[6] Our empirical results are silent on the aggregate equilibrium effects of migration. The lack of aggregate directness of migration that we document, however, helps to explain the long-term stability of the

[5] Hunt (2004) and Fackler and Rippe (2017) provide evidence on the inter-regional mobility of the (newly) unemployed

[6] For example, Niebuhr et al. (2012) and Fendel (2016) analyze the effect of migration on regional unemployment rates and wages. Heise and Porzio (2019) focus on the persistence of spatial wage gaps in the light of migration with frictional labor markets. Card et al. (2013) and Dauth et al. (2019) focus on the determinants and persistence of wage gaps.

interregional wage dispersion.

The article is organized as follows. Section 4.2 describes the data and cohort selection. In section 4.3, the life-cycle regional mobility patterns are documented. Section 4.4 lays out the empirical strategy and presents the results. Section 4.5 concludes.

4.2 Data and cohort selection

Our primary data source is the employment panel of integrated employment biographies (SIAB) provided by the Institute for Employment Research (IAB). The panel is a two percent random sample of all workers who are unemployed or subject to social security contributions and covers the years $1975 - 2014$.[7]

The spatial unit of analysis is the employment agency districts. We merge the three Berlin employment agency districts because the boundaries of the districts have often been modified. Each district covers three municipalities (Kreise, NUTS3), on average. The territory of each EAD is based on the territorial borders of 2014. There are 119 and 35 EDAs in West and East Germany, respectively. The social security data provide us with information on the municipality of an individual's workplace. One limitation of our study is that information on the place of residence for employed is only available from 1999 onward. Hence, we cannot distinguish if a worker, who makes a job-to-job transition, migrates, or decides to commute. To obtain consistent data on regional transitions for the whole observation period, we assign the district of the last employer as the current place of residence when an individual becomes unemployed. Hence, regional unemployment to employment transitions are identified based on the EAD of the previous and the current employer.[8] Similarly, a job-to-job transition is marked as an inter-regional move if the new and old establishments are located in different EADs. As the SIAB data contains information on establishments and not on firms, the inter-regional job-to-job transition might be a transition within one firm.

The panel contains daily information on each employment and unemployment spell. These daily employment histories are aggregated to a monthly frequency. We strictly follow Jung and Kuhn (2014) and use predefined reference weeks to assign labor market states to each worker. Furthermore, we recover monthly flows in and out of unemployment as well as from employment to employment. In general, individuals with no information on their employment status or geographic location are excluded. The classification of labor market states is based on a hierarchical order in which employment ranks above unemployment, which in turn ranks above non-participation. Workers are defined as employed if we observe a full-, part-time, or apprentice employment notifica-

[7]Hence, self-employed workers or civil servants are not included. Overall, the sample covers approximately 80% of the labor force

[8]This limits our analysis and will overstate the fraction of inter-regional movers.

tion. Employment relationships that are inactive (e.g., maternity leave) or that are in marginal employment are excluded if they do not have a parallel unemployment spell. From 2000 onward, a worker is marked as unemployed if he is registered as unemployed at the federal employment agency. For the years before 2000, we use the information on the unemployment benefit recipient status to construct prior histories.

The data contains information on daily wages when employed as well as benefits received during unemployment. A contribution limit to the social security system leads to top-coding of wages. Following Gartner (2005), we impute all censored wages at the social security threshold in each year. Wages and unemployment benefits are adjusted to 2010 prices using the monthly CPI time-series of the Federal Statistical Office.[9]

Cohort selection We construct two samples and track individuals over their life-cycle. The primary analysis is restricted to men. For West Germany, the sample consists of five cohorts aged 25 within 1980 and 1984. We will refer to this sample as the 1980s cohorts. Here, we track individuals up to age 54. We exclude individuals with spells in East Germany or Berlin from this sample. For reunited Germany, we follow three cohorts of workers aged 25 within 1994 and 1996 − the 1990s cohorts.[10] Here, we follow individuals from age 25 up to age 44. We start at age 25 because male university graduates will most likely enter the labor force at this age. Moreover, individuals in apprenticeships are severely limited in mobility for three years, and the earnings are highly regulated (cf., Johnson and Schulhofer-Wohl, 2019, and Lehmer and Ludsteck, 2011).

Additional data Information on offered rents on the county level comes from the Federal Institute for Research on Building, Urban Affairs and Spatial Development and IDN ImmoDaten GmbH and covers the period 2004 to 2014. We aggregate the average rent data to the district level using population weights and predict the rent level for the period 1994 to 2003. A detailed description of the estimation procedure and the additional data utilized is provided in the appendix (Section 4.C.3)

[9]"Preise, Verbraucherpreisindizes für Deutschland, Lange Reihen ab 1948", 2018, Statistisches Bundesamt (Destatis). Although wages are top-coded, there are a few cases where the reported wage is above the threshold and hence is not imputed. We drop all wage observations that are higher than 95% of our imputed wages. As a result, we lose approximately 0.04% of our observations.

[10]Cohorts that enter the labor market within 1992 and 1993 are not included because prior data is not reliable. However, information on employment and unemployment spells in 1992 and 1993 is used to impute location of occupation variables of the later cohorts if applicable.

4.3 Inter-regional transitions

This section provides a comprehensive overview of inter-regional migration over the life-cycle in Germany. We focus on migrants that relocate to a different employment agency district in the course of a job change and document the transition frequency, the life-cycle moving probability, and the allocation of movers by distance to their destination. Our primary focus lies on full-time employed workers in their prime age, while we include the unemployed in parts of the following analysis.

4.3.1 Frequency and probability of moving

We document two important margins of job-related relocation: the probability to migrate by age and the number of completed transitions at different ages. While the former describes the overall propensity to migrate at a certain age, the latter captures the intensity of individual relocation.[11]

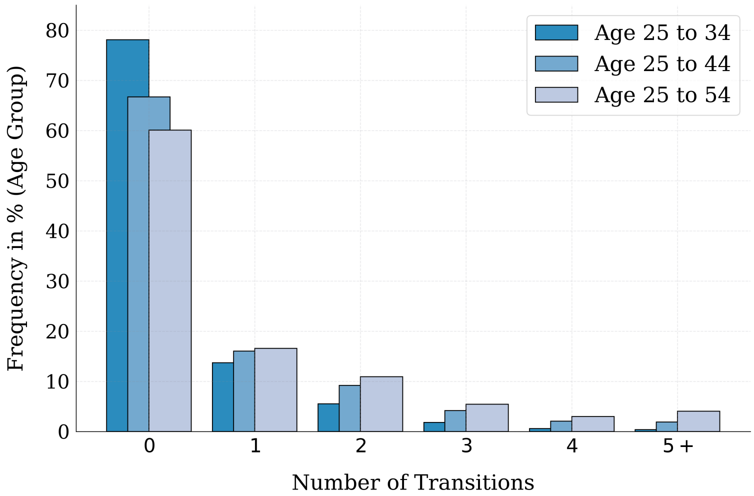
Frequencies of completed migrations Figure 4.2 shows the frequencies of inter-regional transitions into full-time employment over the life-cycle as well as for the first ten and twenty years after the age of 25 for the 1980s cohorts. We sum up the individual number of completed migrations up to ages 34, 44, and 54, and consolidate workers with more than four transitions in the 5+ category.

The resulting frequencies show that between ages 25 and 34, almost 20% of workers relocate to a different employment agency district. Conditional on migrating above three-quarters of workers have completed one or two inter-regional transitions at age 34, where the majority changed regions exactly once. Thus, only a small fraction relocates to a different region more than two times. Furthermore, the fraction of first-time movers reduces substantially by age: Between ages 35 and 44, less than 12% of workers leave their initial district. In the subsequent ten years, this number decreases to 6%. At age 54, almost 40% of workers migrated at least once. Moreover, also the migration intensity decreases as individuals age: More than half of all inter-regional transitions take place up to age 34, and 80% occur between ages 25 and 44. Overall, the majority of movers relocate one or two times.

Workers of the 1990s cohorts are more mobile than those of the 1980s cohorts as slightly less than 40% relocated at least once at age 44. This higher mobility is not due to the addition of East Germany as the frequency pattern between workers in West and united Germany is similar. These findings indicate that the higher aggregate mobility

[11] Individuals with an inter-regional transition − out of employment or unemployment − before age 25 are dropped from the sample. We control for attrition, i.e., individuals must have none missing observations at ages 25, 34, 44, and 54. The unemployeds' place of living is identified as the last recorded place of work.

Fig. 4.2: Frequency of inter-regional transitions (1980s cohorts)



Notes: This figure displays the frequency of completed inter-regional job-to-job or unemployment to employment transitions up to ages 34, 44, and 54 for the 1980s cohorts.

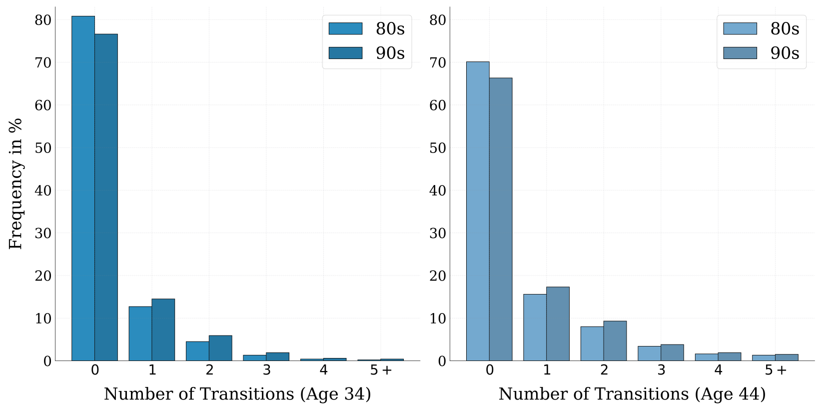
in 1995 compared to 1980, which is documented byHaas (2000), is driven by a larger fraction of movers and a higher frequency of relocations.

We observe considerable differences in completed migrations across educational levels. Individuals with at least a university entrance exam leave their initial region more often: Almost 56% of high skilled workers did move at least once within ages 25 to 54, compared to 35% of the low skilled (Table 4.1). Furthermore, the frequency pattern by education is similar between cohort groups: Almost half of high and 30% of low skilled workers change regions at least once before the age of 45. Nevertheless, the majority of high and low educated workers move one or two times (Table 4.2).

As our main focus lies on inter-regional employment to employment transitions, the next statistics focus on this type of migration. We find that inter-regional job-to-job transitions are an essential component of the gross flows across employers: on average, they account for 37 and 42% of all job-to-job transitions for the 1980s and 1990s cohorts, respectively. This share is almost constant over the life-cycle. Similar results have been reported by Haas (2000), who finds that for the cross-section and at a more aggregated regional level 25% to 31% of employer-changes coincide with an inter-regional relocation of the worker. Moreover, we find that 80% (79%) of all job-related migrants of the 1980s (1990s) cohorts switch employers. Thus, migration out of employment is the most crucial driver of job-related relocation.

Figure 4.3 displays the observed frequencies of inter-regional job-to-job transitions for the 1980s as well as the 1990s cohorts. The profile is similar compared to all job-related migrations as employer-to-employer transitions account for the majority of moves. Overall, both cohort groups show the same pattern: Most workers move one or two times, and most inter-regional job changes take place within the first ten years of working life.

Fig. 4.3: Frequency of inter-regional job-to-job transitions (1980s and 1990s cohorts)



Notes: This figure displays the frequency of completed inter-regional job-to-job transitions up to ages 34, 44, and 54 for 1980s (West Germany) and 1990s cohorts (Germany).

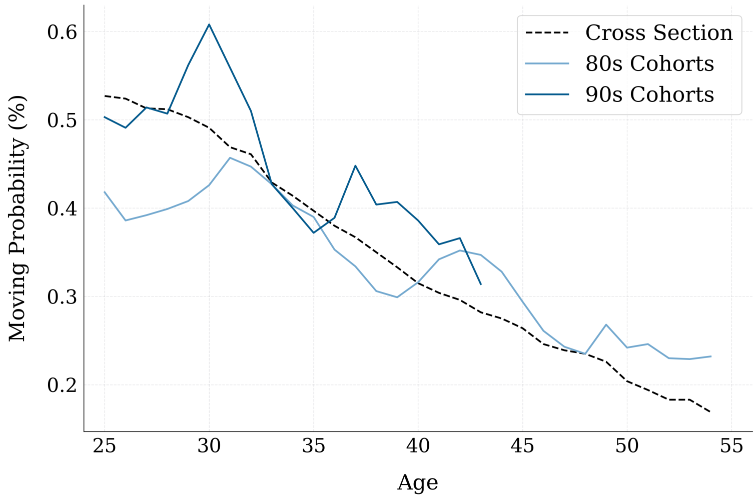
In general, we find that a sizable fraction of workers stay in their initial employment agency district throughout their working lives. Caliendo et al. (2017) state that almost 70% of prime-aged workers have not left their birth state. Concerning the frequencies of completed residential migration (i.e., including within region relocations) Bonin et al. (2008) find that 77% of workers did change their place of residence at least once after leaving the parental home and the majority changed dwellings two to four times. Our results suggest that inter-regional job-related relocation plays a minor role in overall residential migration as individuals relocate 0.86 times, on average, (1980s cohorts at age 54).

Probability to move Beyond the individual intensity of relocation, we compute the probability of moving by age. Figure 4.4 displays the results. We plot the annual average of the monthly moving probability of full-time employed workers of the 1980s and 1990s cohorts as well as the cross-section. In general, the probability pattern of the cohorts tracks closely that of the cross-section. Similar to the results presented before, mobility is higher for workers of the 1990s compared to the 1980s cohorts. Moreover, the probability of moving declines by age. The declining tendency to relocate by age is consistent with the literature (Greenwood, 1997). Similarly, Bauer et al. (2019) find that 18 to 29-year-olds constitute the largest share of migrants.

Concerning difference by education, we find that highly skilled workers move more often, compared to low skilled. Also, the moving probability by education declines over the life-cycle for both education groups and cohort groups (Figure 4.2). Higher geographical relocation of more educated workers is documented by Balgova (2018) for the U.S..

As unemployment to employment transitions only account for a minor share of inter-

Fig. 4.4: Inter-regional moving probability by age



Notes: This figure displays the annual average of monthly inter-regional job-to-job transition probability for full-time employed workers: 1980s and 1990s cohorts as well as West German cross-section (1980-2014).

regional transitions, the inclusion of the unemployed leads to a slight rise in the average inter-regional moving probability (Figure 4.1). However, the fact that we observe an increase at all suggests that unemployed are more likely to relocate, which is in line with previous literature (e.g., Hunt, 2004).

4.3.2 Distance to destination

Individual inter-regional migration decisions are known to be influenced by the distance to the destination due to, e.g., moving costs or regional attachment. We compute the distance between the centers of the origin and destination district and sort the job-to-job transitions into four distance classes. Table 4.5 displays the share of movers for each class. In general, individuals predominantly move to districts close to their previous location. The average median distance is 84 and 97 km for the 1980s and 1990s cohorts, respectively. For the 1990s cohorts, we observe a decrease in the lowest distance class. This decrease is due to the addition of East German workers, which, on average, relocate to farther away regions. The median distance for West German workers is almost constant between cohorts (Table 4.5). Similarly, Bauer et al. (2019) find that migration flows decrease by distance.

We employ a logit model to rationalize the distribution of movers over distance classes, given the wage distribution. The known part of the utility that a mover – employed in district i – obtains from a transition to district k is denoted by V_{ik}. We assume that choices depend solely on average wages $V_{ik} = w_k$. The logit choice

Table 4.1: Proportion of movers by distance to the destination

	< 65 km	65-129 km	130-259 km	≥ 260 km
1980s cohorts	31%	30%	17%	22%
1990s cohorts	27%	31%	17%	26%

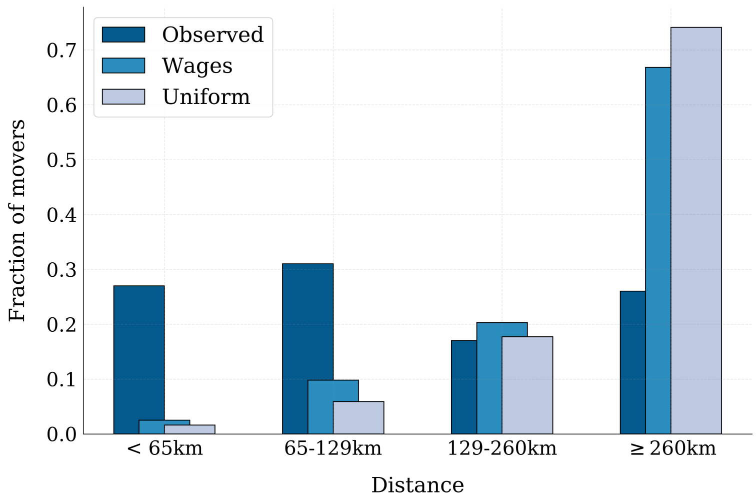
Notes: This table displays the average proportion of inter-regional job-to-job transitions by distance class for the 1980s (aged 25 to 54) and 1990s (age 25 to 44) cohorts.

probability is

$$P_{ik} = \frac{e^{V_{ik}}}{\sum_j e^{V_{ij}}}.$$

Moreover, we compute the distribution of movers when all districts are selected with equal probability. The respective choice probabilities are weighted by employment. Figure 4.5 displays the resulting distributions. A uniform choice probability implies that more than 70% of transitions are to a destination farther away than 260 km. When we account for wage differences, the share of moves in the highest distance category decreases only slightly. Thus, wage differentials alone cannot explain the large share of moves to a close-by destination.

Fig. 4.5: The distribution of inter-regional transitions by distance



Notes: This table displays the observed and predicted distributions of movers across distance classes. The observed values are based on the 1990s cohorts. "Uniform" indicates that each district is selected with equal probability. The distribution that takes into account regional wage differences wages is predicted using the 1990s cohorts' wages at age 25.

We present additional evidence on wage gains by distance and differences in distance moved between East and West German workers in the appendix (Section 4.B). The gains from inter-regional transitions are positive and decrease over the life-cycle for all distance classes. However, immediate wage gains do not differ much conditional on the distance between origin and destination district. This result indicates that the distance moved is not informative for the size of wage gains.

In general, the results suggest that wage differentials cannot explain the allocation of migrants across space. While the wage distribution predicts an allocation of workers that is substantially more skewed to far away districts, the absence of wage gains that increase with the distance indicates that relocation costs could potentially account for a substantial fraction of migrations to close regions.

4.4 Earnings gains from inter-regional transitions

Most studies identify wage levels as one of the principal explanatory factors for migration movements. The fruits of targeted mobility – in the form of wage gains – and their dependence on the wage level of origin and destination district of movers are the focus of this section. First, we document the immediate wage gains from moving. Second, we lay out the methodology applied to estimate long-term earnings gains from mobility. Third, we analyze the relationship between individual relocation behavior, gains from mobility, and regional wage differentials.

4.4.1 Immediate wage gains

In this subsection, we document the immediate gains from an employer to employer job-change and distinguish between transitions within and between regions. Table 4.2 shows the average (log) difference in wages earned in the month before and after the job-change. The resulting wage gains decreases by age for intra- and inter-regional transitions. We find a small mobility premium for workers that change jobs and regions: Between job wage growth is higher, or similar, for inter-regional moves up to the age of 44 for both the 1980s and 1990s cohorts. In contrast, the oldest worker group gain 0.6% more from an intra- than an inter-regional transition. At this age, wage growth between jobs, however, is very low.

Concerning the direction of a move, the proportion of movers to a district with higher average wages is almost 50% (last row). Thus, we observe no tendency to move to a better-paying region. This finding is striking, as the immediate wage gains are higher when workers migrate to a district with a higher wage level: Workers' wages increase by about 3% more compared to a downward movement (Table 4.3).[12] Given these substantial differences in immediate gains, we would expect that a larger share of migrants relocates to districts with higher wages. Moreover, destination We investigate the relationship of earnings gains and wage differentials between origin and destination districts in the following subsections. The immediate gain measure, however, is not

[12]This result is not driven by later migration decisions that prove to be motivated by other factors as the immediate gains are similar when we analyze the first transition of a mover.

Table 4.2: Immediate wage gains: Inter- and intra-regional job changes

	Job-to-Job transitions			
	80s cohorts		90s cohorts	
Age	Inter-regional	Intra-regional	Inter-regional	Intra-regional
25 to 34	0.10	0.08	0.09	0.07
35 to 44	0.07	0.06	0.05	0.05
45 to 54	0.02	0.03	–	–
Upward transitions	49.6%		49.5%	

Notes: This table displays average immediate (log) wage gains of inter- and intra-regional job-to-job transitions for the 1980s and 1990s cohorts. An upward transition represents a job-change to a district with a higher average wage. The values are rounded

suited for this analysis as it does neither capture the persistence of earnings gains nor account for the prospective earnings in the case of not moving.

4.4.2 Long-term earnings gains: Methodology

This section describes the methodology applied to estimate the long-term returns to job-related migration. We follow the literature on earnings losses of displaced workers (Jacobson et al., 1993, and Couch and Placzek, 2010) and measure gains from inter-regional job-to-job transitions as the difference between an individual's earnings during a period following and before the moving event. In the benchmark analysis, the periods before and after the moving event are set to 12 months and ten years, respectively. Our focus lies on the long-term earnings gains of the first inter-regional job-to-job transition. This focus is motivated by the fact that most individuals relocate only once, and subsequent relocations might be driven by other motives (e.g., return migration (Kennan and Walker, 2011)).

We control for the mover's counterfactual earnings path (i.e., given no migration would take place) using a non-parametric matching approach: Each mover is paired with a group of workers that never leave their initial district ("stayers"). In our benchmark-setting, this comparison group is of the same age, gender, and belongs to the same cohort. Furthermore, movers and stayers are full-time employed in the same district at the time of the move.

To ensure the reliability of the earnings measure, we apply the following restrictions. A mover is full-time employed for 12 months before the moving event — moreover, the mover transitions to a full-time job in a different employment agency district. We require non-missing observations for the entire period, before or after. For the period following the transition, we include income out of part-time employment or unemployment in the earnings measure. Also, we allow for subsequent mobility. Therefore, the estimated gains capture the returns to initial mobility but may include gains (or losses) from multiple transitions.

Similar restrictions apply to the group of stayers: Stayers are full-time employed for the 12 months before the moving event. They are of the same age, employed in the same district, and belong to the same cohort as the respective mover. We require non-missing observations for the 11 years of comparison and include earnings out of part-time employment or unemployment in the stayers' accumulated incomes during the years following the mover's reallocation.

Let $t = \delta$ denote the month before the transition, and ξ the joint index of age a, district j, cohort c, and δ. For each mover i, we compute the accumulated monthly earnings during 12 months before (4.1) and during ten years following (4.2) the month of the moving event:

$$e_{i,\xi}^{p,m} \equiv \sum_{t=\delta}^{T=\delta-11} w_{i,t} \qquad (4.1) \qquad\qquad e_{i,\xi}^{f,m} \equiv \sum_{t=\delta+1}^{T=\delta+120} w_{i,t} \qquad (4.2)$$

Subsequently, we match each mover to the group of stayers s that are of the same age a, belong to the same cohort c, and are at the time of the move δ full-time employed in the mover's origin district j for at least the past 12 months. The average accumulated income over the past twelve months (following ten years) across all stayers in district j, of age a, from cohort c at time δ (i.e., ξ) is denoted by $\bar{e}_{\xi}^{p,s}$ ($\bar{e}_{\xi}^{f,s}$). To calculate this average, we require a minimum of three stayers.

The earnings difference between a mover and the corresponding group of stayers before mobility is given by

$$g_{i,\xi}^{p} = \ln\left(e_{i,\xi}^{p,m}\right) - \ln\left(\bar{e}_{\xi}^{p,s}\right). \qquad (4.3)$$

This measure captures the difference of observable and unobservable income determinants between a mover and the group of stayers during the year before the mover leaves his initial district.

The difference in accumulated income compared to stayers during 10 years after the move is

$$g_{i,\xi}^{f} = \ln\left(e_{i,\xi}^{f,m}\right) - \ln\left(\bar{e}_{\xi}^{f,s}\right). \qquad (4.4)$$

For example, earnings differences following the transition of $g_{i,\xi}^{f} = 10\%$ suggest, for average accumulated earnings of stayers of 240,000 Euro, that the average monthly difference in earnings between the mover and the average stayer amounts to 240 Euro.

The difference between the mover's relative accumulated earnings before and after the move determines the long-term return to mobility

$$\Delta g_{i,\xi} = g^{f}_{i,\xi} - g^{p}_{i,\xi}. \tag{4.5}$$

This measure captures the joint premium from regional and job-related mobility. It explicitly accounts for earnings differences before the move (i.e., the unobservable differences between mover and the average stayer) and for the counterfactual earnings path in the event of no mobility. When we present aggregate versions of $\Delta g_{i,\xi}$, we first take averages across individuals that are of the same age, leave the same district, and belong to the same cohort (i.e., relocate the same year). Then, we calculate the averages over cohorts and obtain the long-term mobility gain by district and age. Subsequently, we take averages over districts and, last, over age.

4.4.3 Long-term earnings gains: Results

In this section, we show that variations in average wages across districts lead to considerable differences in long-term mobility gains. These gains positively depend on the average wage differential between origin and destination district. The measure utilized here controls for the mover's counterfactual earnings path, given no relocation takes place. We approximate this counterfactual path by the earnings path of individuals that never leave their initial district. g^p captures the differences in earnings between mover and stayers before the transition. g^f measures the difference in accumulated earnings between mover and stayers after the movers relocation. The long-term, permanent earnings gain measure Δg, thus, captures the gain from inter-regional job-to-job mobility.

In the subsequent analysis, we focus on the 1990s cohorts and follow movers aged 25 to 34 for ten years after their first inter-regional transition.[13] We find substantial returns to mobility: while the average mover earns 2% less than stayer before migrating, the accumulated earnings difference during the ten years after amounts to 12%. Thus, the long-term return to mobility is, on average, 14%.

To illustrate the gain pattern, we group districts based on two characteristics: average wages and location. First, we sort districts into quartiles that are determined by the district's rank in the distribution of average wages.[14] For example, the quartile "1" includes the 25% of districts with the lowest wages, while quartile "4" includes the 25% of districts with the highest wages. Second, we divide Germany in East and West and analyze gains from inter-regional transitions within and between these two regions.

[13]Note that 34 is the last age where we observe movers for at least 10 subsequent years.

[14]Specifically, the district's rank in the distribution of averages wages of full-time employees across all spatial units in the year of each transition. Hence, we allow districts to change quartiles over time. This should not influence the results as wages are highly persistent.

Regional wage differentials Table 4.3 displays the differences in accumulated earnings prior and after the first inter-regional job-to-job transition (g^p, g^f) and the associated long-term gain (Δg) of workers aged 25 to 34.

Column 2 presents the earnings measures conditional on the mover's origin district's rank. In general, the average long-term gain (upper panel) is substantial and amounts to at least 13%. First-time movers that start in the lowest quartile exhibit higher gains, compared to movers from other quartiles. However, we observe no correlation between earnings gains and wage quartiles. The earnings difference before the move (g^p, center panel) is slightly positive for the lowest quartile and negative for the other three. Differences in accumulated earnings within ten years after the move are displayed in the lower panel and show a similar pattern as the long-term gains. We conclude that the information value of the district of origin about the size of the prospective earnings gains is low.

The earnings gains conditional on the destination districts' rank in the wage distribution are displayed in Column 3. We find increasing long-term gains as the quartile of the destination district increases: A move to a district with relatively high average wages, unconditional on the origin district, leads to higher permanent wage increases. Thus, the district component plays a significant role in the determination of the size of the long-term gain. Long-term gains range from 4% to 16% when the transition is to a district ranked in the lowest or highest quartile, respectively. The income differences before the move show that almost all movers earned less than the average stayer. The exception is movers to the highest quartile as their relative earnings prior to the move are higher. Similar as the long-term gains, the earnings differences after the move increase in the destination districts' rank. They range from 2 to 20%.

To document the gains pattern conditional on the wage differential, we group movers into four quartiles according to the average wage difference between their destination (i) and origin (j) district ($\Delta\bar{w}_{i,j} = \bar{w}_i - \bar{w}_j$) at the time of the move. $\Delta\bar{w}_{ij}$ is negative (positive) if the average wage in the destination (origin) district is higher as the wage in the origin (destination) district. Column 4 of Table 4.3 presents the results. As approximately 50% of inter-regional transitions are downward moves, the quartiles 3 and 4 are associated with a move up, while 1 and 2 correspond to a move down. Independent of whether the move is up- or downward, the permanent gain Δg is at least 6%, relative to not moving. It increases up to 18%, as the wage difference between origin and destination district increases. The relative earnings difference before the move increases, as the individual moves to better-paying districts. The differences afterward show an even steeper increase, compared to the long-term gains, ranging from 3 to 24%. Overall, the results show that the direction of the transition within the average wage distribution − or the wage differential between origin and destination district − plays a major role for individual returns to mobility.

Table 4.3: Earnings gains from inter-regional transitions

Percentile	Origin district	Destination district	Up or down
	Δg, permanent wage gain		
1	0.17	0.04	0.06
2	0.13	0.10	0.12
3	0.13	0.12	0.13
4	0.13	0.16	0.18
	g^p, prior to the move (1 year)		
1	0.04	−0.02	−0.03
2	0.00	−0.03	−0.01
3	−0.00	−0.03	−0.00
4	−0.00	0.04	0.06
	g^f, following the move (10 years)		
1	0.21	0.02	0.03
2	0.13	0.07	0.11
3	0.12	0.09	0.13
4	0.13	0.20	0.24

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year before the transition. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain – compared to not moving – is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

East-West differences Table 4.4 presents the results for transitions within and between East and West Germany. Movers within the West or East (first two rows) experience a permanent wage gain of 14 and 12%, compared to staying in their initial district. The relative earnings of the movers within the year before the transition are 1% lower compared to stayers, on average.

Furthermore, while relocating from West to East leads to a permanent income loss of 4%, a job-to-job transition in the other direction increases the wage by 17% (last two rows). The persistent lower earnings during the ten years after the move drive the permanent loss of west-to-east movers. On the other hand, movers from eastern to western districts earn 26% more than the average stayer in the ten years after the move. This vast earnings difference is well reflected by the average wage gap between

East and West Germany. However, the permanent gain amounts to only 17%, as the movers' earnings were substantially higher during the period before the move. The latter results indicate positive self-selection of East-to-West migrants (cf., Brücker and Trübswetter, 2007). The estimated returns are comparable to the results of Heise and Porzio (2019).[15]

Table 4.4: Earnings gains of movers within and between East and West Germany

	Earnings differences		
Direction	after (g_f)	before (g_p)	long-term wage gain (Δg)
W-to-W	0.13	−0.01	0.14
E-to-E	0.11	−0.01	0.12
W-to-E	−0.06	−0.02	−0.04
E-to-W	0.26	0.09	0.17

Notes: The table presents the accumulated monthly earnings differences between movers and stayers between and within West and East Germany for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34.

In general, we find that inter-regional job-to-job mobility leads to substantial long-term earnings gains − relative to not moving. Furthermore, these gains increase with the wage differential between origin and destination district. While we find no clear pattern in earnings gains by the rank of the origin district, the gains substantially increase in the destination districts' rank. Our results thus suggest a vital role for the district component concerning mobility gains. We perform several sensitivity checks to ensure the validity of these results. They are briefly discussed in the following paragraphs.

Jointly controlling for age, distance, and education We confirm the gain pattern across quartiles by estimating a linear least squares model where we control for education, the distance between origin and destination, and age. The estimated premium pattern is almost exactly similar to the results presented above (Table 4.6).

[15] Heise and Porzio (2019) estimate wage growth differences between movers across West and East Germany using local projections. Despite differences in the estimation approach and their focus on short-term wage growth, our results are of a similar magnitude concerning job-to-job transitions within West or East Germany. With respect to moves across regions, our estimates do differ. In part because they estimate cross-region movers' wage growth conditional on the birth region. Concerning East-to-West transitions, east (west) born workers gain 45% (25%). For West-to-East moves, east born workers' gains are negative after one year, but a worker born in West Germany experiences a 30% wage gain.

Concerning education, we find a substantial mobility premium for high skilled workers. Moreover, we include the log distance between origin and destination district. Similar to the results concerning the immediate wage gains, we find that distance has no impact on the long-term gains. Furthermore, corresponding to the prior literature, mobility gains decrease by age (Lehmer and Ludsteck, 2011).

Comparison to the 1980s cohorts The earnings gains for movers of the 1980s cohorts aged 25 to 34 (Table 4.16) show a similar pattern: They increase in the rank of the destination district and the wage differential between districts. The qualitative ordering of the permanent gains is similar, although the magnitude is lower than for the 1990s cohorts. We find negative accumulated earnings differences before the move. Furthermore, we observe no conditional gain pattern for movers within ages 35 and 44. However, the long-term gains for this age group are positive, on average (Table 4.17).

Timing changes The premium pattern is not based on different wage growth effects. The gain pattern five years after the moving event are similar (Section 4.C.2) to those after ten years. Furthermore, when we shorten the comparison period before the transition to six months, we find slightly more negative earnings differences. Overall, the estimated long-term gain is similar.

Rent adjusted gain measure The long-term differences in earnings could reflect the differences in housing costs, as movers relocate to districts with high average rents. Therefore, we utilize data on average rents per square feet to adjust individual wages. In general, the results do not change. Foremost, the differences in accumulated income between movers and stayers are increasing in the destination districts' rank. We observe the most substantial changes for transitions between West and East Germany: Moving to the east is associated with losses that are lower by 1 percentage point, while the gains from moving to West Germany are 6 percentage points lower. These outcomes reflect the lower rent level in the east. A detailed description of the results and the rent adjustment is provided in the appendix (Section 4.C.3). Our results indicate that rent prices have little influence on the size or pattern of earnings gains from mobility. These results correlate with those of Bauer et al. (2019), who find that regional rent prices have a small effect on inter-regional migration flows.

Intra-regional job-to-job transitions as comparison group The analysis of earnings gains of inter- relative to intra-regional job-to-job movers allows us to differentiate between the return from changing jobs and the return from moving to another region. A detailed description is provided in the appendix (Section 4.C.5). We find substantial long-term gains for inter-regional migrants, compared to intra-region job-to-job changers. The gains are of the same magnitude as the those relative to all stayers. The key difference is that inter-regional movers' monthly earnings before the transition are almost always higher, compared to workers that change jobs within their current region, and not lower as in the comparison to all stayers. However, during the ten years following the move, the relative permanent gains are higher, which leads to only slightly lower Δg. Hence, job-to-job returns do not solely drive the gain pattern, as the gains from mobility are substantial. The results are in line with Lehmer and Ludsteck (2011), who find higher gains for inter- compared to intra-regional job-to-job movers.[16]

Including moves out of unemployment Our results are qualitatively robust to the inclusion of transitions out of unemployment. Specifically, we include short-term unemployed in both stayer and mover groups. The adjusted results are presented in section 4.C.4. They differ only slightly from the benchmark results. The long-term wage gains Δg show an increasing/decreasing pattern by destination/origin district's rank in the average wage distribution. Arntz (2010) provides evidence that the wage level is a more important determinant for job-to-job compared to unemployment-to-job movers.

Controlling for education The gain pattern holds for individuals with no tertiary education (Section 4.C.6). The overall gain for low-skilled workers lies below the average gain presented above due to lower relative wage gains after the move.

Return migration The fact that a large share of inter-regional job-to-job movers return to their origin district is well documented (e.g., Kennan and Walker, 2011). As our measure is based on the first inter-regional transition of a worker and we do allow for additional moves, the long-term gain estimates could be influenced by return migrants. When we control for returnees, we find mostly higher permanent gains (Section 4.C.8). This finding is in line with the results of De la Roca (2017), who shows that return migrants exhibit lower earnings, compared to non-returnees.

4.4.4 Relocation across the wage distribution

We refine our analysis and estimate conditional transition probabilities and long-term gains for each destination quartile conditional on the origin districts' rank in the wage distribution. We uncover substantial permanent earnings losses for a small fraction of movers. The conditional moving pattern across districts reveals that inter-regional movers, indeed, take into account the size of the prospective earnings gains.

[16] Their focus lies on transitions between rural and urban locations and they do not control for the timing of the move.

The conditional moving probabilities and the associated permanent gains for the 1990s cohorts are presented in Table 4.5. Our results show that regional mobility leads to substantial earnings gains for most movers (left panel). We also observe significant long-term losses for movers to districts ranked in the lowest quartile. As this quartile predominantly includes East German EADs, these results mirror the differences in gains observed for moves between East and West Germany. The long-term earnings gain is increasing in the destination districts' rank, independent of the origin district: A transition to a higher-ranked EAD leads to a higher long-term gain. Conversely, a relocation to a lower quartile indicates lower earnings increases. However, these gains differ by origin district. In most cases, a lower rank of the origin district implies a higher earnings gain from an inter-regional job-to-job transition. These results, again, confirm that region-specific wage differentials are an important determinant of the returns to mobility.

Table 4.5: Transition and gain matrix (1990s cohorts)

		Δg						$Pr(j\|i)$			
		Destination						Destination			
		1	2	3	4			1	2	3	4
Origin	1	0.13	0.16	0.16	0.22	Origin	1	0.53	0.15	0.12	0.20
	2	-0.04	0.09	0.17	0.14		2	0.11	0.20	0.28	0.41
	3	-0.03	0.07	0.12	0.14		3	0.04	0.21	0.27	0.48
	4	-0.07	0.06	0.10	0.14		4	0.03	0.13	0.21	0.63

Notes: This table displays the average permanent wage gain Δg and the relative conditional moving probability $Pr(j|i)$ of first time movers aged 25 to 34 for the 1990s cohorts.

The probability of moving to quartile j given the current district is in quartile i is denoted by $Pr(j|i)$ and displayed in the right panel. Our prior results stated that almost half of all transitions are downward. The quartile view, however, provides a more detailed picture of moves across the average wage distribution: The major share of inter-regional transitions is either upward (31%) or to a district ranked in the same quartile (45%) where the latter category also includes transitions to districts with lower wages. Thus, the destination of a downward transition is more similar to the origin district, in terms of wages, compared to the destination EAD of an upward move. The lowest quartile constitutes a somewhat unique case. While the propensity is increasing in destination districts' rank for the other three quartiles, most individuals that start in the lowest quartile do not move up. This result is consistent with the findings of Heise and Porzio (2019). They document a strong home-bias: Individuals born in East or West Germany are more likely to relocate to a district in their home region.

While Mitze and Reinkowski (2011) and Bauer et al. (2019) provide evidence on the importance of wage levels for aggregate migration in- and outflows, we document how individual destination choices depend on a districts' average wage. Their results state that aggregate migration inflows increase and outflows decrease with a districts' wage. Our findings are complementary as we show that if a worker leaves his initial region, he will most likely relocate to a district with similar or higher average wages.

These results are robust to the exclusion of return migrants, adjustment for rent differentials, and to timing adjustments.[17] Also, the results for the 1980s cohorts are comparable. Here, the permanent gains for the 25 to 34-year-old workers show an increasing pattern in the destination districts' rank and a move to a lower-ranked district always leads to lower gains. Long-term earnings gains of first-time movers aged $35-44$ are positive but show no clear pattern across quartiles (Table 4.18). The tendency of West (East) German workers to stay in the West (East) is visible if one compares the transition probabilities presented in Tables 4.5 and 4.18. Eastern districts are predominantly ranked in the lowest quartile of the wage distribution. While for the 1980s cohorts almost 75% of inter-regional job-to-job transitions out of the lowest quartile were upward, only 45% of the 1990s cohorts movers left the lowest quartile. Furthermore, moves into the lowest quartile are more frequent for the 1980s cohorts. A comparison of transitions probabilities for workers with only spells in West Germany reveals that the relocation pattern of western workers did not change: More than 75% of movers of both cohort groups leave the lowest quartile. Transitions to the lowest quartile are also substantially higher (See appendix Table 4.19).

In general, the low inter-regional mobility in Germany, combined with the undirected average migration pattern, provides suggestive evidence concerning the causes of the persistence of regional wage differentials. Low mobility and the large share of relocations to neighboring districts are clear indicators for a hampered adjustment process. While the allocation of movers across the wage distribution presented in this subsection indicates that movers at least partially take into account the effect of the wage differential between origin and destination on their earnings gains, a large fraction still relocates to a district with lower wages. These results indicate that upward mobility might be too low to mitigate wage differences between regions. However, it is difficult to determine how an more efficient allocation of movers across the wage distribution should look like, too many or too skilled upward movers might even increase wage disparities.

[17]We ease our requirements substantially, exclude return migrants, and adjust wages for differences in housing costs. Specifically, we compare the differences in wage growth of movers and stayers from the age of 25 to ages 34 and 44. A detailed description is relegated to the appendix (Section 4.C.7). The relative moving probabilities are not subject to large changes. The gain pattern across origin and destination districts' quartiles is comparable. Similar to the previous section, the exclusion of return migrants leads to, on average, higher gains, but the overall pattern does not change (Table 4.15).

The underlying sources of this low upward mobility, however, can be manifold and will, most likely, be interdependent. In general, theory predicts that individuals compare the costs and returns to migration and relocate if it is profitable (Sjaastad, 1962). We document large mobility gains but observe that only a few workers decided to migrate at all, and the first relocation is, in most cases, the last one. This indicates that relocation opportunities are either, e.g., rare due to search frictions, infeasible due to migration costs, or wage differentials are captured by differences in local rent prices or amenities (Albouy et al., 2013, Kennan and Walker, 2011, and Schmutz and Sidibé, 2019). While we do not account for differences in amenities, our results suggest that rent-prices capture only a minor degree of the wage disparities.[18] The substantial long-term gains to mobility that we find depend on the destination's wage level even when we account for the variations in rent-levels. If wage differentials are captured completely in the local price of housing, the gains from mobility should show no dependency. Moreover, monetary moving costs might render the expected gains as to low to migrate. On the one hand, costs that depend on the distance to the destination could explain why movers mostly relocate to near districts, especially as we do not find any significant impact of distance on earnings gains. On the other hand, this pattern could also be due to region-specific preferences of a worker. Furthermore, higher gains from upward moves despite no average directionality of migration decisions could be explained by search frictions. Workers might not be able to obtain an offer from a high wage location at every point in time, or searching for jobs in far-away regions is more difficult. We investigate the importance of potential sources of the low mobility in a companion paper, which utilizes a structural model.

4.5 Conclusion

This paper studies the characteristics of individual inter-regional migration decisions of employed workers and the individual wage gains from mobility. We use an extensive German administrative panel data set, which allows us to follow individual workers over relatively fine-grained work locations and the life-cycle. We find that inter-regional mobility is low as the majority of workers never relocates to a different district. Conditional on moving, most workers change their work region only once or twice throughout working life and primarily to regions that are within 100 kilometers of their starting region. Mobility also decreases with age and distance.

Our results concerning the returns to mobility show that mobile workers do realize long-term wage gains relative to stayers of 12%. To estimate these permanent income

[18] In standard regional equilibrium models, workers are indifferent between staying and moving across regions where, typically, local rent prices play the role of making workers indifferent (Roback, 1982).

effects, we compare accumulated earnings of movers to those of individuals that never leave their initial region. We follow the literature on earnings losses of displaced workers to account for the movers' counterfactual earnings path and the heterogeneity between movers and stayers. We find that average regional wages are an important determinant of mobility gains. Our results show that earnings gains from moving increase with the destination districts wage level and the wage difference between origin and destination district. While these gains increase in the average wage level of a mover's destination, a significant fraction of moves are directed to regions that have a lower wage compared to the origin location.

The limited cross-regional mobility despite considerable moving gains hints at the presence of frictions. These frictions may manifest in the form of, e.g., costs of moving, preferences for the home location, or search frictions (Kennan and Walker, 2011, Schmutz and Sidibé, 2019, and Heise and Porzio, 2019). Given mobility frictions, there is a strong motive to insure employees against location risk. This raises the question of whether the current income tax system is an optimal response to spatial dispersion given endogenous mobility responses? Shall we allow the tax code to differ across regions (as suggested by Eeckhout and Guner (2015)) to fight the dispersion in living standards, or shall the government rely on the market? We answer these questions in a companion paper, which provides a simple theory of optimal relocation and redistribution across space.

Appendix

4.A Frequency and probability of moving

This section provides additional results of the inter-regional moving behavior.

Table 4.1 displays the frequencies of inter-regional transitions by education and gender for the 1980s cohorts at age 54. In general, the results are similar across groups: The majority of workers do not move, and if they do, they change regions one or two times. An exception is high skilled men, of which more than 50% relocate to another district. Men move more often compared to women. Including women into the sample increases the share of individuals that never leaves their origin district by almost 10%.

Table 4.1: Frequency of moves by education and gender (1980s cohorts, Age 54)

	Job-to-Job transitions			
Moves	Male + Female	Male	Male High skilled	Male Low skilled
0	73.8%	64.0%	48.4%	68.7%
1	13.2%	16.6%	21.5%	15.2%
2	6.7%	9.5%	13.5%	8.3%
3	3.1%	4.6%	7.7%	3.7%
4	1.5%	2.4%	3.8%	2.0%
5	1.7%	2.8%	5.1%	2.1%
	Job-to-Job and unemployment to employment transitions			
Moves	Male + Female	Male	Male High skilled	Male Low skilled
0	70.1%	60.1%	44.1%	64.8%
1	13.8%	16.6%	21.7%	15.1%
2	8.0%	10.9%	14.4%	9.9%
3	3.7%	5.4%	8.4%	4.6%
4	1.9%	3.0%	4.5%	2.5%
5	2.4%	4.0%	7.0%	3.1%

Notes: This tables displays the frequency of moves for the 1980s cohorts at age 54. Full-time to full-time job-to-job transitions as well as unemployment to full-time employment transitions coinciding with regional change. Values are rounded.

Table 4.2 shows the frequencies of inter-regional transitions by education and gender for the 1990s cohorts at age 44. Overall, the results are similar, compared to the 1980s cohorts. The majority of individuals stay in their origin district, and if they move, they move one or two times. High educated men, as well as all men, change jobs and regions more often compared to low skilled men or women. Most individuals move one or two times. The frequencies of men with no spell in East Germany (Male (West)) are similar, compared to those of all men.

Table 4.2: Frequency of moves by education and gender (1990s cohorts, Age 44)

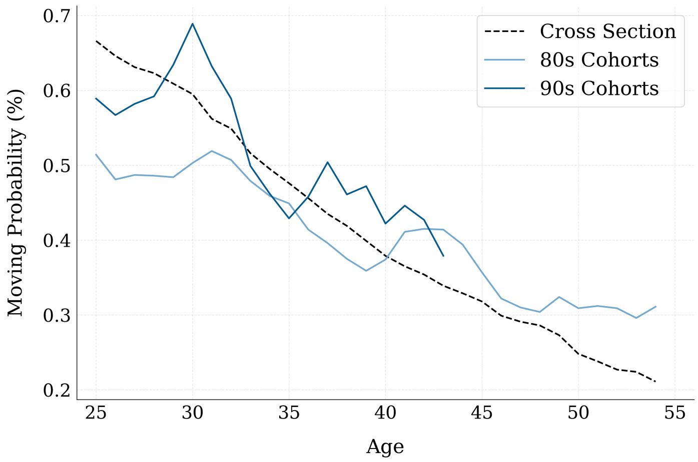
	Job-to-Job transitions				
Moves	Male + Female	Male	Male (West)	Male High Skilled	Male Low Skilled
0	73.8%	65.4%	66.1%	51.0%	71.1%
1	14.2%	17.5%	17.3%	23.4%	15.2%
2	7.0%	9.5%	9.4%	13.1%	8.1%
3	2.7%	4.0%	3.8%	6.7%	2.9%
4	1.2%	1.9%	1.9%	3.0%	1.5%
5	1.1%	1.8%	1.5%	2.8%	1.4%
	Job-to-Job and unemployment to employment transitions				
Moves	Male + Female	Male	Male (West)	Male High skilled	Male Low skilled
0	69.6%	60.3%	61.9%	45.7%	66.1%
1	14.9%	17.7%	17.4%	23.5%	15.4%
2	8.2%	10.9%	10.7%	14.4%	9.6%
3	3.7%	5.4%	5.1%	8.1%	4.4%
4	1.8%	2.8%	2.5%	4.2%	2.2%
5	1.8%	2.8%	2.4%	4.0%	2.4%

Notes: This tables displays the frequency of moves for the 1990s cohorts at age 44. Full-time to full-time job-to-job transitions as well as unemployment to full-time employment transitions coinciding with regional change. The values are rounded.

Figure 4.1 shows the average monthly moving probability by age for the 1980s and 1990s cohorts for both, employment and unemployed.

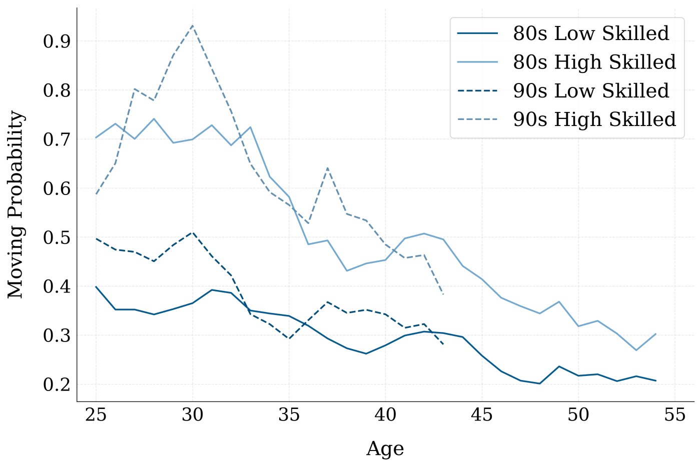
Figure 4.2 displays the average monthly moving probability by age for the 1980s and 1990s cohorts. High skilled men are almost two times more likely to relocate to a different districts.

Fig. 4.1: Moving probability by age (Job-to-job and unemployment to employment transitions)



Notes: Probability of inter-regional job-to-job and unemployment to employment transitions: 1980s and 1990s cohorts as well as West German cross-section (1980-2014).

Fig. 4.2: Moving probability by age and education



Notes: Probability of inter-regional job-to-job transitions for the 1980s and 1990s cohorts by education.

4.B Immediate wage gains: Additional evidence

We observe no clear direction of transitions concerning average wages. Individuals, who move to a district with a higher average wage, account for approximately 50% of all inter-regional job-to-job transitions. The share is almost constant over the life-cycle. However, immediate wage gains do differ. Table 4.3 shows the average wage gain of individuals who transition to a district with a lower ("Down") and higher ("Up") average wage. Wage differences between destination and origin district are computed based on wages of full-time employees at the time of the move. Immediate gains decrease by age. Inter-regional transitions to a higher paying district almost always lead to an additional wage gain compared to a transition to a district with lower average wages. On average, the gains of individuals moving up are approximately 3% higher.

Table 4.3: Immediate wage gains and direction

	1980s cohorts		1990s cohorts	
Age	Up	Down	Up	Down
25 to 34	0.12	0.08	0.10	0.06
35 to 44	0.07	0.04	0.06	0.03
45 to 54	0.02	0.02	–	–
Share	49.6%	50.4%	49.5%	50.5%

Notes: Average instant (log) wage gains for the 1980s and 1990s cohorts from inter-regional full-time to full-time job-to-job transitions.

Moreover, we compute the distance between the centers of the origin and destination district and sort the job-to-job transitions into four distance classes. Table 4.4 presents the results. Overall, the distribution of moves over distances is quite stable over the life-cycle. The distributions are similar across cohorts and cross-sections. The table also shows the instant average gains for distance and age groups. The gains from inter-regional transitions are positive and decrease over the life-cycle for all distance classes. Young workers have the largest instant gain of around 10% for the 1980s and 9% for the 1990s cohorts. Middle-aged men gain $4 - 7\%$ if they migrate to another district, and individuals aged 45 and 54 obtain only a small immediate wage gain. Overall, the instant wage gains of full-time employed males do not differ much over distance classes.

The average distance of moves differs when we compare inter-regional transitions between East and West German workers. This difference is partly due to differences in average distances to other districts: An individual starting in West Germany faces

Table 4.4: Immediate wage gains: Distance to destination

Age	< 65 km	65-129 km	130-259 km	≥ 260 km
	80s Cohorts			
25 to 34	0.10	0.09	0.11	0.11
35 to 44	0.05	0.06	0.07	0.04
45 to 54	0.02	0.02	0.02	−0.01
Share	31.0%	30%	17%	22%
	90s Cohorts			
25 to 34	0.09	0.08	0.09	0.10
35 to 44	0.05	0.04	0.05	0.05
45 to 54	-	-	-	-
Share	27%	31%	17%	26%

Notes: Average instant (log) wage gains for the 1980s and 1990s cohorts. Inter-regional full-time to full-time job-to-job transitions. The share is calculated on all regional job-to-job transitions aver all age classes. Values are rounded.

a 397 km average distance to another location. In contrast, starting in East Germany, the average distance is 425 km. Table 4.5 displays the shares by distance class for the cross-section within 1994 and 2014. On average, West and East German workers transition to a district located approximately 82 and 162 km away from their origin district, respectively.

Table 4.5: Proportion of movers by distance to the destination (East and West Germany)

< 65 km	65-129 km	130-259 km	≥ 260 km	Med. Distance
		1990s West		
30%	30%	16%	23%	82 km
		1990s East		
15%	35%	18%	32%	162 km

Notes: Share of inter-regional full-time to full-time job-to-job transitions by distance classes for the *cross-section* in 1994 to 2014. Values are rounded.

4.C Long-term earnings gains: Robustness

4.C.1 Long-term earnings gains: Regression results

We estimate the following model using linear least squares

$$y = I_d + \ln(distance) + educ + age + \epsilon, \tag{4.6}$$

where I_d is an indicator function representing the origin or destination district's rank in the wage distribution or the respective rank in the wage differential distribution. $\ln(distance)$ is the log distance between origin and destination district. educ indicates whether the individual is high or low skilled. Age is age in years and ϵ represents the white noise error term. The results are presented in Table 4.6. The estimates of the indicator categories are measure in terms of the lowest quartile (0).

Table 4.6: Earnings gains from inter-regional transitions (Regression results)

Quartile	Origin District	Destination District	Up or Down
	Dependent Variable: Δg		
1	-	-	-
2	$-$ 2%	5%**	7%***
3	$-$ 4%*	7%***	9%***
4	$-$ 6%**	9%***	12%***
Age	$-$ 2%***	$-$ 2%***	$-$ 2%***
Distance	$-$0%	0%	$-$0%
Education	15%***	14%***	15%***
Number of Movers		2627	

Notes: "'***"' indicates significance at the 1%, "**" at the 5% level, and "**" at the 10% level.

4.C.2 Robustness to timing changes

We compute the same gain measure as in the main analysis but follow individuals for six months prior and five years after the moving event. Table 4.7 displays the results. The results are similar to those of the main analysis. One difference is that the earnings differences before the move are more negative compared to one year before. Hence, wage differences between movers and stayers increase the closer we get to the moving event.

Table 4.7: Earnings gains from inter-regional transitions (Timing adjustment)

Percentile	Origin District	Destination District	Up or Down
	Δg, permanent wage gain		
1	0.18	0.04	0.06
2	0.13	0.10	0.12
3	0.12	0.12	0.14
4	0.12	0.16	0.19
	g^p, prior to the move (6 months)		
1	0.04	-0.03	-0.04
2	-0.00	-0.04	-0.01
3	-0.01	-0.03	-0.01
4	-0.01	0.04	0.06
	g^f, following the move (5 years)		
1	0.22	0.02	0.03
2	0.13	0.07	0.11
3	0.11	0.09	0.13
4	0.12	0.20	0.25

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within 6 months prior to the move. g^f measures the difference in accumulated earnings within the 5 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

4.C.3 Rent adjusted earnings gains

We use rent data provided by the Federal Institute for Research on Building, Urban Affairs and Spatial Development (BBSR) and IDN ImmoDaten GmbH and adjust wages using a housing costs share of 24%.[19] The rent data is available on the county level (NUTS3) and is aggregated to the EAD level using population weights.[20]

Individual wages are then adjusted:

$$\widehat{w}_{jt} = w_{jt}\Big(0.76 + 0.24\frac{1}{m_{jt}}\Big),$$

where m_{jt} is the relative rent $\frac{rent_{jt}}{\frac{1}{J}\sum_k^J rent_{kt}}$ in district j.

Our data on average rent prices for newly rented units per square-feet covers the period 2004 to 2014. We approximate m for the 1994 to 2003 by predicting the average rent based on other observable data. The predicted values are based on the coefficients β estimated for the period 2004 to 2014 and based on the following model:

$$ln(rent_{jt}) = \beta X_{jt} + \epsilon_{jt}.$$

The vector of observables X_{jt} includes the logarithm of primary income per capita, the share of working age population, the share of 18 to 25 as well as 50 to 65 year-olds, the unemployment rate and population density (population divided by total surface). The regression explains about 75% of the variation in average rent prices across districts ($R^2 = 0.75$). Average rent values are then predicted for the prior years and used to adjust wages for average housing costs.[21]

The unemployment rates are calculated using SIAB data. Data on population shares are taken from the BBSR.[22] We obtain data on total surface from the federal statistical offices.[23] Primary income [24] and population[25] data stem from the Working Group on National Accounts of the Länder.

Table 4.8 and 4.9 display the results.

[19]The average housing cost share is taken from: "Regionaler Preisindex", Berichte, Band 30, Hrsg.: BBSR, Bonn 2009.

[20]Results are similar if the rent data is aggregated by taking means across counties.

[21]There are missing primary incomes values for 8 districts in each year during 1995-1999: All respective rent values are set to 2000 values. Furthermore, rent values for all districts in 1993-1994 are set 1995 average rents.

[22]Indikatoren und Karten zur Raum- und Stadtentwicklung. INKAR. Ausgabe 2017. Hrsg.: Bundesinstitut für Bau-, Stadt- und Raumforschung (BBSR).

[23]Bodenfläche nach Art der tatsächlichen Nutzung - Stichtag 31.12. - Kreise und kreisfr. Städte, (449-01-4). Statistische Ämter des Bundes und der Länder.

[24]Einkommen der privaten Haushalte in den kreisfreien Städten und Landkreisen der Bundesrepublik Deutschland 1995 bis 2015, Reihe 2, Band 3, Arbeitskreis Volkswirtschaftliche Gesamtrechnungen der Länder.

[25]Bruttoinlandsprodukt, Bruttowertschopfung in den kreisfreien Städten und Landkreisen der Bundesrepublik Deutschland 1992 und 1994 bis 2015, Reihe 2, Band 1, Arbeitskreis Volkswirtschaftliche Gesamtrechnungen der Länder.

Table 4.8: Earnings gains from inter-regional transitions (Rent adjusted)

Percentile	Origin District	Destination District	Up or Down
	Δg, permanent wage gain		
1	0.12	0.06	0.08
2	0.14	0.10	0.12
3	0.12	0.13	0.12
4	0.11	0.14	0.14
	g^p, prior to the move (1 year)		
1	0.04	-0.01	-0.02
2	-0.00	-0.02	-0.00
3	-0.02	-0.03	0.01
4	0.01	0.04	0.06
	g^f, following the move (10 years)		
1	0.16	0.05	0.05
2	0.14	0.08	0.12
3	0.10	0.10	0.13
4	0.12	0.18	0.18

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain $-$ compared to not moving $-$ is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

Table 4.9: Earnings gains by region (Rent adjusted)

	Earnings differences		
Direction	after (g_f)	before (g_p)	permanent wage gain (Δg)
W-to-W	0.12	−0.00	0.13
W-to-E	−0.03	−0.01	−0.02
E-to-W	0.20	0.09	0.11
E-to-E	0.11	−0.01	0.11

Notes: The table presents the accumulated monthly earnings differences for movers and stayers between and within West and East Germany for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34.

4.C.4 Earnings gains including moves out of unemployment

We compare inter-regional movers to stayers. Both types can be unemployed or employed at the time of the transition. We restrict the comparison to individuals that are unemployed for at most 12 months and have been employed during the spell before unemployment. The income comparison before the transition is based on the last wage observed when employed. The comparison group of stayers consists of all employed and everyone in unemployment for less than 12 months. After the move, we again compare incomes earned during the next ten years out of full- and part-time employment as well as unemployment, requiring at least ten years of consecutive observations. The inter-regional transition has to be to a full-time job.

Table 4.10: Earnings gains from inter-regional transitions (including unemployed)

Percentile	Origin District	Destination District	Up or Down
	Δg, permanent wage gain		
1	0.16	0.03	0.05
2	0.13	0.09	0.12
3	0.13	0.12	0.12
4	0.12	0.15	0.17
	g^p - prior to the move (1 year)		
1	0.05	-0.03	-0.03
2	0.01	0.00	-0.01
3	-0.00	-0.03	0.01
4	0.01	0.05	0.09
	g^f - following the move (10 years)		
1	0.13	-0.01	-0.01
2	0.11	0.08	0.07
3	0.10	0.10	0.10
4	0.10	0.18	0.23

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

4.C.5 Intra-regional job-to-job transitions as comparison group

In this section, we analyze the earnings gains of inter- relative to intra-regional job-to-job movers, which allows us to differentiate between the return from changing jobs and the return from moving to another region. Establishment movers, who stay within their origin region, are the apparent comparison group.

We compute the same earnings measures as in the main analysis, now relative to individuals that change jobs within a region but never leave their origin district. Movers and non-movers have to be full-time employed for at least 12 months before the move. The transition has to be from full-time to full-time employment. We match each mover to all job-changing stayers of the same age a, cohort c, and employed in the same district j. In contrast to the comparison to all stayers, we do not require the job-to-job transition of the stayer to happen in the same month as the transition of the mover. The stayer has to be of the same cohort and age, i.e., the transitions need to take place in the same year. The results are presented in Table 4.11. We find substantial relative permanent gains for inter-regional movers (Δg). They are slightly lower but of the same magnitude compared to the gains relative to all stayers. The key difference is that the mover's monthly earnings before the move are almost always higher compared to workers changing jobs within their current region and not lower as in the comparison to all stayers. However, the relative permanent gains are, on average, higher during the ten years following the move leading to comparable overall gains Δg.

Table 4.11: Earnings gains relative to intra-regional job-to-job transitions

Percentile	Origin District	Destination District	Up or Down
	Δg, permanent wage gain		
1	0.13	0.01	0.01
2	0.15	0.04	0.05
3	0.06	0.09	0.11
4	0.06	0.11	0.15
	g^p - prior to the move (1 year)		
1	0.15	0.00	0.01
2	0.06	0.04	0.10
3	0.05	0.01	0.01
4	0.08	0.11	0.12
	g^f - following the move (10 years)		
1	0.28	0.01	0.01
2	0.21	0.08	0.14
3	0.11	0.09	0.12
4	0.14	0.22	0.27

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

4.C.6 Education and long-term gains

Table 4.12: Earnings gains from inter-regional transitions by education (Low skilled)

Percentile	Origin District	Destination District	Up or Down
	Δg, permanent wage gain		
1	0.14	0.04	0.00
2	0.09	0.06	0.06
3	0.08	0.09	0.11
4	0.04	0.10	0.15
	g^p - prior to the move (1 year)		
1	0.01	-0.04	-0.06
2	-0.06	-0.09	-0.05
3	-0.07	-0.05	-0.05
4	-0.04	-0.03	-0.01
	g^f - following the move (10 years)		
1	0.15	0.01	-0.05
2	0.03	-0.03	0.01
3	0.02	0.04	0.06
4	0.01	0.07	0.13

Notes: The table presents the accumulated monthly earnings differences for low educated movers and stayers for the 1990s cohorts. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain – compared to not moving – is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

4.C.7 Less restrictive selection

We modify the gain measure presented in Section 4.4 the following way: We follow cohorts of male workers from age 25 over the next 10 and 20 years. Wages of movers and stayers are compared for ages 25 and 34, as well as 25 and 44. Hence, we compute the difference in log wages for movers and stayers at ages 34 or 44 and subtract the initial wage difference (at age 25).

The group of movers is defined as all individuals in full-time employment at 25, 34, and 44, with at least one inter-regional job-to-job transition within the respective age interval. We pair each mover with a group of at least three stayers that were employed in the same district as the mover at the beginning of the age interval and did not relocate within the following 10 or 20 years. Hence, a stayer at age 34 is allowed to move within the next ten years. Stayers have to be full-time employed at the point of comparison. We exclude return migrants.

Origin and destination districts are the places of employment of the movers at age 25 as well as 34 and 44. As we do not observe the time of the move, we sort destination districts into wage quartile at the end of each age interval. Table 4.13 displays the resulting differences in log wages between movers and stayers. The pattern is consistent with the results presented in Table 4.5: Individuals that transition to a higher-ranked district gain more. The relative moving probabilities are also similar.

The modified measure allows us to present a long-term gain measure for the 1990s cohorts at age 44 (right panel). The change in log wage differences is lower, compared to age 34. This result is driven by relatively higher wage growth for the group of stayers (not printed) and by compositional changes as a fraction of stayers become movers. However, the relative moving probabilities are not subject to significant changes.

Table 4.13: Transitions and gain matrix (Less restrictive selection)

Age 25 to 34

Wage differences

		Destination			
		1	2	3	4
Origin	1	0.04	0.11	0.22	0.20
	2	-0.16	0.06	0.10	0.19
	3	-0.00	0.04	0.14	0.17
	4	-0.17	0.03	0.00	0.18

$Pr(j|i)$

		Destination			
		1	2	3	4
Origin	1	0.50	0.14	0.12	0.24
	2	0.11	0.22	0.27	0.40
	3	0.06	0.22	0.27	0.45
	4	0.07	0.14	0.24	0.55

Age 25 to 44

Wage differences

		Destination			
		1	2	3	4
Origin	1	0.03	0.05	0.19	0.14
	2	-0.14	0.06	0.10	0.16
	3	-0.20	0.00	0.10	0.13
	4	-0.20	0.02	0.04	0.13

$Pr(j|i)$

		Destination			
		1	2	3	4
Origin	1	0.44	0.24	0.12	0.20
	2	0.16	0.24	0.26	0.34
	3	0.05	0.20	0.34	0.41
	4	0.05	0.16	0.25	0.53

Notes: This table displays the relative difference in *wage growth* between movers and stayers and conditional moving probabilities $Pr(j|i)$ for the 1990s cohorts. Wages are adjusted for regional differences in rent prices. Return migrants are excluded.

4.C.8 Return migration and long-term gains

In this section, we present the transition probabilities conditional on the origin district's rank in the average wage distribution for the 1990s cohorts without return migrants. The gains are higher, on average.

Table 4.14: Earnings gains from inter-regional transitions (without return migrants)

	Earnings differences: Up or down		
Quartile	after (g_f)	before (g_p)	permanent wage gain (Δg)
1	0.04	-0.03	0.07
2	0.14	0.01	0.13
3	0.15	-0.01	0.16
4	0.28	0.07	0.20

Notes: The table presents the accumulated monthly earnings differences for movers and stayers without return migrants for the 1990s cohorts. g^p measures the difference in accumulated earnings within 1 year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The long-term earnings gain − compared to not moving − is measured by Δg. The inter-regional job-to-job transitions took place at ages 25 to 34. The quartiles 3 and 4 indicate an upward move in the average wage distribution.

Table 4.15: Transition and gain matrix (without return migrants)

	Δg			
	Destination			
	1	2	3	4
Origin	0.15	0.18	0.22	0.23
	-0.11	0.15	0.22	0.16
	-0.01	0.10	0.15	0.15
	-0.09	0.13	0.10	0.16

	$Pr(j\|i)$			
	Destination			
	1	2	3	4
Origin	0.47	0.16	0.13	0.24
	0.10	0.20	0.27	0.42
	0.04	0.21	0.26	0.49
	0.04	0.11	0.21	0.64

Notes: This table displays the long-term earnings gains Δg and conditional moving probabilities $Pr(j|i)$ of first time movers aged 25 to 34 for the 1990s cohorts. Return migrants − movers that, again, reside in their origin district at age 34 − are excluded. Values are rounded.

4.C.9 Earnings gains of the 1980s cohorts

This section presents the long-term earnings gains for the 1980s cohorts. The restrictions are the same as for the 1990s cohorts. Additionally, we exclude individuals with a spell in East Germany or Berlin. We sort districts into quartiles of the average wage distribution, to analyze the moving and gain pattern. Districts are assigned to a quartile based on the average wages of male full-time employees. Table 4.16 shows the results for first-time movers aged 25 to 34.

The long-term earnings gains of movers conditional on the mover's previous district's rank are presented in Column 2. Although movers from the lowest quartile exhibit the highest permanent gains, we observe no decreasing pattern as the origin districts' rank increases (upper panel). Hence, the information value of the district of origin about the size of the prospective earnings gains is low. Movers earn less than stayers in the period before the move (center panel). The differences in the period after the moving event show the same structure as the permanent gains (lower panel).

The earnings gains differ conditional on the destination district's rank in the wage distribution (Column 3). These gains differ notably and show a clear pattern: A move to a higher ranked district leads to larger permanent earnings gains (upper panel). The difference in earnings also mirrors the increasing structure after the move (lower panel). In the period before the moving event, movers exhibit, on average, lower earnings, compared to stayers (center panel).

In Column 4, we group movers into four quartiles according to the average wage differences between destination and origin districts. The permanent gains are positive, independent of whether the move is up- or downward. However, a move upward is associated with higher average gains. In the period before the transition, the mover's earnings are lower or similar, compared to stayers. The relative earnings gain after the move increases from 2 to 10% as the individual moves up in the wage distribution.

Table 4.17) presents the results for first-time movers aged 35 to 44. Although the permanent gains are mostly positive, we observe no clear pattern conditional on the direction of the move.

Reallocation across the wage distribution Table 4.18 displays the results for the 1980s cohorts. Here, we analyze transition probabilities and gains for $25-34$ (upper panel) and $35-44$ (lower panel) year-old movers. The return to mobility is almost always positive and substantial. The permanent gain for the 25 to 34-year-old workers shows an increasing pattern in the destination districts' rank: A move to a lower-ranked district always leads to lower gains. In general, long-term earnings gains of first-time movers aged $35-44$ are positive but show no clear pattern across quartiles. The transitions probabilities are similar for both age groups (right panel). Movers out of the lowest quartile have an almost uniform probability across destination districts. For

Table 4.16: Earnings gains from inter-regional transitions (1980s cohorts, 25-34)

Quantile	Origin District	Destination District	Up or Down
		Δg, permanent wage gain	
1	0.11	0.04	0.03
2	0.05	0.05	0.06
3	0.06	0.07	0.09
4	0.08	0.11	0.11
		g^p, prior to the move (1 year)	
1	-0.01	-0.03	-0.01
2	-0.01	-0.02	0.00
3	-0.02	-0.01	0.00
4	0.01	0.00	-0.01
		g^f, following the move (10 years)	
1	0.11	0.00	0.02
2	0.03	0.03	0.06
3	0.04	0.06	0.09
4	0.09	0.11	0.10

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1980s cohorts. The inter-regional job-to-job transitions took place at ages 25 to 34. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

the other three quartiles, the probabilities are similar to those of the 1990s cohorts: More than half of the inter-regional transitions are either within the same or to a higher wage quartile.

Table 4.17: Earnings gains from inter-regional transitions (80s cohorts, 35-44)

Quantile	Origin District	Destination District	Up or Down
		Δg, permanent wage gain	
1	0.03	0.03	0.03
2	0.00	-0.01	0.02
3	0.03	0.05	0.01
4	0.04	0.05	0.06
		g^p, prior to the move (1 year)	
1	-0.03	-0.03	-0.00
2	-0.02	-0.04	0.01
3	0.05	0.02	0.01
4	0.02	0.04	0.05
		g^f, following the move (10 years)	
1	0.00	0.01	0.02
2	-0.01	-0.04	0.03
3	0.08	0.08	0.02
4	0.06	0.09	0.10

Notes: The table presents the accumulated monthly earnings differences for movers and stayers for the 1980s cohorts. The inter-regional job-to-job transitions took place at ages 35 to 44. g^p measures the difference in accumulated earnings within one year prior to the move. g^f measures the difference in accumulated earnings within the 10 years following the move. The permanent wage gain − compared to not moving − is measured by Δg. In Column 4, the quartiles 3 and 4 indicate an upward move in the average wage distribution.

Table 4.18: Transition and gain matrix (1980s cohorts)

Age 25 to 34		Δg Destination 1	2	3	4			$Pr(j\|i)$ Destination 1	2	3	4
Origin	1	0.10	0.12	0.10	0.18	Origin	1	0.26	0.25	0.19	0.30
	2	-0.05	0.02	0.08	0.12		2	0.15	0.22	0.29	0.34
	3	0.03	0.06	0.07	0.11		3	0.11	0.21	0.21	0.48
	4	0.05	0.06	0.07	0.10		4	0.09	0.13	0.26	0.51
Age 35 to 44		**Δg Destination 1**	**2**	**3**	**4**			**$Pr(j\|i)$ Destination 1**	**2**	**3**	**4**
Origin	1	0.06	-0.21	0.08	0.08	Origin	1	0.26	0.25	0.25	0.25
	2	0.06	-0.16	0.05	0.12		2	0.14	0.26	0.32	0.28
	3	-0.04	0.01	0.02	0.03		3	0.08	0.21	0.27	0.44
	4	0.03	0.04	0.03	0.06		4	0.08	0.14	0.26	0.53

Notes: This table displays the relative permanent earnings gains Δg and conditional moving probabilities $Pr(j|i)$ of first time movers aged 25 to 34 or 35 to 44 for the 1980s cohorts.

4.C.10 Transition probabilities in West Germany

This section presents the conditional transition probabilities for West Germany. For the 1990s cohort, we exclude individuals with spells in East Germany. The probabilities for the 1980s cohorts are the same as those presented in the main text.

Table 4.19: Conditional moving probabilities (1980s and 1990s cohorts, West Germany)

		1980s cohorts						1990s cohorts			
		Destination						Destination			
		1	2	3	4			1	2	3	4
Origin	1	0.26	0.25	0.19	0.30	Origin	1	0.23	0.21	0.24	0.32
	2	0.15	0.22	0.29	0.34		2	0.16	0.20	0.30	0.33
	3	0.11	0.21	0.21	0.48		3	0.14	0.14	0.22	0.49
	4	0.09	0.13	0.26	0.51		4	0.07	0.12	0.22	0.59

Notes: This table displays the conditional moving probabilities $Pr(j|i)$ of first time movers aged 25 to 34 for the 1990s and 1980s cohorts. We consider only individuals with spells in West Germany. The values are rounded.

The results show that the conditional moving probabilities are similar for the 1980s and 1990s cohorts in West Germany: Approximately 75% leave the lowest quartile. Transitions into the lowest quartile are substantially higher, compared to the results for Germany.

4.D Classification

Table 4.20: District Classification

Code	District	Code	District	Code	District
030	AA Greifswald	373	AA Paderborn	729	AA Fürth
031	AA Neubrandenburg	375	AA Recklinghausen	735	AA Nürnberg
032	AA Rostock	377	AA Rheine	739	AA Regensburg
033	AA Schwerin	381	AA Siegen	743	AA Schwandorf
034	AA Stralsund	383	AA Meschede – Soest	747	AA Schweinfurt
111	AA Bad Oldesloe	387	AA Wesel	751	AA Weiden
115	AA Elmshorn	391	AA Solingen – Wuppertal	759	AA Würzburg
119	AA Flensburg	411	AA Bad Hersfeld – Fulda	811	AA Augsburg
123	AA Hamburg	415	AA Darmstadt	815	AA Deggendorf
127	AA Heide	419	AA Frankfurt	819	AA Donauwörth
131	AA Kiel	427	AA Gießen	823	AA Freising
135	AA Lübeck	431	AA Hanau	827	AA Ingolstadt
139	AA Neumünster	433	AA Bad Homburg	831	AA Kempten – Memmingen
211	AA Braunschweig – Goslar	435	AA Kassel	835	AA Landshut – Pfarrkirchen
214	AA Bremen – Bremerhaven	439	AA Korbach	843	AA München
221	AA Celle	443	AA Limburg – Wetzlar	847	AA Passau
224	AA Emden – Leer	447	AA Marburg	855	AA Rosenheim
231	AA Göttingen	451	AA Offenbach	859	AA Traunstein
234	AA Hameln	459	AA Wiesbaden	863	AA Weilheim
237	AA Hannover	511	AA Bad Kreuznach	035	AA Cottbus
241	AA Helmstedt	515	AA Kaiserslautern – Pirmasens	036	AA Eberswalde
244	AA Hildesheim	519	AA Koblenz – Mayen	037	AA Frankfurt (Oder)
251	AA Lüneburg – Uelzen	523	AA Ludwigshafen	038	AA Neuruppin
257	AA Nordhorn	527	AA Mainz	039	AA Potsdam
261	AA Oldenburg – Wilhelmshaven	535	AA Montabaur	922	AA Berlin Süd
264	AA Osnabrück	543	AA Landau	955	AA Berlin Nord
267	AA Stade	547	AA Neuwied	962	AA Berlin Mitte
274	AA Vechta	555	AA Saarland	041	AA Bernburg
277	AA Nienburg – Verden	563	AA Trier	042	AA Dessau-Roßlau – Wittenberg
311	AA Aachen – Düren	611	AA Aalen	043	AA Halberstadt
315	AA Bergisch Gladbach	614	AA Balingen	044	AA Halle
317	AA Bielefeld	617	AA Freiburg	045	AA Magdeburg
321	AA Bochum	621	AA Göppingen	046	AA Weißenfels
323	AA Bonn	624	AA Heidelberg	047	AA Sangerhausen
325	AA Brühl	627	AA Heilbronn	048	AA Stendal
327	AA Coesfeld	631	AA Karlsruhe – Rastatt	093	AA Erfurt
331	AA Detmold	634	AA Konstanz – Ravensburg	094	AA Altenburg – Gera
333	AA Dortmund	637	AA Lörrach	095	AA Gotha
337	AA Düsseldorf	641	AA Ludwigsburg	096	AA Jena
341	AA Duisburg	644	AA Mannheim	097	AA Nordhausen
343	AA Essen	647	AA Nagold – Pforzheim	098	AA Suhl
345	AA Gelsenkirchen	651	AA Offenburg	071	AA Annaberg-Buchholz
347	AA Hagen	664	AA Reutlingen	072	AA Bautzen
351	AA Hamm	671	AA Waiblingen	073	AA Chemnitz
353	AA Herford	674	AA Schwäb. Hall – Tauberb.	074	AA Dresden
355	AA Iserlohn	677	AA Stuttgart	075	AA Leipzig
357	AA Köln	684	AA Ulm	076	AA Oschatz
361	AA Krefeld	687	AA Rottweil – Villingen-Schw.	077	AA Pirna
364	AA Mettmann	711	AA Ansbach – Weißenburg	078	AA Plauen
365	AA Mönchengladbach	715	AA Aschaffenburg	079	AA Riesa
367	AA Ahlen – Münster	723	AA Bayreuth – Hof	080	AA Freiberg
371	AA Oberhausen	727	AA Bamberg – Coburg	092	AA Zwickau

Employment agency districts according to territorial status in 2014.

Chapter 5
Concluding Remarks

This thesis presents three essays on local labor markets, local taxation, and migration. Chapter two investigates the relative impact of two distinct local business tax on wages. This chapter demonstrates that variation in local tax rates leads to spatial variations in wages. The results suggest that governments should take into account that different tax instruments led to a different pass-through of the tax burden onto workers. Chapter three documents the development of the local labor market disparities in Germany and stresses that policymakers should take into account the hiring and separation margin when designing regional policies. Foremost, region-specific labor market policies can lead to large welfare gains. Chapter four investigates the within-country migration behavior of employed and documents substantial earnings gains from relocation that depend on the average wage level of the destination district. The lack of aggregate direction in the mobility of workers helps to explain the persistent wage disparities in Germany.

To summarize, this thesis contributes to the understanding of regional disparities in Germany.

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