Job Loss and Political Entry*

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Abstract

The supply of politicians affects the quality of democratic institutions. Yet, the role of individual economic shocks for political selection is largely understudied. This paper investigates how experiencing a job loss affects individuals' decision to enter politics. Using administrative data on the universe of formally employed individuals, local political candidates, and party affiliates in Brazil, and relying on mass layoffs for causal identification, we show that job loss increases the probability that individuals join a political party and run for office in local elections. Heterogeneity analyses reveal a more pronounced increase in candidacies among laid-off individuals with higher financial incentives from office holding and larger predicted income losses. In addition, we find that being eligible for unemployment benefits after job loss also increases party memberships and candidacies. These results are consistent with the reduction in private-sector opportunity costs and the increased time resources explaining the rise in political entry. Moreover, we document that layoff-induced candidates are positively selected in various competence measures, suggesting that economic shocks may improve the quality of politicians.

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1 Introduction

What are the political consequences of experiencing economic shocks? Research on this question has grown rapidly, not least due to the rise of populism, political polarization, and right-wing extremism observed in many countries over the last few years. In this literature, studies typically focus on the "demand-side" of politics, asking how negative income shocks affect the electoral behavior and political attitudes of voters (see Margalit (2019) and Guriev and Papaioannou (2022) for recent reviews of this literature). At the same time, we know relatively little about how economic shocks impact the "supply-side" of politics. Politicians are perhaps the most important agents affecting the quality of democratic institutions. They aggregate citizens' preferences in legislatures and influence which policies are enacted and how well they are implemented. Empirical work shows that political leaders strongly matter for the quality of policymaking (Pande, 2003; Jones and Olken, 2005; Besley *et al.*, 2011; Clots-Figueras, 2012). Therefore, it is important to understand how individual circumstances, such as economic shocks, influence the selection of those seeking a political office.

In this paper, we study how experiencing a job loss affects individuals' decision to enter political life. Using rich administrative data from Brazil, we focus on party membership and candidacy for the local council, the lowest-level entry positions in Brazilian politics. We rely on individual-level information on the universe of party members and local candidates between 2000 and 2020, which we match to employer-employee data on all formally employed workers between 2004 and 2018. The comprehensiveness of the data allows us to investigate effect heterogeneity along individual, party, and institutional characteristics and explore the consequences of eligibility for unemployment benefits after the job loss. In addition, we document the implications of job loss for the selection of local politicians. Taken together, our results document novel insights on the mechanisms behind individuals' decision to engage with politics following a layoff, as well as the consequences of these economic shocks for the political system.

To estimate the effects of job loss on entry into politics, we exploit firm-level mass layoffs as plausibly exogenous variation for individual-level job loss. Mass layoffs should depend neither on the political behavior of a specific worker nor on other worker-level shocks that simultaneously affect dismissals and political behavior. We compare individuals laid off in a mass layoff firm to a matched control group of similar workers from other firms who were not dismissed in the same year. The richness of our data allows us to match treated and control workers exactly on several characteristics, such as gender, education, age, wage, tenure, and firm size, and to control for fine-grained local- and industry-specific shocks that may drive mass layoffs or political behavior.

Importantly, accounting for changes in the local political environment ensures that our results capture individual supply responses and are not influenced by changes at the demand side. We use a difference-in-differences (DiD) design to estimate dynamic treatment effects over up to three election cycles after job loss. The underlying parallel-trends assumption is supported by the absence of different trends in political outcomes between treated and control individuals before the layoff and we provide evidence that the results are not driven by selection into layoffs - even within mass layoffs.[1]

Our main results show that experiencing a job loss increases the probability that individuals newly register as a party member by more than 12% compared to the control mean. This increase in membership is long-lasting, starting in the election cycle after job loss and remaining statistically significant in the two subsequent election cycles. For the probability of running as a candidate in a local election, the layoff effect is increasing over time, ranging from 6% in the first election cycle after layoff up to 28% in the two subsequent election cycles. The dynamic effect observed for candidacies is consistent with the fact that local councilors have no term limit and often re-run for office after their first candidacy. Additionally, the process of deciding to become a candidate and preparing an election campaign requires a substantial amount of time and planning, potentially explaining the delayed effects. Importantly, our results are robust to alternative mass-layoff definitions, different sets of fixed effects, and varying sample restrictions. Moreover, we show that the effects of job loss extend beyond the realms of party memberships and councilor candidacies, encompassing various facets of political participation. Specifically, we also find positive effects on the likelihood of running for mayor, donating to a political candidate, and working for a political campaign.

What explains the positive effects of job loss on political entry? The size and granularity of our data enable us to examine heterogeneity in treatment effects and shed light on the channels driving the political responses to job loss. Laid-off individuals may be successfully mobilized through a change in party support. A job loss may spark individuals into action as they seek to overturn the incumbent who is blamed for the change in their own economic circumstances (e.g., Tilley *et al.*, 2018; Ahlquist *et al.*, 2020). Moreover, individuals can be activated by parties promising to better represent their new economic policy preferences. Evidence suggests that more extreme parties, both on the left and the right, are often more successful in mobilizing voters who have experienced economic hardship (e.g., Algan *et al.*, 2017; Autor *et al.*, 2020). Another mobilizing force may arise from labor unions which often take part in mass layoff ne-

[1] Our empirical strategy closely follows recent papers studying the effects of mass layoffs in Brazil on crime (Britto *et al.*, 2022a), domestic violence (Bhalotra *et al.*, 2021), health (Amorim *et al.*, 2023), and children's education (Britto *et al.*, 2022b).

gotiations and have been shown to exert influence over the political behavior of employees (e.g., Kim and Margalit, 2017; Matzat and Schmeißer, 2023). To disentangle which mechanisms are at play, we implement several heterogeneity analyses. We differentiate memberships and candidacies for the party of the local incumbent mayor versus other non-incumbent parties. The results indicate positive participation effects for memberships and candidacies in both incumbent and non-incumbent parties. We then examine ideological differences by classifying parties as left, center, and right. We find that job loss increases memberships and candidacies for parties across the ideological spectrum, with larger increases in membership with right-wing parties in the 2020 cycle, the term of office of far-right president Jair Bolsonaro. Finally, we look at the role of union affiliation by identifying parties historically aligned with labor unions. We document that following layoffs, there is an increase in individuals' likelihood of becoming members and running for office both at union-affiliated and non-union-affiliated parties. Overall, the heterogeneity analyses indicate that incumbent, center, and non-union-affiliated parties also experience increased memberships and candidacies following dismissals, suggesting that mobilization and politicization may not be the sole mechanisms driving the rise in participation.

Economic theories of political selection model the decision to run for office as a trade-off between the benefits from holding office (multiplied by the chance of winning the election) and the costs of running (Black, 1972; Caselli and Morelli, 2004). A layoff decreases the opportunity costs of running by lowering the private-sector outside options and increasing the time available for political engagement. In order to examine whether this mechanism is at play, we start by analyzing effect heterogeneity along individual characteristics, which likely reflect differences in opportunity costs. The positive effects of job loss on party memberships and candidacies are remarkably pervasive across different socio-demographic groups. At the same time, we document substantially larger effects among individuals with university education. These results are consistent with the interpretation that highly educated individuals experience larger reductions in opportunity costs following layoffs, given the expected higher salaries in the private sector. We also investigate the role of financial incentives by examining effect heterogeneity depending on councilor salaries—which vary discontinuously across Brazilian municipalities (Ferraz and Finan, 2011)—and individuals' predicted income loss. The former captures how attractive a political position is, while the latter captures the predicted severity of the economic shock based on worker and location observable characteristics. We find larger layoff effects on candidacies in municipalities with higher wage caps, with no corresponding effects for memberships. Consistently, the increase in candidacies is highest among individuals who are predicted to face the greatest income loss due to a layoff. Together, these results show the importance of financial incentives for individuals' decision to enter politics following a job loss.

In order to explore the role of time resources, we analyze the effects of eligibility for benefits from unemployment insurance. During benefit receipt, individuals have lower incentives to find new employment and thus more time available for engaging with politics. We conduct a regression discontinuity analysis exploiting the fact that benefit eligibility varies discontinuously with the time since the last layoff (a minimum of 16 months is required for claiming a new benefit). We find that benefit eligibility does not only reduce employment but also increases individuals' likelihood of becoming a party member or a local political candidate. The effects are only driven by individuals with higher education and point towards the importance of time resources for political entry after a layoff.

A natural question that arises is what are the consequences of layoffs for the selection of politicians. Our results suggest that individuals' decision to engage with politics following layoffs is at least partly explained by economic rather than ideological motives. If the economic incentives vary across individuals' skill levels, economic shocks may also impact the average competence of politicians and, thereby, the functioning of the political system. To study implications for the selection of political participants, we compare the characteristics of treated and control individuals who become party members and candidates after 2012 (i.e., after the mass-layoff period). We document that treated members and candidates are positively selected in terms of previous earnings, education, and political experience as compared to those who do not experience a layoff.[2] Second, to study aggregate-level implications in a descriptive exercise, we use municipality data and investigate the correlation between changes in unemployment rates and changes in candidates' average education. We find a positive correlation between local unemployment and candidates' education, which is in line with the patterns we find at the individual level. Overall, the positive selection of politicians following layoffs and changes in local unemployment rates suggests that economic downturns may entail better politicians.

Our paper contributes to two main strands of research. First, it adds to the literature on the political consequences of economic shocks. A growing number of papers documents that voting patterns are responsive to negative macro shocks, such as import competition (Autor *et al.*, 2020; Dippel *et al.*, 2022), economic crisis (Barros and Santos Silva, 2019), automation (Anelli *et al.*, 2019), and austerity reforms (Fetzer, 2019). Similarly, at the individual level, a large set of studies document that changes in economic conditions affect individuals' policy preferences, such as their support for redistribution and welfare provision (e.g. Margalit, 2013; Martén, 2019;

[2]We acknowledge that higher levels of education and wages may not necessarily imply more qualified politicians. Yet, a number of papers in the literature use these variables to construct measures of politician's competence (e.g., Besley *et al.* (2017)).

Ahlquist *et al.*, 2020; Ballard-Rosa *et al.*, 2021). In terms of political participation, the literature has focused mostly on turnout and has found mixed results, ranging from turnout increasing with local unemployment rates (Charles and Stephens, 2013; Burden and Wichowsky, 2014), decreasing after individual job loss (Emmenegger *et al.*, 2017; Östermann and Lindgren, 2023), or not being affected at all by income changes (Jungkunz and Marx, 2022; Geys and Sørensen, 2024). Our contribution to this strand of literature is to show that economic conditions also affect the supply side of politics by changing individuals' propensity to enter the political system, either as members of a political party or as candidates in local elections. In addition, our mass-layoff identification strategy, together with the size and granularity of our data, allows us to circumvent the identification challenges associated with the endogeneity of personal changes in economic conditions and study the mechanisms behind individuals' decisions to engage with politics after a layoff.

Second, we speak to the literature studying the determinants of political selection.[3] In the tradition of the citizen-candidate framework (Osborne and Slivinski, 1996; Besley and Coate, 1997), this literature highlights the importance of endogenous entry into politics. Individuals' decision to run for office is modeled as a cost-benefit calculation, where the expected benefits of holding office are weighed against the costs of running (Black, 1972; Caselli and Morelli, 2004). On the benefit side, Ferraz and Finan (2011) and Gagliarducci and Nannicini (2013) document that increasing politicians' salaries attracts higher-quality candidates and improves the performance of incumbent politicians. In addition to salaries, holding a public office may lead to other non-salary earnings while in office (whether legal or not) and enhance private sector opportunities after leaving office (Eggers and Hainmueller, 2009; Fisman *et al.*, 2014). On the cost side, running for office may involve not only monetary, time, and opportunity costs, but also psychological and family burdens (Hall, 2019). Beyond individual factors, political selection is shown to be affected by institutional constraints, such as electoral rules (Arora, 2022), gender quotas (Besley *et al.*, 2017), or campaign spending limits (Avis *et al.*, 2022). Our primary contribution to this line of research is to show that adverse economic shocks increase individuals' likelihood to enter politics, either as party members or as candidates. The effects increase with several measures of individual competence, suggesting that layoffs can improve candidate quality. In addition, we show that unemployment benefits, by changing individuals' time constraints after layoffs, impact political selection, leading to an increased entry of better-educated individuals into politics.

This paper is structured as follows. Section 2 provides background information on party

[3] See Dal Bó and Finan (2018) and Gulzar (2021) for recent reviews of this literature.

members and local council candidates in Brazil. In Section 3, we present the datasets used in the analyses, and in Section 4, we outline our empirical strategy. Section 5 presents the main results and heterogeneity analyses for the effect of job loss. In Section 6, we investigate the role of UI benefit eligibility. Section 7 discusses our findings and potential implications for political selection, and Section 8 concludes.

2 Institutional Background

In Brazil, a significant portion of the population is engaged in politics. As mandated by the 1988 Brazilian constitution, voting is compulsory for all literate citizens aged 18 to 69.[4] Apart from voting, more than 16 million Brazilians were registered as members of a political party in 2020. This corresponds to about 11% of the electorate, one of the highest shares across democracies. In addition, between 2000 and 2020, about 1.7 million individuals have run for office in local elections in Brazil. In this section, we describe the institutional environment in which individuals make decisions about their involvement as party members or candidates in local politics.

2.1 Party members

Brazil has a multi-party political system, encompassing 33 registered parties as of 2024. The effective functioning of political parties relies heavily on an engaged membership base. According to the Brazilian electoral law, a party must have a minimum number of members equal to 0.5% of the votes cast in the previous national election. Challenging the notion that political affiliation in Brazil is solely driven by clientelistic motives, Ribeiro and Do Amaral (2019) show that most political parties are well-organized and have a strong base of activists. Members carry out the day-to-day organizational work of parties, participate in electoral campaigns, raise funding, attend party meetings, help maintain ties with the electorate outside of election years, and serve as the pool from which parties recruit their candidates and officials.

Registration as a party member is, in principle, open to all eligible voters. Nevertheless, the rules for registration and membership vary substantially across political parties. In addition to the opportunity costs related to registration and effective participation, in some parties, members have to pay registration fees and monthly dues.[5] Some parties use interviews and formal tests for the selection of members, while others, especially more recently, allow a simple

[4]The average turnout rate in Federal Elections in the last 20 years was about 80% (International Institute for Democracy and Electoral Assistance, 2024).

[5]For example, NOVO requires members to contribute a minimum of 460 BRL (93 USD in 2024 values) yearly, with a fee waiver for low-income individuals. PT has sliding-scale fees, with a minimum yearly fee of 30 BRL (6 USD) for low-income individuals and a maximum of 12% of members' net salary for those individuals earning more than 6 times the minimum wage. PSDB, on the other hand, does not charge any fee for their members.

online registration.

Membership recruitment in Brazil is predominantly a local phenomenon, fueled by internal competition at parties' municipal conventions. Local elites, aiming to secure the party's nomination, attract supporters by enlisting them in the party to build a loyal voting base at the convention and signal their potential as candidates (Sells, 2020; Frey, 2022). Supporters who become members of a winning political party are rewarded with employment opportunities within the government bureaucracy (Brollo *et al.*, 2017; Barbosa and Ferreira, 2023). The importance of members for local politics also becomes evident when looking at the distribution of new affiliations over time: as we show in Appendix Figure A.1, parties significantly expand their ranks in the year before local elections.

2.2 Local councilors

Brazil consists of 5,568 municipalities, each of which is governed by a mayor (*prefeito*) and a council of local legislators (*vereadores*).[6] Every four years, during the municipal elections, citizens cast one vote for a mayoral candidate and one vote for a council candidate.[7] Local councilors are elected based on a system of open-list proportional representation and serve four-year terms, with no term limits. The size of local councils varies depending on population, with a minimum of 9 councilors elected in municipalities with up to 15,000 inhabitants and a maximum of 55 councilors elected in municipalities with more than 8 million inhabitants. Local legislators are responsible for drafting laws, voting bills, proposing public works, and monitoring the actions of the executive. Brazil is one of the most decentralized countries in the world, with local governments managing substantial budgets to provide public services, such as local infrastructure, transportation, education, and health care. Local councils must approve the municipal budgets and thereby have great influence over the delivery of these public goods.

Working as a local councilor can be financially rewarding. The wages of councilors are substantial on average – amounting to about 2.5 times the average wage in their municipalities. Councilor wages also vary strongly by municipality size – being capped at 20% of state deputies' wages in municipalities with less than 10,000 inhabitants and at 75% in municipalities with more than 500,000 inhabitants. Importantly, serving as a councilor is not a full-time activity and allows legislators to retain their private-sector jobs after being elected. On average, legislators

[6] This number excludes the insular district of Fernando de Noronha and the Federal District, which are excluded from local elections.

[7] Voting for a party instead of an individual candidate for councilor is also allowed. Local elections are held at the same time for all municipalities across the country, usually in October. Elections for the President, governors, and federal and state legislators also take place every four years but are staggered by two years relative to local elections.

need to be present in the council only four days per month (Ferraz and Finan, 2011).

To become a candidate for local council, individuals must be affiliated with a political party one year prior to the election.[8] Each political party (coalition) nominates up to 1.5 (2) times the total number of available seats in the local council. Political parties are required to meet a gender quota, with at most 70% of candidates being of the same gender. The nomination of political candidates occurs at the parties' municipal conventions which take place between July and August of each electoral year. Due to the large number of available seats and parties' incentives to fill as many candidacies as possible, running for local council is relatively common, even for individuals without prior political experience.[9]

How large are the financial costs associated with candidacies for local council? To finance their campaigns, candidates rely on party funds, financial contributions from individuals and firms, as well as own resources.[10] We collected data from the *Tribunal Superior Eleitoral* on all campaign contributions received by councilor candidates in the electoral cycles 2004 to 2020. Appendix Table A.1 presents detailed summary statistics for each year. Across all years, the median candidate raised a total of 1,587 BRL (392 USD), with only 168 BRL (34 USD) coming from own resources. Only about half of all candidates use any private funds for the campaign. Thus, the majority of resources financing candidates' campaigns do not originate from their own money suggesting that the financial barriers to entering politics as a councilor candidate are limited.

3 Data

In our analysis, we combine detailed information on the universe of party members and local political candidates with administrative data on the population of workers employed in the Brazilian formal labor market. This section describes the data sources as well as the steps we follow to construct and merge our estimation sample. By having identified records on individuals' employment histories as well as political outcomes, we can uniquely identify whether individuals' political behavior changes after they lose their jobs. Additionally, detailed information on individual, party, and municipality characteristics allows us to disentangle potential mechanisms.

[8]Additionally, candidates must be Brazilian citizens, aged at least 18 years old, have residency in the municipality where they are running for at least one year, possess full political rights, and possess a certificate of completion of military service (for men).

[9]The number of seats the party wins depends on the total number of votes it receives, creating an incentive for parties to fill as many candidatures as possible.

[10]Since 2015, campaign spending limits have been introduced in Brazilian elections (Avis *et al.*, 2022). Besides banning financial contributions from private firms, the electoral law established spending limits for political campaigns depending on municipalities' population. Importantly, since 2019, candidates for local council are restricted to finance only up to 10% of the spending limit for their municipality with their own resources to ensure more equitable running conditions.

3.1 TSE

To measure individuals' decision to enter politics, we use two datasets provided by the *Tribunal Superior Eleitoral* (TSE). Information on the universe of political members comes from the *Sistema de Filiação Partidária* (FILIA), a national registry containing records of all affiliations to political parties, including individuals' names, affiliated party, as well as date and municipality of registration. The data cover the period 2000 to 2020. Second, we gather data on the universe of individuals running for local council in the six election cycles of 2000, 2004, 2008, 2012, 2016, and 2020. The data contain information on candidates' taxpayer registry numbers (*Cadastro de Pessoas Físicas*, henceforth CPF), socio-demographic characteristics, party of candidacy, whether they have been elected, and municipality of candidacy.

3.2 RAIS

We obtain matched employer-employee data covering the universe of formally employed workers between 2004 and 2018 from the *Relação Anual de Informações Sociais* (RAIS), an administrative database collected annually by the Ministry of Labor and Employment (MTE).[11] The database was created in 1975 for statistical and administrative purposes and contains detailed information such as hiring and separation date, average monthly earnings, contracted hours, occupation, firm's location and industry, and worker's socio-demographic characteristics. For our analysis, we make use of the identified dataset, containing individuals' names, taxpayer registry numbers (CPF), as well as firms' tax identification numbers (*Cadastro Nacional da Pessoa Jurídica*, henceforth CNPJ). This enables us to identify individuals' complete employment histories over time and across firms.

3.3 Merging political and labor market data

We merge the information on political candidates with the employment data using individuals' taxpayer registry number (CPF). As for party membership, we merge this information to the RAIS data based on each individual's full name. To minimize measurement error when exactly matching on names, we restrict our sample to individuals who have a unique name in Brazil. Individuals with unique names cover roughly half of the population, given that Brazilians commonly have multiple surnames, including at least one from both their father and mother.[12]

[11] In 2015, formally employed individuals accounted for approximately 55% of the Brazilian labor force (Derenoncourt *et al.*, 2021).

[12] To identify unique names, we build a dataset of names in Brazil. We collect the names of all workers who appear in the RAIS data in any year between 2004 and 2018, and supplement this list with information on the universe of beneficiaries in the social programs *Bolsa Família* and *Benefício de Prestação Continuada* between 2013 and 2021 and on all party members between 2000 and 2020. The resulting list covers about 90% of the Brazilian

Appendix Table A.2 compares the characteristics of laid-off individuals in our estimation sample who have a unique and non-unique name, respectively. Those with a unique name are slightly positively selected (9% higher earnings, 8% more years of education, 0.9 percentage points more likely to be a manager), and are more likely to be female (8.1 percentage points) and white (5.3 percentage points). Given that we merge our second political outcome - local candidacies - using the CPF, we have this information for individuals with and without a unique name. Before the layoff, we do not observe differences in the likelihood of running for councilor across both groups. While we focus on the sample of individuals with unique names (in which we have merged political members and candidates) in our baseline specification, we also analyze the external validity of the results by estimating layoff effects on candidacies in the sample of individuals with non-unique names.

4 Empirical Strategy

The main difficulty in estimating the effect of layoffs on individuals' entry into politics is the fact that layoffs are not random. On the one hand, individual shocks could simultaneously influence both political participation and the likelihood of losing a job, leading to a spurious relationship. On the other hand, individuals' political behavior could also affect their probability of being dismissed. Colonnelli *et al.* (2022) show, for instance, that political discrimination affects individuals' labor market outcomes in Brazil. To address these identification challenges, we exploit the timing of firm-level mass layoffs as a source of plausibly exogenous variation in individual-level job loss using a difference-in-difference design. The timing of mass layoffs is arguably orthogonal to individual workers' political behavior and other confounding worker-level shocks, and has been widely used to estimate the effects of job loss on various outcomes. In particular, our empirical strategy closely follows recent papers that study the effect of dismissals in Brazil (Bhalotra *et al.*, 2021; Britto *et al.*, 2022a,b; Amorim *et al.*, 2023). This section starts by describing the steps we follow to construct our estimation sample and then proceeds to present our estimation model.

4.1 Sample selection

We restrict our sample to workers aged 25 to 50 with a full-time (i.e. at least 30 hours per week), open-ended, private-sector job of at least six months tenure. Among those, the treatment group

adult population. Not observing the names of a small part of the population will introduce some measurement error in our merged outcome variables which is, however, not expected to be related to the likelihood of being laid-off.

comprises all workers displaced without just cause from a mass layoff firm between 2009 and 2011.[13] This treatment period allows us to study long-run effects over three election cycles (12 years) after the layoff, and also to check for pre-trends over two election cycles (8 years) before the layoff. In our baseline specification, we define a mass layoff firm as a firm with at least 30 workers dismissing at least 30% of its workforce in a given year, and later we show that our results are robust to various thresholds of firm size and layoff share.[14]

We then match each treated individual with a control worker who is not dismissed in the same year and who works in a non-mass layoff firm.[15] To obtain a comparable control group, we perform an exact matching on gender, education (5 categories), age (deciles), monthly earnings (deciles), tenure (deciles), firm size (deciles), and state (27 categories).[16] For 79.5% of the workers in the initial treatment pool, we are able to find a comparable control worker. We end up with an estimation sample that consists of nearly one million treated and matched control workers.

Table 1 presents summary statistics for treated and control individuals and shows that both groups are similar in socio-demographic, job, and firm characteristics. For all considered variables, the standardized mean difference between the two groups is well below the threshold of 0.20 recommended by Imbens and Rubin (2015). Importantly, treated and control units are not only balanced in terms of the matching variables but also in terms of pre-treatment political outcome variables that were not part of the matching process. In both groups, approximately 7% of individuals were members of a political party in 2008. Nearly 2% newly registered as a member in the election cycle between 2005 and 2008, and 0.1% ran for local councilor in the 2008 election.[17] The validity of our DiD design does not require that treated and control individuals are similar in levels, but covariate balance enhances the likelihood that both groups exhibit similar outcome trends, making the required parallel-trends assumption more plausible.

[13]We exclude individuals who are recalled by their initial employer within five years after the layoff. In cases where individuals experienced multiple layoffs between 2009 and 2011, we keep only the first one.

[14]We always drop firms from the mass layoff sample if they reallocate under a new identifier, identified as those firms in which at least 30% of their dismissed workers transition to the same new firm identifier in the following year.

[15]Note that control workers are not laid off in the matching year but may be laid off in subsequent years. Given that we estimate effects for up to 12 years after the layoff, restricting control workers to those who were continuously employed throughout the whole observation period would result in a small and highly selected pool of control workers. Defining control workers without conditioning on future employment follows a large share of papers in the job loss literature (e.g., Britto *et al.*, 2022a; Bertheau *et al.*, 2023; Schmieder *et al.*, 2023). In addition, we do not impose any restrictions on potential layoffs prior to the matching year, other than the minimum six-month tenure requirement. Given the high turnover rates in Brazil, stricter tenure requirements would yield a substantially selected sample. In Appendix Figure B.1, we show that treated and matched control individuals were not differently exposed to mass layoffs before the treatment period.

[16]We allow control individuals to be matched to more than one treated individual. In cases where more than one control individual was matched to one treated individual, we randomly selected one control individual.

[17]Appendix Table A.3 also shows that there are no differences in the ideology, local incumbency, and union affiliation of the parties that treated and control observations affiliated with or ran for before the layoff.

Table 1: Summary statistics

	(1) Treated	(2) Control	(3) Std Diff
Socio-demographic characteristics			
Male	65.78	65.78	0.00
White	57.32	60.00	0.05
Age	33.84	33.85	0.00
Years of education	10.72	10.77	0.02
Job and firm characteristics			
Earnings (per month, BRL, CPI 2018)	2158.15	2140.37	-0.01
Tenure (months)	32.62	32.85	0.01
Manager	2.70	2.20	-0.03
Firm size	610.80	625.65	0.01
Political outcomes before layoff			
Party membership (2008, %)	7.47	7.30	-0.01
New party membership (2005-2008, %)	1.99	1.89	-0.01
Candidate (2005-2008, %)	0.11	0.10	-0.00
Observations	944,214	944,214	

Note: The table reports the average characteristics of treated workers who are displaced from a mass layoff firm (column (1)) and matched control workers who are not displaced in the same calendar year (column (2)), and the standardized difference between the two groups (column (3)). The sample only includes workers with a unique name within the country. Appendix Table **??** reports results for additional pre-layoff political outcomes split by party characteristics.

4.2 Estimation model

Using the sample of treated and matched control individuals, we estimate the following dynamic DiD (event-study) model:

$$Y_{ic} = \alpha + \beta Treat_i + \sum_{c=2000, c\neq 2008}^{2020} \delta_c (Treat_i \times Cycle_c) + \mu_{m(i)k(i)c} + \epsilon_{ic}, \tag{1}$$

where i denotes workers and c denotes four-year municipal election cycles (2000, 2004, 2008, 2012, 2016, 2020). Our main outcomes, Y_{ic}, are binary variables capturing new party memberships and candidacies for local councilors in a given election cycle.[18] The effects of interest are captured by δ_c which are the coefficients for interactions between a treatment group dummy, $Treat_i$, and dummies for each election cycle, $Cycle_c$. For $c = \{2012, 2016, 2020\}$, the parameters identify the dynamic treatment effects (relative to the baseline period $c = 2008$), and for $c = \{2000, 2004\}$

[18] Although we have information on the exact date individuals affiliate with a political party, we focus on cycle-level information since new affiliations are highly concentrated around election years (see Appendix Figure A.1). Later on, we present a year-level analysis, which allows us to test for pre-trends closer to the layoff event and to assess political responses in the shorter run.

they check whether outcomes evolved in parallel in the pre-treatment period.[19] We also estimate a static version of the DiD model:

$$Y_{ic} = \alpha + \beta Treat_i + \delta_{DiD}(Treat_i \times \mathbb{1}[c \geq 2012]) + \mu_{m(i)k(i)c} + \epsilon_{ic}. \tag{2}$$

which summarizes the average treatment effects over all periods in the coefficient δ_{DiD}. In both models, we include period-specific municipality $\times$ 2-digit industry fixed effects, $\mu_{m(i)k(i)c}$, which flexibly control for local and industry shocks that may simultaneously cause mass layoffs and changes in political behavior. Importantly, the municipality fixed effects also hold constant the local political environment and allow us to separate individual-level supply-side effects from potential effects of mass layoffs on the demand for politicians among the local electorate.

5 Effects of Job Loss on Political Entry

5.1 Main results

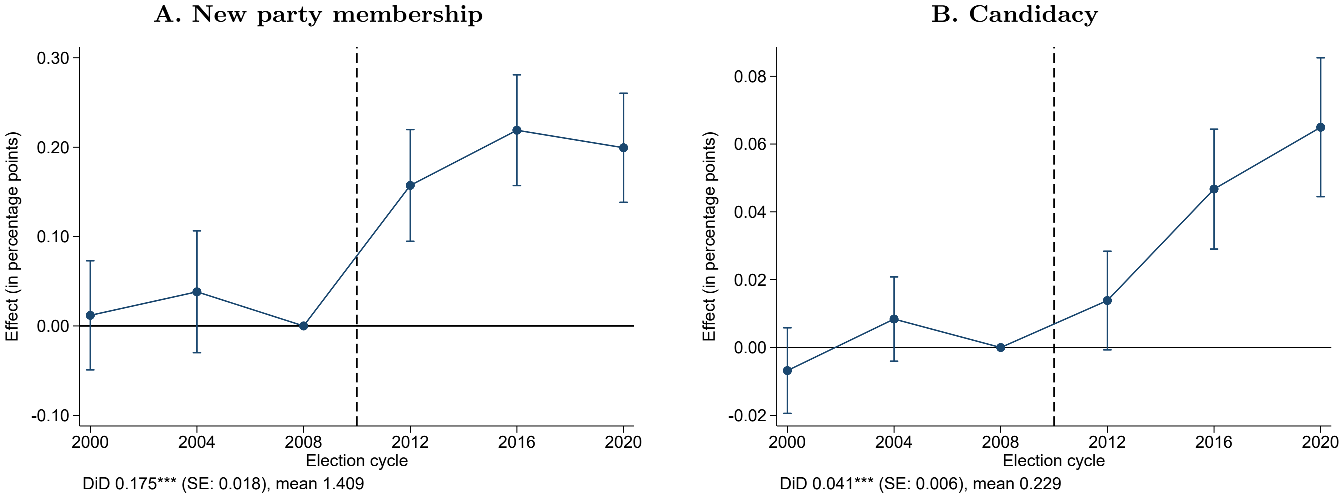
Figure 1 reports the results from models (1) and (2) for the effect of job loss on newly registering as a party member and running for local council. We observe that both outcomes do not develop significantly differently across treated and control individuals before the layoff, supporting the parallel-trends assumption. After the layoff, outcomes sharply diverge. We estimate that job loss increases the likelihood of registering as a party member by about 0.20 percentage points, which is a 12% increase relative to the control group mean.[20] The effect occurs in the election cycle immediately after the job loss and remains at a similar level for the following two cycles. For the probability of running as a candidate for local council, we also observe a significantly positive effect that increases from 0.01 percentage points (6% relative to the control mean) in the cycle of the layoff to about 0.06 percentage points (28%) two cycles after. The dynamic effect observed for candidacies may be explained by the preparation time required for election campaigns, coupled with the fact that local councilors have no term limit and often seek re-election after their initial candidacy. In sum, the results indicate that job loss has a strong and long-lasting positive effect on individuals' probability of engaging with politics, either as members of political parties or as local candidates.

The persistence in the effect of job loss on political behavior is consistent with its effects

[19] Note that our estimation model relies on a single treatment period, as all treated workers were laid off in the 2012 cycle. Thus, the model is not susceptible to weighting issues that arise in staggered difference-in-differences models with heterogeneous treatment effects (de Chaisemartin and D'Haultfœuille, 2020; Goodman-Bacon, 2021; Sun and Abraham, 2021).

[20] We compare effect sizes to the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020).

Figure 1: Effect of job loss on party membership and running for local councilor



Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). The treated group contains workers who were displaced from a mass layoff firm between 2009 and 2011, and the matched control group comprises workers who were not displaced in the same calendar year and who worked in a non-mass layoff firm. $N = 11,317,632$ individual-cycle observations. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

on labor market outcomes. In Appendix Figure B.1, we show, using labor market information from RAIS, that employment and earnings effects of job loss are also long-lasting. Our estimates indicate that seven years after the layoff (the longest period we can observe with the RAIS data), dismissed workers still have a 15 percentage points lower probability of being formally employed and receive about 5,000 BRL (1,014 USD) fewer earnings per year.

Robustness and external validity. Under our baseline definition of mass layoffs, there may still be considerable scope for firms to select which 30% of workers to dismiss. Appendix Table B.1 evaluates the robustness of our main findings when applying more stringent mass layoff definitions. Irrespective of whether we increase the minimum share of dismissed workers, use plant closures, increase the minimum firm size, or focus on firms with stable employment growth before the mass layoff, we still find similarly large positive effects of job loss on party memberships and candidacies. These results also mitigate concerns regarding the external validity of our analyses. The effects of mass layoffs could differ from those of regular layoffs, for example, due to spillover effects across laid-off workers or increased attention to mass layoffs by politicians or the media. The absence of differential effects by the share of displaced workers suggests that the size of a mass layoff is not an important factor driving our results. In Appendix Table B.2 we also explore the sensitivity of our results when including different sets of municipality and industry fixed effects in the DiD model. We show that our results are largely unaffected by accounting for local and temporal variation, suggesting that our matching strategy already compares treated and control workers in similar labor markets. Finally, for the candidacy outcome, we check the external validity of our results for individuals without a unique name. As shown in Appendix Figure B.2, the increase in the probability of running for office following job loss is very similar across individuals with and without unique names. Altogether, the robustness checks reduce concerns that our estimated effects are driven by a particular mass-layoff definition, model specification, or sample restriction.

Effect timing. Next, we take a closer look at the temporal dynamics in the estimated treatment effects. Given that deciding to become a candidate and preparing for an election campaign requires time and planning, we start by examining whether our results differ by proximity to the next local election.[21] Appendix Figure B.3 shows results when splitting the sample by the layoff year (2009, 2010, or 2011).[22] We find that the positive effects of job loss on party memberships and candidacies are not driven by one specific treatment year, in particular not by the temporal proximity to the local election. This dynamic also speaks in favor of the external validity of our results. For instance, there is a concern that political responses following layoffs in 2009 could differ from other layoffs, given the proximity to the 2008 financial crisis. However, our results show similar effects for layoffs occurring in 2010 and 2011, once the crisis had already faded out in Brazil (Barbosa, 2010).

In addition, we conduct a yearly-level analysis exploiting more detailed information on the exact timing when individuals become members of a political party.[23] We estimate the following event-study model for new party memberships measured in each year relative to the layoff:

$$Y_{it} = \alpha + \beta Treat_i + \sum_{k \neq -1} \delta_k (Treat_i \times \mathbb{1}[t = t^* + k]) + \mu_{m(i)k(i)t} + \lambda_{tt^*} + \epsilon_{it}, \tag{3}$$

where i, t, and t^* refer to workers, calendar years, and layoff years, respectively. δ_k capture the dynamic treatment effects for each event year after the layoff k.[24] Results are reported in

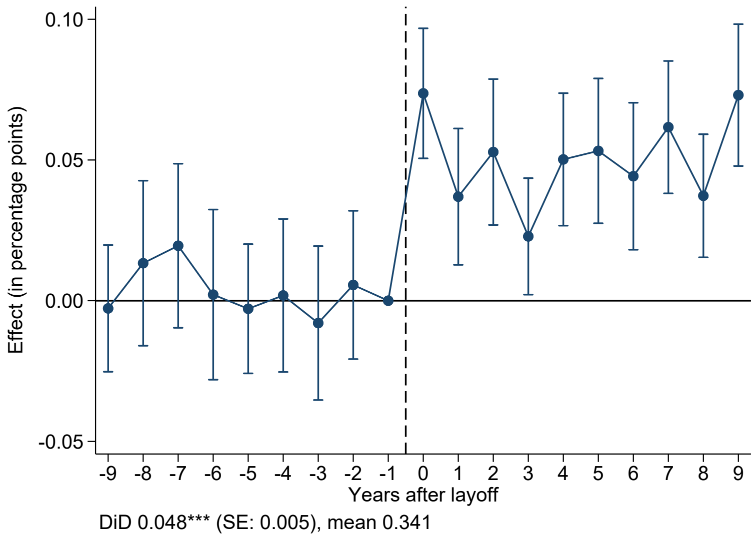
[21] As mentioned in Section 2.2, in order to become candidates, individuals need to be affiliated to a political party one year prior to the election. Additionally, parties chose their candidates during party conventions, which occur a few months before the election.

[22] For this analysis and the subsequent year-level event-study model, we assign control workers the layoff year of its matched treated worker.

[23] We also rely on the year-level model when investigating the labor market consequences of layoffs (that we discussed before) as well as the consequences of layoffs for individuals' engagement with political campaigns (which will be presented in the following paragraph).

[24] The model includes calendar year × layoff year fixed effects, λ_{tt^*}, to ensure that the identifying variation only comes from comparing treated to control individuals who have been matched in the same year. This stacking approach avoids the "forbidden comparisons" between late and early-treated individuals that can cause negative weights under heterogeneous treatment effects (Cengiz *et al.*, 2019; de Chaisemartin and D'Haultfœuille, 2020; Goodman-Bacon, 2021; Sun and Abraham, 2021).

Figure 2: Effect of job loss on new party membership - yearly-level analysis



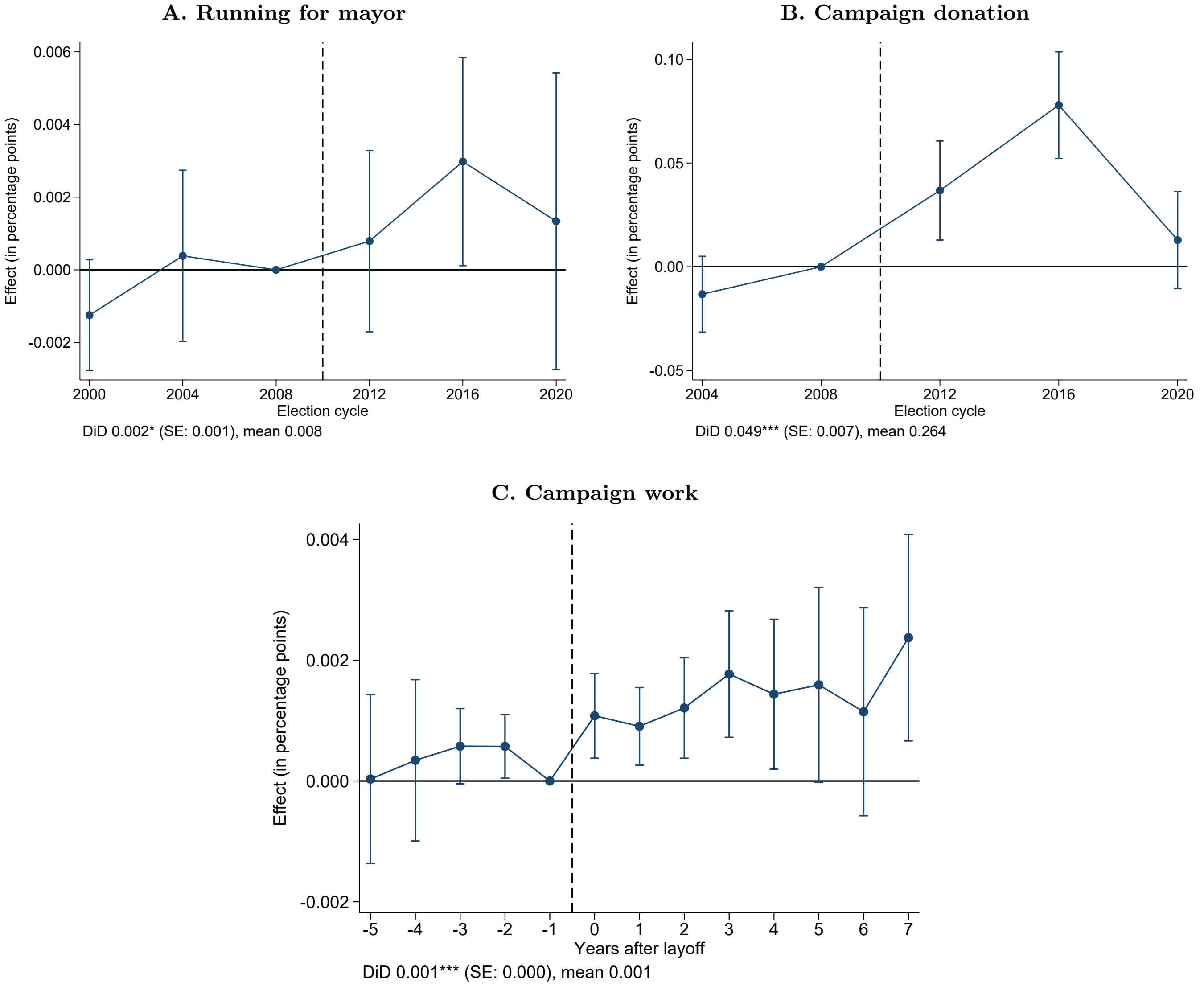
Note: The figure reports event-study coefficients δ_k, estimated in model (3), for the effect of job loss on the likelihood of newly registering as a party member. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. $N = 35,880,132$ individual-year observations. Below the graph, the DiD coefficient from a static version of model (3), its standard error, and the mean of the control group in the post-treatment period are reported.

Figure 2. We find positive and statistically significant effects on individuals' probability of newly registering as a party member that start in the year directly following the layoff and persist up to 9 years after the shock. Reassuringly, we do not detect pre-trends before the layoff, neither in the year immediately preceding the layoff nor in earlier years. Taken together, the short-term dynamics are in line with the patterns we document at the cycle level.

Additional participation outcomes. In addition to estimating the effects on new party memberships and candidacies for local council, in Figure 3 we also investigate the consequences of job loss for three additional forms of political participation. First, we consider candidacies for municipal mayor. Compared to councilor candidacies, running for mayor requires a more elaborate election campaign to win the support of party members during the party's municipal convention. Moreover, mayoral candidacies typically involve greater resources and party investments, increasing the constraints associated with individual candidacies. Nevertheless, we also find positive effects of job loss on the likelihood of running for mayor. Given the rarity of mayoral candidacies, the results are not estimated very precisely, however. The positive effect is highest and significant only in the medium term (in the 2016 cycle).

Second, we examine the effect of job loss on donations to political candidates. In Brazil, individuals are allowed to donate up to 10% of their annual income to a political campaign. For candidates, individual contributions serve as an important source for their campaign spending.

Figure 3: Effect of job loss on additional participation outcomes



Note: The figure shows results for the effect of job loss on the likelihood of running for municipal mayor (Panel A), donating to a political campaign (Panel B), and working for a political campaign (Panel C). Information on mayoral candidates and campaign donations are obtained from the *Tribunal Superior Eleitoral.* The data on campaign donations is only available starting from the 2004 election cycle. Following Bazzi and Labanca (2023), campaign workers are identified in RAIS as individuals who have a full-time contract at a firm in industry *Atividades de Organizações Políticas* (CNAE 1.0: 91928, CNAE 2.0: 94928) or with legal nature *Partido Político* (3123) or *Candidato a Cargo Político Eletivo* (4090). Panels A and B report cycle-specific event-study coefficients δ_c from model (1), while Panel C reports year-specific event-study coefficients from model (3). The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below the graph, the DiD coefficient from static versions of models (1) and (3), its standard error, and the mean of the control group in the post-treatment period are reported.

Consequently, political candidates frequently reward their supporters by offering jobs in the public sector (Colonnelli *et al.*, 2020). Our findings reveal a positive effect on campaign donations in the two cycles after job loss.[25] The estimates are both highly significant and substantial in magnitude. The DiD coefficient indicates a 0.049 percentage point increase in individuals' probability of donating to a political campaign, corresponding to an 18% increase relative to the

[25] Data on campaign contributions is also obtained from the *Tribunal Superior Eleitoral.* We exclude contributions to individuals' own campaigns to ensure that the results are not only driven by own candidacies.

control mean. The results suggest that the financial setback resulting from job layoffs does not hinder individuals from participating in political endeavors that necessitate monetary contributions. This also aligns with the observed positive impact on party memberships and candidacies, both of which may entail financial investments in the form of membership fees or campaign funding.

Third, we study the possibility of working for a political campaign. Bazzi and Labanca (2023) show that working for a winning mayoral campaign in Brazil yields substantial private income gains from improved access to jobs in the local bureaucracy. We follow the authors and identify dedicated salaried campaign staff in the RAIS data using information on the industry and legal nature of the employers.[26] Our results show an increase in individuals' probability of working for a political campaign after job loss. Despite also being a rare event in our sample, the estimates are highly significant and suggest roughly a doubling in the likelihood of becoming a campaign worker. The effects are also highly persistent up to seven years after the layoff which is in line with our findings on party memberships and councilor candidacies.

5.2 Party disaggregations

Job loss may not only affect overall political participation but may also lead to changes in individuals' support of different political parties. We, therefore, study heterogeneities in the effects on memberships and candidacies by party characteristics, examining party ideology, local incumbency of the party, and party's alignment with labor unions. Estimates from the static DiD model (2) are presented in Table 2, while the dynamic event-study results from (1) are shown in Appendix Figures B.4, B.6, and B.5. We also report results from model (2) separately for each of Brazil's major parties in Appendix Figure B.7.

Ideology. To analyze whether layoffs lead to ideological shifts, we distinguish left, center, and right parties, following the classification used by Colonnelli *et al.* (2022).[27] Our results, shown in columns (2) to (4) of Table 2, indicate that job loss leads to an increase in party memberships and candidacies across the ideological spectrum. Relative to the control mean, we find an increase of 8.4%, 11.0% and 17.6% in the likelihood of individuals newly affiliating with left, center and right parties, respectively. The effects on candidacy reveal a similar pattern, with likelihood increases of 15.1%, 17.3%, and 21.3% for running for office in a left, center, or right-wing party. Although we observe larger effects for right-wing parties, our results show that center and left-wing parties

[26] Specifically, we define campaign workers as individuals who have a full-time contract at a firm in industry *Atividades de Organizações Políticas* (CNAE 1.0: 91928, CNAE 2.0: 94928) or with legal nature *Partido Político* (3123) or *Candidato a Cargo Político Eletivo* (4090).

[27] The ideological classification of parties is shown in Appendix Table B.3.

also receive more members and candidates following layoffs. Interestingly, the event-study results presented in Appendix Figure B.4 show that the higher increase in affiliations with right-wing parties is driven only by the 2020 cycle, the first local election after Bolsonaro took power in 2019. The rise in right-wing support following economic hardships, which is well documented in the literature (e.g., Barros and Santos Silva, 2019; Autor *et al.*, 2020; Dippel *et al.*, 2022), appears to be contingent on the current political climate.

Local incumbency. Next, we check if individuals punish the party of the local incumbent government following layoffs. Workers may blame incumbent politicians for local economic policies that can result in mass layoffs, or for insufficient support policies that mitigate economic hardship after a layoff (e.g., Tilley *et al.*, 2018; Ahlquist *et al.*, 2020). In columns (7) and (8) of Table 2, we compare the effects for the party of the elected mayor in each individual's municipality before the layoff in 2008 and all other non-incumbent parties. The results show that job loss leads to increased participation in both incumbent and non-incumbent parties. Relative to the control mean, the effects are considerably stronger for incumbent parties, with an increase of 24.0% in the probability of being affiliated with the party of the incumbent and an increase of 43.7% in the likelihood of running for office in the party of the incumbent (as opposed to 11.6% and 16.7% in non-incumbent parties). The results do not support incumbent punishment as a main driver of our results.

Union affiliation. Lastly, we investigate whether there are heterogeneities depending on parties' alignment with labor unions. Unions can influence employees' political behavior through various activities, such as educating, informing, and mobilizing their members, endorsing selected candidates and policies, and organizing get-out-the-vote campaigns (e.g., Kim and Margalit, 2017; Matzat and Schmeißer, 2023). In Brazil, labor unions have historically been linked to left and centrist parties and have played an important role in advocating for worker interests in domestic politics. Until 2017, unions were required to participate in negotiations to formalize dismissals in the context of mass layoffs.[28] We define union-affiliated parties following the classification of Ogeda *et al.* (2021) and compare them to all other left-wing or center parties that have no historical ties with labor unions in Brazil.[29] Our results show an increase in the probability of individuals affiliating or becoming local candidates in both union-affiliated and non-union-

[28]In 2017, the Brazilian labor reform (Law 13,467/17) established that it was no longer mandatory for unions to take part in collective bargaining agreements during mass layoffs.

[29]Union-affiliated parties are PT, PDT, PSB, PCB, PSD, and MDB, while non-union left and center parties are CIDADANIA, PCDOB, PV, PMN, PSOL, PSTU, PCO, SD, PROS, PPL, PMB, REDE, UP, PSDB, PTB, and AVANTE.

Table 2: Heterogeneity by party characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	All	Ideology			Local incumbency		Union affiliation	
		Left	Center	Right	Yes	No	Yes	No
[A] Outcome: new party membership								
δ_{DiD}	0.175***	0.045***	0.039***	0.091***	0.024***	0.151***	0.033***	0.051***
	(0.018)	(0.011)	(0.010)	(0.011)	(0.006)	(0.018)	(0.011)	(0.010)
Control mean	1.407	0.537	0.353	0.517	0.099	1.308	0.442	0.448
Relative effect	12.4%	8.4%	11.0%	17.6%	24.0%	11.6%	7.5%	11.3%
Observations	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632
[B] Outcome: candidacy								
δ_{DiD}	0.041***	0.012***	0.010***	0.019***	0.005***	0.036***	0.010***	0.012***
	(0.006)	(0.003)	(0.003)	(0.004)	(0.001)	(0.006)	(0.003)	(0.003)
Control mean	0.229	0.078	0.060	0.090	0.013	0.216	0.066	0.072
Relative effect	18.1%	15.3%	17.2%	21.1%	43.1%	16.6%	15.5%	16.7%
Observations	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632	11,317,632

Note: The table reports DiD coefficients, estimated in model (2), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). Column (1) presents the baseline results for all parties. Columns (2) to (4) differentiate left, center, and right parties, following the ideological classification of Colonnelli *et al.* (2022) which is shown in Appendix Table B.3. Columns (5) and (6) distinguish memberships and candidacies in the party of the elected mayor in the individual's municipality in 2008 and all other non-incumbent parties. Columns (7) and (8) show results separately for parties historically affiliated with labor unions (PT, PDT, PSB, PCB, PSD, and MDB) and all other left or center parties. All coefficients, standard errors, and control means have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

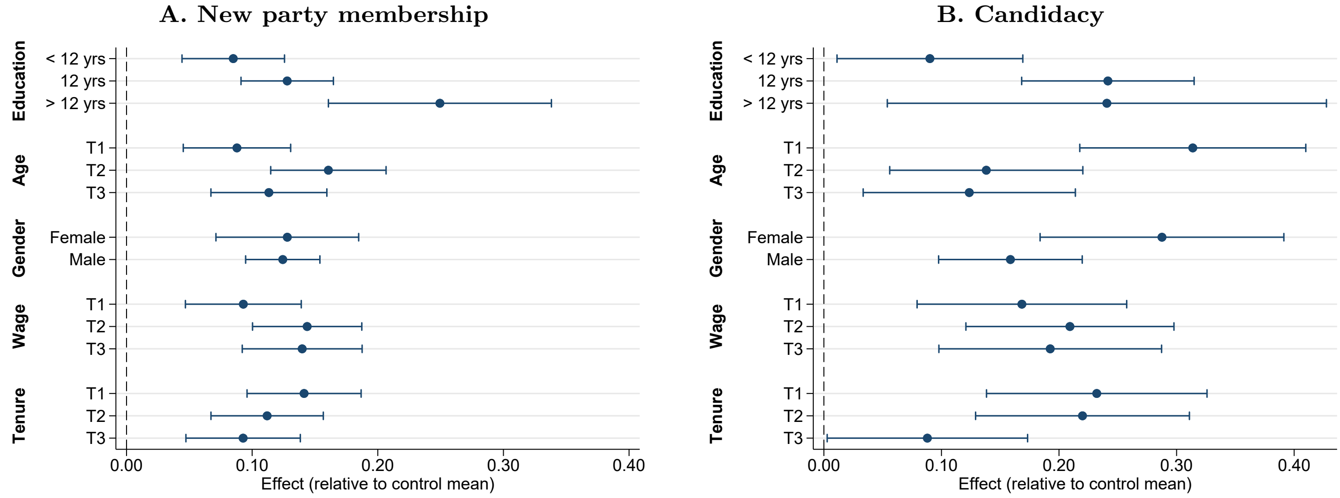
affiliated parties (columns (5) and (6) of Table 2). In quantitative terms, we estimate a 7.5% (11.3%) increase in the probability of newly affiliating with union-aligned (non-union aligned) parties and a 15.4% (16.7%) increase in candidacy probabilities in union-affiliated (non-union affiliated) parties following mass layoffs. Thus, we do not find evidence for a large role of unions in politically mobilizing laid-off individuals.

5.3 Individual heterogeneity

As a next step, we analyze heterogeneity in the effects of job loss along individual characteristics. We test for differential effects by the pre-layoff monthly earnings, tenure, age, education, and gender. For the first three variables, we differentiate between sample terciles, while for education, we distinguish less than 12 years (no school, elementary school, or middle school), 12 years (high school), and more than 12 years (university education). For each group, we separately estimate our DiD model (2) and scale the treatment effect by the group-specific control mean. Results are shown in Figure 4. One striking pattern that emerges is that the effects of job loss on party memberships and candidacies are remarkably pervasive, being significantly positive in all sub-samples considered.

Beyond the overall positive effects, we find substantial differences across educational groups

Figure 4: Heterogeneity by individual characteristics



Note: The figure reports DiD coefficients, estimated in model (2), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B) for different subgroups of individuals. For education, we differentiate no degree, elementary school, or middle school (< 12 years), high school (12 years), and university education (> 12 years). For age, pre-layoff monthly wage, and tenure, the sample is split by terciles of each variable. The horizontal bars depict 95% confidence intervals based on standard errors clustered at the individual level. All coefficients and standard errors are scaled by the group-specific mean of the control group in the post-treatment period (average across 2012, 2016, and 2020).

with higher effects on party memberships and candidacies for better-educated individuals. In quantitative terms, we estimate that individuals with university education experience an increase of about 25% in both the probability of newly affiliating with a political party and running for office, while for individuals with lower schooling levels, the effects on both outcomes are smaller than 10%. One potential explanation may be found in differential shocks to the opportunity costs of entering politics after job loss. Highly educated individuals can generally be expected to have better earnings prospects in the private sector. Being displaced due to a mass layoff may thus imply a larger shock to their opportunity costs in terms of increased time availability and reduced labor income, making a political career more attractive. These results may also have important implications for the competence of individuals who choose to enter politics, a point we will discuss in more detail in Section 7.

For the effects on new party memberships, we do not observe a clear gradient for any of the other individual characteristics considered. In contrast, for the impact on candidacies, we find higher effects for younger and low-tenured individuals and for women. The larger increases for young workers are consistent with the "impressionable years" hypothesis according to which political attitudes are more malleable at younger ages (e.g., Krosnick and Alwin, 1989). Moreover, the labor market careers of younger individuals are less entrenched, making it easier to start a political position after the layoff. Given the large gender gaps in labor market opportunities, careers are also less entrenched for women. This potentially explains women's higher likelihood of transitioning to a political career after a layoff.[30] The results are also consistent with the larger political responses following female economic shocks, as compared to male shocks, documented by Barros and Santos Silva (2019) for Brazil.

5.4 Monetary incentives

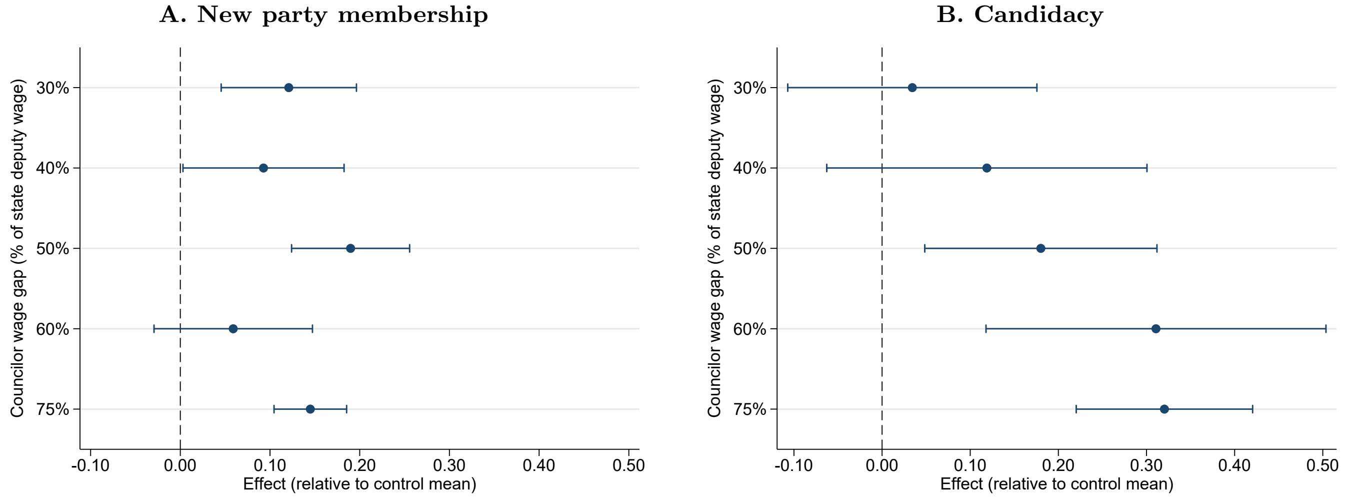
Holding a political position often brings monetary benefits. Running for political office may thus be an appealing strategy for mitigating the income loss resulting from being laid off in the private sector. To examine the role of financial incentives in shaping the decision to enter politics after a job loss, we conduct two heterogeneity analyses. First, we study the attractiveness of local councilor jobs by exploiting variation in councilor wages across Brazilian municipalities. Second, to capture changes in private-sector opportunity costs, we construct a measure of the predicted income loss that arises from job loss and test whether it relates to differences in job loss effects on political participation.

Councilor wages. As documented by Ferraz and Finan (2011), since 2004 a constitutional amendment has set caps to councilor salaries that depend on the population size of the municipality. In municipalities with less than 10,000 inhabitants, councilors can receive a maximum of 20% of state deputy salaries. This share increases for larger municipalities, up to 75% in municipalities with more than 500,000 inhabitants. State deputy salaries, in turn, can get at most 75% of the federal deputy salary. Appendix Table B.4 reports the maximum allowed councilor salary for each population size group, as estimated by Ferraz and Finan (2011). In absolute terms, maximum councilor wages vary between 1,927 BRL (391 USD) and 7,227 BRL (1,466 USD) per month, demonstrating that there is substantial variation in the financial incentives to run for local councilor.

In Figure 5, we examine heterogeneity in the effects of job loss by the maximum wage that councilors can receive in an individual's municipality. We find a strong gradient in the effects on running for councilor. In municipalities with the lowest wage caps, job loss does not significantly affect local candidacies. In contrast, in municipalities in which councilors can earn up to 75% of state deputy salaries, job loss is estimated to increase the likelihood of running for councilor by more than 30% relative to the control mean. At the same time, we do not find similar patterns for the effects on party memberships. As party members do not receive wages, this result is

[30] Even though in Brazil the political system is still heavily male dominated, policies have been implemented to increase the share of female politicians. Since 1997, parties have been restricted to nominating a maximum of 70% of candidatures of the same gender. In addition, starting from 2021, parties were mandated to allocate at least 30% of electoral funds to support female candidacies and reserve 30% of free advertisement time for female candidates.

Figure 5: Heterogeneity by councilor wages



Note: The figure reports DiD coefficients, estimated in model (2), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). We split the sample by the maximum monthly wage that councilors can earn (in % of state deputy salaries) in the municipality that the individual has worked in before the layoff. The wage cap groups are determined by municipality population counts of 2008 (see cutoffs in Appendix Table B.4). We pool the two lowest population groups, referring to a maximum councilor wage of 20% and 30%, as the lowest population group is too small to yield meaningful results. The horizontal bars depict 95% confidence intervals based on standard errors clustered at the individual level. All coefficients and standard errors are scaled by the group-specific mean of the control group in the post-treatment period (average across 2012, 2016, and 2020).

expected and gives us confidence that we are capturing the role of financial incentives, rather than other differences between small and large municipalities, in influencing the decision to enter politics after being laid off.

Predicted income loss. Next, we examine whether predicted income losses due to the layoff are related to the effects of job loss on political entry.[31][32] The income effects of job loss are predicted as follows. Similar to Schmieder *et al.* (2023), we first construct for each laid-off worker in our sample an individual-specific measure of the relative earnings losses after layoff by exploiting the fact that we have matched each treated worker to a similar control worker:

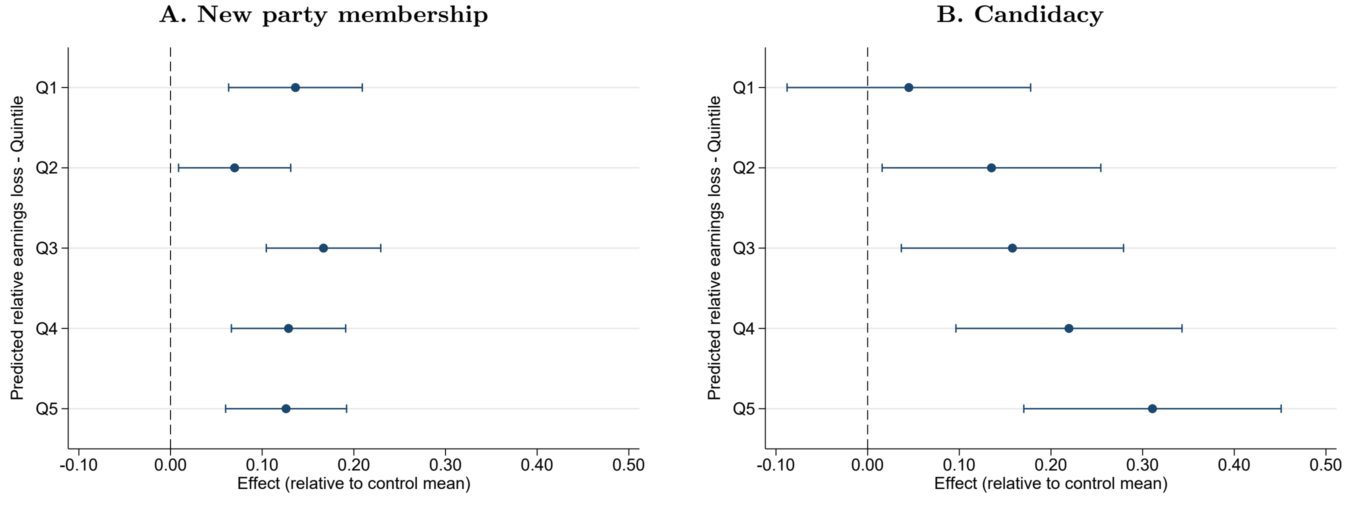
$$\Delta_{dd} y_{it} = \frac{\bar{y}_{i,post} - \bar{y}_{i,pre}}{\bar{y}_{i,pre}} - \frac{\bar{y}_{i',post} - \bar{y}_{i',pre}}{\bar{y}_{i',pre}}$$

where $\bar{y}_{i,post}$ is the average earnings in the 5 years after the layoff of treated worker i, or her/his matched control worker i'. $\bar{y}_{i,pre}$ is the average earnings in the 5 years before the layoff. This double difference can be thought of as the share of earnings that each worker loses in the medium run because of the layoff. We then regress $\Delta_{dd} y_{it}$ on a rich set of pre-layoff characteristics and compute the predicted relative earnings loss. The regression includes the following individual

[31] Hilger (2016) and Britto *et al.* (2022b) conduct similar exercises for the effect of parental job loss on children's outcomes.

[32] We focus on predicted, rather than actual earnings loss, since the latter would be directly affected by individual employment decisions after the layoff (for e.g., becoming a politician).

Figure 6: Heterogeneity by predicted earnings loss



Note: The figure reports DiD coefficients, estimated in model (2), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). We split the sample by quintiles of the treated individuals' predicted relative earnings loss due to the layoff. See the text for details on how earnings losses are predicted. Control workers are assigned to their matched treated workers in each sub-sample. The horizontal bars depict 95% confidence intervals based on standard errors clustered at the individual level. All coefficients and standard errors are scaled by the group-specific mean of the control group in the post-treatment period (average across 2012, 2016, and 2020).

covariates: gender, education, age decile, earnings decile, tenure decile, firm size decile, and layoff year. We also add municipality × industry dummies to capture variation in labor market conditions. Finally, we re-estimate the effects of job loss on new party memberships and candidacies in model (2), by splitting the sample into quintiles of the predicted earnings loss distribution.

Results are shown in Figure 6. The predicted income losses are clearly correlated with the estimated effects of job loss on councilor candidacies. In the lowest quintile of earnings losses, we find null effects, whereas in the highest quintile job loss is estimated to raise the probability of running for councilor by more than 30%. Again, we do not find heterogeneity in the effects on new party memberships. These results offer further support to the idea that the reduction in private-sector opportunity costs is an important mechanism in explaining the positive effect of job loss on pursuing a political office that comes with financial benefits.

6 Effects of Unemployment Insurance on Political Entry

The results presented in the previous section show that job loss significantly increases individuals' likelihood of becoming a party member and running for local office. In addition, the heterogeneity analyses suggest that the rise in political engagement can—at least partly—be explained by the implied reduction in private-sector opportunity costs following a layoff. The opportunity cost for displaced individuals may also vary depending on their eligibility for unemployment insurance (UI) benefits. In Brazil, workers who are dismissed without just cause can receive UI benefits for 3 to 5 months, with an average replacement rate of about 80%.[33] Benefit receipt mitigates the income losses from layoffs and may thereby reduce the need to seek salary from a political position. At the same time, during benefit receipt dismissed workers have lower incentives to find reemployment.[34] The negative employment effects of UI eligibility decrease labor earnings, which may offset the benefit payments, and increase individuals' time availability. The increased time resources, in turn, may enable individuals to engage with politics. In this section, we thus investigate if eligibility for UI benefits affects laid-off individuals' decision to enter politics. This analysis can help shed light on the mechanisms driving the effect of job loss on political entry.

6.1 Empirical strategy

We conduct a regression discontinuity (RD) analysis exploiting the fact that eligibility for UI benefits varies discontinuously with the time since the last layoff. Specifically, workers who are displaced without just cause from a formal job with at least six months tenure are eligible for UI benefits if at least 16 months have elapsed between the worker's current layoff date and the most recent layoff date used to claim UI benefits in the past. We compare workers displaced shortly before and shortly after the 16-month cutoff by estimating local linear regressions of the following form:

$$Y_i = \alpha + \beta D_i + \gamma_1 X_i + \gamma_2 D_i X_i + \epsilon_i, \tag{4}$$

where Y_i is an outcome of worker i. X_i denotes the time since the last layoff (re-centered around the 16-month cutoff), and D_i is a dummy for being eligible for UI benefits, i.e., $D_i = 1(X_i \geq 0)$. β is the coefficient of interest that identifies the effect of UI eligibility.[35] To ensure that eligible and non-eligible workers are comparable, our main results are based on a local linear model with a narrow bandwidth of 60 days around the cutoff.[36] In Appendix Tables C.1 and C.2 we show that our main results are robust to alternative polynomial orders and bandwidth choices, including the optimal bandwidth proposed by Calonico *et al.* (2014). In Appendix Figure C.4 we also perform permutation tests that compare our main estimates to the distribution of RD estimates obtained from placebo cutoff points.

[33]The length of UI payments depends on the length of employment in the 36 months before the layoff, and the replacement rate depends on the previous earnings (varying between 100% for individuals earning the minimum wage and 67% at the benefit cap, which is at 2.65 times the minimum wage).

[34]The negative effects of UI eligibility on formal employment have been widely documented, including for Brazil (Gerard *et al.*, 2020; Gerard and Gonzaga, 2021; Britto *et al.*, 2022a). Britto (2022) also shows that an extension of the UI eligibility period in Brazil decreased total employment when taking into account the positive effect on informal employment.

[35]As we lack information on UI take-up, our focus is on benefit eligibility. Consequently, our estimates should be interpreted as an intent-to-treat effect. According to Britto *et al.* (2022a), the probability of UI take-up jumps by about 60 percentage points at the 16-month eligibility cutoff.

[36]This choice follows the main specifications in Britto *et al.* (2022a), Britto *et al.* (2022b), and Amorim *et al.* (2023).

As in the previous section, the estimation sample consists of all workers aged 25 to 50 displaced from a full-time, open-ended, private-sector job of at least six months tenure. To increase statistical power, we extend our layoff period to cover the years 2006 to 2014.[37] This allows us to measure political outcomes over two election cycles after the layoff and labor market outcomes over up to 4 years after the layoff.[38] In addition, we exclude from the sample all workers who were displaced on the first or the last day of each month in their last layoff. Appendix Figure C.1 shows that the number of dismissals is higher at the very beginning and the very end of the month. Dropping these observations ensures that our RD cutoff does not coincide with spikes in the density of dismissals that occur every 30 days, i.e., that are not specific to the 16-month cutoff.

Appendix Figure C.2 shows that, when using the restricted sample, there is no evidence of discontinuity in the density of the running variable at the 16-month cutoff. We also confirm the continuity using the bias-robust test by Cattaneo *et al.* (2018, 2020). In addition, in Appendix Figure C.3 we show that individuals around the cutoff are similar in a rich set of pre-determined covariates, including gender, education, age, earnings, tenure, and sector shares. Overall, these results provide strong support for the continuity assumption of the RD approach.

6.2 Results

Results from model (4) are presented in Table 3, and the corresponding RD plots are shown in Figure 7. As a first step, we estimate the effects of UI eligibility on labor market outcomes (Panel A, columns (1) and (2) of Table 3). We find a significant negative effect on employment that amounts to 0.68 fewer months employed over the four years after the layoff. Moreover, labor earnings decrease by 2,192 BRL. The drop in labor earnings is likely compensated by the UI payments that eligible individuals receive.[39] Thus, the results suggest that UI eligibility leads to a rise in time availability following the job loss without strongly affecting the total available income.

In the second step, we examine the effects of UI benefits on individuals' entry into politics. In columns (4) and (6), we show the effects on the likelihood of newly registering to a political party and running for local council in the two cycles after the layoff. Columns (3) and (5) show that

[37] We start in 2006 because this allows us to observe past layoffs up to 24 months prior to the current layoff (our employment data begin in 2004) and to measure placebo political outcomes in the two cycles prior to the current layoff (our political data begin in the 2000 cycle). We stop in 2014 because of the numerous changes introduced to the UI system after that year.

[38] While we must restrict the sample to individuals with unique names for the party membership outcome, we do not impose this restriction for the candidacy outcome, which is less common and thus requires larger sample sizes.

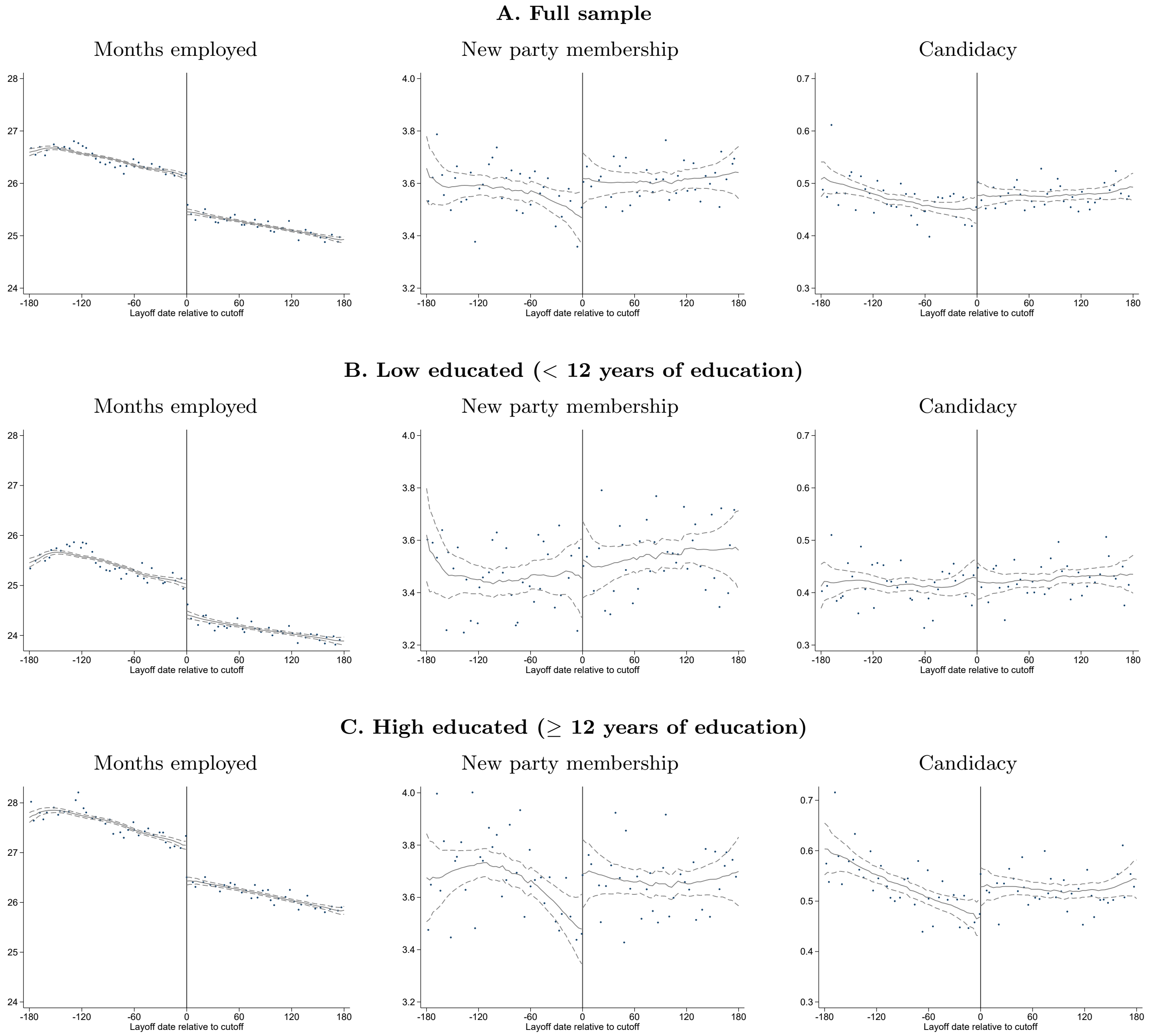
[39] Britto *et al.* (2022a) find that individuals barely meeting the 16-months cutoff receive additional UI benefits of about 2,100 BRL.

both outcomes were balanced in the two cycles before the layoff, providing additional support for the RD design. After the layoff, we find that UI eligibility increases the likelihood of becoming a party member by 0.15 percentage points (4.3% relative to the control mean). For candidacies, we find an increase of 0.03 percentage points (6.2% relative to the control mean). However, the estimate is not statistically significant due to the limited power that we have for the candidacy outcome.[40]

Given that earlier we have documented differential effects of job loss on political entry by individuals' education, Panels B and C of Table 3 also show the effects of UI benefits separately for low-educated workers ($<$ 12 years of schooling) and high-educated workers ($\geq$ 12 years of schooling). We find positive effects on party memberships and candidacies among the highly educated but no significant impact among the low educated. The effects on workers with high levels of education are statistically significant and meaningful in size, indicating a 6.2% increase in the likelihood of becoming a party member and a 13.0% increase in the likelihood of running for office (both relative to the control mean). These patterns are consistent with the larger effects of layoffs on individuals with higher levels of education that we presented earlier. Eligibility for UI benefits reduces individuals' reemployment incentives, thereby increasing the time resources that they have available to engage in politics. As argued before, this shock to individuals' opportunity costs is likely to be more relevant for highly educated workers, which may explain the larger effects of both job loss and UI eligibility on their decision to enter politics.

[40] As shown in Appendix Table C.2, the effect is significant for some alternative bandwidth and polynomial order choices.

Figure 7: Effect of UI eligibility



Note: The figure plots the average outcomes of workers around the 16-month cutoff date for eligibility for unemployment benefits, for different groups of workers. Outcomes are the number of months employed in the four years after the layoff, the probability of newly registering as a party member in the two election cycles after the layoff (in %), and the probability of running for local councilor in the two election cycles after the layoff (in %). Panel A shows results among all workers in the sample, and Panels B and C distinguish workers with less than vs. at least 12 years of education. Dots show averages for 5-day bins. The lines represent a local linear polynomial smoothing using a 60-day bandwidth, together with 95% confidence intervals.

Table 3: Effects of UI eligibility

	(1)	(2)	(3)	(4)	(5)	(6)
Outcome:	Months employed	Labor earnings	New party membership before layoff	after layoff	Candidacy before layoff	after layoff
[A] Full sample						
β	-0.679***	-2,192.5***	-0.002	0.150**	-0.011	0.028
	(0.041)	(149.9)	(0.080)	(0.072)	(0.013)	(0.019)
Control mean	26.138	56,281.8	4.330	3.467	0.236	0.449
Relative effect	-2.6%	-3.9%	-0.1%	4.3%	-4.6%	6.2%
Observations	2,048,674	2,048,674	1,047,683	1,047,683	2,048,674	2,048,674
[B] Low educated (< 12 years of education)						
β	-0.617***	-1,822.6***	0.071	0.066	-0.015	-0.006
	(0.057)	(166.1)	(0.121)	(0.109)	(0.019)	(0.026)
Control mean	25.029	47,526.6	4.264	3.454	0.221	0.428
Relative effect	-2.5%	-3.8%	1.7%	1.9%	-6.9%	-1.4%
Observations	986,755	986,755	454,665	454,665	986,755	986,755
[C] High educated ($\geq$ 12 years of education)						
β	-0.708***	-2,311.2***	-0.059	0.216**	-0.007	0.061**
	(0.059)	(240.6)	(0.106)	(0.097)	(0.019)	(0.027)
Control mean	27.143	64,213.1	4.381	3.476	0.250	0.468
Relative effect	-2.6%	-3.6%	-1.3%	6.2%	-2.7%	13.0%
Observations	1,061,919	1,061,919	593,018	593,018	1,061,919	1,061,919

Note: The table presents results from model (4) for the effect of unemployment benefit eligibility on the number of months employed and the total formal labor earnings in the four years after the layoff (columns (1) and (2)), on the likelihood of newly registering as a party member in the two election cycles before and after the layoff (columns (3) and (4)), and on the likelihood of running for local councilor in the two election cycles before and after the layoff (columns (5) and (6)). The sample includes all workers displaced within a bandwidth of 60 days around the cutoff date that determines benefit eligibility, which is 16 months since the last layoff. Standard errors clustered at the individual level are in parentheses. The table also reports the control mean of the outcome at the cutoff and the effect sizes scaled by the control mean. All coefficients, standard errors, and control means in columns (3) to (6) have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

7 Discussion and Implications

Discussion of mechanisms. Our main findings show that job loss increases individuals' likelihood of becoming a party member or a local political candidate. We now discuss how our results can help shed light on potential mechanisms that explain individuals' increased political engagement after experiencing a layoff. First, displaced individuals may be successfully mobilized through a change in party support. On the one hand, job loss could induce further political engagement in case workers perceive the incumbent government as the one responsible for their economic conditions. Retrospective voting has been widely documented to rationalize the behavior of voters following economic shocks (e.g., Tilley *et al.*, 2018; Ahlquist *et al.*, 2020). On the other hand, individuals may be more likely to align with parties whose political programs directly address workers' economic interests, i.e., via redistributive policies or alignment with labor unions. Even in the absence of economic platforms, extreme parties may be successful in exploiting individuals' resentment with the establishment following economic downturns (e.g., Algan *et al.*, 2017; Autor *et al.*, 2020). Our results show that job loss increases party memberships and candidacies across the party spectrum, including also incumbent, centrist and non-union affiliated parties. This suggests that mobilization and politicization may not be the only mechanisms explaining the increase in participation.

An alternative explanation focuses on individuals' opportunity costs of engaging with politics. In economic models of political selection, an individual runs for office if her benefits from winning office B (weighted by the expected probability of winning P) are larger than the costs of running C: $PB - C \geq 0$ (Black, 1972; Caselli and Morelli, 2004). A layoff reduces the costs of getting involved with politics by reducing the private-sector outside options and increasing time availability. Our results are supportive of this mechanism. First, we show larger layoff effects among highly educated individuals, who arguably experience larger reductions in opportunity costs following layoffs. Second, when investigating alternative outcomes, we find that following job loss, individuals are more likely to work for a political campaign, illustrating the trade-off between working in the private and in the public office. Third, with respect to the financial incentives associated with individuals' decision to run for office, we find that both the severity of the income shock and the potential wages in political office correlate strongly with individual responses. Specifically, we find substantially larger positive effects on running for office for individuals who live in places with higher councilor wages and who face larger predicted income losses. Fourth, unemployment insurance reduces work incentives and thus increases the time individuals can spend on politics. We find that eligibility for UI benefits has a positive effect on political entry, especially among the highly educated. As a whole, our results suggest that individuals' decision to enter political life after job loss is related to economic motives rather than purely ideological ones and can likely be rationalized by the reduction in opportunity costs.[41]

Implications of job loss for political selection. An important question that follows is how these new entrants affect the political system. The increased number of party members and candidates means that parties have a larger pool from which they can select competent politicians for leadership roles. In addition, layoffs may drive individuals with different motivations and different skill levels into politics, and thereby change the average quality of politicians. In particular, if layoffs are a larger shock to the opportunity costs of more competent individuals, they may improve the average competence of the candidate pool.

To shed light on the implications of layoffs for the selection of political entrants, we compare characteristics of treated and control individuals who become party members or candidates after 2012, i.e., after our layoff period. The idea of this simple comparison is as follows. Due to our matching procedure, in the full sample, treated and control individuals have identical pre-layoff characteristics (see Table 1). Any difference that we see in the pre-layoff characteristics of treated and control individuals who enter politics after the layoff period should then only be driven by the effects of job loss varying across individuals. The comparison can thus tell us how individuals who enter political life due to a layoff are different from all other individuals entering political life in our sample.

We study an individual's qualification for a political job using three different measures. First, we consider years of education as a standard measure of competence. Second, we examine individuals' monthly earnings before the layoff, which, following Dal Bó *et al.* (2013), can be treated as a measure of skill if the private sector rewards highly skilled workers with higher earnings. Third, we follow Besley *et al.* (2017) and Dal Bó *et al.* (2017) and consider residuals from a Mincer earnings regression. Specifically, we regress log monthly earnings on dummies for gender, education categories, age deciles, tenure deciles, firm size deciles, and municipality $\times$ 2-digit industry categories. The standardized residual from this regression measures the earnings premium that a worker receives relative to similar (on observable characteristics) workers and thus likely reflects the worker's ability.

[41]Note that the benefits from holding office B may include not only monetary returns but also policy payoffs. In the original citizen-candidate models (Osborne and Slivinski, 1996; Besley and Coate, 1997), political candidates represent particular policies and care about winning *per se*. As Hall (2019) and Thomsen (2017) argue, the policy returns may be higher than for ideologically extreme candidates. In this case, a reduction in the costs of running C, such as a drop in private-sector opportunity costs, may lead to an increase in the share of moderate candidates entering politics. This mechanism may counteract potential mobilization dynamics that favor extreme ideologies and explain why we see few ideological differences in the effects of job loss on party memberships and candidacies.

Table 4: Characteristics of new party members and candidates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	New party members				Candidates			
	Treated	Control	Std Diff	P-Val	Treated	Control	Std Diff	P-Val
Years of education	10.79	10.66	-0.05	0.00	10.79	10.67	-0.05	0.01
Earnings	2169.33	1981.66	-0.10	0.00	2145.66	1956.10	-0.11	0.00
Earnings residual	0.03	-0.02	-0.05	0.00	0.04	0.00	-0.04	0.03
Previous candidacy	3.23	2.53	-0.04	0.00	12.93	11.09	-0.06	0.00
Observations	46,828	39,901			7,914	6,477		

Note: The table reports the average pre-layoff characteristics of treated and control workers who newly register as a party member and run for local councilor in any post-layoff cycle (2012, 2016, or 2020). "Earnings" are monthly formal labor earnings in BRL (CPI 2018) before the layoff. "Earnings residual" is the standardized residual from a regression of log monthly earnings on dummies for gender, education categories, age deciles, tenure deciles, firm size deciles, and municipality × 2-digit industry categories. "Previous candidacy" is a dummy for whether an individual ran for municipal councilor in any pre-layoff cycle (2000, 2004, or 2008). Columns (3) and (6) report standardized differences in characteristics between treated and control workers. Columns (4) and (8) report p-values for the t-tests for equality of means.

Table 4 shows the results. We observe that among individuals who become a party member or councilor candidate after the layoff, those from the treatment group exhibit higher qualifications than those from the control group. This is consistently found across all three competence measures. For example, individuals in the treatment group who run for office have earned about 9.5% more than control individuals running for office. We also consider an individual's political experience, which we measure by whether the individual has already run for councilor in any previous election cycle (2000, 2004, or 2008). We find that among those running for office after the layoff, the share of individuals who have already run for office before the layoff is 1.9 percentage points (17%) higher in the treated compared to the control group. In sum, these results suggest that job loss increases political entry more strongly among more competent and politically experienced individuals.

Labor market shocks and political selection at the aggregate level. Finally, we provide descriptive evidence on the aggregate relationship between economic conditions and the selection of politicians. It is not clear *a priori* whether the individual-level results translate into changes at the aggregate level. For instance, if new entrants after a layoff simply crowd-out politicians with similar levels of education, the share of highly educated individuals would remain unchanged. To gain a better understanding of these aggregate dynamics, we gather municipality-level data on unemployment rates from the 2000 and 2010 Brazilian censuses and on the average years of education of council candidates in the 2000 and 2012 municipal elections. Figure 8 shows a positive correlation between changes in local unemployment rates and changes in candidates' education level.[42] This result is in line with our findings at the individual level. Taken together, the evidence suggests that economic downturns might have positive effects for the selection of local politicians.

8 Conclusion

The quality of democratic institutions largely depends on individuals' willingness to engage in politics. However, while most papers investigate the consequences of institutional constraints for political selection, there is a lack of evidence on the causal effects of personal experiences in affecting individuals' decision to actively take part in politics. This paper fills this research gap by investigating if experiencing a job loss increases the probability of entering politics, either by becoming a member of political parties or by running for office in local elections.

Answering this research question is not trivial, given the fact that job losses are highly endogenous and researchers often face a rare-event problem when studying entry into politics. We circumvent these challenges by combining detailed information on the universe of party members and local candidates with matched employer-employee data covering all formally employed workers in Brazil. Our identification strategy relies on mass layoffs as a source of plausibly exogenous individual job losses. The richness of our data allows us to exactly match treated and control workers along several characteristics and then compare their political entry over time using DiD estimations.

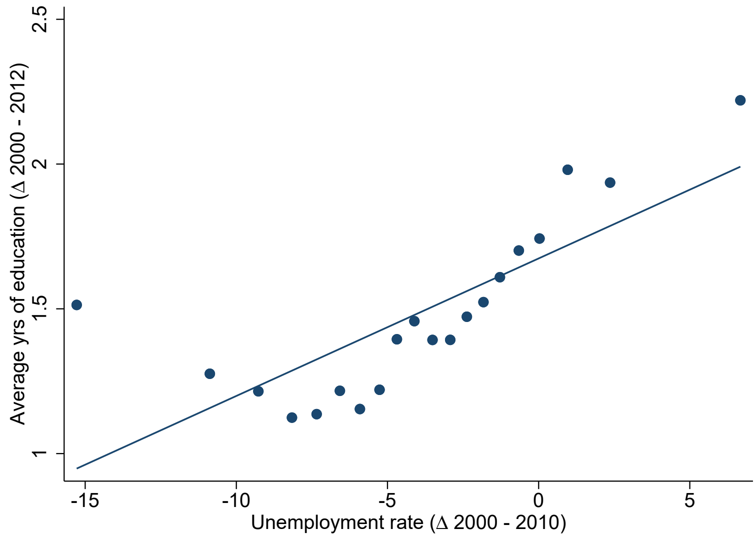
We find that job loss substantially increases the likelihood that an individual becomes member of a political party and runs for office in local elections. These positive effects are highly persistent, remaining statistically significant three election cycles (about 12 years) after the layoff. Moreover, they are prevalent across a wide range of individual and party characteristics, and they extend to alternative forms of political participation, including running for mayor, donating to political campaigns, and working on a political campaign.

Several of our results suggest that the increase in political entry can be explained by an opportunity-cost mechanism, in which job loss reduces individuals' alternative options in the private sector and increases their time availability. Most importantly, we observe larger increases in candidacies among individuals with higher financial incentives – measured either as wages in the public office or the predicted income loss due to the layoff. In addition, we find that eligibility for UI benefits, by increasing the time outside of employment, positively affects political entry.

In terms of the implications of layoffs for the selection of political entrants, we document that

[42] The corresponding regression results, reported in Appendix Table D.1, show that the relationship is highly significant and robust to including state fixed effects and weighting municipalities by their population size.

Figure 8: Municipality-level unemployment rate and candidate quality



Note: The figure depicts a binned scatterplot (with 20 bins) for the municipality-level relation between councilor candidates' average years of education (change between elections in 2000 and 2012) and unemployment rates (change between 2000 and 2010). N = 5,476 municipalities. The corresponding regression results are reported in Appendix Table D.1.

treated individuals who become party members or candidacies after their layoff had, on average, higher wages prior to layoff, higher education levels, and more political experience, as compared to control candidates. This positive selection is also evident at the aggregate level, where we find a positive correlation between municipal changes in unemployment rates and changes in candidates' average years of education. Taken together, our paper provides novel evidence that layoffs increase the likelihood that highly capable individuals enter politics.

We welcome future research that studies the effects of job loss not only on political selection but also on the behavior of politicians after being elected to office. How do the individual economic circumstances of politicians affect their performance in office, including their legislative productivity, the extraction of personal rents, and the effective provision of local public goods? Moreover, how does job loss affect the policy focus set by politicians and the representation of voters' preferences in the policy-making process? By answering these questions, we could shed more light on the broader consequences of economic shocks for the political system.

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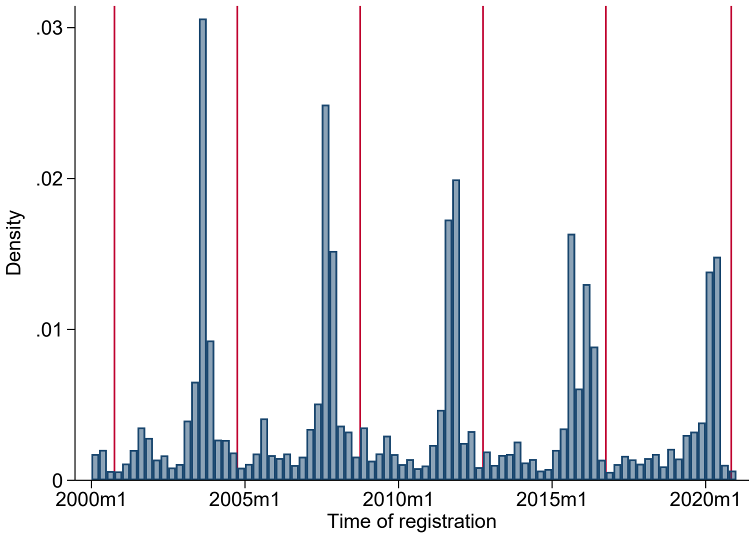
Appendix

A Appendix to Sections 2 to 4

Figure A.1: Distribution of new party memberships over time



Note: The figure reports the density of new party memberships over time. We include all affiliations in Brazil between January 2000 and December 2020. Red lines indicate local election months.

Table A.1: Campaign donations of councilor candidates

	2004	2008	2012	2016	2020	All
Total donations						
Mean	3,715.55	4,861.25	5,649.81	3,427.54	3,467.26	4,200.86
P10	202.69	211.58	212.38	230.70	214.83	213.61
P25	520.92	617.11	613.78	534.02	519.33	552.17
P50	1,403.30	1,833.70	2,123.80	1,468.55	1,327.48	1,586.85
P75	3,506.12	4,752.17	5,932.49	3,741.29	3,276.06	4,169.63
P90	7,994.55	10,655.85	13,199.48	7,435.77	7,450.96	9,300.73
Donations from own resources						
Share>0	0.61	0.68	0.55	0.56	0.46	0.56
Mean	1,632.70	1,718.96	1,704.98	1,225.80	624.07	1,313.14
P10	0.00	0.00	0.00	0.00	0.00	0.00
P25	0.00	0.00	0.00	0.00	0.00	0.00
P50	318.93	439.03	196.81	213.61	0.00	168.13
P75	1,811.53	1,819.59	1,792.28	1,281.64	681.86	1,275.73
P90	4,297.07	4,557.80	4,955.54	3,226.50	1,494.49	3,539.67
N	202,534	284,031	359,753	395,179	415,691	1,657,188

Note: The table shows summary statistics for campaign donations received by candidates for local council in the elections 2000 to 2020. We report total donation values and the value of donations that come from candidates' own resources. All values are in BRL (CPI 2018) and are winsorized at the 1% level. The data is obtained from the *Tribunal Superior Eleitoral.*

Table A.2: Summary statistics: workers with and without unique names

	(1) Unique	(2) Non-unique	(3) Std Diff
Socio-demographic characteristics			
Male	65.78	73.86	0.18
White	57.32	52.03	-0.11
Age	33.84	34.52	0.10
Years of education	10.72	9.97	-0.25
Job and firm characteristics			
Earnings (per month, BRL, CPI 2018)	2158.15	1977.51	-0.09
Tenure (months)	32.62	32.95	0.01
Manager	2.70	1.77	-0.06
Firm size	610.80	619.27	0.01
Political outcomes before layoff (2005-2008, %)			
Candidate	0.11	0.12	0.00
Left party	0.04	0.04	0.00
Center party	0.03	0.03	-0.00
Right party	0.04	0.05	0.00
Incumbent party	0.01	0.01	-0.00
Not incumbent party	0.10	0.11	0.00
Union-affiliated party	0.03	0.03	0.00
Not union-affiliated party	0.04	0.04	-0.00
Observations	944,214	937,353	

Note: The table reports the average characteristics of displaced workers in our estimation sample who have and who do not have a unique name within Brazil, and the standardized difference between the two groups.

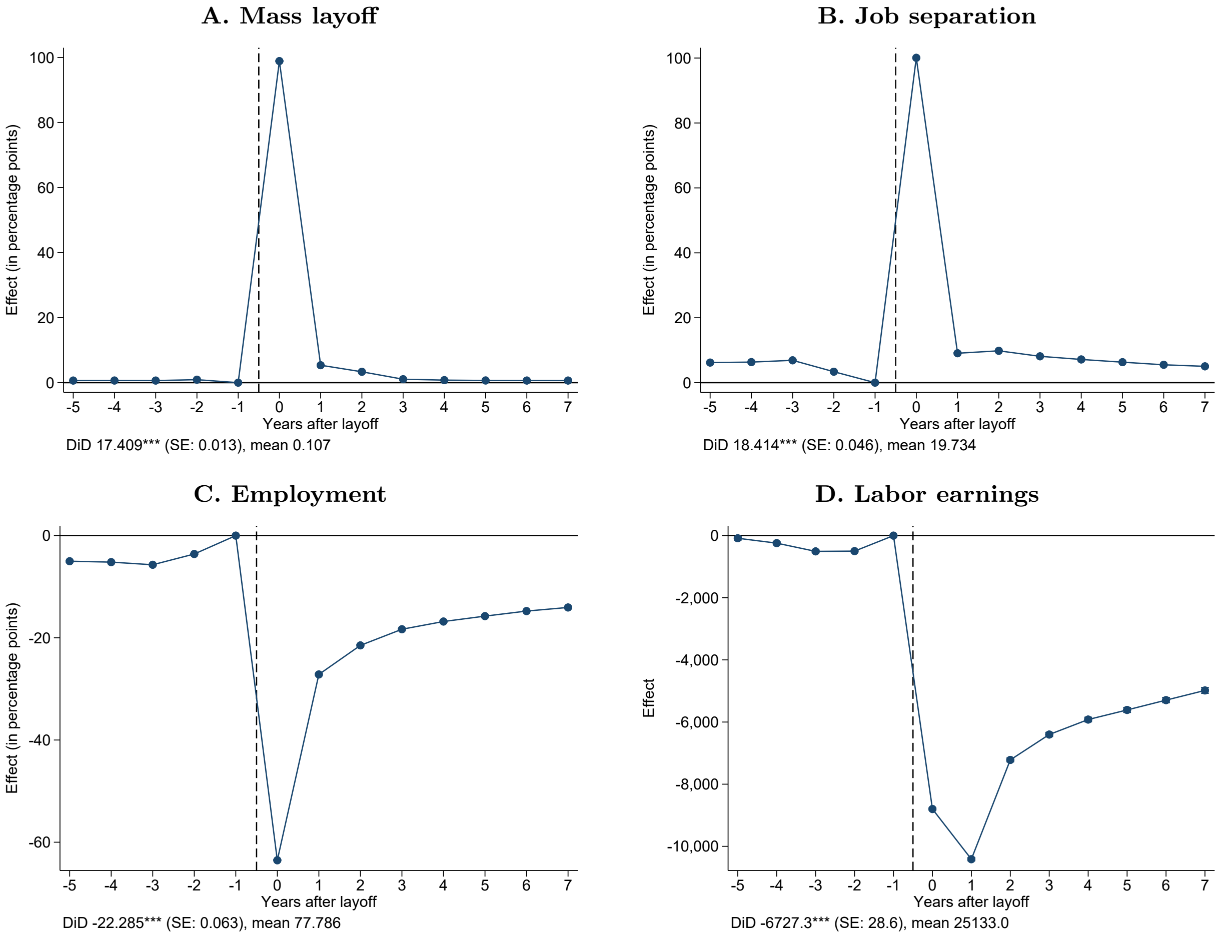
Table A.3: Summary statistics: party characteristics

	(1) Treated	(2) Control	(3) Std Diff
Party membership (2008, %)	7.47	7.30	-0.01
Left party	2.66	2.52	-0.01
Center party	2.40	2.40	-0.00
Right party	2.41	2.38	-0.00
Incumbent party	0.78	0.82	0.00
Not incumbent party	6.69	6.49	-0.01
Union-affiliated party	2.81	2.74	-0.00
Not union-affiliated party	2.25	2.18	-0.00
New party membership (2005-2008, %)	1.99	1.89	
Left party	0.73	0.67	-0.01
Center party	0.57	0.58	0.00
Right party	0.68	0.65	-0.00
Incumbent party	0.23	0.24	0.00
Not incumbent party	1.76	1.65	-0.01
Union-affiliated party	0.71	0.67	-0.00
Not union-affiliated party	0.59	0.57	-0.00
Candidate (2005-2008, %)	0.11	0.10	-0.00
Left party	0.04	0.04	-0.00
Center party	0.03	0.03	-0.00
Right party	0.04	0.03	-0.01
Incumbent party	0.01	0.01	-0.00
Not incumbent party	0.10	0.09	-0.01
Union-affiliated party	0.03	0.03	0.00
Not union-affiliated party	0.04	0.03	-0.00
Observations	944,214	944,214	

Note: The table reports the average pre-layoff political outcomes of treated workers who are displaced from a mass layoff firm and matched control workers who are not displaced in the same calendar year, and the standardized difference between the two groups. The sample only includes workers with a unique name within the country. The ideological classification into 'left', 'center' and 'right' parties follows Colonnelli *et al.* (2022) and is shown in Appendix Table B.3. The 'incumbent' party is the party of the elected mayor in the individual's municipality in 2008, and 'not incumbent' refers to all other parties. The 'union-affiliated' parties are PT, PDT, PSB, PCB, PSD, and MDB, and 'not union-affiliated' refer to all other left or center parties.

B Appendix to Section 5

Figure B.1: Labor market outcomes



Note: The figure reports event-study coefficients δ_k, estimated in model (3), for the effect of job loss on the likelihood of being displaced in a mass layoff (Panel A), the likelihood of having any job separation (Panel B), the likelihood of being full-time employed at the end of the year (Panel C), and yearly labor earnings (Panel D). $N = 24,549,564$ individual-year observations. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from a static version of model (3), its standard error, and the mean of the control group in the post-treatment period are reported.

Table B.1: Effect of job loss: alternative mass layoff definitions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Layoff share:	$\geq$ 30%	$\geq$ 50%	$\geq$ 75%	closure	$\geq$ 30%	$\geq$ 30%	$\geq$ 30%	Excl.
Firm size:	$\geq$ 30	$\geq$ 30	$\geq$ 30	$\geq$ 30	$\geq$ 50	$\geq$ 100	$\geq$ 250	volatile
[A] Outcome: new party membership								
δ_{DiD}	0.175***	0.143***	0.135***	0.201***	0.171***	0.141***	0.147***	0.129***
	(0.018)	(0.029)	(0.046)	(0.072)	(0.022)	(0.027)	(0.039)	(0.026)
Control mean	1.407	1.444	1.454	1.452	1.413	1.417	1.423	1.382
Relative effect	12.4%	9.9%	9.3%	13.9%	12.1%	10.0%	10.3%	9.3%
Observations	11,317,632	5,681,304	2,850,012	1,506,516	9,085,176	6,652,320	4,298,994	5,446,026
[B] Outcome: candidacy								
δ_{DiD}	0.041***	0.050***	0.055***	0.053**	0.039***	0.041***	0.044***	0.034***
	(0.006)	(0.009)	(0.015)	(0.023)	(0.007)	(0.009)	(0.013)	(0.008)
Control mean	0.229	0.237	0.233	0.233	0.227	0.225	0.226	0.223
Relative effect	18.1%	21.2%	23.5%	22.8%	17.4%	18.2%	19.3%	15.5%
Observations	11,317,632	5,681,304	2,850,012	1,506,516	9,085,176	6,652,320	4,298,994	5,446,026

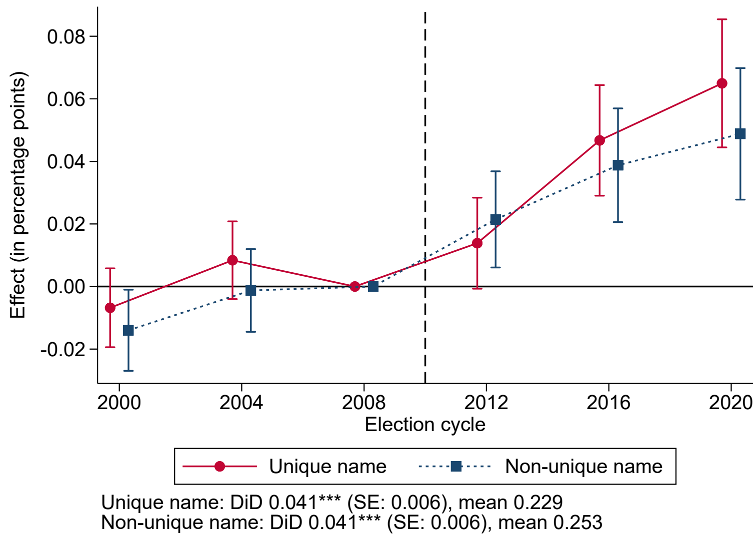
Note: The table presents robustness checks for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B), as estimated from model (2) using different definitions of mass layoffs. Columns (1) to (3) progressively increase the minimum layoff share, and column (4) restricts the treated group to individuals laid off in a plant closure (defined as a minimum layoff share of 95%). Columns (5) to (7) increase the minimum size of firms used to define mass layoffs. Column (8) excludes establishments with an employment growth of more than 30% in at least one of the two last years before the mass layoff (using the minimum layoff share and firm size as in our baseline definition in column (1)). Standard errors clustered at the individual level are in parentheses. All coefficients, standard errors, and control means have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B.2: Effect of job loss: alternative specifications

	(1)	(2)	(3)	(4)
Fixed effects:	Cycle	Municipality × cycle	2dgt industry × cycle	Municipality × 2dgt industry × cycle
[A] Outcome: new party membership				
δ_{DiD}	0.151***	0.150***	0.178***	0.175***
	(0.015)	(0.016)	(0.016)	(0.018)
Control mean	1.409	1.408	1.409	1.407
Relative effect	10.7%	10.7%	12.6%	12.4%
Observations	11,330,568	11,329,716	11,330,568	11,317,632
[B] Outcome: candidacy				
δ_{DiD}	0.033***	0.041***	0.041***	0.041***
	(0.005)	(0.005)	(0.006)	(0.006)
Control mean	0.229	0.229	0.229	0.229
Relative effect	14.5%	18.1%	18.1%	18.1%
Observations	11,330,568	11,329,716	11,330,568	11,317,632

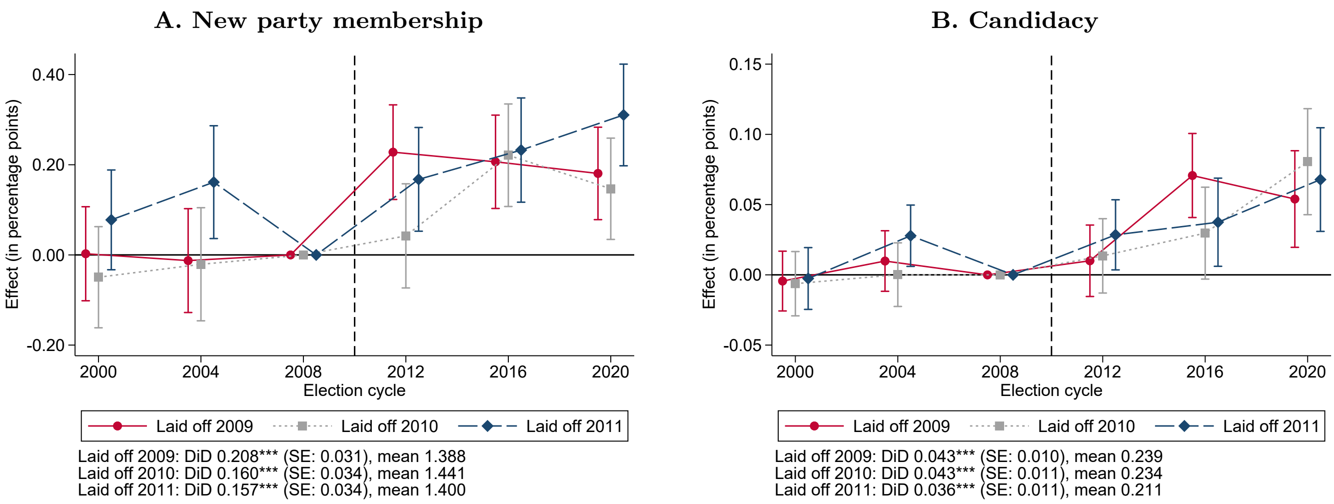
Note: The table presents robustness checks for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B), as estimated from model (2) using different definitions of mass layoffs. Columns (1) to (3) progressively increase the minimum layoff share, and column (4) restricts the treated group to individuals laid off in a plant closure. Columns (5) to (7) increase the minimum size of firms used to define mass layoffs. Column (8) excludes establishments with an employment growth of more than 30% in at least one of the two last years before the mass layoff (using the minimum layoff share and firm size as in our baseline definition in column (1)). Standard errors clustered at the individual level are in parentheses. All coefficients, standard errors, and control means have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure B.2: Effect of job loss on running for local councilor: unique vs. non-unique names



Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of running for local councilor, separately for individuals with unique and non-unique names in Brazil. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below the graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

Figure B.3: Heterogeneity by layoff year



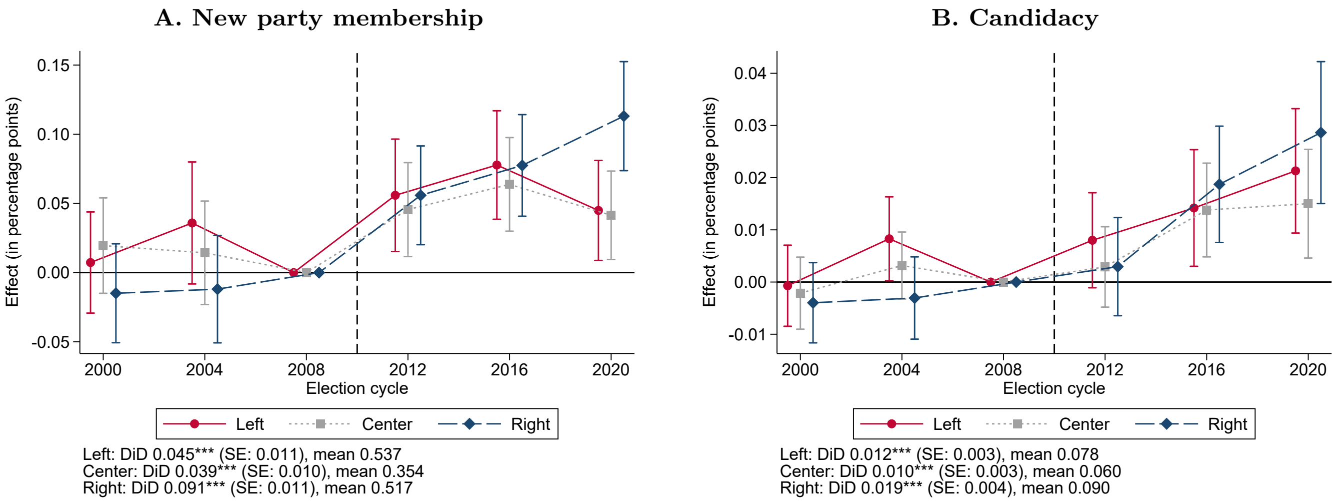
Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B), separately for individuals laid off in 2009, 2010, and 2011 (and their matched control individuals). The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

Table B.3: Ideological classification of political parties

Left	Center	Right
PT	MDB	PP
PDT	PSDB	DEM
PSB	PTB	PL
CIDADANIA	AVANTE	PSC
PCDOB	PSD	REPUBLICANOS
PV		PSL
PMN		PTC
PSOL		DC
PCB		PODE
PSTU		PRTB
PCO		PRP
SD		PHS
PROS		PATRIOTA
PPL		NOVO
PMB		
REDE		
UP		

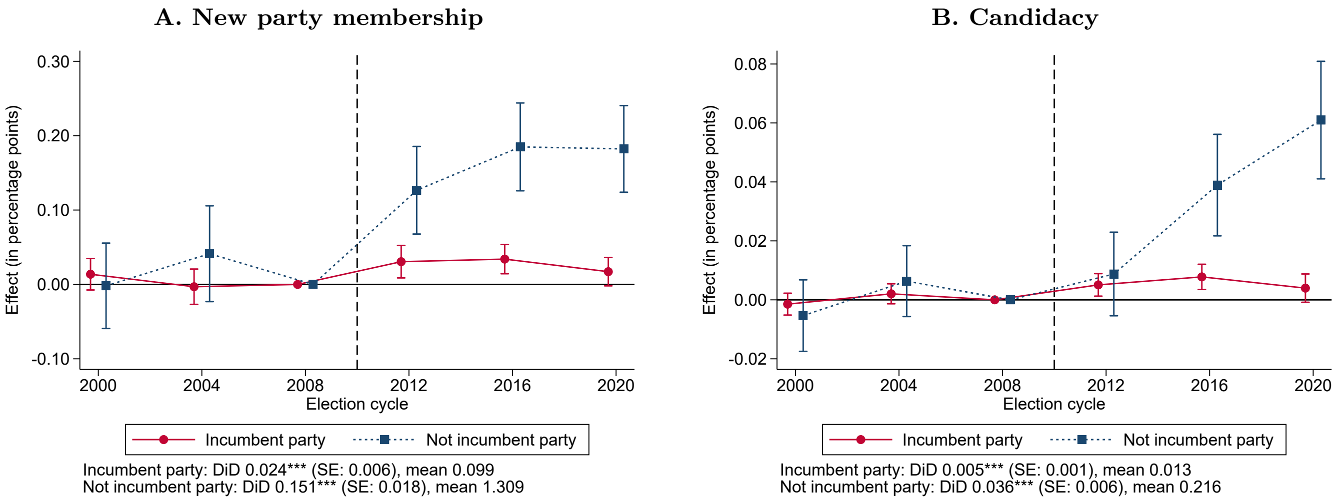
Note: The table shows the ideological classification of political parties, based on Colonnelli *et al.* (2022).

Figure B.4: Heterogeneity by party ideology



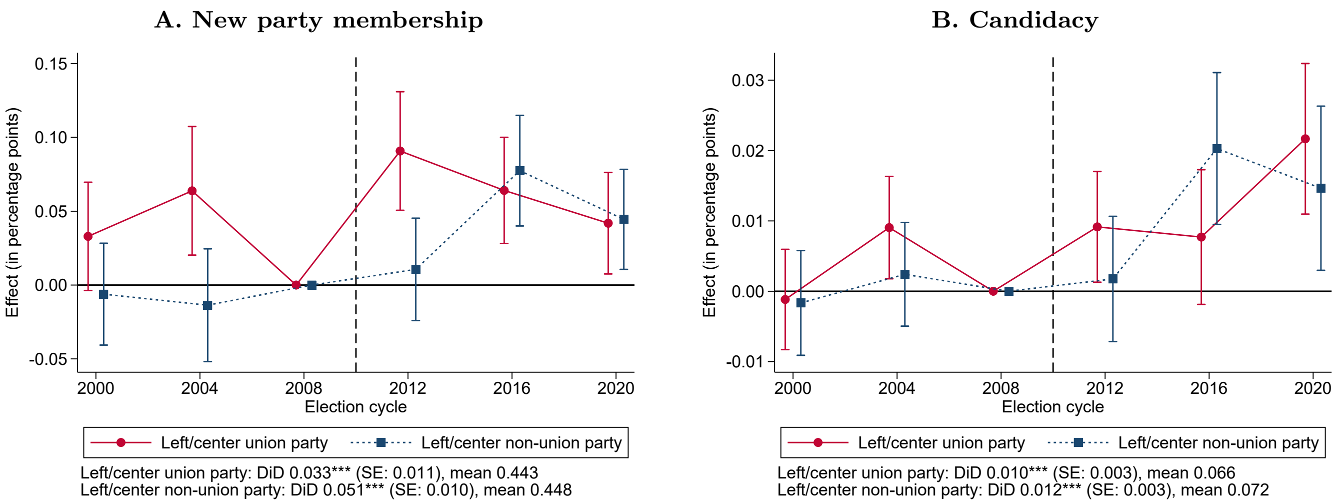
Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B) in left, center, and right parties. The ideological classification of parties follows Colonnelli *et al.* (2022) and is shown Table B.3. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

Figure B.5: Heterogeneity by party local incumbency



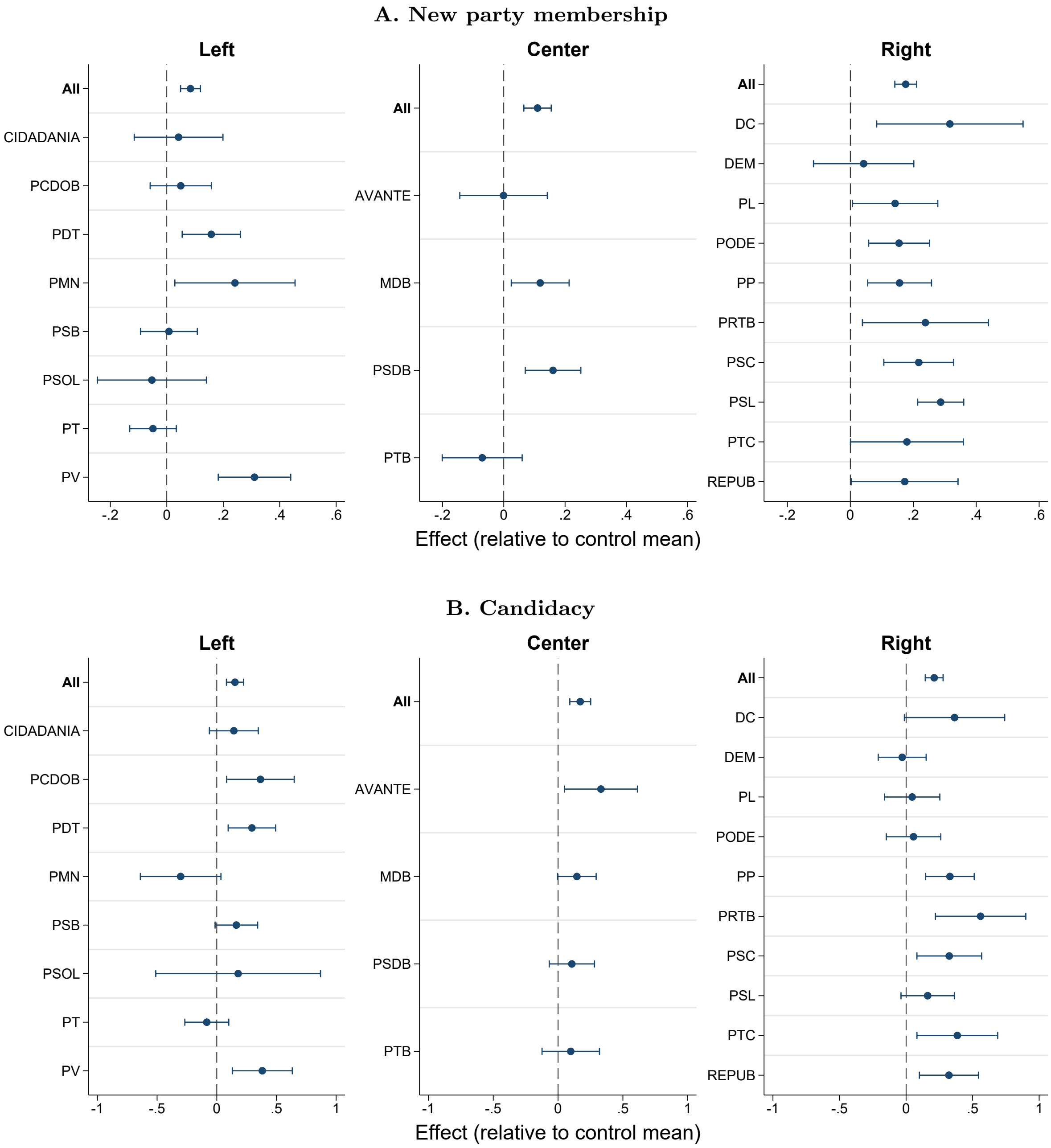
Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). Results are shown separately for memberships and candidacies in the party of the elected mayor in 2008 and in other non-incumbent parties. The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

Figure B.6: Heterogeneity by party union affiliation



Note: The figure reports event-study coefficients δ_c, estimated in model (1), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B). Results are shown separately for memberships and candidacies in parties historically affiliated with labor unions (PT, PDT, PSB, PCB, PSD, and MDB) and all other left or center parties (defined in Table B.3). The vertical bars depict 95% confidence intervals based on standard errors clustered at the individual level. Below each graph, the DiD coefficient from model (2), its standard error, and the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020) are reported.

Figure B.7: Heterogeneity by party



Note: The figure reports DiD coefficients, estimated in model (2), for the effect of job loss on the likelihood of newly registering as a party member (Panel A) and running for local councilor (Panel B) in each of Brazil's major parties. The ideological classification of parties follows Colonnelli *et al.* (2022). The horizontal bars depict 90% confidence intervals based on standard errors clustered at the individual level. All coefficients and standard errors are scaled by the mean of the control group in the post-treatment period (average across 2012, 2016, and 2020).

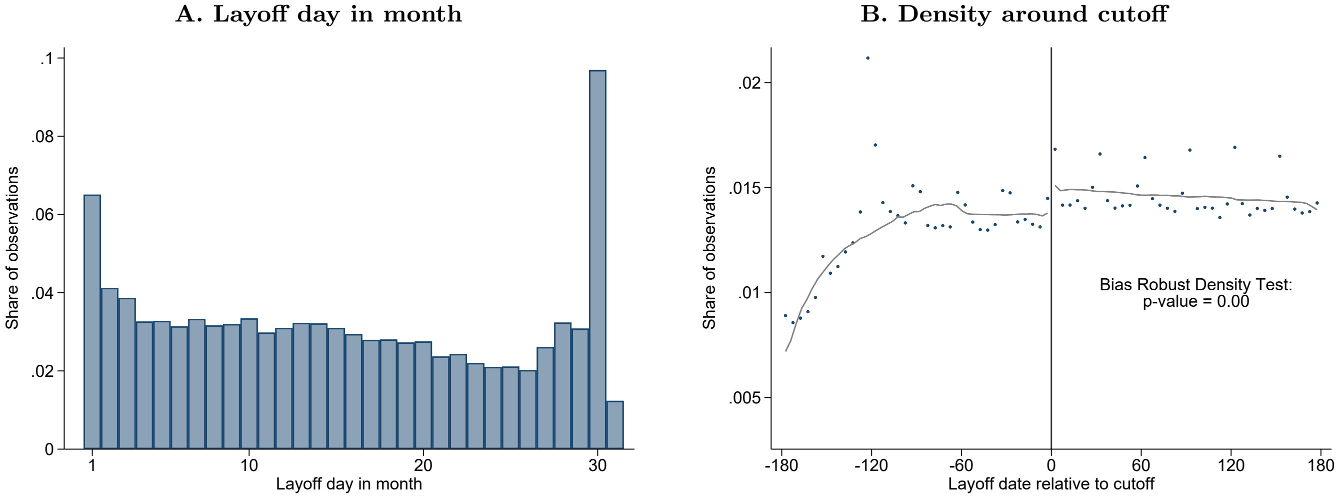
Table B.4: Councilor wage caps

Population	Maximum wage in % of state deputy salary	Maximum wage in BRL, 2004
1-10,000	20%	1927.1
10,001-50,000	30%	2890.6
50,001-100,000	40%	3854.2
100,001-300,000	50%	4817.7
300,001-500,000	60%	5781.2
above 500,000	75%	7226.6

Note: The table shows the maximum monthly wage that councilors can earn by municipality population size. Wage caps are shown in % of state deputy salaries, as defined by the Constitutional Amendment No 25. in 2000, and in absolute values, as estimated by Ferraz and Finan (2011) based on the 2004 federal deputy salary.

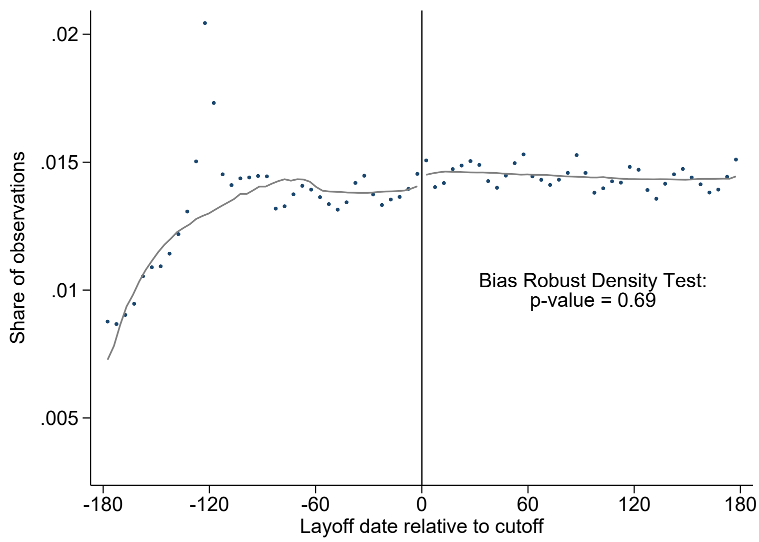
C Appendix to Section 6

Figure C.1: Layoff date cyclicality



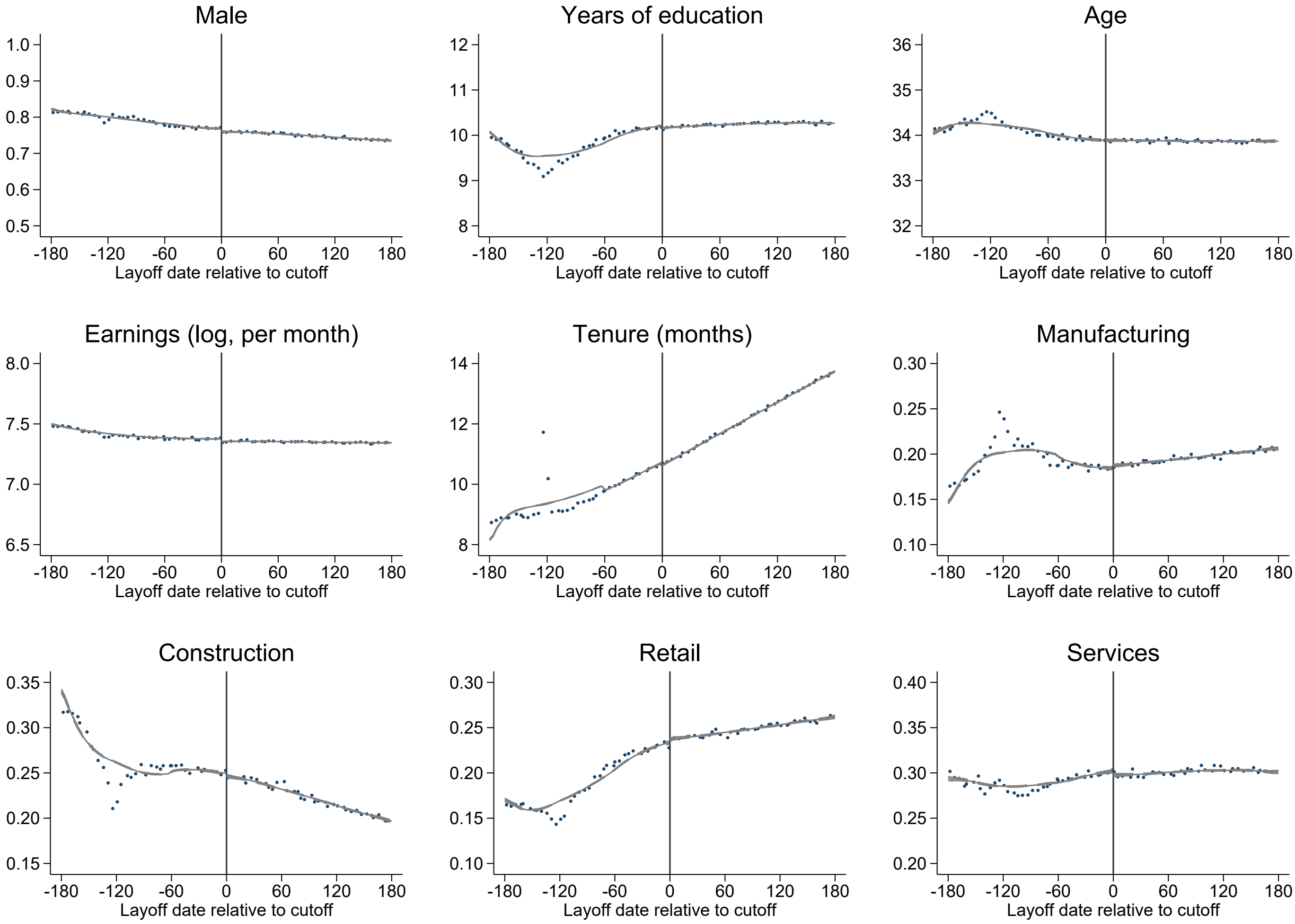
Note: Panel A reports a histogram for the distribution of layoff dates by calendar day in each month. Panel B reports the density of time between the current and the last layoff date around the 16-month cutoff date for eligibility for unemployment benefits, based on the initial sample that includes all layoff dates. The line represents a local linear polynomial smoothing using a 60-day bandwidth. Results for the bias-robust density test proposed by Cattaneo *et al.* (2018, 2020) are also shown.

Figure C.2: Layoff date density



Note: The figure reports the density of time between the current and the last layoff date around the 16-month cutoff date for eligibility for unemployment benefits, based on our estimation sample that excludes all workers who were displaced on the first or the last day of each month in their last layoff. The line represents a local linear polynomial smoothing using a 60-day bandwidth. Results for the bias-robust density test proposed by Cattaneo *et al.* (2018, 2020) are also shown.

Figure C.3: Balance of covariates near the UI eligibility cutoff



Note: The figure plots the average characteristics of workers around the 16-month cutoff date for eligibility for unemployment benefits. All characteristics are measured just before the layoff. Dots show averages for 5-day bins. The lines represent a local linear polynomial smoothing using a 60-day bandwidth, together with 95% confidence intervals.

Table C.1: Effect of UI eligibility on party membership: alternative specifications

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Bandwidth (days):	30	opt	30	60	90	opt	90	120	150	opt
Polynomial order:	0	0	1	1	1	1	2	2	2	2
[A] Full sample										
β	0.092*	0.097**	0.200**	0.150**	0.122**	0.162**	0.174*	0.153**	0.150**	0.173*
	(0.051)	(0.049)	(0.102)	(0.072)	(0.059)	(0.075)	(0.089)	(0.077)	(0.069)	(0.090)
Control mean	3.500	3.493	3.463	3.467	3.479	3.454	3.465	3.461	3.458	3.450
Relative effect	2.6%	2.8%	5.8%	4.3%	3.5%	4.7%	5.0%	4.4%	4.3%	5.0%
Observations	529,098	565,362	529,098	1,047,683	1,560,276	976,135	1,560,276	2,083,546	2,590,023	1,507,096
[B] Low educated (< 12 years of education)										
β	0.033	0.017	0.128	0.066	-0.031	0.091	0.100	0.008	0.016	0.112
	(0.077)	(0.074)	(0.153)	(0.109)	(0.089)	(0.112)	(0.133)	(0.115)	(0.104)	(0.136)
Control mean	3.479	3.473	3.417	3.454	3.487	3.449	3.471	3.483	3.469	3.432
Relative effect	0.9%	0.5%	3.8%	1.9%	-0.9%	2.6%	2.9%	0.2%	0.5%	3.3%
Observations	229,191	244,890	229,191	454,665	682,272	430,963	682,272	922,954	1,157,173	650,326
[C] High educated (≥ 12 years of education)										
β	0.137**	0.137**	0.255*	0.216**	0.244***	0.231**	0.230*	0.262**	0.259***	0.214*
	(0.068)	(0.068)	(0.136)	(0.097)	(0.080)	(0.094)	(0.119)	(0.104)	(0.093)	(0.121)
Control mean	3.517	3.517	3.498	3.476	3.467	3.459	3.462	3.446	3.443	3.460
Relative effect	3.9%	3.9%	7.3%	6.2%	7.0%	6.7%	6.6%	7.6%	7.5%	6.2%
Observations	299,907	299,907	299,907	593,018	878,004	641,120	878,004	1,160,592	1,432,850	848,505

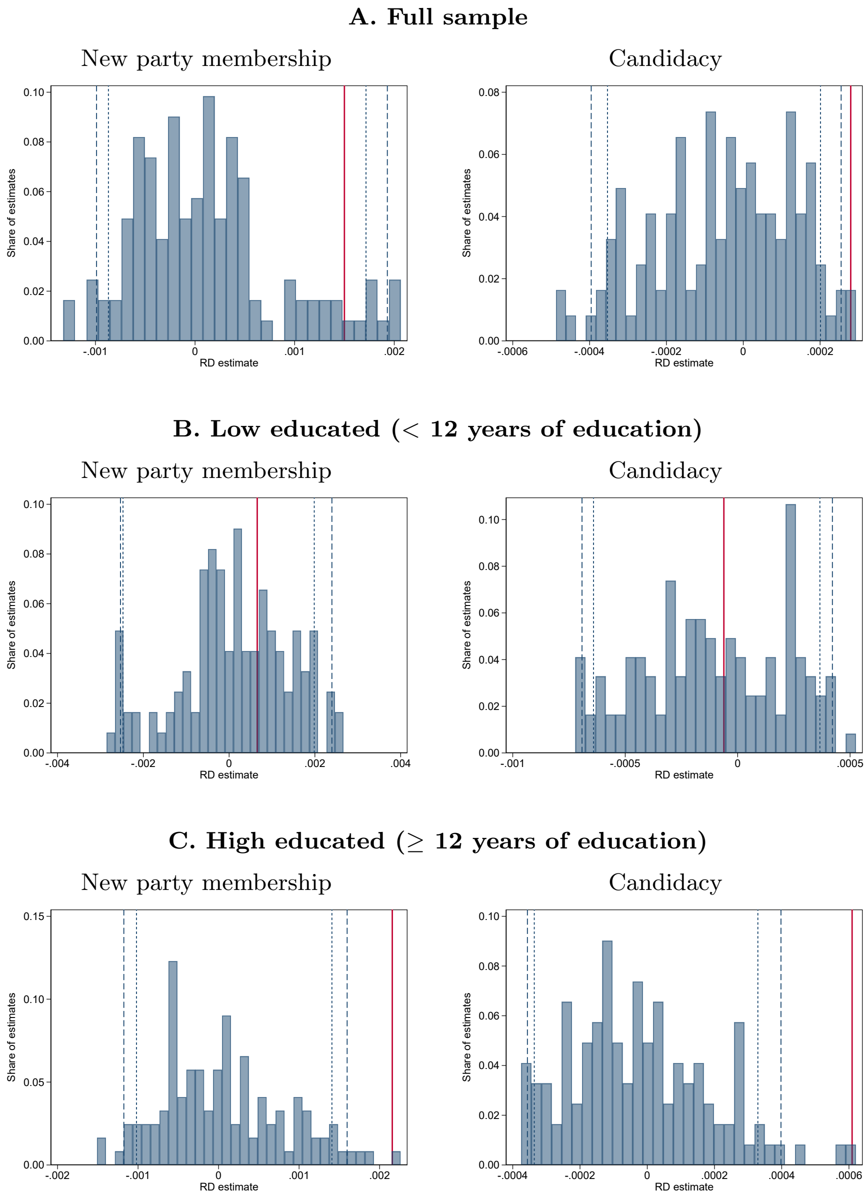
Note: The table presents robustness checks for the effect of unemployment benefit eligibility on the likelihood of newly registering as a party member when using different bandwidths and polynomial orders of the running variable. 'opt' refers to the optimal bandwidth proposed by Calonico *et al.* (2014). Standard errors clustered at the individual level are in parentheses. The table also reports the control mean of the outcome at the cutoff and the effect sizes scaled by the control mean. All coefficients, standard errors, and control means have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table C.2: Effect of UI eligibility on running for local councilor: alternative specifications

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Bandwidth (days):	30	opt	30	60	90	opt	90	120	150	opt
Polynomial order:	0	0	1	1	1	1	2	2	2	2
[A] Full sample										
β	0.025^{*}	0.020	0.055^{**}	0.028	0.030^{*}	0.030	0.031	0.024	0.025	0.030
	(0.013)	(0.013)	(0.027)	(0.019)	(0.016)	(0.020)	(0.023)	(0.020)	(0.018)	(0.024)
Control mean	0.447	0.453	0.431	0.449	0.444	0.447	0.450	0.446	0.451	0.447
Relative effect	5.6%	4.3%	12.9%	6.2%	6.7%	6.7%	7.0%	5.3%	5.6%	6.6%
Observations	1,035,237	1,105,868	1,035,237	2,048,674	3,058,455	1,908,388	3,058,455	4,101,059	5,107,090	2,953,238
[B] Low educated (< 12 years of education)										
β	0.001	-0.007	0.035	-0.006	-0.004	-0.003	-0.002	-0.008	-0.001	-0.004
	(0.018)	(0.018)	(0.037)	(0.026)	(0.021)	(0.027)	(0.032)	(0.028)	(0.025)	(0.033)
Control mean	0.419	0.424	0.396	0.428	0.417	0.427	0.435	0.427	0.423	0.426
Relative effect	0.3%	-1.6%	8.9%	-1.4%	-0.9%	-0.6%	-0.5%	-1.9%	-0.2%	-1.0%
Observations	497,269	531,075	497,269	986,755	1,485,874	934,928	1,485,874	2,021,245	2,540,083	1,414,964
[C] High educated (≥ 12 years of education)										
β	0.047^{**}	0.047^{**}	0.075^{**}	0.061^{**}	0.063^{***}	0.072^{***}	0.063^{*}	0.054^{*}	0.053^{**}	0.063^{*}
	(0.019)	(0.019)	(0.038)	(0.027)	(0.023)	(0.026)	(0.034)	(0.029)	(0.026)	(0.034)
Control mean	0.473	0.473	0.463	0.468	0.465	0.461	0.463	0.464	0.473	0.462
Relative effect	10.0%	10.0%	16.2%	13.0%	13.6%	15.6%	13.5%	11.6%	11.2%	13.6%
Observations	537,968	537,968	537,968	1,061,919	1,572,581	1,148,350	1,572,581	2,079,814	2,567,007	1,519,603

Note: The table presents robustness checks for the effect of unemployment benefit eligibility on the likelihood of running for local councilor when using different bandwidths and polynomial orders of the running variable. 'opt' refers to the optimal bandwidth proposed by Calonico *et al.* (2014). Standard errors clustered at the individual level are in parentheses. The table also reports the control mean of the outcome at the cutoff and the effect sizes scaled by the control mean. All coefficients, standard errors, and control means have been scaled by 100, such that effects are interpreted in terms of percentage points. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure C.4: Effect of UI eligibility: permutation tests



Note: The figures compare RD estimates for the effect of UI eligibility on the likelihood of newly registering as a party member and running for local councilor at the true 16-month cutoff (vertical red line) with the distribution of estimates at all possible placebo cutoffs within 180 days away from the true cutoff. The dashed (dotted) lines indicate the 2.5 (5) and 97.5 (95) percentiles of the distribution of placebo cutoffs. Estimates are based on the local linear regression model (4) with a bandwidth of 60 days. Panel A shows results among all workers in the sample, and Panels B and C distinguish workers with less than vs. at least 12 years of education.

D Appendix to Section 7

Table D.1: Municipality-level unemployment rate and candidate quality

	Candidates' average years of education (Δ 2000-2012)			
	(1)	(2)	(3)	(4)
Unemployment rate (Δ 2000-2010)	0.0461*** (0.00374)	0.0330*** (0.00352)	0.104*** (0.00733)	0.0888*** (0.00833)
Observations	5,476	5,476	5,476	5,476
State FE		✓		✓
Weighting			✓	✓

Note: The table presents regression results for the municipality-level relation between councilor candidates' average years of education (change between elections in 2000 and 2012) and unemployment rates (change between 2000 and 2010). Columns (2) and (4) include state fixed effects, and in columns (3) and (4) municipalities are weighted by their population size. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table D.2: Characteristics of new party members, councilor candidates, and elected candidates

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	New party members				Councilor candidates				Elected councilors			
	Treated	Control	Std Diff	P-Val	Treated	Control	Std Diff	P-Val	Treated	Control	Std Diff	P-Val
Years of education	10.79	10.66	-0.05	0.00	10.79	10.67	-0.05	0.01	11.15	11.17	0.01	0.92
Earnings	2169.33	1981.66	-0.10	0.00	2145.66	1956.10	-0.11	0.00	2360.78	2202.11	-0.09	0.16
Earnings residual	0.03	-0.02	-0.05	0.00	0.04	0.00	-0.04	0.03	0.16	0.08	-0.07	0.26
Previous candidacy	3.23	2.53	-0.04	0.00	12.93	11.09	-0.06	0.00	28.37	32.72	0.09	0.12
Elected					7.22	7.60	0.01	0.38				
Observations	46,828	39,901			7,914	6,477			571	492		