Essays on the Political Economy of Public Spending: Evidence from Brazil

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I, Fernanda Senra de Moura, confirm that the work presented in this thesis is my own. Where information has been derived from other sources, I confirm that this has been indicated in the work.

Abstract

In this thesis, I study the relationship between political institutions and the provision of public goods. The chapter titled "The Role of Parties in the Distribution of Public Funds" studies the role of political parties in explaining inequalities in access to public funds for pork-barrel by federal legislators in Brazil. I find that differences in individual heterogeneity are the main drivers of the observed inequality, which, in turn, is limited by negative sorting of low influence politicians into political parties with higher collective bargaining premiums. In the chapter titled "Gender Gaps in Parliament: Access to Public Funds by Legislators in Brazil", I show that this pattern is particularly important to explain the outcomes of female legislators. More specifically, I document a gender gap in access to funding and use the estimates from the model to decompose the gender gap into components associated to politician and party heterogeneity. The results suggest that the segregation of women in parties with higher party premiums partially offsets the gender gap. I find evidence that negative sorting is associated to party ideology: left wing parties are more likely to have an advantage in party premiums and members with lower individual influence, while members of centrist and right wing parties have, on average, higher individual fixed effects but lower collective bargaining premiums. The chapter titled "The Heterogeneous Effects of Reelection Incentives: Evidence from Brazil" exploits term limits on mayoral elections in Brazil to study the heterogeneous effects of reelection incentives on the provision and maintenance of water wells in the Brazilian Semiarid area. Using mixed incumbent-challenger close elections, I find causal evidence that mayors running for reelection provide more water wells to citizens as long as the frequency of dry years is sufficiently high.

Impact Statement

This project studies the role and effect of political institutions on policy outcomes relevant to development, equality and, more generally, to social welfare. Firstly, my findings show the important contribution of political parties as institutions that are able to organize individuals for collective action in the competition for public funds, a feature that is particularly relevant for females and, possibly, other minorities.

This study also documents gender gaps in access to a type of funding that is relevant for the reelection outcomes of incumbent politicians, therefore contributing to the discussion on the low pace at which the participation of females in politics has been increasing and the progression of women's political careers.

As the political representation of underprivileged groups is key to reduce inequalities, this work contributes to the discussion on how political institutions can facilitate the access to public resources by underrepresented groups. More generally, this work contributes to the debate on whether political parties should be strengthened and if independent candidacies should be encouraged.

The discussion on the effects of reelection incentives on the provision of public goods may also inform policy makers and development institutions in public policy design, for example.

In academia, my work brings methods well-established in fields such as Labour and Education Economics to address questions pertaining to the Political Economy and Gender Economics fields and to Political Science disciplines. To the best of my knowledge, it is the first to quantify the contribution of parties to inequalities in access to funding by elected incumbents. Moreover, the methods presented in this paper could be used in other settings in which politicians switch parties often.

The findings and methods in this project may be disseminated through scholarly journals, specialist or mainstream media, and as an input to applied microeconomics and political economy courses, for example. Finally, while I focus on evidence from Brazil, the methods and findings from this project could contribute to similar studies in any democratic political system.

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Contents

List of Figures

List of Tables

Chapter 1

Introduction

The specific features of a political institution's design may have profound implications on the behaviour of voters and politicians, affecting major economic outcomes and social welfare. A large body of literature studies the origins, stability, and effects of political institutions on difference outcomes. This project studies the relationship between political institutions and the provision of public goods. I study how political parties and reelection incentives affect the behaviour and outcomes of politicians in office.

In Chapter 2, I begin by studying the distributional role of parties in explaining inequalities in access to public funds by federal deputies. In Brazil, all federal deputies are entitled to the same amount in Individual Budgetary Amendments (IBAs) each fiscal year, but the actual implementation of the projects they propose is subject to the approval of the Executive Power (President and Ministers). Specifically, I use a two-way fixed effects model to decompose the total variance of log disbursements into three main components: (1) party-specific heterogeneity, captured by party fixed effects, (2) segregation and (3) sorting of politicians into parties.

The first main result is that individual characteristics are the main drivers of the observed inequality in funding, but parties have a role in offsetting the variance of log disbursements through a composition effect measured by a negative covariance between the party fixed effects and individual heterogeneity. Moreover, a detailed decomposition of the variance shows that while politicians with higher individual fixed effects tend to segregate in different parties than those with lower individual fixed effects (the variance of the average politician effects explains 29% of the total variance), they are likely to do so in parties with lower party premiums. Converted to correlations (-0.55), the estimates are indicative of negative assortativity. In other words, as politicians with lower individual influence tend to sort into parties with higher party premiums, the overall contribution of parties is to reduce inequalities in access to public funding.

Next, I ask what are the implications of the findings above for the outcomes of women. Female representation in politics has been increasing over the years, but given the slow pace at which the number of women in politics has been growing, understanding the bottlenecks hindering the political careers of women is still of paramount importance. In Chapter 3 I document gender differences in access to IBAs and estimate the contribution of political parties in explaining the observed gender gaps. I focus on the execution of IBAs because signaling competence and rewarding voters through the provision of public goods is an important channel available to incumbents seeking to advance their political careers.

I use the estimates from a two-way AKM-type model with party and politician fixed effects and then apply the method in (Cardoso et al. 2016, Gelbach 2016) to decompose the cross-sectional gap. I find evidence of moderate negative sorting on party and individual fixed effects associated to party ideology and that this pattern of sorting plays an important role in explaining the observed gender gaps, since women tend to segregate in left wing parties in the period considered. On average, the estimated party fixed effects of females are higher than for males, while the average individual fixed effects estimated for women are lower than those for men. That is, while there is a female discount in individual premiums, it is partially offset by an advantage in party premiums, especially for the women in left wing parties.

Chapter 4 studies the relationship between reelection incentives and the provision of water wells in the driest areas of the Northeast and Southeast Brazil. I use a regression discontinuity design with mixed incumbent-challenger close elections and variation in rainfall to identify heterogeneous effects of eligibility for a second consecutive term on the drilling of water wells. By allowing for an interaction between treatment and the frequency of dry years, I show that reelection incentives increase the drilling of water wells as long as the frequency of dry years is sufficiently high and that this effect is significantly larger in the driest areas. Additionally, using data from the Brazilian Underground Water Censuses, I show a negative relationship between the maintenance of public water wells and time to election in municipalities where drought frequency is sufficiently high. The above relationship is explained in the context of a Rogoff-type model (Rogoff, 1990) of political budget cycles, which I modify to allow for the provision of a public good from which voters derive utility only under certain states of nature (state-dependent utility).

This thesis is organized as follows: Chapter 2 studies the role of parties in explaining inequalities in access to public funds by federal legislators in Brazil. In Chapter 3 I document gender differences in access to public funds and study the role of parties in explaining the observed gender gaps. Chapter 4 presents theory and evidence on the effect of reelection incentives on the provision of water wells in the Brazilian Semiarid. Chapter 5 concludes.

Chapter 2

The Role of Parties in the Distribution of Public Funds

2.1 Introduction

The allocation of public funds for the provision of public goods is one of the main responsibilities of elected politicians. Motivated by welfare and reelection incentives, incumbents compete for and allocate funding strategically. Since inequalities in access to public funding by politicians in office may affect their political careers and hinder the active representation of groups of voters, understating the role of political institutions in explaining the distribution of public budgets among elected politicians is a relevant subject.

This paper studies the distributional role of parties in explaining inequalities in access to public funds by federal legislators in Brazil between 2011 and 2018. In Brazil's Chamber of Deputies, Individual Budgetary Amendments (IBAS) are an important source of discretionary funding available to and typically used by federal legislators to deliver public goods to their constituencies (Finan and Mazzocco, 2016). Using data on IBAs, I build measures of access to funding (log disbursements) at different stages of the budgetary process, from planning to actual project implementation, and document that inequality is highest at the disbursement (final) stage.

In Brazil, all federal legislators are entitled to the same amount in IBAs each fiscal year, but the actual implementation of the projects proposed is subject to approval by the Executive Power (President and Ministers). This institutional setting creates a bargaining problem between federal legislators and the leaders of the Executive, the latter with power and incentives to allocate funding strategically. So while there is equal treatment of politicians in the initial allocation of funding (planned spending), throughout the stages of the budgetary execution the negotiations with the Executive Power create inequalities and, as a result, the dispersion at the stage of disbursements (actual spending) is the highest.

To study the role of parties in explaining the observed funding inequalities, I apply methods widely used in the Labour Economics literature in the study of wage inequalities. As in Song et al. (2018), I start with a within- between-party decomposition of the total variance of log disbursements and find that the contribution of the between-party component was relatively low every year, ranging from 4.8% to 8.2% between 2011 and 2018. I also find that differences between parties are more relevant among the smaller parties and for parties in the opposition.

To further decompose the total variance of log disbursements, I fit a public funding equation with log disbursements as the outcome variable and party and politician fixed effects. In labour market applications, two-way fixed effects models were first proposed in Abowd et al. (1999) and Abowd et al. (2002), henceforth AKM, and subsequently used by many others (e.g. Card et al. (2013), Song et al. (2018), and Card et al. (2015)) to estimate wage equations.

In my setting, the politician fixed effect aims to capture the component of a politician's funding outcomes that is fully portable across parties and that may reflect factors such as heterogeneity in individual ability, including bargaining ability, preferences over public spending, and discrimination, for example. The party fixed effect captures a proportional funding premium (or discount) that is common to all party members, which in a context of competition for public funding and bargaining with the president I interpret as party-level collective bargaining component of the politicians' disbursement outcomes. Besides the fixed effects, the model includes time intercepts and a measure of *governism* capturing politicians' support to the Government's legislative agenda, fully interacted with a pre-election indicator. This specification draws on works such as Brollo and Nannicini (2012), Pereira and Mueller (2004), and Baião et al. (2018).

Using the estimates from the model, the contribution of parties in explaining inequalities in access to funding can be decomposed in terms of party-specific heterogeneity, captured by party fixed effects, and of composition components, namely segregation and sorting of politicians into parties. More specifically, segregation aims to capture the extent to which politicians with similar individual fixed effects tend to be members of the same parties, while sorting refers to the covariance between individual and party fixed effects.

The first main result is that individual characteristics are the main drivers of the observed inequality in access to funding, but parties have a role in offsetting the variance of log disbursements through a composition effect driven by a negative covariance between the party fixed effects and individual heterogeneity. The variance of individual fixed effects (0.96) contributes to more than 80% of the total variance of log disbursements (1.17) and the total contribution of the party-related components (-0.32) partially offsets the dispersion arising from differences between individual politicians. More specifically, while the variance of party fixed effects (0.31) corresponds to about 26% of the total variance, the estimated negative covariance (-0.60) between party and person fixed effects is such that negative sorting more than offsets any differences between the party premiums, as sorting corresponds to 77% of the components with a negative contribution to the dispersion in log disbursements.

A detailed decomposition of the variance shows that while politicians with higher individual fixed effects tend to segregate in different parties than those with lower individual fixed effects (the variance of the average politician effects explains 29% of the total variance), they are likely to do so in parties with lower party premiums. Converted to correlations (-0.55), the estimates are indicative of moderate negative assortativity. In other words, as politicians with lower (higher) individual influence tend to select into parties with higher (lower) influence, the overall contribution of parties is to reduce inequality through a composition effect.

Next, I document a relationship between negative assortativity and party ideology. In particular, I find substantial and statistically significant positive differences in mean party fixed effect between left and right-wing parties and between left and centrist parties. Additionally, the average politician fixed effects in center and right-wing parties is higher than in left-wing parties. The differences in means are also substantial and statistically significant. That is, left-wing parties have an advantage in collective bargaining, as measure by the estimated party fixed effects, while politicians in center and right-wing parties have an advantage in individual fixed effects. The differences in means between centrist and right-wing parties are not substantial nor significant.

Regarding the dynamics between legislatures, the total variance of log disbursements was lower in the second period, which was expected given a set of reforms introduced in 2015 to limit the discretion of the Executive Power over the execution of the budgetary amendments. This decline in the total variance is mainly explained by the decreased dispersion in person-related components of the variance. The overall contribution of the party-related components was negative in both terms, but less so in the second as the distributive contribution of the negative sorting was lower. Interestingly, the advantage in party premiums of left-wing parties is observed in both terms, despite the switch in the leadership of the Executive Power from left to center, suggesting that the term-specific party premiums also capture heterogeneity such as party preferences and approach to collective bargaining.

One major drawback of the AKM model is the possibility that the variance and covariance estimates indicate negative sorting even in the presence of true positive assortative matching due to estimation error caused by limited mobility bias (Andrews et al., 2008). To tackle this problem, Kline et al. (2020) propose a framework for unbiased estimation of the variance and covariance components, which I apply as a robustness check. In sum, the bias corrected estimator corroborates the conclusions from the baseline AKM estimator. However, bias correcting requires more strict sample trimming than the baseline AKM model and, given the limited sample size of this application, the number of party fixed effects in the dynamic analysis significantly declines. Therefore, I focus on the baseline AKM estimates since bias correcting does not provide evidence that the estimated negative sorting is solely a result of limited mobility bias.

Besides the standard OLS identifying assumptions, identification of unit fixed effects in a AKM model relies on a set of orthogonality conditions that hold under the assumption of exogenous mobility, that is, if mobility patterns are independent of the error term. One potential source of endogenous mobility is systematic sorting of politicians into parties based on a match effect if, for example, politicians with strong bargaining skills systematically move to parties with a stronger relationship with the executive power in order to perform better in terms of disbursements.

To test for the presence of match effects, I follow Card et al. (2013) and fit a model fully saturated with party-politician interaction terms. Although the root mean square error (RMSE) from the baseline AKM model is slightly higher than the RMSE from the match model, suggesting that the latter fits the data slightly better, the reduction in RMSE is very small (3.5% in the polled model), implying that the magnitude of the match component is small.

This paper relates to various strands of literature. Firstly, I contribute to the literature on the role of parties in the competition for public funds by incumbent politicians. Brollo and Nannicini (2012) show that party affiliation plays a role in the allocation of federal public funds across Brazilian municipalities. They find that in the final two years of the municipal mandate (before municipal elections take place), the amount of transfers to municipalities party-aligned with the President increases, while the transfers to mayors from the opposition decline. Azulai (2018) uses exogenous variation in city-ministry partisan alignment to identify the effect of political connection (measured by party membership) and quantify party-based favouritism in the allocation of funds to Brazilian municipalities. The author shows that when cities become politically connected with a ministry through mayor-minister co-partisanship, the amount of funds they receive from that ministry increases by 15% on average. Within parties, Curto-Grau and Zudenkova (2018) show that in the U.S. House of Representatives discretionary spending increases with party discipline. My contribution is to take advantage of party switching in the Brazilian multi-party system to identify and quantify the contribution of party heterogeneity to inequalities in access to funding, therefore focusing on the role of parties as political institutions as part of the bargaining process between the Legislative and the Executive Powers.

My focus on Budgetary Amendments, which are a specific source of funding within the National Budget, speaks to a literature that studies the broader role of this budgetary mechanism within the context of the political and electoral institutions currently in place in Brazil. Pereira and Mueller (2004) argue that bargaining over the execution of budgetary amendments enables the president to sustain a coalition in Congress at a relatively low cost and they show evidence that legislators who vote in line with the Government preferences more often are rewarded with budgetary amendments execution, while those who vote with the Government less often are punished. Similarly, Raile et al. (2011) show that the President allocates discretionary funding to reward legislative support with higher rates of execution of budgetary amendments. Finan and Mazzocco (2016) examine the within-state allocation of budgetary amendments across municipalities during the 50^{th} Legislature and they document a clear positive relationship between the amount of public funds allocated by incumbent federal legislators and the share of votes they received after running for reelection in 1998. Departing from this motivation, the authors estimate a structural model of allocation of public funds and simulate institutional reforms as alternatives to reduce the distortions caused by reelection incentives of incumbent federal legislators.

Finally, my research relates to a large body of work on the role of firm and worker heterogeneity explaining wage inequalities. I use methods first proposed in Abowd et al. (1999) and Abowd et al. (2002) and further developed in works such as Andrews et al. (2008), Upward (2004), Cornelissen (2008), Kline et al. (2020) and applied by Card et al. (2013), Song et al. (2018), Card et al. (2015), and others.

The remainder of the chapter is organised as follows: Section 2.2 describes the context and institutional setting, followed by a descriptive overview of funding inequalities in Section 2.3. Section 2.4 presents the econometric methods I use to study the role of parties in explaining funding inequalities, followed by the results in Section 2.5. Section 2.6 concludes.

2.2 Institutional setting

2.2.1 Individual Budgetary Amendments

The Chamber of Deputies is the lower house of the Brazilian National Congress. It comprises 513 federal legislators, who serve four-year terms with no term limits and whose primary responsibility is the allocation of public funds (Finan and Mazzocco, 2016) - every year they discuss, amend, and approve the State Budget, known as the Annual Budget Law[1].

As part of the process, legislators are entitled to fixed yearly budgets to finance projects of their choice[2]. These are known as Individual Budgetary Amendments (hereinafter IBAs) and are typically used by federal legislators to allocate funds to their constituencies [3]. Finan and Mazzocco (2016) show that Budgetary Amendments are used by federal legislators strategically and that voters reward them accordingly. Specifically, the authors show that federal legislators allocate more funding to municipalities with more voters and document a positive correlation between vote shares and funding allocation at the municipality level[4]. That is, Budgetary Amendments are an effective channel to improve electoral performance and are

[1] *Lei Orçamentária Anual (LOA)*.

[2] Politicians decide which institutions and projects receive the funds, subject to broad guidelines such as the priorities and Programs (budget lines) set on the Multiannual Plan. Moreover, since 2015 each legislator must allocate 50% of their Budgetary Amendments to projects in the health sector. This rule was introduced as part of the Constitutional Amendment 86/2015, a reform that limited the power of the Executive Branch to delay or cancel Budgetary Amendments. As part of the bill, the Executive Branch negotiated the prioritization of the health sector as a means to support the country's overall mandatory spending on health and to partially offset the impact of the reform on the overall level of mandatory spending.

[3] In Brazil, each state is a constituency with a number of federal legislators that depends on the state's population size.

[4] Finan and Mazzocco (2016) estimate a model of Budgetary Amendment allocation across municipalities to quantify the welfare costs of reelection incentives. They find that social welfare concerns are relevant, but that reelection motives cause substantial distortions relative to the socially optimum.

perceived by federal legislators as such.

2.2.2 Inequalities in access to funding

The Annual Budget Law is a planning institution based on expected revenues and public expenditures, and it is responsibility of the Executive Power[5](hereinafter Government) to ensure that actual revenues and spending are balanced as approved by the National Congress, cutting discretionary spending during the year if necessary.

Therefore, while each year all federal legislators are entitled to issue Budgetary Amendments up to the same amount, these projects may be subject to budget cuts, postponements or cancellations at the discretion of the Executive Power. In practice, budget cuts are routine and competition for funding is expected every fiscal year, so actual project implementation is not guaranteed and inequality in access to funding arises.

Within this context, the Government negotiates the actual implementation of IBAs both with party leaders and directly with individual legislators. These negotiations do not take place through formal institutions, but they are a major practice in the Brazilian National Congress and have been examined extensively in the literature.

Regarding the Government's objectives in such negotiations, Pereira and Mueller (2004) and Baião et al. (2018) provide evidence that Budgetary Amendment execution is a coalition-building tool [6]. The Government may also favour its own party and parties in the Government coalition (Baião et al., 2018).

The Government may negotiate the execution of IBAs with party leaders, who would represent all party members. Records from the Chamber of Deputies illustrate the bargaining process:

> ”Yesterday, after over 12 hours of negotiations among party leaders,

[5]President and Ministers

[6]The idea is that in Brazil's fragmented multiparty system, the Government's party is unlikely to constitute a majority in Congress so, in order to get support for its legislative agenda, it needs to build and maintain a coalition as well as to negotiate with the opposition, and the execution of Budgetary Amendments in one major mechanism through which support is negotiated.

no agreement was reached regarding the votes for two bills...(party leaders) are unhappy with the pace of Budgetary Amendments execution...The opposition, mainly, demands fulfillment of a deal closed last week with the Government's Institutional Relations Secretariat...which included the allocation of R$2.5 million in Budgetary Amendments for each legislator...(Siqueira and Júnior, 2012)"

However, the Government may also approach legislators individually, as illustrated by an interview given by a senator in 2020, in which he claims being approached by a government representative and offered R$30 million in Budgetary Amendments. According to the legislator, not all senators were offered funding and the distribution of funds was not based on objective criteria (Nery, 2020).

That is, the party leadership may coordinate the negotiation of IBAs execution for all members, but individual politicians may also engage directly with representatives of the Government if the party leadership allows.

2.2.3 Recent attempts at reducing inequality

Up until 2015, Budgetary Amendments could be cancelled by the Executive without limit or need for justification. In order to reduce such discretion, in 2015 the National Congress approved a Constitutional Amendment (Nr. 86/2015) determining that the execution of individual budgetary amendments be mandatory[7], subject to the availability of financial resources at the National Treasure and to technical approval[8]. While this reform limited the ability of the Executive Power to cancel projects, through the course of the fiscal year the Budgetary Amendments can still be cancelled if fiscal revenues are below expectations. Prior to the approval of the Constitutional Amendment Nr. 86, similar rules had already been included in the 2013 and 2014 Annual Budget Laws. Pederiva and Pederiva (2015), shows that the actual execution of the Budgetary Amendments issued during those years remained significantly below the amount planned in the Budget, indicating that room

[7]Up to a limit of 1.2% of the net current revenues in the previous year.

[8]There are technical criteria to be followed, such as 50% of the total funding being allocated to the health sector and eligibility requirements for the grantee institution (typically Municipal or State Governments).

for negotiation between the Executive and the Legislative powers remained despite the reforms. Similarly, since 2015 budget cuts still affect Budgetary Amendments, even if to a lesser extent due to the Constitutional Amendment Nr. 86/2015.

2.2.4 The budgetary process

In the lifetime of a project, budget cuts and project cancellations take place within the context of a budgetary process that has various stages.

Authorization stage: In the first stage essentially consists of budget planning based on expected revenues and expenditures for the year. The Executive Power (President and Ministers) proposes a detailed annual budget for the upcoming fiscal year, which is revised, amended, and approved by the Legislative, and then formalized as the Annual Budget Law. Only expenditures included in the Annual Budget Law can be executed by the government, therefore it is crucial that federal legislators negotiate to have their Budgetary Amendments authorized in the annual budget.

Commitment stage: Once the fiscal year begins and tax collection is initiated, the next stage of the budgetary process is the commitment of funds, which effectively means creating financial reserves for projects. At this stage, binding budget constraints already imply additional competition, and not all projects authorized in the annual budget have its funds committed.

Moreover, even after commitment of funds, the actual implementation of a project is not guaranteed. Throughout the year, it is typically necessary to limit commitments and payments in order to keep spending aligned with actual government revenues, so some projects may be cancelled, revised down or delayed after the commitment stage. In 2018, for example, the payments to budgetary amendments proposed by federal legislators amounted to about half the total commitments that year and nearly 25% of the budgetary amendments issued in 2015 still have not been paid. That is, commitment is necessary but not sufficient for actual project implementation and which projects will be delayed or cancelled is a result of extensive negotiations.

Liquidation stage: For the successful projects, after contracting of goods and/or services, the next stage is the verification that goods and/or services have been provided as expected. This stage is known as *liquidation*.

Payment stage: Disbursement of funds to pay contracted providers for actually delivered goods and/services.

Within this context, in order to capture funding inequalities it is important to take into account the different stages of the budget execution, since competition for funding and bargaining between the executive and legislative can take place at all stages. To fully capture inequalities in access to funding, I focus on the last stage of the budgetary execution, and use disbursements as my main outcome variable.

2.2.5 Party switching

One interesting feature of Brazil's institutional setup is the combination of partisan fragmentation with frequent party switches: currently, the country has 33 parties registered at the Brazilian Electoral High Court (TSE), most of which are represented in the National Congress[9] and, historically, it has been among the countries with the highest rates of party switching in the world (Desposato, 2006; O'Brien and Shomer, 2013).

This feature of the context is key because politicians mobility across parties is the variation that allows party and politician fixed effects to be separately identified.

The party switching process involves different sets of rules depending on the period being considered. Elected federal legislators were allowed to switch parties without justification at any point during their terms until late 2007, when the Electoral High Court issued a Resolution (Electoral High Court, 2007) determining that, for politicians elected via party-list proportional representation (federal legislators included), unjustifiable party switches could be sanctioned with loss of the mandate if the origin party claimed the seat[10]. The Resolution considers as acceptable justification (i) party merges or incorporation, (ii) the establishment of a new party, (iii) ideological conflicts due to major changes in the party program, and (iv) personal discrimination.

[9] 30 political parties with elected members in 2018, 28 in 2014, and 22 in 2010).
[10] Other interested parts may also request sanction

Despite this reform, party switches continued to take place and, in 2015, a switching window was formalized into Law (Presidência da República, 2015) to allow elected politicians to switch parties without justification during a designated period of 30 days in each election year.

Desposato (2006) proposes a model for party switching in Brazil's Chamber of Deputies and, using data from the 49^{th} and 50^{th} Legislatures, finds that federal legislators switch parties to join the government coalition, for ideological consistency, and to boost electoral outcomes by joining parties with popular candidates.

Parties, on the other hand, are most interested in receiving politicians with a large voter base, since the most popular candidates may significantly contribute to the election of other party members within the party-list proportional representation system.

Finally, another feature of the institutional setting relevant to this paper concerns the measurement of access to funding. In Brazil, public spending is implemented in four main stages: (i) budget allocation, which essentially consists of budget planning based on the expected revenues and expenditures for the year, (ii) procurement and commitment, which involves project-specific cost assessment and the financial reservation of funds before contracting, (iii) contracting and delivery of goods or services, which are physically verified and formalized in the budget in a step named verification, and (iv) disbursement, which refers to the payment of projects actually contracted and delivered. In order to take into account the planning-execution gap and capture inequalities in access to funding, I focus on disbursements, the last stage of the budgetary execution, as my main outcome variable.

2.3 Data and descriptive overview

2.3.1 Data sources

Data on the budgetary amendments are obtained from the Matching Grants Administrative System (*Sistema de Gestão de Convênios e Contratos de Repasse - SICONV*) from the Brazilian Federal Government. Matching grants are an institutional mechanism for project-specific budgetary execution and the budgetary amendments proposed by federal legislators are formally implemented through this institutional device. The SICONV database is composed of several datasets with detailed information on all matching grants, including daily information on the budgetary execution of the projects financed through the budgetary amendments. Daily data are aggregated as needed. In particular, given that the financial execution of projects is often carried out in a small number of installments, for the main analysis I build measures of access to funding by fiscal year. The database also includes detailed information on several project characteristics, including the authorship of the proposal. I consider only the amendments that can be assigned to a single federal deputy and I use the name of the individual applicant to merge data from the SICONV to other political data, such as party affiliation and individual characteristics. The analysis does not include projects that were proposed by blocs or groups of politicians. In light of the different stages of the budgetary execution, I build measures of access to funding at different stages of spending process: *proposals*, *commitments*, *net commitments* (after *cancellations*), and *disbursements*. The SICONV was launched in 2008, but good availability of data start from 2010. I focus on two whole Legislatures which span from 2011 to 2018.

Data on the identity of elected federal legislators come from the Chamber of Deputies. From the same source I obtain detailed data on all open votes in order to build, for each year and federal deputy, a measure of *governism* that captures the share of the open vote sessions in which a federal deputy has voted in agreement with the government guidance, whenever guidance was given.

From the Chamber of Deputies website I also scrape data on party membership and build a monthly panel in order to track party switches. The main analysis is based on annual data and given that party switches happen within the fiscal year, different criteria to assign party membership in a given year are considered. Party membership in the main analysis is defined as the party in which the politician was a member for the most part of the second half of the fiscal year. Using the second half of the year is more appropriate because of the general seasonality in the execution of public spending. More specifically, it is common that a significant share of the budget is disbursed near the end of the fiscal year after tax revenues have realised and the uncertainty about the primary fiscal balance is reduced. Therefore, in order to account for this seasonality I consider that party membership in the second half of the fiscal when tracking party switches by year[11].

Coalition parties are identified from Figueiredo and Limongi (2000), Garcia (2017), Magna and Rezende (2015), Mauerberg Junior (2016), and Ribeiro (2018). For periods not previously analysed in the literature, participation in the government coalition is inferred based on the appointment of ministers, which is one of the main mechanisms for coalition building in Brazil. (Mauerberg Junior, 2016).

Similarly, party ideology is assigned based on the classification in Faustino et al. (2019) and Miguel and Machado (2007). For the parties not covered in the analysis above, the party's self-declared ideology and/or evidence of alignment with left or right-wing parties was used when available. A small number of parties could not be classified according to ideology.

Party merges and the advent of new parties is tracked based on the official records of Brazilian parties from the Superior Electoral Court (*Tribunal Superior Eleitoral - TSE*), which contain information such new party registration, renaming, and closing of parties. In order to avoid that party renaming is coded as party switches, I harmonize party membership according to the 2018 official party names[12].

Data on individual characteristics such as age, gender, and education are obtained from the TSE (*Tribunal Superior Eleitoral*) election database and merged with the budgetary amendments dataset based on the politician's name. Measures of political experience are also built from data available at the TSE. Specifically, I use the politicians' unique tax payer number, unique voter identification number, name, and date of birth[13] to identify previous elections in which federal legislators

[11] Results are similar when party membership is defined as the party in which the politician was a member for the most part of the whole fiscal year

[12] See Appendix A for details on party evolution.

[13] While the politicians' unique tax payer number is more likely to be unchanged over the lifetime, this information is missing for some politicians. Registration of the unique voter identification

in the sample have participated or won. Finally, state level covariates come from the Brazilian Institute of Geography and Statistics (*IBGE*).

2.3.2 Main sample and descriptive statistics

I use data on the budgetary execution of Budgetary Amendments during two Legislatures, from 2011 to 2018. The final dataset contains 2,750 person-year observations. It includes data on 687 of the 763 federal legislators elected in 2010 and 2014, since I restrict the analysis to politicians with positive total disbursements and registered voting activity in the Chamber of Deputies in any given year. As a result, 29 of the 32 parties represented in the Chamber of Deputies are included in the analysis.

Table 2.1 reports descriptive statistics for the politicians and parties included in the main sample. The main outcome of interest is *Log Disbursements*. As previously discussed, the dispersion of outcome variable in the main sample is lower than in the population of elected federal legislators, but in both cases the standard variation slightly decreases between the first and second terms.

Considering both legislatures, the median age of the federal legislators in the sample is close to 55. The youngest politician was 23 years old when elected and the eldest federal deputy observed while in power is 88 years old. About 80% of the politicians have at least a college degree and, considering that about 8% have at least the high school degree, about 10% did not report having completed at least the high school when elected.

Representation of women in the sample is low, corresponding to only about 9% of the observations. About 34% of the observations on politicians from parties classified as left-wing, while those affiliated to center of right parties correspond to 65.5% of the observations. Only a small fraction of the observations could not be classified by ideology, including those on individuals without party membership.

Finally, the share of politicians in parties participating in the Government Coalition was, on average, 67%. Such composition is observed in both legisla-

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number in the election dataset is mandatory, but this identifier changes over time for a few politicians, possibly due to being reissued after cancellation. Therefore, the politician's name and date of birth are also used to merge election data from different years.

Table 2.1: Descriptive Statistics - Politicians

		Largest Connected Set			All elected politicians		
		54th Leg.	55th Leg.	Pooled	54th Leg.	55th Leg.	Pooled
Politicians Disbursements (Log):	Mean	14.347	14.692	14.537	9.421	12.032	10.721
	Std. Dev.	1.125	1.025	1.084	6.857	5.725	6.452
Age:							
	Min.	23.156	25.773	23.156	22.153	23.770	22.153
	Median	55.438	55.559	55.493	54.412	54.686	54.552
	Max.	85.518	88.529	88.529	85.518	88.529	88.529
Education:							
	College (%)	0.790	0.808	0.801	0.778	0.799	0.788
Gender:							
	Female (%)	0.089	0.095	0.093	0.088	0.099	0.094
	Male (%)	0.911	0.905	0.907	0.912	0.901	0.906
Party Ideology:	Left (%)	0.358	0.327	0.340	0.363	0.319	0.341
	Center (%)	0.278	0.291	0.284	0.276	0.305	0.291
	Right (%)	0.363	0.374	0.370	0.360	0.367	0.363
	Not classified (%)	0.000	0.000	0.000	0.001	0.009	0.005
Government Alignment:	Avg. % in Coal.	0.696	0.658	0.672	0.651	0.656	0.654
	Avg. % in Oppos.	0.304	0.342	0.328	0.349	0.344	0.346
Sample Size:							
	Person-year obs.	1,191	1,545	2,749	2,052	2,052	4,104
	Politicians	449	475	686	513	513	763
	Parties	20	26	28	27	32	32

tures (69% in the first period and 66% in the second period), indicating the attempts at majority building by presidents in power over the period considered.

2.3.3 Dispersion of public funds

While all federal legislators are entitled to the same amount in budgetary amendments every year, the actual execution of the funds is subject to approval by the executive power. The statistics reported in Table 2.2 show, for all federal legislators elected in 2010 and in 2014, the percentiles of public spending at different stages of the budgetary execution process. Each row contains the percentiles of the annual average values of different metrics computed for politicians and parties[14]. The data indicate that the average annual disbursements are significantly lower than the commitments, which in turn are significantly lower than the budgetary amendments approved at the first stage of the negotiations.

[14]Party averages are weighted by the number of politicians in each party.

Table 2.2: Percentiles of Public Funding - All Elected Politicians

		10%ile (1)	25%ile (2)	50%ile (3)	75%ile (4)	90%ile (5)	C.V. (6)
Group	Statistic						
Politicians	Approvals	5.13	6.71	9.11	11.39	13.56	0.37
	Commitments	2.75	4.27	5.97	7.37	9.10	0.42
	Disbursements	0.14	0.74	1.48	2.83	4.34	0.83
Parties	Approvals	6.51	7.38	8.88	9.94	10.54	0.28
	Commitments	1.69	4.09	5.65	6.49	6.83	0.46
	Disbursements	0.10	1.54	1.72	2.46	2.53	0.51

Notes: Percentiles of total annual values for individual politicians and of party annual average amount per member. Funding variables in real 2018 R$ millions. C.V. refers to coefficient of variation.

More importantly, the coefficients of variation (C.V.) reported in Column (6) show that the dispersion of the data increase at each stage of the budgetary execution. For the annual averages of individual politicians, the coefficient of variation of the amounts initially approved is 36.8%, increases to 42.2% at the stage of commitment, and finally reaches 83.3% at the stage of disbursement. The party averages follow a similar pattern, but between the approval and the disbursement stages the coefficient of variation increases from 27.9% to only 51.4%, suggesting that within-party differences explain a larger share of the total variance than differences between parties.

Table 2.3 reports the same statistics for the main sample, which shows the same patterns observed in the population of elected federal legislators: both politician and party average disbursements are significantly lower than the amounts committed and initially approved, and the dispersion also increases between key stages of the budgetary execution. The coefficient of variation for the individual averages increases less in the main sample between approvals and disbursements, from 35.4% to 63.7%. It is worth noting that the disbursement dispersion is expected to be lower in the main sample than in the population of elected politicians because I exclude the observations with disbursement equal to zero in order to fit a log-level model. However, while restricting the sample to positive outcomes limits the observed vari-

Table 2.3: Percentiles of Public Funding - Main Sample

		10%ile (1)	25%ile (2)	50%ile (3)	75%ile (4)	90%ile (5)	C.V. (6)
Group	Statistic						
Politicians	Approvals	5.49	6.99	8.81	11.10	13.21	0.35
	Commitments	3.76	5.61	7.19	8.62	10.12	0.38
	Disbursements	0.82	1.40	2.37	3.65	5.09	0.64
Parties	Approvals	6.57	7.59	8.85	9.74	10.33	0.26
	Commitments	4.95	6.29	7.10	7.72	8.64	0.31
	Disbursements	1.41	2.38	2.72	3.18	3.55	0.32

Notes: Percentiles of total annual values for individual politicians and of party annual average amount per member. Funding variables in real 2018 R$ millions. C.V. refers to coefficient of variation.

ation, the dispersion of disbursements still is significantly higher than than that of approvals and commitments in the main sample. For the party averages, on the other hand, commitments and disbursements have very similar coefficients of variation, which are only slightly higher than the relative dispersion of approvals. These results indicate that the between-party differences have a lower contribution to the total variance of disbursements in the main sample than in the population of elected federal legislators.

2.3.4 Between and within-party variance decomposition

I begin with a decomposition of the cross-sectional variance of log disbursements into measures of between-party and within-party dispersion. Let $y_{i,t}^{j(i)}$ denote log disbursement of politician i, member of party $j(i)$ in year t. As in Song et al. (2018), write

$$y_{i,t}^{j(i)} \equiv \bar{y}_t^j + \left[y_{i,t}^{j(i)} - \bar{y}_t^j \right], \tag{2.1}$$

where $\bar{y}_t^j$ is the average log disbursement of members of party j at time t. The total variance can be decomposed into between-group and within-group components, as shown below:

$$Var\left(y_{i,t}^{j(i)}\right) = \underbrace{Var\left(\bar{y}_t^j\right)}_{\text{Between-party dispersion}} + \underbrace{\sum_j \frac{n_j}{N} Var\left(y_{i,t}^{j(i)} \mid i \in j\right)}_{\text{Within-party dispersion}}. \tag{2.2}$$

For each year t, the first component is the between-party dispersion, measured by the variance of the party average log disbursement and the second term is a weighted sum of the within-party variance of payments to individual politicians. Figure 2.1 shows that in every year from 2011 to 2018 most of the variance in log disbursements was due to within-party differences, as the variance in party averages (between-party component) was low in all years, with a contribution to the total variance between 4.7% and 8.2%.

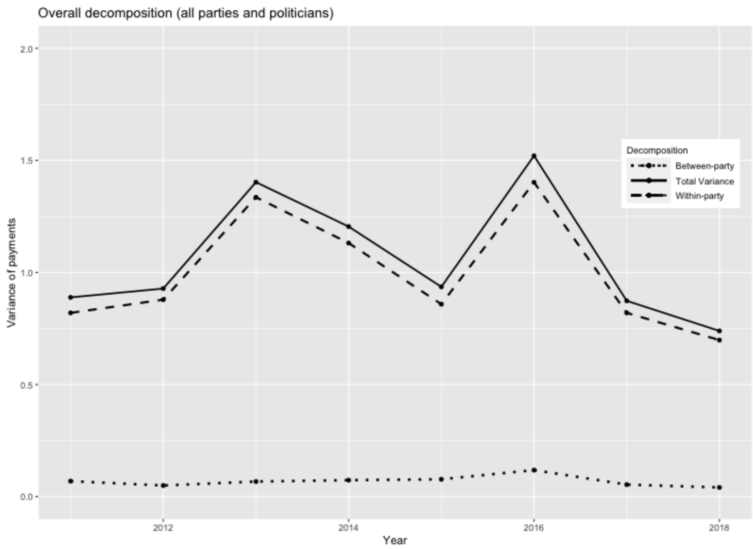


Figure 2.1: Between and within-party variance decomposition: all parties and politicians

The raw party averages between 2011 and 2018 shown in figure 2.2 further illustrates that the average log disbursement per federal deputy does not vary significantly across parties.

Another prominent feature of the variance composition plotted in Figure 2.1 is that, over time, the dispersion of the party average log disbursement is relatively stable, meaning that changes in the total variance are mainly driven by changes in the within-party component.

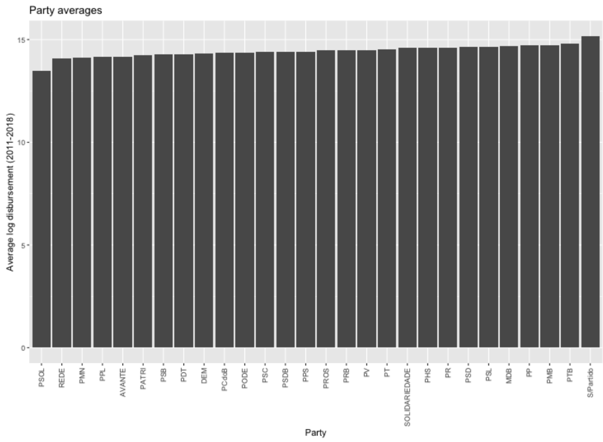


Figure 2.2: Raw party average log disbursements from 2011 and 2018

In Figures 2.4, 2.5, and 2.3 I plot the variance decomposition as above, but separately for politicians of different party types. Overall, the results indicate that in all years the between-party variance corresponds to a small share of the total variance for all subpopulations considered, but it is more relevant for the opposition and small parties.

More specifically, first I examine how the relative importance of the within and between-party variance components varies with the party's ideology. Figure 2.3 shows that the contribution of the between-party component is somewhat higher for left-wing than for center and right-wing parties. On average, the between-party component corresponded to 7.9% of the total variance across left-wing parties, 5% for center, and 4% for right-wing parties. However, it is worth noting that the relevance of between-party differences for left-wing parties is higher during the second term (2015-2018), when the leadership of the executive power was center-right for most of the time, while for right-wing parties the between-party component is relatively more relevant from 2011 to 2014, when the leadership of the executive power was left-wing.

Such dynamics suggests that during the period considered the differences by ideology might be related to membership in the government coalition rather than ideological differences. In Figure 2.4 I examine how the relative importance of the within and between-party variance components varies with the party's coalition status.

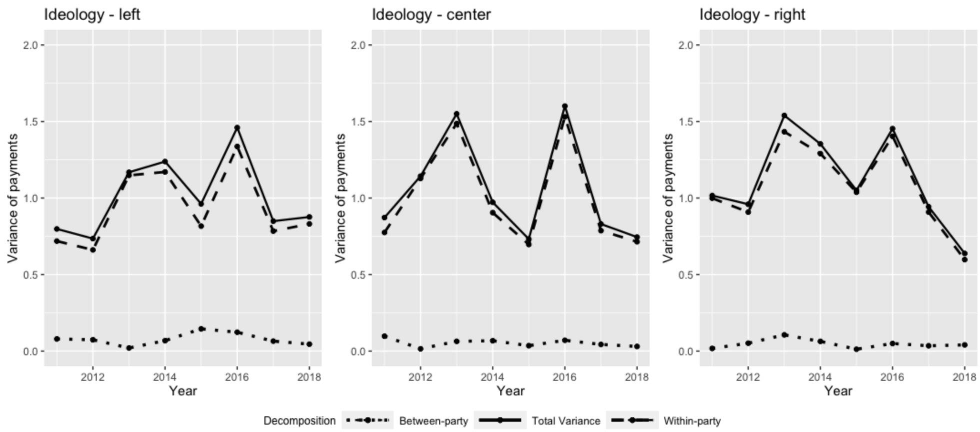


Figure 2.3: Between and within-party variance decomposition, by ideology

While in both cases the contribution of the between-party component is small, relative to the within-party variance the dispersion of the party average disbursements is higher among parties of the opposition. In other words, while the party averages of the parties in the government coalition were very similar, contributing to, on average, 3.7% of the total variance, across the opposition parties the dispersion of the average disbursements was higher and it contributed to, on average, 12.3% of the total variance, ranging from 7% to 17% between 2011 and 2018.

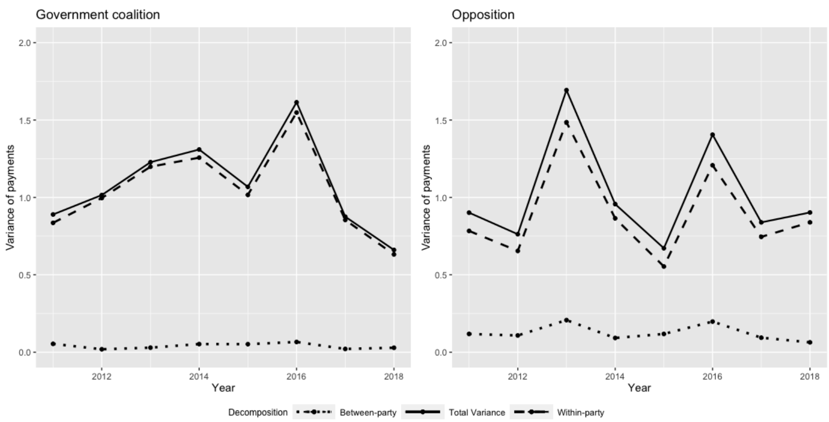


Figure 2.4: Between and within-party variance decomposition, by coalition status

Finally, a similar pattern is shown in Figure 2.5 with the variance decomposition by party size: across large parties the average disbursements are very similar, so the between- component contributed to, on average, only 2% of the total variance. Nevertheless, the dispersion of the party averages was higher in small parties, with an average contribution of 8.8% to the total variance.

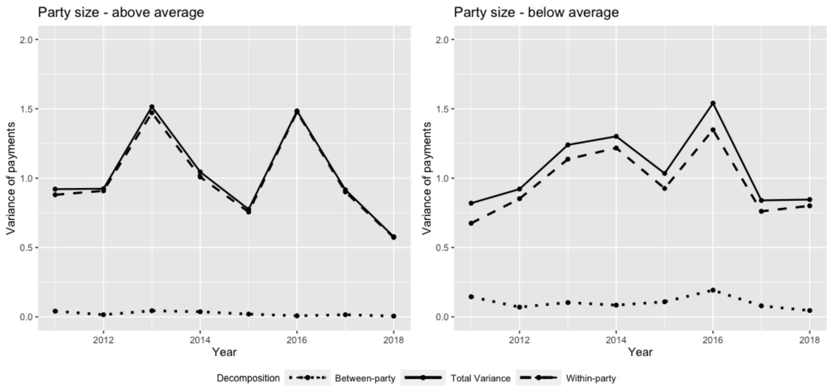


Figure 2.5: Between and within-party variance decomposition, by number of seats

In sum, while the variance decompositions by subpopulations indicate that the relative importance of between-party inequality is higher for certain groups (opposition and small parties), overall, regardless of ideology, size, or participation in the government coalition the within-party variance is by far the most relevant variance component for all parties.

2.4 Modelling party and individual influence

The decomposition above indicates that inequality in disbursement arises mainly from within-party differences. In order to better understand the mechanisms behind the low dispersion of party averages and the large differences within parties I estimate a two-way fixed effects model with party and politician fixed effects and use the model estimates to decompose the variance of log disbursements into components that capture the direct contribution of party and individual characteristics to the variance of log disbursements. The following sections introduce the econometric methods, present its main assumptions, and discuss its application in a political market for public funding.

2.4.1 Model specification

To further decompose the variance of log disbursements into components attributable to party and individual heterogeneity, I proceed in two steps: first, I estimate a two-way fixed effects model with party and politician fixed effects for each

Legislature. Second, I use the model estimates to decompose the total variance of log disbursements during each term.

Specifically, I estimate the following regression, for each term p, in step 1 :

$$y_t^{i,j} = \theta^{i,p} + \psi^{j,p} + X_t^i \beta^p + \varepsilon_t^{i,j}, \tag{2.3}$$

where $y_t^{i,j}$ is log real disbursements for politician $i(i = 1,...,N)$ affiliated to party $j(j = 1,...J)$ at time $t(t = 1,...T_i)$, θ^i is a politician fixed effect, ψ^j is the fixed effect of party j of which politician i is a member, and X_t^i is a set of time-varying controls and year dummies. The effects of the time-varying observable characteristics included in X_t^i is captured by β and any transitory fluctuations in funding is captured by $\varepsilon_t^{i,j}$. There are T_i observations for each politician i and a total of $N^{p*} = \sum_{i=1}^{N} T_i$ person-year observations in each term. Years are indexed by t, ranging from 2011 to 2018, with Term 1 defined as the period between 2011 and 2014 and Term 2 as the period between 2015 and 2018.

The time-varying observables (X_t^i) include a set of year dummies, a dummy variable (*End Term*) equal to one in pre-election years and equal to zero at the beginning of the Legislature, a variable (*Governism*) capturing support to the Government's legislative agenda by politician i, and an interaction term between *Governism* and *End Term* to allow for different returns to government support during pre-election times[15]. More specifically, *pre-elec* is equal to zero during the first two years of each Legislature and is equal to one in the last two years, while *governism* is defined as the share of the open votes in the National Congress in which the federal legislator voted in line with the Government recommendation.

Two-way fixed effects models like the above have been widely used to decompose wage inequality into components related to firm heterogeneity, worker heterogeneity, observable employee characteristics and residual variation. Since the method was first proposed in Abowd et al. (1999) (henceforth, AKM), it has been

[15]Brollo and Nannicini (2012) analyse discretionary transfers from line ministries to Brazilian municipalities and they find that in the two years before municipal elections the amount of transfers to municipalities with mayor-president partisan alignment increases, while the transfers to mayors from the opposition decline. The authors do not find the same effect in the first two years of the municipal mandate.

further developed (e.g, Abowd et al. (2002), Upward (2004), Cornelissen (2008), and Kline et al. (2020)) and used in the study of trends in wage inequality in various contexts (e.g., Song et al. (2018), Card et al. (2013), and Card et al. (2015)).

This paper employs the same statistical framework to study the role of parties and politicians in explaining inequalities in access to public funds. In particular, the model in Equation 2.3 includes additive party ($\psi^{j,p}$) and politician ($\theta^{i,p}$) fixed effects that I estimate and then use to decompose the variance of log disbursements.

The party fixed effect $\psi^{j,p}$ captures a proportional, time-constant party effect on the execution of IBAs that is common to all members of the party. Within the context discussed in Section 2.2, the party fixed effect will capture the relative, average performance of the party leadership in negotiating the execution of IBAs for all of its members during each term. It is worth noting that the relative performance captured by ψ^{j} may reflect both the party preferences for collective bargaining and its ability to negotiate with the Government. Moreover, $\psi^{j,p}$ may also capture the effect of other time constant party characteristics such as party-level preferences over spending and over within-party inequality in access to funding.

The politician fixed effect $\theta^{i,p}$, on the other hand, captures the component of a politician's funding outcomes that is fully portable across parties, such as the effect of a politician's direct connection with the Government, as discussed in Section 2.2. Other time constant factors such as individual preferences over public spending and discrimination are also captured in $\theta^{i,p}$. Contrary to labour market applications, the correlation between the true parameters $\psi^{j,p}$ and $\theta^{i,p}$ may be negative in this political market. Firstly, because parties with high party fixed effects ($\psi^{j,p}$) may still be interested in politicians with low individual influence ($\theta^{i,p}$) over the distribution of IBAs. That would be the case if such individuals had large voter bases or high influence in setting the agenda in Congress, for example. More importantly, parties with a strong preference for collective bargaining could discourage direct, individual negotiations with the Government. Finally, individuals more likely to be discriminated or that have weaker direct links to the Government may also select into parties with higher fixed effects.

In step 2, I follow Song et al. (2018) and I rewrite the standard variance decomposition presented above as follows (ignoring the covariate index $X_t^i\hat{\beta}$ for simplicity):

$$\begin{aligned} var(y_t^{i,j}) = &\underbrace{var(\hat{\psi}^j) + 2cov(\hat{\psi}^j, \bar{\theta}^i) + var(\bar{\theta}^i)}_{\text{Between-party component}} \\ &+ \underbrace{var(\theta^i - \bar{\theta}^j) + var(\hat{\varepsilon}_t^{i,j})}_{\text{Within-party component}}, \end{aligned} \tag{2.4}$$

where the between-party variance is composed by the variance of party effects ($var(\hat{\psi}^j)$), the covariance between politican and party effects ($2cov(\psi^j, \bar{\theta}^j)$), the variance of the average politician effects in each party ($var(\bar{\theta}^j)$), and the terms related to the party-average of the time-varying observables (($var(\bar{X}b^j)$, $2cov(\psi^j, \bar{X}b^j)$, and $2cov(\bar{\theta}^j, \bar{X}b^j)$).

With this detailed variance decomposition the role of party composition can be considered in terms of both sorting and segregation. The sorting component is measured by the covariance between party premiums and the average politician effect in each party and captures the extent to which politicians with high/low individual pay premiums are members of parties with high/low collective pay premiums. Segregation is measured by the variance of the average politician fixed effects, capturing whether politicians with higher individual influence are likely to be members of different parties than whose with lower individual pay premiums.

2.4.2 Estimation and identification

Estimation and identification of the full model has been extensively discussed in the literature since Abowd et al. (1999) first proposed consistent estimators for the parameters in Equation 2.3.

In order to discuss the estimation and identification of the party and politician fixed effects, I rewrite Equation 2.3 in matrix form:

$$Y = D\boldsymbol{\theta} + F\boldsymbol{\psi} + X\boldsymbol{\beta} + \boldsymbol{\varepsilon} \tag{2.5}$$

where Y is a $(N^* \times 1)$ vector of log disbursements, D is $(N^* \times N)$ design matrix of politician effects, $\boldsymbol{\theta}$ is a $(N \times 1)$ vector of politician effects, F is $(N^* \times J)$ design matrix for party effects, $\boldsymbol{\psi}$ is a $(N \times 1)$ vector of party effects, X is a $(N^* \times k)$ matrix of observable time-varying politician characteristics and time dummies, $\boldsymbol{\beta}$ is a $(k \times 1)$ vector of coefficients of the covariates in X, and $\boldsymbol{\varepsilon}$ is a $(N^* \times 1)$ vector of disturbances.

As discussed in Abowd et al. (2002), the least-square solution to the estimation problem solves the normal equations for all estimable effects:

$$\begin{bmatrix} D'X & D'D & D'F \\ F'X & F'D & F'F \\ X'X & X'D & X'F \end{bmatrix} \begin{bmatrix} \boldsymbol{\theta} \\ \boldsymbol{\psi} \\ \boldsymbol{\beta} \end{bmatrix} = \begin{bmatrix} D'Y \\ F'Y \\ X'Y \end{bmatrix} \tag{2.6}$$

Abowd et al. (1999) and Abowd et al. (2002) developed statistical approximations and exact methods to solve the normal equations above in high-dimension applications[16]. In my application the number of politicians and parties is sufficiently low to allow for the estimation of the full model by fixed-effect methods implemented in general purpose software based on the sweep algorithm.

However, the issues concerning the identification of the unobserved party effects still apply. Firstly, party effects ψ^j are identified solely from the politicians who switch parties (movers). This point is clear considering that the solution to the normal equations 2.6 is the same to the solution to a transformed problem that includes dummy variables for the party heterogeneity but sweeps out the politician heterogeneity algebraically by time-deamening the variables (Upward et al., 2005):

$$y_t^{i,j} - \overline{y^i} = \sum_{j=1}^{J} \psi^j (F_t^{i,j} - \overline{F}^{i,j}) + (X_t^i - \overline{X}^i)\beta + \varepsilon_t^{i,j}, \tag{2.7}$$

,

where, for each $j = 1, ..., J$, $F_t^{i,j}$ is an indicator equal to 1 if individual i is a member of party j at time t, and $\overline{z}^i = \sum_t (z_{i,t}/T_i)$ for any variable z. From Equation

[16]Such as those using linked employed-employee data on millions of workers to estimate very high dimension vectors of firm and person effects.

2.7 it is clear that for politicians who do not switch parties $(F_t^{i,j} - \overline{F}^{i,j}) = 0$ for all J dummies. So the identification of ψ^j is driven by the number of movers in each firm j, meaning that ψ^j cannot be identified for parties without movers. Additionally, since

$$\sum_{j=1}^{J} (F_t^{i,j} - \overline{F}^{i,j}) = 0$$

one of the party dummies is dropped since they form a collinear set of variables. This is a standard procedure, but it has implications for an additional issue regarding the identification of the party effects (Upward et al., 2005).

As shown in Abowd et al. (2002), the solution to the identification problem for the party (and politician) fixed effects applies graph theory methods to determine groups of connected parties and politicians within which identification can be determined using conventional methods. The connected sets are groups of individuals and parties containing (i) all politicians who have ever been a member of any of the parties in the group, and (ii) all the parties of which any of the politicians was ever a member and can be constructed by the algorithm[17] presented in Abowd et al. (2002). Within each connected set g, $N_g - 1 + J_g - 1$ politician and party effects are identified, where N_g is the number of individuals and $J_g - 1$ is the number of parties in group g.

As further argued in Abowd et al. (2002), given the G connected sets, the normal equations in 2.6 can be rearranged so that the party and politician effects of each group g are placed in ascending order in the design matrix. Setting the reference party and person effects of each group to zero, the resulting normal equations are shown below:

[17]The algorithm assigns the first party to group 1 then proceeds in two steps: first it adds to group 1 all politicians who have ever been a member of the first party. In the second step all parties in which the politicians in group 1 have ever been a member of are added to group 1. Steps 1 and 2 are repeated until no additional parties or politicians are added to group 1. If there are any remaining parties, the algorithm chooses one and starts the same procedure to construct group 2. This procedure continues until no parties remain.

$$\left[\begin{array}{c|cc|cc|c|cc}
X'X & X'D_1 & X'F_1 & X'D_2 & X'F_2 & \ldots & X'D_G & X'F_G \\
\hline
D_1'X & D_1'D_1 & D_1'F_1 & 0 & 0 & \ldots & 0 & 0 \\
F_1'X & F_1'D_1 & F_1'F_1 & 0 & 0 & \ldots & 0 & 0 \\
\hline
D_2'X & 0 & 0 & D_2'D_2 & D_2'F_2 & \ldots & 0 & 0 \\
F_2'X & 0 & 0 & F_2'D_2 & F_2'F_2 & \ldots & 0 & 0 \\
\hline
\ldots & \ldots & \ldots & \ldots & \ldots & \ldots & \ldots & \ldots \\
\hline
D_G'X & 0 & 0 & 0 & 0 & \ldots & D_G'D_G & D_G'F_G \\
F_G'X & 0 & 0 & 0 & 0 & \ldots & D_G'D_G & D_G'F_G
\end{array}\right]
\left[\begin{array}{c}
\boldsymbol{\beta} \\
\hline
\boldsymbol{\theta}_1 \\
\boldsymbol{\psi}_1 \\
\hline
\boldsymbol{\theta}_2 \\
\boldsymbol{\psi}_2 \\
\hline
\ldots \\
\hline
\boldsymbol{\theta}_G \\
\boldsymbol{\psi}_G
\end{array}\right]
=
\left[\begin{array}{c}
X'Y \\
\hline
D_1'Y \\
F_1'Y \\
\hline
D_2'Y \\
F_2'Y \\
\hline
\ldots \\
\hline
D_G'Y \\
F_G'Y
\end{array}\right] \tag{2.8}$$

That is, if blocked by the connected sets, the normal equations have a diagonal sub-matrix, so the unique solution to the parameter vector is identified within each group. Moreover, if there are G different connected groups in the sample, within each group the party (and person) effects are identified only up to a constant, as discussed above, as G different parties are dropped. More importantly, because the reference party and individual are chosen arbitrarily, the estimates $\hat{\psi}^j$ and $\hat{\theta}^i$ cannot be compared across connected sets (Upward et al., 2005). For this reason, the applications of the model 2.3 typically focus on the largest connected set of individuals and units within their sample or population and I follow the same approach.

Finally, the model in Equation 2.3 is estimated by fixed effects under the key orthogonality assumption $E[F'\varepsilon] = 0$. A sufficient condition is that mobility patterns are independent of ε, known as the exogenous mobility assumption.

Desposato (2006) presents a model of party-membership patterns and estimates the role of different factors behind party mobility of federal legislators. The author finds that federal legislators change parties for ideological consistency, to participate in more advantageous party lists and maximize their probability of re-election, and to join the government coalition, which is interpreted as proxy to access to funding from the Executive Power. Party switching into the government coalition could be a source of endogeneity, but switching into the coalition tends to take place within the first year of the Legislature and I do not use variation from within-year mobility, my estimates are not driven by this particular type of mobility.

2.4.3 Largest connected sets

I use software[18] from Cornelissen (2008) and Kline et al. (2020) to construct the largest connected set for each period. Mobility of politicians across parties is high, so despite the relatively small sample size, estimation on the largest connected sets is feasible and includes most parties represented in Congress: only eight politicians, from five parties, are dropped from the analysis in the first term (2011-2014), and all politicians and parties are connected in the 2015-2018 period (Table 2.4).

Table 2.4: Overall Sample and Largest Connected Set

	Overall Analysis Sample			Largest Connected Set		
Interval	Person-. year obs.	Parties	Politicians	Person- year obs.	Parties	Politicians
2011-2018	2,750	29	687	2,749	28	686
2011-2014	1,205	24	457	1,191	20	449
2015-2018	1,545	26	475	1,545	26	475

Moreover, overall there are no systematic differences in observables between movers and stayers (Table 2.5). The average levels of IBA disbursements, age, education, experience, and access to campaign funding are similar for movers and stayers. The exception is that, in Term 1, movers were more likely to be reelected incumbents.

[18] Both the Stata command *felsdvreg* from Cornelissen (2008) or the Matlab software from Kline et al. (2020) can be used to compute the Largest Connected Set within the overall sample.

Table 2.5: Characteristics of movers and stayers

Panel (a): 2011-2014				
	Movers	Stayers	Movers-Stayers	P-value
Log disbursement	14.14	14.27	-0.13	0.348
Age	56.63	54.21	2.41	0.192
Has at least college degree	0.76	0.77	-53.44	0.903
Times elected	3.00	2.80	0.19	0.167
Won previous election	0.95	0.74	0.20	0.000
Campaign funding (R$ million)	3.361	3.196	0.164	0.623
Observations	43	406	449	
Panel (b): 2015-2018				
	Movers	Stayers	Movers-Stayers	P-value
Log disbursement	14.525	14.551	-0.026	0.793
Age	53.353	54.012	-0.658	0.637
Has at least college degree	0.742	0.815	-53.269	0.140
Times elected	2.856	3.132	-0.277	0.091
Won previous election	0.701	0.735	-0.034	0.509
Campaign funding (R$ million)	4.592	4.622	-0.030	0.907
Observations	97	378	475	

2.5 Results

2.5.1 Detailed variance decomposition

Table 2.6 presents the detailed decomposition of funding inequalities, by Legislature, based on the estimates from Equation 2.3. One limitation of splitting the data by term is the sample size. Another issue is the reduced mobility, since with this specification it is not possible to exploit the party switches of reelected federal legislators who switched parties between terms. Still, as shown in Table 2.6, for the first term (2011-2014) it is possible to identify 20 party fixed effects, while in the second period (2105-2018) mobility is higher (97 movers versus 43 movers in the first period) and 26 party fixed effects can be estimated.

Before discussing the changes over time, it is worth noting that the main findings from the basic decomposition for the whole period are observed in both terms. Firstly, the person-related effects were the main drivers of the dispersion of log disbursements in both Legislatures, with the negative covariance between the individual fixed effects offsetting only a small fraction of the dispersion driven by the large differences in unobserved individual heterogeneity and the time-varying covariates.

The dynamic analysis shows that despite the peak observed in 2016 (see Figure 2.1), the total variance of log disbursements was lower in the second period, which is consistent with the institutional reforms implemented in 2015 to limit the discretion of the executive power in approving the disbursement of budgetary amendments. The main driver of such decline were the person-related components, which decreased more than proportionally to the change in the total variance. In particular, both the variance of the politician fixed effects and of the linear index decreased between the first and second terms, and while the contribution of the negative covariance between observable behaviour and unobservable heterogeneity slightly increased, the overall contribution of the person-related components was smaller in the second than in the first term.

In both terms, when converted to correlations the estimates are indicative of moderate negative sorting. As shown in the first column of Table 2.8, between 2011 and 2014 the correlation between-party and person effects was -0.61 and the negative correlation between-party effects and the covariance index was close to zero (-0.009). During the second Legislature, the correlation between party and person fixed effect was -0.52, while the correlation between the party effects and the covariate index was close to zero as well (-0.04).

2.5.2 Party ideology and negative sorting

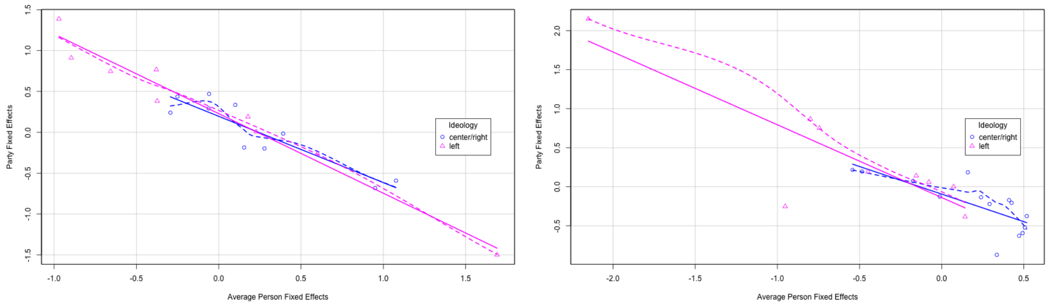
One question that follows is which parties have higher collective bargaining components and which parties high-influence politicians are more likely to select into.

Figure 2.6 suggests that the pattern of negative assortative matching discussed above is correlated with ideology. Panel (a) shows that left-wing parties have a clear advantage in party premiums during the first period, an advantage that persists during the second period, as suggested in panel (b).

Table 2.6: Detailed Variance Decomposition - by Legislature

		2011-2015 Components	2016-2018 Components	Change Components
Total variance	Var (Log Disbursement)	1.265	1.050	-0.215
Between-party	Var (Log Disbursement)	0.035	0.053	0.019
	Var (Average Pol. effect)	0.475	0.324	-0.151
	Var (Party effect)	0.444	0.285	-0.160
	Var (Average Xb)	0.010	0.000	-0.010
	2*Cov (Average Pol. effect, Party effect)	-0.885	-0.555	0.329
	2*Cov (Average Pol. effect, Average Xb)	-0.004	0.006	0.010
	2*Cov (Party effect, Average Xb)	-0.005	-0.006	-0.001
Within-party	Var (Deviation Log Disbursement)	1.231	0.997	-0.234
	Var (Deviation Pol. effect)	0.699	0.646	-0.053
	Var (Deviation Xb)	0.223	0.017	-0.207
	Var (Residual)	0.433	0.337	-0.096
	2*Cov (Deviation Pol. effect, Deviation Xb)	-0.125	-0.003	0.122
	2*Cov (Deviation Pol. effect, Residual)	-0.000	0.000	0.000
	2*Cov (Deviation Xb, Residual)	0.000	0.000	0.000

Notes: Party FE denotes Party fixed effects, Person FE is person fixed effects, and Xb denotes the covariate index. Decomposition of total variance (over all person-year observations). In each interval, party averages are calculated over the whole period.



(a) 2011-2014 **(b)** 2015-2018

Figure 2.6: Party Composition: Segregation and Negative Sorting, by Ideology and Legislature

The overall advantage in party fixed effects of left wing parties could reflect ideological preferences for collective bargaining and for equality in funding outcomes. The results could also reflect the alignment between left wing parties and the Government, since between 2011 and May 2016 the Government leader (President) was from the Labour Party. However, I cannot separately identify the difference components of the party fixed effects.

Table 2.7 shows the difference in group means, by party ideology, of the party and politician fixed effects in each term. Overall, in both periods the estimated party fixed effects are higher for left wing parties. These differences are substantial and statistically significant.

The average politician fixed effects, on the other hand, are higher in center and right wing parties. The results are consistent with center and right wing parties favouring individual bargaining rather than collective negotiations. Moreover, individual politicians from centrist and right wing parties may also have been more likely to negotiate individually with the Government as a means to overcome their party's weaker alignment with the Government's party. The relative contribution of each of the above components, nevertheless, cannot be separately identified.

In sum, I find evidence that negative assortativeness is associated to party ideology, with politicians with low individual influence being more likely to belong to high influence parties, which in turn are more likely to be left wing.

Table 2.7: Party Ideology Differences in Means, by Legislature and Pooled Sample

	Left-Right (1)	Left-Center (2)	Right-Center (3)
Panel (a): 2011-2014			
Log disbursement	-0.080	-0.055	0.025
Party Fixed Effect	0.678***	0.978***	0.300***
Person Fixed Effect	-0.707***	-0.971***	-0.263**
Covariate index (Xb)	-0.048	-0.064	-0.016
Panel (b): 2015-2018			
Log disbursement	-0.153	-0.052	0.101
Party Fixed Effect	0.598***	0.583***	-0.015
Person Fixed Effect	-0.740***	-0.622***	0.119
Covariate index (Xb)	-0.022***	-0.015**	0.007
Panel (c): 2011-2018			
Log disbursement	-0.086	-0.030	0.056
Party Fixed Effect	0.685***	0.848***	0.163***
Person Fixed Effect	-0.714***	-0.796***	-0.082
Covariate index (Xb)	-0.061	-0.085**	-0.024

Notes: Xb denotes the covariate index. Results from Tukey Honest Significant Differences test: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

2.5.3 Validity checks

2.5.3.1 Additivity assumption

As discussed above, a sufficient condition for identification is that mobility patterns are independent of the transitory component $\varepsilon_t^{i,j}$. Let $m^{i,j}$ denote a party-politician match component of $\varepsilon_t^{i,j}$ with mean zero for every politician and every party. Before considering potential sources of endogenous mobility, it is worth discussing the interpretation of the match heterogeneity in the model. As in Card et al. (2013), the match effect $m^{i,j}$ allows for a time-invariant disbursement premium or discount for politician i at party j, in relation to the baseline fixed-effects $\theta^i + \psi^j$. In the political market being considered, complimentary between the skills of the politician and party characteristics could be a source of match heterogeneity. For example, a politician with strong bargaining skills could perform better in terms of disbursements if matched with a party with a stronger relationship with the executive power.

If party switches are systematically related to such match effects, bias could arise. In order to test for match-based sorting, I follow Card et al. (2013) and consider a fully saturated match-effects model. In particular, I include interaction dummies for each party-politician match and compare the fit of the match model to the baseline AKM model. If match effects are relevant, the unrestricted match model should have a significantly better fit than the baseline model.

Table 2.8 shows the estimation results for the baseline and match models in both terms being considered and for the whole period. Firstly, after allowing for match effects the increase in the adjusted R-squared are negligible. The root mean square error (RMSE) from the baseline AKM model is slightly higher than the RMSE from the match model, suggesting that the later model fits the data slightly better than the baseline. However, the reduction in RMSE is very small is all time intervals (1.8% in the first term, 4.6% in the second term and 3.5% in the whole period) and the magnitude of the gap in fit between the baseline and the match models changes only slightly from the first to the second term. Consistently with the hypothesis of random match effects, the improvement in fit in relation to the baseline model is lower in the first period, when the dispersion of party and person fixed effects is higher, than in the second period. Similarly, the standard deviation of the match effects is low and has increased between the first and second term, while the dispersion of log disbursements and of most of its variance components has decreased.

Table 2.8: Estimation Results for AKM Model, by period

	54^{th} Leg. 2011-2014 (1)	55^{th} Leg. 2015-2018 (2)	Pooled 2011-2018 (3)
Person and Party Parameters:			
Number of party effects	20	26	28
Number of person effects	449	475	686
Movers	43	97	132
Summary of Parameter Estimates:			
SD of person effects	1.084	0.985	0.982
SD of party effects	0.666	0.533	0.558
SD of Xb	0.483	0.130	0.498
Correlation of person/party effects	-0.612	-0.528	-0.551
Correlation person effects/Xb	-0.124	0.012	-0.191
Correlation party effects/Xb	-0.009	-0.046	-0.043
RMSE of AKM	0.658	0.581	0.682
R-squared	0.657	0.679	0.604
Adjusted R-squared	0.430	0.522	0.462
Comparison Match Model:			
RMSE match model	0.646	0.554	0.658
R-squared	0.670	0.708	0.631
Adjusted R-squared	0.428	0.526	0.464
SD of Match Effect	0.124	0.174	0.181
Additional information:			
SD of log disbursement	1.125	1.025	1.084
Person-year observations	1,191	1,545	2,749

In sum, allowing for match effects does not significantly improve the fit of the model in any of the time intervals considered and the dynamics of the estimated match effects is consistent with the assumption that the match effects are random.

2.5.3.2 Pooled sample

Since splitting the data by term reduces the overall mobility, in Table 2.9 I present the results of the detailed variance decomposition in Equation 2.4 for the pooled sample.

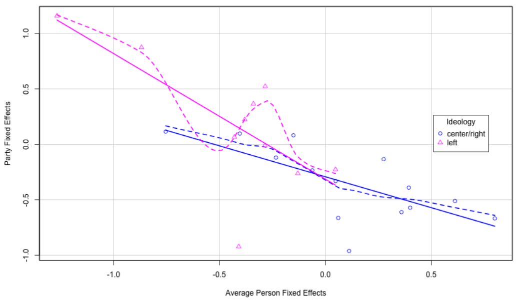
Table 2.9: Detailed Variance Decomposition (2011-2018)

		2011-2018
Total variance	Var (Log Disbursement)	1.175
Between-party	Var (Log Disbursement)	0.036
	Var (Average Pol. effect)	0.347
	Var (Party effect)	0.311
	Var (Average Xb)	0.008
	2*Cov (Average Pol. effect, Party effect)	-0.604
	2*Cov (Average Pol. effect, Average Xb)	-0.002
	2*Cov (Party effect, Average Xb)	-0.024
Within-party	Var (Deviation Log Disbursement)	1.139
	Var (Deviation Pol. effect)	0.617
	Var (Deviation Xb)	0.240
	Var (Residual)	0.466
	2*Cov (Deviation Pol. effect, Deviation Xb)	-0.184
	2*Cov (Deviation Pol. effect, Residual)	-0.000
	2*Cov (Deviation Xb, Residual)	0.000

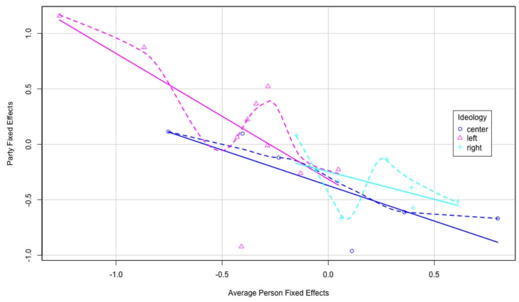
Notes: Party FE denotes party fixed effects, Person FE is person fixed effects, and Xb denotes the covariate index. Decomposition of total variance (over all person-year observations).

As previously discussed, the dispersion of party average log disbursements is low, and the between-party component contributed to just over 3% of the total variance. In line with the analysis by, the detailed decomposition indicates that such low contribution of the between-party differences to the overall dispersion is essentially a result of negative sorting, as politicians with low (high) individual pay premiums tend to segregate in parties with high (low) collective pay premiums. More specifically, the variance of the party premiums (0.31) and of the average politician fixed effect (0.35) add up to about 56% of the total variance, but with the negative sorting component (-0.60), the final composition is such that differences between party averages are minimum and the between-party component is 0.036.

Therefore, both the polled baseline model and the analysis by legislature suggest that the contribution of party fixed effects to the total variance of log disbursement is low compared to the influence of individual characteristics, but that moderate negative sorting contributes to reduce the variation in access to funding.



(a) Left x Center or Right



(b) Left x Center x Right

Figure 2.7: Party Composition: Segregation and Negative sorting by Ideology

Similarly to the findings from the dynamics analysis, Figure 2.7 illustrates the relationship between negative sorting and ideology. Panel (a) plots the party average individual fixed effects against the estimated party premiums of left-wing and other (center or right) parties and shows that high-influence individuals are more likely to segregate in center and right-wing parties, which, in turn, tend to have lower party premiums. Panel (b) replicates the same exercise but tracking center and right parties separately. The same general pattern emerges as expected, but showing a narrow intersection between the range of party and person fixed effects of right and left-wing parties, while the parties in the center have a range of party and person fixed effects largely overlapping with the intervals of estimated effects for both left and right-wing parties.

2.5.3.3 Limited mobility bias correction

One well known drawback of the AKM model is the possibility of limited mobility bias arising as a large number of unit-specific parameters are identified only from individuals who move across units (Bonhomme et al., 2020). Bias correction methods have been developed (Andrews et al. (2008), Kline et al. (2020)) and Table 2.10 presents the baseline AKM and the bias corrected leave-out estimates of the variance of person and party effects proposed in Kline et al. (2020), as well as of the covariance between party and person fixed effects.

The bottom section of Table 2.10 shows that, in relation to the largest connected set, leave-out pruning decreases the sample size for the 2011-2018 period by less than 3% and the number of identified party fixed effects from 28 to 25. Similarly, only three movers from the largest connected set are excluded from the leave-out sample. The variance and the mean of log disbursements are nearly the same in both samples. In other words, mobility in this political market is such that the leave-out sample is very similar to the largest connected set.

Table 2.10: Variance Decomposition and Bias Correction (2011-2018)

	AKM (1)	Plug-in (2)	Leave Out (3)
Variance of log Disbursements	1.175	1.152	1.152
Party and Person Parameters:			
Var(Person FE)	0.965	0.946	0.776
Var(Party FE)	0.311	0.335	0.326
2*Cov(Person FE,Party FE)	-0.604	-0.655	-0.636
Correlation between Party and Person Effects	-0.551	-0.582	-0.633
Additional Information			
Person-year observations	2,749	2,674	2,674
Movers	132	129	129
Party effects	28	25	25

Column (1) presents the baseline AKM estimates on the largest connected set discussed and shown in Table 2.9 above. Column (2) shows the naive plug-in estimates (AKM estimator on the leave-out connected sample) and Column (3) shows the bias corrected leave-out estimates from Kline et al. (2020). Overall, the bias corrected estimates and the baseline model are only slightly different and lead to the same conclusions. In particular, compared to the baseline AKM model the leave-out estimator in Column (3) yields only slightly higher estimates for the variance of party effects. The estimated covariance between party and person fixed effects is slightly lower than the baseline AKM estimator. The estimated variance of the person effects decreases from 0.965 to 0.776 after bias correction, which does not change the main conclusion from the baseline model that individual characteristics are the main drivers of the observed inequality in disbursements. Finally, the negative correlation between person and party fixed effects decreases from -0.55 in the baseline AKM model to -0.63, so the results based on both estimators are suggestive of a negative moderate correlation between individual and party fixed effects.

In sum, the bias corrected estimator corroborates the conclusions from the baseline AKM estimator. Therefore, given the sample size limitations in this application the main results and analysis by subpopulations and time intervals in the following sections are based on the standard AKM estimator using all observations in the largest connected set.

2.6 Concluding Remarks

Inequalities in access to public funds for the provision of public goods can have relevant consequences for the career progression of politicians in office. This paper studies the role of parties in explaining the variance in disbursements of funds by federal legislators in Brazil. I find that the differences in individual heterogeneity are the main driver of the observed inequalities and that parties have a distributive contribution thorough a composition effect as politicians with lower individual influence sort into parties with higher funding premiums. I also find evidence that this sorting pattern is associated to party ideology.

Future work will address some remaining concerns. For example, individuals who did not disburse any positive amount in a given year were not included in the analysis, so extending the model to accommodate the occurrence of zero-valued outcomes is part of a future research agenda. Additionally, the analysis is restricted to separate models for the estimation of time-constant and term-specific party heterogeneity. Future work will include extending the model to allow for the joint estimation of time-constant and term-specific heterogeneity as to fully exploit politician mobility in the sample while considering the requirements for the identification of comparable parameters.

Finally, the framework presented in this chapter could be used in the study of inequalities in other relevant outcomes and in various settings in which party switching during term is frequent.

Chapter 3

Gender Gaps in Parliament: Access to Public Funds by Legislators in Brazil

3.1 Introduction

Women's Representation in national parliaments has been increasing over the last 30 years, but at a slow pace. What are the factors explaining such slow progression? How do these factors change over time?

Various channels through which women may be disadvantaged in a political career have been studied in the literature. I show a new channel: gender differences in access to a type of funding that can improve reelection outcomes of politicians in office. Building on the results from Chapter 2, this chapter (i) documents gender gaps in access to IBAs and (ii) studies the role of partisan and individual influence in explaining these gaps. Further, I quantify how the contribution of these sources change over time.

Interestingly, I find that at the beginning of the budgetary process men and women have similar average outcomes, but that after a competitive process that involves project cancellations and delays, women disburse lower amounts on average. I document substantial and statistically significant gender gaps at the stage of project execution between 2011 and 2018, and a large increase from term 1 (2011-2014) to term 2 (2015-2018).

As previously discussed, the actual execution of Budgetary Amendments may involve individual and party-level effort and in Chapter 2 I show that, in general, selection into parties may partially offsets inequalities in access to funding by legislators. What are the implications of these findings to the outcomes of women?

In line with the broader pattern of negative assortative matching discussed in Chapter 2, I find that women tend to have a disadvantage in individual influence, but are more likely to be in parties that provide higher levels of support to its members.

Specifically, I combine the results from Chapter 2 with the method in Cardoso et al. (2016) and Gelbach (2016) to decompose the gender gaps into components attributable to differences in party and individual heterogeneity. The decomposition shows that the allocation of women into parties contributed to 25% of the adjusted gap (-13.8 log points) in term 1. But the most interesting results are from term 2, when the gap is significantly larger (-29.8 log points). This difference is mainly driven by differences in individual influence of men and women (-52.2 log points), but allocation of women into parties partially offsets this gap, as women have a large advantage of 22.4 log points in the party component.

This pattern of negative assortative matching is correlated with ideology. Women are more likely to be members of left-wing parties, which in turn have an advantage in party premiums. Therefore, ideology-based sorting is a mechanism behind the role of parties in partially compensating for the gender gaps in individual influence. In particular, I find that women in all parties lost individual influence between the first and the second period, but the the disadvantages in individual influence of women in left wing parties contributed the most (65%). On the other hand, it is mainly support from left wing parties that explains the positive party component of the gender gap in term 2, at least partially compensating for the loss of individual influence of their female members.

Overall, in a context of deterioration of the influence and performance of women, selection into parties with higher levels of support contributed to improve their outcomes on average. This is especially the case for women in left wing parties who experienced the loss of individual influence after the impeachment of the left-wing female president.

This paper contributes to the literature on gender gaps in political representation. Women in office may be less available to around-the-clock work and receive less private donations than their male counterparts (Rosenbluth et al., 2015). They may also be punished by voters for behaviour perceived as power-seeking (Okimoto and Brescoll, 2010). In turn, women are less likely to run for reelection, compete for higher offices, and to be reelected (Brown et al., 2019; Brollo and Troiano, 2016). I add to the discussion by documenting disadvantages in competition for funding and showing that support from political parties may contribute to partially offset such disadvantages. More generally, I contribute to the literature on the role of political parties in the allocation of public funds (Brollo and Nannicini, 2012; Baião et al., 2018; Curto-Grau and Zudenkova, 2018), and I do so by applying well-established methods from the Labour Economics literature to a political market (Abowd et al., 1999; Card et al., 2013; Song et al., 2018; Cardoso et al., 2016; Gelbach, 2016).

The remainder of the chapter is organized as follows: Section 3.2 presents the context and a descriptive overview of gender gaps in access to IBAs. Section 3.3 presents the methods to quantify the sources of the gender gap and the main results. I show the relative contribution of party support and individual influence to the gaps and discuss the role of ideology in explaining the observed patterns. Section 3.4 concludes.

3.2 Descriptive Overview

3.2.1 Context

Women are still underrepresented in National Parliaments, and various channels through which women may be disadvantaged in a political career have been studied in the literature[1]. I show a new channel: gender differences in access to a type of

[1] Women are less likely to run for reelection, compete for higher offices, and to be reelected. Brown et al. (2019) find that male state legislators are twice as likely as women to compete for a Congressional seat and that their probability of winning is five times larger. Women, on the other hand, are more likely to rerun for the same seat. In Brazil, Brollo and Troiano (2016) show that female mayor's probability of being reelected is lower. Female incumbents may be less available to

funding that can improve reelection outcomes of politicians in office (IBAs)[2].

Overall, the institutional setting is the same as discussed in Section 2.2: each year federal legislators are entitled to Individual Budgetary Amendments (IBAs) on the same amount, but the actual implementation of the projects they propose is subject to approval by the Executive branch. Delays and cancellations may take place, and politicians may also negotiate extra funding for their projects. As a result, at the project execution level inequality arises, including, gender inequality.

Within this context, politicians negotiate the implementation of their projects individually and/or via the party leadership. Given that access to IBAs may contribute an incumbent's reelection outcomes, systematic gender gaps in project implementation is a potential mechanism slowing down the political careers of women.

3.2.2 Main sample and descriptive statistics

I build from Chapter 2 and carried out the analysis on the Largest Connected Set. As shown in Table 3.1, during the 54th Legislature 40 female (9.7%) and 409 male politicians are in the largest connected set, totalling 1,191 person-year observations. In the 55th Legislature, the largest connected set contains 1,545 person-year observations from 48 female (11.2%) and 427 male federal legislators. The largest connected set in the pooled sample (from 2011 to 2018) contains 2,749 person-year observations from 71 female (11.5%) and 615 male individuals.

Overall, men and women do not have systematic differences in observables. They have similar age profiles, most have completed at least a college degree and have had previous experience in office, indicating that seats in the parliament tend to be won by senior, experienced politicians[4].

around-the-clock work and receive less private donations than their male counterparts (Rosenbluth et al., 2015). Women in office may also be punished by voters for behaviour perceived as power-seeking (Okimoto and Brescoll, 2010).

[2]Pork-barrelling has been found relevant for the electoral outcomes of incumbents in various contexts. In Brazil, Finan and Mazzocco (2016) provide evidence that voters reward incumbent federal legislators based on the public resources their municipality receives[3]. They find a positive relationship between the amount of IBAs incumbents allocated to a municipality and the share of votes they received.

[4]Gender differences in observable characteristics such as age and education are more pronounced among entry level politicians (mayors and city councilors).

Table 3.1: Differences in means, by gender

Panel (a): 2011-2014				
	Women	Men	Difference	P-value
Age	54.69	54.42	0.26	0.89
Has at least college degree	0.72	0.78	-0.05	0.46
Times elected	2.67	2.84	-0.16	0.40
Won previous election	0.72	0.77	-0.05	0.55
Campaign funding (R$ million)	2.81	3.25	-0.44	0.24
Mover	0.03	0.10	-0.07	0.01
Observations	40	409		
Panel (b): 2015-2018				
	Women	Men	Difference	P-value
Age	53.82	53.88	-0.07	0.97
Has at least college degree	0.85	0.79	0.06	0.28
Times elected	2.58	3.13	-0.55	0.02
Won previous election	0.69	0.73	-0.05	0.52
Campaign funding (R$ million)	4.20	4.66	-0.46	0.19
Mover	0.15	0.21	-0.06	0.24
Observations	48	416		
Panel (c): 2011-2018				
	Women	Men	Difference	P-value
Age	53.27	54.15	-0.61	0.55
Has at least college degree	0.77	0.78	-0.02	0.95
Times elected	2.49	2.89	-0.64	0.01
Won previous election	0.70	0.72	-0.04	0.73
Campaign funding (R$ million)	3.62	3.98	-0.37	0.22
Mover	0.11	0.20	-0.09	0.03
Observations	71	615		

Notes: Statistics computed across individuals in the largest connect set of each time interval.

Some differences are worth noting. Men were more likely to switch party during the first period, but the difference is smaller and no longer significant in Term 2, when mobility is higher, and it is reassuring that mobility increases for both men and women. Overall, because the number of women in office does not allow for the estimation of gender-specific party fixed effects (as is done is Card et al. (2015)), the model assumes that the true party fixed effects are gender-invariant. In the presence of within-party gender discrimination, the party fixed effects of women could be biased since most movers in the sample are men.

Men have more experience, on average, and campaign contributions to male politicians were larger than for women in both periods[5]., although the gap was smaller in the second period. The share of men with at least a college degree was about 80% in both terms, but increased from 72% to 85% for women.

Table 3.2 reports the mean values of party characteristics for women and men in the largest connected set. The average party size, as measured by the number of seats in the Chamber of Deputies at election, was similar for men and women in both periods.

Table 3.2: Party Characteristics - Mean Values, by Gender

	54$^{\text{th}}$ Legislature (2011-2014)		55$^{\text{th}}$ Legislature (2015-2018)		Pooled Sample (2011-2018)	
	Females	Males	Females	Males	Females	Males
Count seats	49.22	51.34	38.16	38.75	42.48	44.07
Share seats	0.10	0.10	0.07	0.08	0.08	0.09
% females (at election)	0.15	0.08	0.17	0.09	0.16	0.09
Ideology:						
Right	0.20	0.38	0.21	0.39	0.21	0.39
Center	0.23	0.28	0.33	0.29	0.28	0.28
Left	0.58	0.34	0.46	0.31	0.51	0.32
Observations	106	1,085	147	1,398	255	2,494

Notes: Statistics computed across person-year observations in the largest connect set of each time interval.

However, the data show relevant differences in the composition of the parties in which men and women are members. Firstly, the data suggest segregation of women into certain parties. More specifically, women are more likely than men to have female party-colleagues. The share of female party members was, on average, 15% in the parties of female politicians and 8% in the parties of men politicians in

[5] Brollo and Troiano (2016) finds causal evidence that Brazilian female mayors receive less campaign contributions when running for reelection.

the first period. Similarly, the average in the second period was 17% for women and 9% for men. Moreover, segregation of women could be related to party ideology. In the 54^{th} Legislature, 58% of observations on female politicians were from left-wing parties, versus 34% for male politicians. Conversely, 20% of observations are on women in right-wing parties, versus 38% for men. A similar pattern of gender-based segregation is observed in relation to centrist parties. In the 55^{th} Legislature women are also segregated in left-wing parties (46% of observations), but less so, as the share of observations on women in centrist parties increased from 23% to 33%. The share of right, center and left-wing parties in the observations on men was similar between terms and across party ideology, indicating that male politicians do not segregate based in ideology.

3.2.3 Measuring gender inequalities in access to public funds

This section reports the outcomes of men and women at different stages of the budgetary process (Table 3.3). At the commitment stage, the average annual budget is virtually the same for men and women. However, budget cuts are higher for projects from female legislators, as measured by the cancellations. As a result, net of cancellations the average amount of funds annually committed to projects proposed by women are lower.

Table 3.3: Gender Differences in Funding (2011-2018)

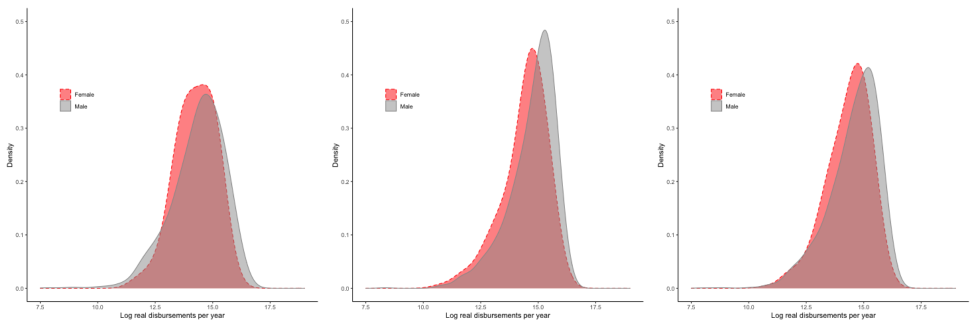
	All (1)	Females (2)	Males (3)	Diff. (4)	P-value (5)
Avg. Annual Commitments	7.079	7.093	7.078	0.016	0.94
Avg. Annual Cancellations	1.308	1.518	1.287	0.231	0.06
Avg. Annual Net Commitments	5.771	5.575	5.791	-0.215	0.33
Avg. Annual Disbursements	3.116	2.322	3.195	-0.873	0.00

Notes: Project amounts in R$ millions, real 2018 values. Statistics calculated across person-year observations of all elected federal deputies.

The largest difference in outcomes is observed at the disbursement stage: from 2011 to 2018, the average annual disbursements for women was 27% lower than the average for men and the difference in annual disbursements (R$ 0.87 million) corresponds to nearly two times the average project size (R$ 0.47 million).

3.2.4 Graphical evidence

Figure 3.1 shows the density of annual disbursements for men and women in the main sample, in logs. Panels (b) and (c) plot the distribution of all observations for politicians in power in the first (2011-2014) and second terms (2015-2018), respectively. Panel (c) plots the distribution of all observations in the pooled sample (2011 to 2018). Overall, the distribution of log disbursements for women is shifted to the left, with lower averages. Interestingly, despite the reform implemented in 2015 to reduce the overall inequality in payments, the density plots by term suggest that the gender differences are larger in the second period, a pattern I will discuss throughout the paper.



(a) 2011-2014 **(b)** 2015-2018 **(c)** 2011-2018

Figure 3.1: Density of Log Disbursements, by Gender.

In sum, men and women start the budgetary process with similar average outcomes, but after cancellations and delays the actual execution of projects proposed by women is lower. At the end of the budgetary process, the difference in the outcomes of men and women is substantial, statistically significant and arises mainly in the second term, despite reforms that aimed at reducing the overall inequalities in disbursements of Budgetary Amendments.

3.2.5 Raw and adjusted gender gaps: methodology

The results above suggest that in the competition for funding in Parliament female politicians are disadvantaged. In this section I document the raw and adjusted gender gaps in access to funding, showing that these differences are substantial and statistically significant. I follow the standard approach and, for each period of interest p, the raw gender gap is defined as the coefficient γ in Equation 3.1 below:

$$\ln y_t^i = g^i \gamma^p + \upsilon_t^i, \tag{3.1}$$

where, for each politician i, y_t denotes log real disbursements at year t, g is an indicator for female, and υ_t^i is the error term. Similarly, the adjusted gender gap is the coefficient γ in Equation 3.2:

$$\ln y_t^i = g^i \gamma^p + X_t^i \beta^p + \varepsilon_t^i, \tag{3.2}$$

where, X_t is a set of time-varying observables with the associated vector of coefficients β, and ε_t is the error term. I include year dummies, a measure of support to the Government's legislative agenda in the Chamber of Deputies (*Governism*), an indicator for the two final years of each legislature (*End of Term*), and an interaction between the two.

Governism is the share of open-vote Parliamentary sessions in which a politician voted in line with the government recommendation. As previously discussed, disbursement of Budgetary Amendments can be used by the President as a tool to build majorities in Congress. If there are systematic gender differences in support to government, omitting this control could bias the estimates of γ. I control for *End of Term* to account for electoral cycles in overall spending observed in the data[6]. The interaction between the two allows the return to government support to differ in pre-election years.

3.2.6 Raw and adjusted gender gaps: main results

Table 3.4 shows the raw and adjusted gender gaps in log disbursements for the whole period (columns 1 and 2) and in each legislature (columns 3 to 6).

Firstly, between 2011 and 2018 the gender gap was substantial at 24 log point

[6] Brollo and Nannicini (2012) analyse discretionary transfers from line ministries to Brazilian municipalities and they find that in the two years before municipal elections the amount of transfers to municipalities with mayor-president partisan alignment increases, while the transfers to mayors from the opposition decline. The authors do not find the same effect in the first two years of the municipal mandate.

in both the raw and adjusted specifications. However, the most striking result is that the difference significantly increases from 14 to 30 log points between periods, despite the fact that the overall dispersion of log disbursements is lower in the second term following the 2015 reform.

It is worth noting that the coefficient on *End of Term* is significant, and that the change of sign between periods is expected given that 2017 marks the beginning of a period of fiscal adjustment aimed at reducing non-mandatory spending, which includes Budgetary Amendments. The coefficients on *Support to Government* suggest that the returns to government support are higher in pre-election periods, but within this sample the effects are not statistically different from zero.

3.3 The sources of the gender funding gaps

What is the role of political parties in explaining the different outcomes of men and women? Do women systematically allocate into disadvantageous parties ?

In this section I investigate the sources of the gender gaps in Equation 3.2. More specifically, for each period p I decompose γ^p into components attributable to gender differences in party membership and to gender gaps in individual heterogeneity. The decomposition is obtained in two main steps: first I augment the model in Equation 3.2 to include party and individual fixed effects. In the second step I use the estimates from step 1 to decompose γ^p using the methodology in Gelbach (2016) that I discuss below.

3.3.1 Gender differences in party and person heterogeneity

As discussed above, the estimates from model 2.5 suggest that the observed inequality in log disbursements is mainly driven by differences in individual heterogeneity, as allocation into parties contributes to reduce the inequality overall. Below I show that this pattern of negative assortative matching has implications for the outcomes of women.

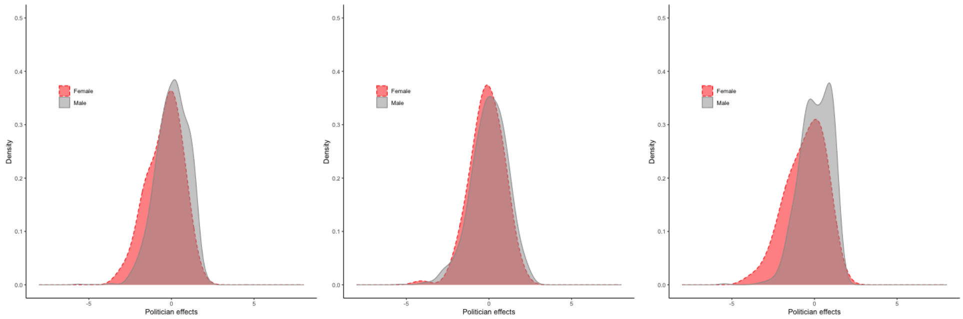
Figure 3.2 plots the distribution of individual fixed effects, by gender, in each Legislature and in the pooled sample. Panel (a) shows that the distribution of politician effects for women is shifted to the left relative to the distribution for men,

Table 3.4: Raw and Adjusted Gender Gap

	Dependent variable:					
	Log real disbursements					
	Pooled Sample (2011-2018)		54^{th} Legislature (2011-2014)		55^{th} Legislature (2015-2018)	
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.242^{***}	-0.249^{***}	-0.110^{*}	-0.138^{**}	-0.331^{***}	-0.298^{***}
	(0.083)	(0.083)	(0.064)	(0.060)	(0.088)	(0.087)
End of Term		1.093^{***}		1.050^{***}		-0.452^{***}
		(0.125)		(0.171)		(0.132)
Governism		0.159		-0.021		0.284
		(0.145)		(0.259)		(0.184)
End of Term x Governism		0.090		0.054		0.045
		(0.178)		(0.312)		(0.227)
Year dummies?	No	Yes	No	Yes	No	Yes
Obs. in the LCS	2,749	2,749	1,191	1,191	1,545	1,545
R^2	0.004	0.096	0.001	0.096	0.009	0.052

Notes: *p<0.1; **p<0.05; ***p<0.01. *End of Term* is an indicator for the last two years of the legislature. *Governism* is a measure of support to the Government's legislative agenda in the Chamber of Deputies.

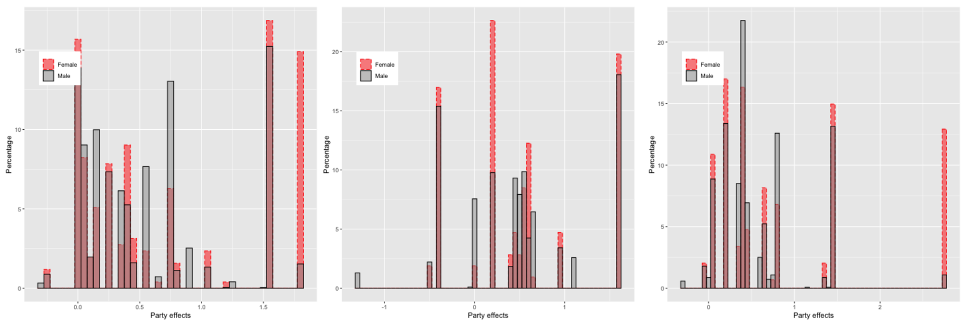
suggesting that unobserved heterogeneity correlated with gender is relevant in explaining the observed gender gaps in disbursements. Panels (b) and (c) indicate that these differences arise mainly in the second period.



(a) 2011-2018 **(b)** 2011-2014 **(c)** 2015-2018

Figure 3.2: Density of Politician Fixed Effects, by Gender

The gender differences in the empirical distribution of the estimated party effects, on the other hand, suggest a concentration of women in parties with higher fixed effects on average.



(a) 2011-2018 **(b)** 2011-2014 **(c)** 2015-2018

Figure 3.3: Density of Party Fixed Effects, by Gender

Table 3.5 shows the differences in means, by gender, of the estimated party and person fixed effects. Panels (a) and (b) shows that the average party fixed effects of female politicians is larger than of men, while, conversely, the average of person effects is lower for women. That is, the pattern of assortative matching described above is correlated with gender, as women with low individual influence allocate in parties with high levels of support.

Table 3.5: Difference in Means by Gender

	Females	Males	Female-Male	P-value
Panel (a): 54^{th} Legislature				
	Females	Males	Female-Male	P-value
Log disbursement	14.132	14.276	-0.144	0.258
Party FE	0.327	0.281	0.045	0.679
Person FE	-0.351	-0.128	-0.223	0.237
Xb	14.157	14.123	0.034	0.392
Observations	40	409	449	
Panel (b): 55^{th} Legislature				
	Females	Males	Female-Male	P-value
Log disbursement	14.281	14.575	-0.294	0.015
Party FE	0.196	-0.024	0.220	0.071
Person FE	-0.608	-0.083	-0.525	0.004
Xb	0.178	0.167	0.011	0.108
Observations	40	435	475	
Observations	40	646	686	

Notes: Party FE is party fixed effects, Person FE is person fixed effects, and Xb denotes the covariate index. Statistics calculated across person-year observations.

3.3.2 The role of parties in explaining the gender gap

3.3.2.1 Gender gap decomposition: methodology

The analyses above is not directly connected to the gender gap γ as defined in Equation 3.2. To decompose γ into party and person related components while partialling out the effects of covariates, I follow Cardoso et al. (2016) and use the politician ($\hat{\boldsymbol{\theta}}$) and party ($\hat{\boldsymbol{\psi}}$) fixed effects estimates from the augmented model in Equation 2.5 and apply the decomposition in Gelbach (2016).

More specifically, the adjusted gender gap is the OLS estimate $\hat{\gamma}$ from the baseline model

$$Y = X\boldsymbol{\beta} + G\boldsymbol{\gamma} + \boldsymbol{\varepsilon} \tag{3.3}$$

which I decompose into components associated with the fixed effects included in the full model

$$Y = X\boldsymbol{\beta} + D\boldsymbol{\theta} + F\boldsymbol{\psi} + \boldsymbol{\varepsilon}, \tag{3.4}$$

where the design matrices D groups the politician fixed effects and, similarly, F groups the party fixed effects.

The decomposition is an exact, order-invariant approach to measuring the effect of adding covariates to a model on the coefficients on some variables of interest, while taking into account time-varying observables (*Support to Government*, *End of Term* and year dummies). The decomposition in Gelbach (2016) is linear in the contribution of each covariate included in the full model (individual and party dummies in our case), so it is possible to estimate the party and politician components of the gender gap using only two auxiliary regressions.

Let $\hat{\boldsymbol{\theta}}$ and $\hat{\boldsymbol{\psi}}$ be the estimates of the fixed effects in the AKM-type model in Equation 2.5. Let $\tilde{X} \equiv [X, G]$ and let g be the gender dummy column index in $\tilde{X}$. Then, the politician fixed effect component of the adjusted gender gap $\hat{\gamma}_\theta$ is obtained from the g^{th} row in

$$C_\theta \equiv (\tilde{X}'\tilde{X})^{-1}\tilde{X}'D\hat{\boldsymbol{\theta}}, \tag{3.5}$$

which is a $((k+1) \times 1)$ vector that contains the politician fixed effects components of each of the $k+1$ coefficients estimated from Equation 3.2. Similarly, for the party fixed effects, the g^{th} row in

$$C_\psi \equiv (\tilde{X}'\tilde{X})^{-1}\tilde{X}'F\hat{\boldsymbol{\psi}}, \tag{3.6}$$

contains the party fixed effect component $(\hat{\gamma}_\psi)$ of the adjusted gender gap.

Gelbach (2016) shows that

$$\hat{\gamma} = \hat{\gamma}_\theta + \hat{\gamma}_\psi \tag{3.7}$$

so the adjusted gender gap is unambiguously decomposed into components associated with the party and politician fixed effects estimated from the specification in Equation 2.3.

Cardoso et al. (2016) show that $G^* \equiv MG$, where $M \equiv [I - X(X'X)^{-1}X']$ is such that pre-multiplying the terms $D\hat{\boldsymbol{\theta}}$ and $F\hat{\boldsymbol{\psi}}$ by $(G^{*'}G^*)^{-1}G^{*'}$ also gives $\hat{\gamma}_\theta$ and $\hat{\gamma}_\psi$ such that

$$\hat{\gamma} = \hat{\gamma}_\theta + \hat{\gamma}_\psi. \tag{3.8}$$

The authors further discuss the intuition behind this methodology by presenting a decomposition of the raw gender gap γ below:

$$Y = G\boldsymbol{\gamma} + \boldsymbol{\varepsilon} \tag{3.9}$$

based on the party and person fixed from the augmented model:

$$Y = D\boldsymbol{\theta} + F\boldsymbol{\psi} + \boldsymbol{\varepsilon}. \tag{3.10}$$

In this simpler setting, $G^* \equiv MG$, where $M \equiv [I - 1(1'1)^{-1}1']$. When multiplied by a vector, $(G^{*'}G^*)^{-1}G^{*'}$ gives the difference in gender means of such variable. Therefore, without covariates in the model, $(G^{*'}G^*)^{-1}G^{*'}$ decomposes the raw gender gap into the gender difference in mean politician fixed effects and the gender difference in mean party fixed effects. With covariates, the procedure is the same but using $M \equiv [I - X(X'X)^{-1}X']$ to partial out the effects of time-varying covariates.

3.3.2.2 Main results

Table 3.6 reports the decomposition results based on the methodology discussed above.

In the first period, the gender gap (-0.138) is mainly explained by the individual heterogeneity component (-0.104), and women's disadvantage in individual influence does not vary significantly across ideology. Selection into low support parties (-0.034) explains a small share of the gender gap and reflect the participation of women in low support right wing and centrist parties. Support from left wing parties, on the other hand, contributed to slightly narrow down the overall gap.

Table 3.6: Conditional Decomposition of the Adjusted Gender Gaps

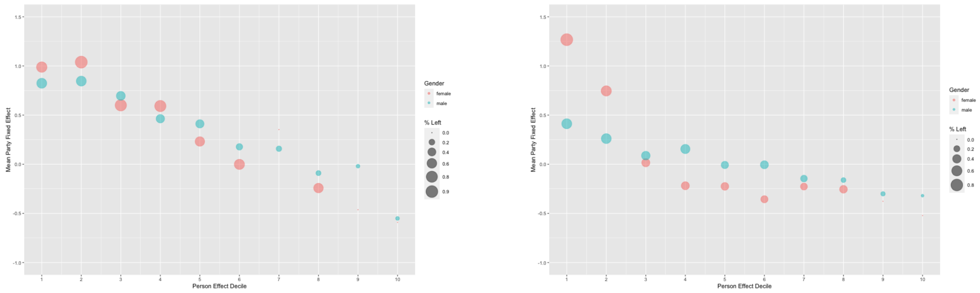
	54^{th} Legislature (2011-2014)	55^{th} Legislature (2015-2018)
Adjusted Gap	-0.138	-0.298
Politician Effects	-0.104	-0.522
Party Effects	-0.034	0.224
Observations	1,191	1,545

The gender gap was significantly larger (-0.298) in the second period. This large increase takes place in a context of decline in the total variance of log disbursements. That is, while the overall funding inequality was lower in the second term, the gender gap nearly doubled. What explains the worse outcomes of women despite the reforms implemented to reduce competition and discrimination in the execution of IBAs?

In period 2, the larger gap is explained solely by differences in the individual influence component (-0.52), which could include differences in preferences, beliefs, and discrimination. Selection into parties, on the other hand, actually contributed to significantly narrow down the differences between men and women (by 0.22). That is, before controlling for parties the gender gap is underestimated because women are more likely to select into high support parties. A natural follow up question is what types of parties provide support to women?

3.3.2.3 Gender gaps and ideology

Panels (a) and (b) in Figure 3.4 show that women at the bottom of the person effects distribution were, on average, in parties with higher levels of support when compared to men. That is, at least for a subset of women party support partially compensates for their gap in individual influence. Figure 3.4 also shows that this gender-related negative assortativeness is related to ideology, as most women in parties with high support are members of the left.



(a) 2011-2014 **(b)** 2015-2018

Figure 3.4: Mean Party Fixed Effects by Deciles of Person Effects - Gender and Ideology, by Term

Table 3.7 shows how each ideology groups contributes to the person and party components of the gender gap. In both periods, politicians in all ideology groups contributed to the person effects component of the gap. However, the contribution of persons in left wing parties is higher and significantly increased from about 44% to 61.2% between terms.

Table 3.7: Conditional Decompositions of Adjusted Gender Gaps - By Ideology

	First Period	%	Second Period	%
Adjusted Gap	-0.138		-0.298	
Politician Effects	-0.104		-0.522	
Party Effects	-0.034		0.224	
Party Effects by Ideology:				
All	-0.034		0.224	100.00
Left	0.012	-35.36	0.234	104.75
Center	-0.015	44.54	-0.016	-7.36
Right	-0.031	90.82	0.008	3.48
Not Classified	0.000	-0.00	-0.002	-0.87
Person Effects by Ideology:				
All	-0.104		-0.522	100.00
Left	-0.046	44.06	-0.320	61.20
Center	-0.030	28.68	-0.112	21.52
Right	-0.028	27.27	-0.090	17.27
Not Classified	0.000	-0.00	0.000	-0.00
Observations	1,191	-	1,545	-

Interestingly, it is party support to women in left wing parties that contributes to partially offset their loss of individual influence, specially in the second period, when it contributed to narrow down the gender gap by 23.4 log points.

I cannot pin down within party mechanisms potentially related to gender-based segregation of women in the left. For example, if left wing parties have a preference for collective action and are more gender-inclusive, women in the left would benefit from higher levels of party support than women in other parties simply. But a different mechanism could be at play: women could benefit from gender-based segregation in left wing parties as a result of their own influence, as a subpopulation, on the party approach to collective action. It is also plausible that both mechanisms be at play, but I cannot decompose the contribution of each.

Still, within a context of overall deterioration of the individual influence and performance of women, selection into left wing parties with higher levels of support contributed to partially offset such deterioration and improve their outcomes on average.

Such deterioration reflects mainly the loss of influence of left wing women rather than an large improvement in the outcomes of men in left wing parties, as I discuss below.

Table 3.8 shows the contributions of men and women to the person effects component of the gap, by ideology group, and relative to the contribution of men in centrist parties. Columns 1 and 3 present the contribution of each ideology x gender sub-group to the person effects component. Negative values add to the gender gap, while positive values indicate that the group's contribution was to narrow down the overall gender differences. For example, women in left wing parties are more likely to be at the bottom of the distribution of persons effects. As a result, their total contribution was to enlarge the gap by 0.28 units in the first period and by 0.52 units in the second. Men in the left wing parties also tend to be at the bottom of the distribution of persons effects, but their relative low influence contributed to narrow the overall gender gap by 0.24 units in term 1 and by 0.20 units in term 2. Men in center and right wing parties, in turn, contributed to increase the gender differences in person effects in both periods. Women in the center contributed to improve the

Table 3.8: Composition of the Person Effects Component - by Ideology and Gender

	First Period	Contribution relative to Center-Men	Second Period	Contribution relative to Center-Men
	(1)	(2)	(3)	(4)
Adjusted Gap	-0.14	-	-0.30	-
Politician Effects - All	-0.10	-	-0.52	-
Left-All	-0.05		-0.32	
Left-Women	-0.28	1.94	-0.52	9.10
Left-Men	0.24	-1.63	0.20	-3.46
Center-All	-0.03		-0.11	
Center-Women	0.12	-0.80	-0.06	0.98
Center-Men	-0.15	1.00	-0.06	1.00
Right-All	-0.03		-0.09	
Right-Women	0.04	-0.25	0.03	-0.54
Right-Men	-0.07	0.45	-0.12	2.13

average of women in the first period, but not in the second, while the individual influence of right wing women contributed to narrow the gender gaps in both terms.

These estimates cannot be directly compared between periods unless they are normalized. As a simple normalization I compute the contribution of each group in relation to the outcomes of men in centrist parties (Columns 2 and 4). The most striking results is that, while in the first period left wing women's contribution to the gap was about twice as large as the component attributable to men in centrist parties, this ratio significantly increased to more than nine times in the second period. The contribution of individual influence of men in the left, on the other hand, was to narrow the gender gap by 1.3 times the "Center-Men" component in term 1, and by 3.4 times in period 2.

These results indicate that the negative assortative matching indicated in the total variance decomposition analysis in Chapter 2 contributed to narrow the differences in outcomes of men and women, at least partially and especially for women in left wing parties during the second term.

3.4 Concluding Remarks

Female representation in politics has been increasing but at a slow pace worldwide. A large body of research has pointed out to various factors that hinder the participation and progression of women in political careers. This paper documents a gender gap in access to public funds by women legislators in Brazil and decomposes this gap into components attributable to party and individual heterogeneity. I find that the differences in politician effects are the main drivers of the observed gender gaps in access to funding, but that segregation of women in parties with higher fixed effects partially offsets the gap. I also find evidence that the negative sorting is associated to party ideology.

Finally, the framework for the estimation of party fixed effects and the decomposition of the gender gaps presented in this paper could be used to study other types of between-group differences in access to funding, like race, or in other contexts in which party switching is common.

Chapter 4

The Heterogeneous Effects of Reelection Incentives: Evidence from Brazil

4.1 Introduction

The effects of reelection incentives on the behaviour of incumbent politicians have been widely studied in the literature, both theoretically and empirically. The general idea is that incumbents can use policy to signal their type to voters and distinguish themselves from challengers. However, whether specific policies are affected by reelection concerns of incumbents should depend on how voters value the policy in question.

This paper studies the relationship between reelection incentives and the provision of water wells in dry areas of the Northeast and Southeast Brazil. I use a regression discontinuity design with mixed incumbent-challenger close elections and variation in rainfall to identify heterogeneous effects of eligibility for a second consecutive term on the drilling of water wells. By allowing for an interaction between treatment and the frequency of dry years, I show that reelection incentives increase the drilling of water wells as long as the frequency of dry years is sufficiently high and that this effect is significantly larger in the driest areas.

I take advantage of term limits on mayors of Brazilian municipalities, who are elected for a mandate of 4 years and allowed to run for one second consecutive term. Elected challengers, therefore, have reelection incentives, while reelected incumbents face term limits and cannot run for a third consecutive term. A simple comparison between municipalities with first and second-term mayors could be biased if reelection of an incumbent is correlated with unobserved municipality-specific determinants of the outcome of interest. For example, in a context of uncertain precipitation, risk preferences of citizens could affect both their voting decisions when an incumbent is running and their preferences for the provision of water wells. In order to tackle this issue, I use close elections between challengers and incumbents who run for reelection to identify the effect of reelection incentives on the provision of water wells with a regression discontinuity design. In order to allow for heterogeneous effects, I interact treatment status with a measure of the frequency of dry years, which is built from municipality-level precipitation data produced in Rocha and Soares (2012).

The semiarid is the driest area in Brazil, with a historical precipitation average at about half the average for the rest of the country and recurrent droughts that have been a major source of vulnerability. The concentration of rainfall in rainy seasons, combined with the topography and the temperature profile of the area, results in water losses due to quick evaporation and high levels of salinity and low quality of the surface water remaining for consumption ((Rocha and Soares, 2012),(Bobonis et al., 2019)). Underground water, while being more resilient to high temperatures, is typically saline and, for this reason, more relevant as a water source when surface water is scarce. The context in the Brazilian Semiarid, therefore, provides an opportunity for the evaluation of the heterogeneous effects of reelection incentives on the provision of public goods that can mitigate the impacts of negative shocks such as droughts. Additionally, using data from the Brazilian Underground Water Censuses, I show a negative relationship between the maintenance of public water wells and time to election in municipalities where drought frequency is sufficiently high.

The above relationship is explained in the context of a Rogoff-type model (Rogoff, 1990) of political budget cycles, which I modify to allow for the provision of a public good from which voters derive utility only under certain states of nature (state-dependent utility). In the context of the Brazilian semiarid, the shock refers to dry years and the public goods to the drilling of water wells.

This paper relates to a literature that extends the standard explicit incentives models and applies the career concerns principal-agent framework to political agency problems. This approach allows for situations in which performance is observable, but not contractible, as it is usually the case in political mandates with terms that cannot be conditioned on observed performance and reelection depends on the voters assessment of the incumbent's ability.

In particular, this paper relates to the theory and evidence on political budget cycles, in which Rogoff-type models and those based on Holmstrom (1999) have been widely used. Drazen and Eslava (2010) present a model in which incumbent politicians target voters by changing the composition of spending. Evidence from Colombian municipalities supports the conclusions from the model and shows that voters responded to targeting. Cole (2009) reconciles theories of political budget cycles with tactical electoral redistribution and shows that government-owned bank lending, specially agricultural credit, tracks the electoral cycle and that the same pattern is not observed among private banks. Brender and Drazen (2005) show that in a large cross-section of countries a political deficit cycle is driven by recent democracies in both developed and developing economies.

The effects of electoral accountability go beyond fiscal policy. Additional cross-country evidence is presented in Block (2002) using panel data of African countries, where both fiscal and monetary policy variables track a political cycle. In Brazil, Ferraz and Finan (2011) find causal evidence that reelection incentives have significant impacts on corruption: in municipalities where mayors can be reelected, misappropriation of resources is 27% lower than in those where mayor are in their last term.

The remainder of this chapter is organized as follows. Section 2 describes the theoretical framework. Context and data are presented in section 3. Section 4 discusses the empirical strategy. The main results are reported in Section 5, followed by a presentation of additional results in section 6. Section 7 concludes.

4.2 Theoretical Framework

4.2.1 The Model

This section presents a theoretical framework for the provision of public goods by incumbent politicians with reelection motives. I modify the Equilibrium Political Cycle model in Rogoff (1990) to allow for the provision of public goods from which voters derive utility only under certain states of nature, but have to bare the costs regardless of the realization of the shock. In the context of the chosen application, the public good refers to the provision of water wells in drought-prone municipalities of the Brazilian semiarid. The basic idea is that in dry years citizens must use underground water as a replacement for scarce surface water, therefore deriving utility from the water wells. In normal or wet years, voters use other sources of drinking water and do not derive utility from the availability of underground water. The economy is composed of a continuum of agents and the representative voter cares about the expected value of their lifetime utility from time t onward, which is denoted by $E_t^P(\Gamma_t)$, where P denotes the public's information set and Γ_t is given by

$$\Gamma_t = \sum_{s=t}^{T} \{pU(c_s, g) + (1-p)H(c_s) + V(k_s) + \eta_s\}\beta^{s-t}. \tag{4.1}$$

The consumption of private goods is denoted by c, g represents the public good, and k represents a public "investment" good that takes one period to materialize and be perceived by voters. The utility functions U, H, and V are assumed to be regular strictly concave functions. β is a discount factor and T is the time horizon of the representative agent.

In each period, voters preference's depend on the realization of a state of nature. The state space is given by $R = \{d, w\}$, with $p \equiv Pr(r = d)$, $1 - p \equiv Pr(r = w)$ where d refers to a bad state of nature (a dry year, in our application). If d is realized, then voters derive utility from the positive provision of g and disutility if $g = 0$. But if $r = w$, then g generates no welfare and the resources applied to its provision are perceived by voter as a tax with no utility counterpart. Assuming that $U_1(c,0) < H'(c)$ for all c formalizes the idea that the citizens are "punished" if $r = d$ and $g = 0$ (increased consumption of the private good will not have the same welfare effect as when $r = w$).

An exogenous endowment y of an storable good is given to each citizen in each period and can be privately consumed or used as an input to the provision of public goods. The provision of public goods is decided by an incumbent politician who establishes lump-sum taxes τ to raise the resources for the provision of g. The resources constraint is given by

$$c_t = y - \tau_t, \tag{4.2}$$

and the technology for the production of the public good is described by

$$g_t + k_{t+1} = \tau_t + \varepsilon_t, \tag{4.3}$$

where ε denotes the ability of the incumbent politician. This specification means that a politician with high ability is able to provide the same amount of public goods charging lower taxes. The timing of production of the public investment good k differs from that of g. The amount of k available at t is invested in period $t-1$. As in Alesina and Tabellini (2008), competence is assumed to evolve according to

$$\varepsilon_t^i = \alpha_t^i + \alpha_{t-1}^i \tag{4.4}$$

where, as in Rogoff (1990), α is observed by the incumbent before the choice of policy and is an independent (across agents and across time) variable drawn from a Bernoulli distribution with $\rho \equiv Pr(\alpha = \alpha^H)$ and $1 - \rho \equiv Pr(\alpha = \alpha^L)$, where $\alpha^H > \alpha^L > 0$. Besides the competency shock, the model allows for an ideology bias that follows the process below:

$$\eta_t^i = q_t^i + q_{t-1}^i \tag{4.5}$$

where q is drawn from a continuous distribution on a closed interval [- $\bar{q}$,$\bar{q}$], $\bar{q} > 0$.

The expected utility of the incumbent politician is given by

$$E_t^I(\Gamma_t) + \sum_{s=t}^{T} \beta^{s-t} X \pi_s, t, \tag{4.6}$$

where $E_t^I(\Gamma_t)$ represents the expected utility the politician derives from the provision of g itself, which could be either because the politician derives utility from g and private consumption since they are an ordinary citizen or because they put some weight on social welfare. The term $\sum_{s=t}^{T} \beta^{s-t} X \pi_{s,t}$ captures the utility from holding office, where X represents "ego rents" and $\pi_{s,t}$ the probability of being in office in period t, estimated at time s.

Voters behaviour is modelled as follows: elections are held every other period and, in deciding their vote, the representative agent compares their expected utility in case the incumbent is reelected with their expected utility if the challenger wins. Let $v = 1$ denote a vote for the incumbent. Then,

$$v_t = \begin{cases} 1 & E_t^P(\Gamma_{t+1}) \geq E_t^P(\Gamma_{t+1}^O) \\ 0, & otherwise. \end{cases} \tag{4.7}$$

Where $E_t^P(\Gamma_{t+1})$ denotes the representative voter's expected utility if the incumbent is reelected and $E_t^P(\Gamma_{t+1}^O)$ represents the voter's expected utility in case the challenger wins the election.

4.2.2 Equilibrium Under Full Information and Drought Risk

The equilibrium under full information (voters observe α, p and r prior to voting) is discussed below. In short, in this case the politician cannot use their choice of policy to influence the voters inference about their future ability and, therefore, the term $\sum_{s=t}^{T} \beta^{s-t} X\pi_{s,t}$ becomes exogenous to the incumbent. As a result, the incumbent's choice reduces to the same as the representative voter, which can be written as a sequence of static problems:

$$\max_{\tau_t, c_t, g_t, k_{t+1}} \{pU(c_t, g_t) + (1-p)H(c_t) + \beta V(k_{t+1})\} \tag{4.8}$$

subject to (2), (3), k,c,g ≥ 0 and $k_{T+1} = \bar{k}$ for all t$\geq T$.

Let $\widetilde{U}(c_t, g_t; p) \equiv pU(c_t, g_t) + (1-p)H(c_t)$. This optimization can be rewritten as

$$\max_{\tau, g} W(g, \tau, \varepsilon; p) \equiv \widetilde{U}(y-\tau, g; p) + \beta V(\tau + \varepsilon - g) \tag{4.9}$$

such that g, y-τ, $\tau + \varepsilon - g \geq 0$.

The first-order conditions with respect to τ and g are given by

$$\widetilde{U}_c(y-\tau, g; p)(-1) + \beta V'(\tau + \varepsilon - g) = 0 \tag{4.10}$$

$$\widetilde{U}_g(y-\tau, g; p) + \beta V'(\tau + \varepsilon - g)(-1) = 0 \tag{4.11}$$

$$\Longleftrightarrow$$

$$pU_1(y-\tau, g) + (1-p)H'(y-\tau) = \beta V'(\tau + \varepsilon - g) \tag{4.12}$$

$$pU_2(y-\tau, g) = \beta V'(\tau + \varepsilon - g) = 0 \tag{4.13}$$

Assuming that U, H and V are strictly concave utility functions and given that the constraint set is convex, there exists a unique $[g^*(\varepsilon, p), \tau^*(\varepsilon, p)]$ satisfying equations 4.10 and 4.11, which is a global maximum. Moreover, equations 4.10 and 4.11 define a function $\tau = \phi(g; y, p)$ (income expansion path) which is decreasing in g and will be useful in the discussion on the equilibrium under asymmetric information. Briefly, equations 4.10 and 4.11 imply that

$$F(g,\tau) \equiv pU_1(y-\tau,g) + (1-p)H'(y-\tau) - pU_2(y-\tau,g) = 0, \tag{4.14}$$

and, assuming that $U(c,g)$ is separable and concavity of U and H, by the implicit function theorem it is possible to show that

$$\frac{\partial \tau}{\partial g} = -\frac{\frac{\partial F}{\partial g}}{\frac{\partial F}{\partial \tau}} = -\frac{-pU_{22}}{-pU_{11} - (1-p)H''} < 0. \tag{4.15}$$

Finally, it is possible to show that $g^*(\varepsilon, p)$ is increasing in p depending on the preferences of the voters on g and c. The following proposition summarizes the conditions under which this holds and sketches a proof.

Proposition 1. *If*

1. *$U(c,g)$ is separable and*

2. *$U_1 \leq H'$,*

then, the solution $[g^*(\varepsilon, p)]$ to (9) is increasing in p.

Proof. From the first-order conditions in 4.10 and 4.11, define

$$F(g,t;\varepsilon,p) \equiv pU_1(y-\tau,g)(-1) + (1-p)H(y-\tau)(-1) + \beta V'(\tau+\varepsilon-g) = 0 \tag{4.16}$$

$$G(g,t;\varepsilon,p) \equiv pU_2(y-\tau,g)(1) + \beta V'(\tau+\varepsilon-g)(-1) = 0. \tag{4.17}$$

By total differentiation and applying the Cramer's rule

$$\frac{\partial g^*(\varepsilon,p)}{\partial p} = \frac{-\frac{\partial G}{\partial \tau}\frac{\partial F}{\partial p} + \frac{\partial F}{\partial \tau}\frac{\partial G}{\partial p}}{\frac{\partial G}{\partial \tau}\frac{\partial F}{\partial g} - \frac{\partial G}{\partial g}\frac{\partial F}{\partial \tau}} \equiv \frac{A}{B}. \tag{4.18}$$

Let $K \equiv -\beta V''(\tau-\varepsilon-g)$. Then, it is possible to write B as

$$\begin{aligned} B = p^2(U_{12}U_{21} - U_{11}U_{22}) - pU_{12}K - pU_{21}K \\ +pU_{11}K - (1-p)H''pU_{22} + (1-p)H''K + KpU_{22}. \end{aligned} \tag{4.19}$$

Assuming U_{11}, U_{22}, and $H'' < 0$ and $U_{12} \geq 0$, then the only term in 4.13 possibly greater or equal to zero is $p^2(U_{12}U_{21} - U_{11}U_{22})$. Therefore, if $p^2(U_{12}U_{21} - U_{11}U_{22}) \leq 0$ (which is the case if $U(c,g)$ is separable) then $B \leq 0$ for all p.

Next, write

$$A = [U_1 - H'][K - pU_{21}] + [pU_{11} + (1-p)U'' - K]U_2 \tag{4.20}$$

.

Assuming U_{21}=0, then $A \leq 0$ if, and only if,

$$\frac{[U_1 - H'][K]}{U_2} \leq -[pU_{11} + (1-p)U'' - K], \tag{4.21}$$

where the *RHS* is positive. Therefore, $U_1 \leq H'$ is a sufficient condition for $A \geq 0$. That is, if the marginal utility of private consumption is the same in both states of the world or lower if a bad state is realized, then the optimal provision of the public good increases with p for all p. The intuition is that if the realization of the bad state does not increase or reduces welfare derived from private consumption, then the provision of the public good will increase with p. For a given value of U_2, this result can also be interpreted in terms of restrictions on the marginal rate of substitution of private consumption for the public good. If $\left(\frac{U_1}{U_2}\right)$ is sufficiently low (that is, if voters would give up sufficiently high amount of private consumption for the provision of g), then the LHS in 4.15 is negative and $A \leq 0$. $\square$

In the context of the application to the Brazilian semiarid, the model implies that, the higher the frequency of dry years, the higher the provision of water wells. On the other hand, it also implies that if there is a chance that the water wells will not be needed, effort will be lower than it would if the public good g were valuable in all states of nature, even under full information.

Because when voters decide whether to vote for the incumbent or for the challenger at a time t they only care about the politician's choice of policy and ability from $t+1$ onward, if voters observe α_t, p, and the realization of d, then as in the standard case the incumbent's policy prior to the election cannot affect the voter's inference about their competency in the future. Therefore, the incumbent has no incentives to deviate from the optimal (risk management) policy choice $[g^*(\varepsilon, p), \tau^*(\varepsilon, p)]$.

Formally, $v = 1$ (the incumbent is reelected) if

$$E_t^P[W^*(\varepsilon_{t+1}; p) \mid \alpha_t] - E_t^P[W^*(\varepsilon_{t+1}^O; p)] + q_t - q_t^O \geq 0, \tag{4.22}$$

where

$$\begin{aligned} E_t^P[W^*(\varepsilon_{t+1}^O); p] = \rho^2 W^*(2\alpha^H) + 2\rho(1-\rho)W^*(\alpha^H + \alpha^L; p) \\ + (1-\rho)^2 W^*(2\alpha^L; p) \equiv \Omega^O(p), \end{aligned} \tag{4.23}$$

$$E_t^P[W^*(\varepsilon_{t+1}; p) \mid \alpha_t = \alpha^H] = \rho W^*(2\alpha^H; p) + (1-\rho)W^*(\alpha^H + \alpha^L; p) \equiv \Omega^H(p), \tag{4.24}$$

and,

$$E_t^P[W^*(\varepsilon_{t+1}; p) \mid \alpha_t = \alpha^L] = \rho W^*(\alpha^H + \alpha^L; p) + (1-\rho)W^*(2\alpha^L; p) \equiv \Omega^L(p). \tag{4.25}$$

Since $\Omega^H(p) > \Omega^O(p) > \Omega^L(p)$ for all p, the incumbent is reelected only if voters observe they are of the high ability type.

Next section analyzes under which conditions a political cycle arises in an equilibrium with asymmetric information and what is the relationship between the political cycle and p.

4.2.3 Equilibrium under Asymmetric Information and Drought Risk

4.2.3.1 The Voters' and Politician's Problems

If ability is not observable, then the voters form beliefs about α_t based on their observations of (g_t, τ_t). Let $\hat{\rho}(g,\tau)$ denote such beliefs. That is, $\hat{\rho}(g,\tau)$ denotes the probability that the voter assigns to the event $\alpha_t = \alpha_H$. Then, the incumbent wins the election if

$$\hat{\rho}\Omega^H(p) + (1-\hat{\rho})\Omega^L(p) - \Omega^O(p) + q - q^o \geq 0 \tag{4.26}$$

Given such system of beliefs, an incumbent of type i calculates their probability of reelection according to

$$\pi(\hat{\rho}) \equiv P(v=1 \mid g,t;p) = 1 - G\big(\Omega^O(p) - \Omega^H(p) - (1-\hat{\rho})\Omega^L(p)\big), \tag{4.27}$$

and chooses (g,τ) to solve

$$\max_{\tau,g} \left[\mathscr{X}^i \pi\big(\hat{\rho}(g,\tau)\big) + \widetilde{U}(y-\tau,g;p) + \beta V(\tau+\varepsilon-g)\right] \tag{4.28}$$

such that

$$g, y-\tau, \tau+\varepsilon^H - g \geq 0,\ i = H,L, \tag{4.29}$$

where,

$$\mathscr{X}^i \equiv \beta[X(1+\beta) + \Omega^i(p) - \Omega^O(p)], \tag{4.30}$$

and $G(.)$ is the CDF of $q-q^O$.

4.2.3.2 Sequential Equilibria

In the following subsections, attention is restricted to equilibria in pure strategies and the definition of sequential equilibrium is presented below. Let (g^i,τ^i) describe a strategy for an incumbent politician of type i and $v\big(\hat{\rho}(g,\tau),q-q^O\big)$ denote a voters' strategy.

Definition 1. A pair $\{(g^i,\tau^i),v\big(\hat{\rho}(g,\tau),q-q^O\big)\}$ is a sequential equilibrium if
(i) the voting rule is set according to 4.26;
(ii) the incumbent policy choice satisfies 4.28 and 4.29; and
(iii) voters have Bayes-consistent beliefs in the sense that: if $(g^L,\tau^L)\neq(g^H,\tau^H)$, then $\hat{\rho}(g^L,\tau^L)=0$ and $\hat{\rho}(g^H,\tau^H)=1$. If $(g^L,\tau^L)t=(g^H,\tau^H)$, then $\hat{\rho}(g^L,\tau^L)=\hat{\rho}(g^H,\tau^H)=\rho$.

In a separating equilibrium $(g^L,\tau^L)\neq(g^H,\tau^H)$ and therefore the low type chooses their first-best optimal policy. The necessary conditions for a separating equilibrium are that (i) the low type does not benefit from pretending to be of the high type, and that (ii) the high type is better off by separating themselves from the low type than if they chooses their first-best policy and is perceived by voters as a low type[1]. Formally, let $Z(g,\tau,1,\varepsilon^L;p)\equiv\mathscr{X}^i\pi\big(\hat{\rho}(g,\tau)\big)+W(g,\tau,\varepsilon;p)$. If off-equilibrium-path beliefs are such that $\hat{\rho}(g,\tau)=0\ \forall(g,\tau)\neq(g^H,\tau^H)$, then the necessary conditions of a separating equilibrium can be written as

$$(g^H,\tau^H)\in\{(g,\tau)\mid Z(g,\tau,1,\varepsilon^L;p)\}\leq Z\big(g^*(\varepsilon^L),\tau^*(\varepsilon^L),0,\varepsilon^L;p\big)\}\equiv\mathscr{A},\quad(4.31)$$

and

[1] As in Rogoff (1990), these restrictions on the voters' system of beliefs are insufficient to rule out the possible pooling equilibria. However, by refining the equilibrium concept it is possible to exclude all possible pooling equilibria and preserve the separating equilibria. In the model with state-dependent utility the argument discussed in the baseline model applies as well. For details, see Rogoff (1990).

$$(g^H,\tau^H) \in \{(g,\tau) \mid Z(g,\tau,1,\varepsilon^H;p)\} \geq Z\big(g^*(\varepsilon^H),\tau^*(\varepsilon^H),0,\varepsilon^H;p\big)\} \equiv \mathscr{B}. \quad (4.32)$$

.

The following proposition, detailed in Rogoff (1990), summarizes the argument above.

Proposition 2 (Rogoff)**.** *The set of all separating equilibria is nonempty and is characterized by* $(g^L,\tau^L) = \big(g^*(\varepsilon^L),\tau^*(\varepsilon^L)\big)$*, and* $(g^H,\tau^H) \in \mathscr{A} \cap \mathscr{B}$*.*

By imposing that off-path beliefs are such that $\hat{\rho}(g,\tau) = 1\ \forall (g,\tau) \in \mathscr{A} \cap \mathscr{B}$, and not just for $(g,\tau) = (g^H,\tau^H)$ it is possible to reduce the range of separating equilibria to a single point. If voters have such beliefs, then in the unique separating equilibrium, (g^H,τ^H) solves

$$\max_{\tau,g} W(g,\tau,\varepsilon^H;p) \quad (4.33)$$

$$\text{such that } g, y-\tau, \tau+\varepsilon^H - g \geq 0 \quad (4.34)$$

$$\text{and } (g,\tau) \in \mathscr{A} \quad (4.35)$$

Because under such system of beliefs $\hat{\rho} = 1$ for any $(g,\tau) \in \mathscr{A}$, adding the constraint makes the term $\mathscr{X}^i \pi\big(\hat{\rho}(g,\tau)\big)$ irrelevant to the optimization and, therefore, the incumbent's problem can be written as shown above. The associated Lagrangian is given by

$$\begin{aligned}\mathscr{L} \equiv\ & \widetilde{U}(y-\tau,g_t) + \beta V(\tau+\varepsilon^H+g) + \\ & \lambda[\pi(0)\mathscr{X}^L + W_L^* - \pi(1)\mathscr{X}^L - \widetilde{U}(y-\tau,g_t) - \beta V(\tau+\varepsilon^L - g)]\end{aligned} \quad (4.36)$$

and the Kuhn-Tucker conditions are

$$-\widetilde{U}_c + \beta V_H' + \lambda[\widetilde{U}_c - \beta V_L'] \leq 0 \qquad (4.37)$$

$$\tau\{\widetilde{U}_c + \beta V_H' + \lambda[\widetilde{U}_c - \beta V_L']\} = 0 \qquad (4.38)$$

$$\widetilde{U}_g - \beta V_H' + \lambda[-\widetilde{U}_g - \beta V_L'] \leq 0 \qquad (4.39)$$

$$g\{\widetilde{U}_g - \beta V_H' + \lambda[-\widetilde{U}_g - \beta V_L']\} = 0 \qquad (4.40)$$

$$\pi(0)\mathscr{X}^L + W_L^* - \pi(1)\mathscr{X}^L - \widetilde{U}(y-\tau, g_t) - \beta V(\tau + \varepsilon^L - g) \geq 0 \qquad (4.41)$$

$$\lambda[\pi(0)\mathscr{X}^L + W_L^* - \pi(1)\mathscr{X}^L - \widetilde{U}(y-\tau, g_t) - \beta V(\tau + \varepsilon^L - g)] = 0 \qquad (4.42)$$

which for $(\tau, g) \neq (0,0)$ reduce to

$$\widetilde{U}_c - \beta V_H' = \lambda[\widetilde{U}_c - \beta V_L'] \qquad (4.43)$$

$$\widetilde{U}_g - \beta V_H' = \lambda[\widetilde{U}_g - \beta V_L'] \qquad (4.44)$$

$$\pi(0)\mathscr{X}^L + W_L^* - \pi(1)\mathscr{X}^L - \widetilde{U}(y-\tau, g_t) - \beta V(\tau + \varepsilon^L - g) \geq 0 \qquad (4.45)$$

Note that, by equations 4.43 and 4.44,

$$\widetilde{U}_c = \widetilde{U}_g \qquad (4.46)$$

$$\leftrightarrow \qquad (4.47)$$

$$pU_1 + (1-p)H' = pU_2 \qquad (4.48)$$

which is also among the first-order conditions for the first best (equilibrium with full information) and, therefore, defines the same function $\tau = \phi(g; p)$ as in the first-best. Equations 4.43 and 4.44 also show that if the constraint is not biding ($\lambda = 0$), then the first best equilibrium is attained. Whether or not this is the case depends on the parameters of the problem. For example, if the ability of the high type is sufficiently high, then they would be able to chose their first-best policy and still separate themselves. If the difference between ε^H and ε^L is not too high, then the constraint is biding, $\lambda > 0$, and two solutions to the above systems emerge.

The second order conditions for a maximum, however, hold only for the solution at which

$$U_2 - \frac{\beta V_H'}{p} = \frac{\beta\lambda(V_H' - V_L')}{p(1-\lambda)} < 0. \quad (4.49)$$

Assuming that U is a concave function, this means that under asymmetric information the solution g^H to Equation4.49 is higher than in the first-best optimal policy $g^*(\varepsilon^H)$, which solves $U_2 = \frac{\beta V_H'}{p}$, as shown in figure 4.1 below.

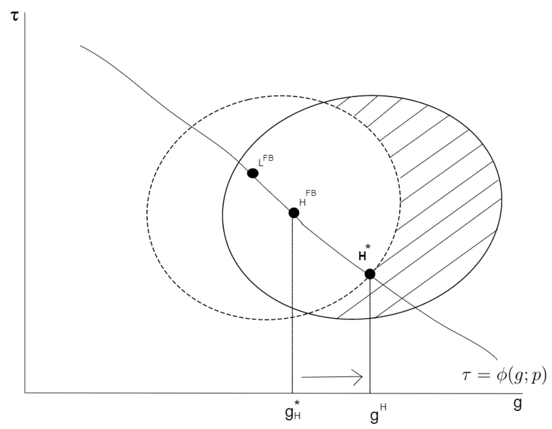


Figure 4.1: Separating Equilibrium

Note that if $p = 1$ these conditions reduce to the same as in the baseline model and if $\lambda = 0$ the first best is attained (for all p $\in [0,1]$). That is, the political cycle result in Rogoff (1990) holds in the application of that framework to a state-dependent utility function and, moreover, the magnitude of the cycle will depend on the parameter p.

Because the equilibrium value of λ also depends on p, Equation4.49 shows that the magnitude of the cycle will depend on p via two channels: through the effect of p on the set $\mathscr{A}$, as captured by the relationship between λ and p, and also through the effect of p on the risk management problem itself, as captured by the function $\phi(g;p)$.

Whether or not the magnitude of cycle increases with p depends on the relationship between $U(c,g)$ and $H(c)$. As for the effect of p on the set $\mathscr{A}$, recall that

$$\mathscr{A} \equiv \{(g,\tau) \mid Z(g,\tau,1,\varepsilon^L;p)\} \leq Z\big(g^*(\varepsilon^L),\tau^*(\varepsilon^L),0,\varepsilon^L;p\big)\} \tag{4.50}$$

so that

$$\mathscr{A}^c \equiv \{(g,\tau) \mid Z(g,\tau,1,\varepsilon^L;p)\} > Z\big(g^*(\varepsilon^L),\tau^*(\varepsilon^L),0,\varepsilon^L;p\big)\} \tag{4.51}$$

where, by definition,

$$Z\big(g^{L^*},\tau^{L^*},0,\varepsilon^L;p\big) = \mathscr{X}^L\pi\big(0\big) + pU(y-\tau^{L^*},g^{L^*}) + (1-p)H(y-\tau^{L^*}) + \beta V(\tau^{L^*}+\varepsilon-g^{L^*}) \tag{4.52}$$

and $g^*(\varepsilon^L) \equiv g^{L^*}$ and $\tau^*(\varepsilon^L) \equiv \tau^{L^*}$ for ease of notation. The set $\mathscr{A}^c$ contains the points at which, if voters' beliefs allow, a low ability incumbent is better off by mimicking the high ability type than by choosing their first-best policy and revealing their type to voters and is represented in figure 4.2 as the points within the small dashed ellipse.

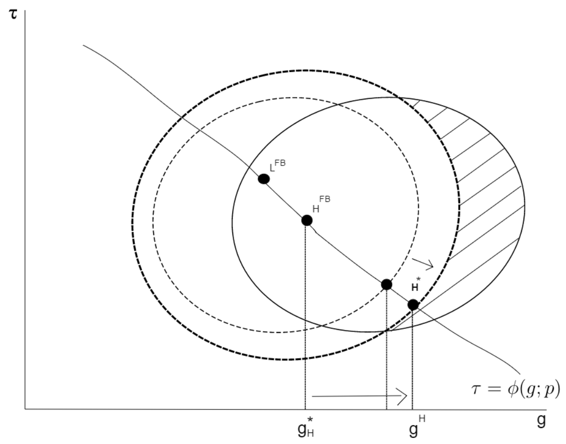


Figure 4.2: Effect of p on the Constraint Set

Note that, if $U(y-\tau^{L^*},g^{L^*}) < H(y-\tau^{L^*})$, then $Z\big(g^{L^*},\tau^{L^*},0,\varepsilon^L;p\big)$ is decreasing in p and a higher value of p expands the solid ellipse, as shown in figure 4.2.

Under these conditions, the set of points at which a low type prefers to mimic a high ability is larger the higher the value of p and, in order to separate themselves from a low ability type, a high ability incumbent needs to increase g even more in election years.

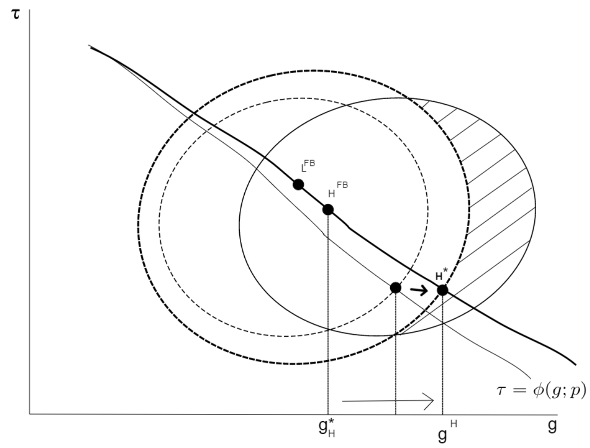


Figure 4.3: Effect of p on $\phi(g;p)$

The effect of p on the magnitude of the cycle through the function $\phi(g;p)$ goes in the same direction and with no additional restrictions besides concavity of U and H. From Equation4.15 it is possible to obtain

$$\frac{\partial\tau(g;p)}{\partial q\partial p} = -\frac{H''U_{22}}{\left(-pU_{11}-(1-p)H''\right)^2} < 0, \tag{4.53}$$

which means that the slope of ϕ decreases as p increases, as shown in figure 4.3. As previously mentioned, depending on the parameters (including p) of the problem it is possible that a high ability incumbent chooses their first-best option and still is able to separate themselves from a low ability type. For given values of ε^H and ε^L, a sufficiently low p could imply a set $\mathscr{A}^c$ small enough and $\phi(g;p)$ sufficiently steep to guarantee that the first-best policy of a high type lies outside the small dashed ellipse, in which case even under asymmetric information no cycle is observed. In other words, the parameters of the problem could be such that a political cycle is observed only if p is sufficiently high.

Finally, it is worth noting that within this framework reelection motives are

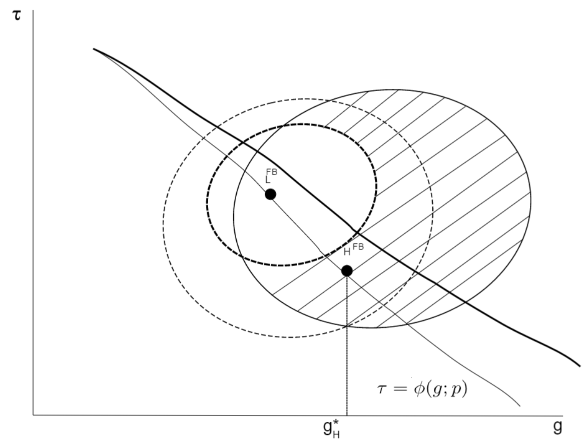


Figure 4.4: An Equilibrium with Low p and no Cycle

account for when voters assess the expected utility of electing a challenger. Therefore, as long as voters interpret the policy as a signal of the politician's underlying ability, incumbents have an opportunity to separate themselves if the parameters of the problem allow. In other words, as long as the expected utility of electing a low ability challenger (who will have reelection incentives when in power) is lower than the expected utility of electing a high ability incumbent for their final term, a separating equilibrium may arise.

4.3 Context and Data

4.3.1 Droughts in the Brazilian Semiarid

The Brazilian Semiarid area comprises ten states, nine in the Northeast and one in the Southeast region, including 1,262 of the 5,570 Brazilian municipalities, according to the Brazilian Institute of Geography and Statistics (*IBGE*). This area is disproportionately poor and rural, and the driest in the country with an historical precipitation average at about half the average for the rest of the country and recurrent droughts that have been a major source of vulnerability and exposure to shocks in the past decades. The Semiarid is characterized by rainy and dry seasons each year, and the concentration of rainfall, combined with its topography and temperature, results in water losses due to quick evaporation and high levels of salinity and low quality of the surface water remaining for consumption (Rocha and Soares, 2012; Bobonis et al., 2019).

The impacts of water scarcity and of interventions to mitigate its effects have also been studied in the literature. For example, Rocha and Soares (2012) document a strong relationship between negative rainfall shocks and shorter gestation periods, lower birth weight, and higher infant mortality, while Da Mata et al. (2021) show that access to cisterns during early pregnancy increased birth weight.

Given the wide range of impacts from recurrent droughts, various policies have been adopted - and addressed during political campaigns (Silva, 2017) - to mitigate such impacts in the region: drilling of water wells, installation of desalination equipment, installation of rainwater catchment units (*cisterns*), the distribution of water by water trucks during emergencies.

The provision of water wells, specifically, has been used strategically by contesting politicians, with anecdotal evidence that drilling of water wells may take place before the election accompanied by the promise that the full installation of the units will follow once the politician is elected (Silva, 2020; Rebouças, 1997). That is, while technically the installation of water wells can be relatively speedy, it is most relevant that the timing of installation can be exploited by contesting politicians.

The prominence of water wells in electoral platforms is not surprising considering that the share of households reliant on water wells is substantial, with a median of 21% in the main sample. And while I do not observe turnout rates for subpopulations nor information about the political platforms of the candidates, at the municipality level I find no evidence that turnout is lower in municipalities with higher shares of households reliant on water wells.

In sum, as an aid to water scarcity water wells do not offer long-term autonomy to voters but its provision can influence short-term voting behaviour since individuals value this technology but, in general, are not able to install and/or operate such units without government support[2].

[2]Cisterns, on the other hand, are much better evaluated by users for the autonomy and long-run reliability provided (Silva, 2020). Consistently with the points above, Bobonis et al. (2019) provide experimental evidence that the installation of cisterns reduces economic vulnerability of voters and, as a result, decreases clientelism and votes for incumbent mayors.

4.3.2 Reelection Rules and Electoral Accountability

In order to study the effect of reelection incentives on the provision of public goods, I take advantage of term limits to Brazilian municipal governments. Mayors of Brazilian municipalities are elected for a mandate of 4 years and are allowed to run for a second consecutive term since the municipal elections in 2000. That is, elected challengers face reelection incentives, while reelected incumbents face term limits and cannot run for a third consecutive term.

Ferraz and Finan (2011) find causal evidence that reelection incentives have significant impacts on corruption: in municipalities where mayors can be reelected, misappropriation of resources is 27% lower than in those where mayors are in their last term. I follow the authors and interpret that second-term mayors, on average, behave as if serving their last term. The argument is that, while the majority of incumbents of Brazilian municipalities run for reelection, less than half are reelected, of whom only a small fraction returns to office after the cooling-off period or run for higher offices after their mayoral mandates. Moreover, even if politicians in their second term still have political career concerns, a comparison between first and second-term mayors would provide a lower bound for the effects of reelection motives.

4.3.3 Data Sources

The outcome variables come from the Underground Water Information System (*Sistema de Informações de Águas Subterrâneas - Siagas*) and from the Water Wells Censuses, both from the Geological Survey of Brazil (*Serviço Geológico do Brasil - CPRM*). The water wells censuses were carried out between 1998 and 2005 and produced cross-sectional, detailed data on individual water wells surveyed in eight estates in the Brazilian Semiarid [3]. The Siagas, on the other hand, is a system in which water wells are registered on an ongoing basis and I use a scrapping program to collect data from the *SIAGAS WEB* and build a municipality level panel data on

[3] Between 1998 and 2005, the censuses covered the States of Ceará, Alagoas, Bahia, Piauí, Paraíba, Pernambuco, Rio Grande do Norte, and Minas Gerais.

the drilling of water wells[4].

Election and politicians data are available at the Electoral Data Repository, from the Superior Electoral Court (Tribunal Superior Eleitoral - TSE). Other municipality characteristics are from the Brazilian Institute of Geography and Statistics (Instituto Brasileiro de Geografia e Estatística - IBGE). I use municipality-level precipitation data produced in Rocha and Soares (2012) based on the *Terrestrial Air Temperature and Terrestrial Precipitation: 1900-2010 Gridded Monthly Time Series*, versions 3.01 and 3.02, respectively[5].

4.3.3.1 Provision of Water Wells

The main outcome of interest is the drilling of water wells. I also use maintenance data to provide additional evidence of a relevant relationship between reelection motives and the working conditions of water wells.

To build measures of construction, I scrape data on individual wells from the *SIAGAS WEB* and, based on the date of drilling, build a monthly series on the number of water wells drilled in each municipality. Monthly data are then aggregated by year and political term as needed. I use population data to compute the number of water wells drilled per 10 thousand population and take the log to obtain the main outcome variable (*Log Water Wells*).

For maintenance, I use data from the Water Wells Censuses carried by the same institution. I build a measure of the share of public water wells in good working condition at the time of the survey. This alternative measure of effort (E) used in the empirical analysis is obtained according to the following:

$$E_i = \frac{\text{\# public wells in use}_i}{\text{\#total}_i - \text{\# abandoned}_i} \tag{4.54}$$

The censuses contain information on the working condition, ownership, and

[4]Between 2001 and 2012, municipalities in all Brazilian states reported drilling of water wells on Siagas. I use data from 9 dry states: Ceará, Alagoas, Bahia, Piauí, Paraíba, Pernambuco, Rio Grande do Norte, Minas Gerais, and Sergipe.

[5]Refer to Rocha and Soares (2012) for details on the construction of monthly series of precipitation by municipality.

use of each equipment surveyed, so it is possible to distinguish the wells installed in public areas used by the community from the private ones, and to observe whether the wells are operational or not. The operational status of each water well is *in use*, *not installed*, *not in use*, or *abandoned*, where *abandoned* units were either dry or blocked, and therefore considered unrecoverable and excluded from the analysis. Water wells are considered public if they are located in state-owned land or if they are located in private land but for public use. The remaining water wells were coded as private.

4.3.3.2 Frequency of dry years

The frequency of dry years is built from municipality-level precipitation data produced in Rocha and Soares (2012) using the *Terrestrial Air Temperature and Terrestrial Precipitation: 1900-2010 Gridded Monthly Time Series*, versions 3.01 and 3.02, respectively.

The precipitation series starts in 1938, so the frequency of dry years at election year t and municipality i (DR_{it}) is defined as the number of years in which annual precipitation was below the long-term annual average, from 1938 until the election year t, divided by the total number of years between 1938 and t. Formally,

$$DR_{it} = \frac{\sum_{1938}^{t} [\mathbb{1}(prec_{is} < cn_i)]}{t - 1938 + 1} \tag{4.55}$$

where $prec_{is}$ denotes the annual precipitation at municipality i and year s and cn_i denotes the municipality-level climate normals, a long-term annual average as defined by the *World Meteorological Organization*. Specifically, climate normals are 30-years averages updated every ten years. I use 1961-1990 climate normals because this period precedes all election years in the main sample, but the results are virtually unchanged if I use 1971-2000 climate normals or the average yearly rainfall from 1938 to the election year of each political term.

4.3.3.3 Political data

Mayors of Brazilian municipalities are elected for a mandate of 4 years and are allowed to run for a second consecutive term since the municipal elections in 2000.

Availability of data on the drilling of water wells declines after 2014, so the main sample includes observations from 2001 to 2012 (aggregated by political term) and covers the 2000, 2004, and 2008 elections.

To determine whether in municipality i the mayor is in their first or second mandate at term t, I use data on the mayoral election in term $t-1$ and compare the politician identifier for the winner in each election. More specifically, depending on data availability, I match winners using the candidate's taxpayer unique identification number, voter identification number, or full name[6]. If the incumbent at term t is different from the winner at term $t-1$, then they are eligible to run for a second term.

For our identification strategy it is also necessary to identify all municipalities in which the incumbent run for reelection in order to select the incumbent-challenger mixed elections. Using the same matching procedure outlined above, for each municipality I compare incumbent mayors elected for term t to all candidates who run the election for term $t+1$. If the incumbent mayor at term t is identified as one of the candidates for term $t+1$, then the election at $t+1$ is an incumbent-challenger mixed election.

Once all incumbent-challenger mixed elections have been identified, I compute the vote share of the previous incumbent and the vote share of the challenger with most votes. The vote share of candidate j is defined as the number of valid votes for candidate j divided by the total number of valid votes in municipality i. As the main purpose of this paper is to identify the effect of having a first-term mayor in power, treatment status is a function of the margin of victory of the challenger, which is defined as the difference between the vote share of the challenger and the vote share of the previous incumbent. As will be discussed below, in a regression discontinuity (RD) design, the margin of victory of the challenger is the forcing variable that determines treatment status: if the margin of victory of the challenger is positive, the municipality is assigned a first-term mayor with reelection incentives.

[6] While the taxpayer number is an unique identifier, this information is missing for a large number of candidates. Additionally, the voter identification number may differ between elections. So matching names is still necessary for completion and validation. Name strings are processed for removal of all special characters, upper cases, and spaces.

Otherwise, the municipality is assigned a second-term mayor who cannot run for a third consecutive term.

Finally, data on gender, age and education of all candidates is also available at the Electoral Data Repository. Missing data on gender are imputed with a program based on the candidate's first name and census data developed in Meireles (2018). Age at time of election is computed from the date of birth. For educational attainment, I compute an indicator equal to one if the candidate has at least a college degree.

4.3.4 Main sample and descriptive statistics

Our analysis comprises two samples and sets of results. The main empirical strategy uses data from the Siagas on the drilling of water wells between 2001 and 2012 as the outcome of interest. Data from the Water Censuses on the working condition of public water wells between 1998 and 2005 is used in an additional analysis.

Table 4.1 presents the difference in means between municipalities with first and second-term mayors included in our main analysis. The sample includes municipalities in dry states in which a mixed incumbent-challenger election took place. Overall, we do not observe statistically significant or economically relevant differences between the two groups.

Table 4.1: Differences in Means - Municipalities with First- and Second-term Mayors

	All	2nd-term mayors	1st-term mayors	Diff	P-Value
College	0.38	0.38	0.39	0.011	0.731
Female	0.07	0.06	0.07	0.011	0.494
Age	48.72	49.06	48.24	-0.815	0.270
Log Population	9.76	9.76	9.75	-0.015	0.799
Log per capita GDP	8.01	8.03	7.98	-0.056	0.210
Access to Water Networks (%)	0.56	0.56	0.57	0.006	0.647
Frequency of dry years	0.55	0.55	0.54	-0.007	0.147
Log Water Wells	0.77	0.78	0.75	-0.033	0.665
Observations	1,081	678	403	-	-

Table 4.2 presents descriptive statistics for the variables and municipalities included in the additional empirical analysis, which focuses on the relationship between the time to election and the good maintenance of public water wells and includes 1,096 municipalities in the Brazilian Semiarid with detailed information about the timing of measurement of the dependent variable.

Table 4.2: Summary Statistics - Municipalities

Statistic	Obs.	Mean	St. Dev.	Min	Max
% of public water wells in use	1,096	0.63	0.26	0.00	1.00
Months to election	1,096	16.31	7.47	0	34
Drought risk	1,096	0.62	0.07	0.35	0.75
Log municipal per capita GDP	1,096	0.43	0.47	−1.05	3.28

Figure 4.5 shows the distribution of the share of water wells in good work conditions with the observations at the extreme values of zero and one. On average, 63% of available public water wells were in use at the time of measurement in the municipalities in the sample. While in 10% of the municipalities in the sample (113) all the public water wells were in good working conditions, the empirical strategy aims at explaining the heterogeneity observed across the remaining units. I keep the municipalities where none or all water wells were in use are included in the sample and use a fractional logit model to account for the domain of the outcome variable.

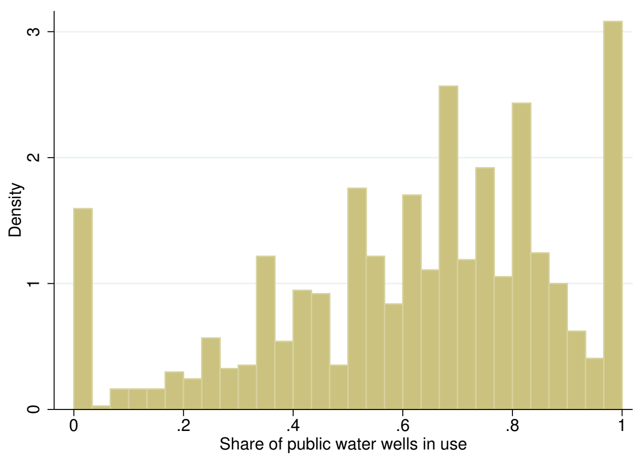


Figure 4.5: Distribution of Share of Working Water Wells

4.4 Empirical Strategy

4.4.1 Econometric model

The main objective of this paper is to study the relationship between reelection incentives and the provision of public goods. More specifically, we investigate whether first-term mayors, who are allowed to run for reelection, provide more water wells than second-term mayors with no reelection incentives.

However, a simple comparison between municipalities with first and second-term mayors could be biased if the election of challengers is correlated with unobserved municipality-specific determinants of the outcome of interest.

In order to control for unobserved characteristics of the municipalities, I carry out a regression-discontinuity (RD) analysis (Lee, 2008; Lee and Lemieux, 2009) and, as in Ferraz and Finan (2011), use close elections to compare municipalities in which a challenger barely won or lost an election after running against an incumbent mayor. More specifically, $\forall$ pairs (i,t) such that $MV_{it} \in [c-h, c+h]$, I use the regression discontinuity (RD) specifications below to estimate the treatment effects of interest:

$$y_{it} = \alpha + \boldsymbol{W}(FT_{it}, DR_{it})'_{it}\boldsymbol{\beta} + \boldsymbol{X}'_i\gamma + \boldsymbol{Z}_t + f(MV_{it}) + \varepsilon_{it}, \tag{4.56}$$

where i indexes municipalities and t is an index for term, y_{it} is the outcome variable of interest, FT_{it} is the treatment and indicates whether the mayor is in their first term, DR_{it} is a measure of the frequency of dry years at the beginning of term t, $\boldsymbol{W}(\cdot,\cdot)$ is a set of functions of FT_{it} and DR_{it}, $\boldsymbol{X}'_i$ denotes a set of relevant municipal and mayor characteristics, $\boldsymbol{Z}_t$ is a set of term dummies, MV_{it} is the challenger margin of victory in municipality i in term t, which is defined as the vote share of the challenger candidate minus the vote share of the previous incumbent, and ε_{it} are unobserved municipal and mayor characteristics that affect the outcome of interest in municipality i during term t.

Treatment status of municipality i (FT_{it}) equals to 1 if the mayor is in their first term. Within the subset of municipalities in which the incumbent mayor ran for reelection against challengers, treatment status is determined based on the margin of victory of the challenger candidate with most votes. That is, treatment assignment is based on the forcing variable MV_{it} and on a cutoff $c = 0$, where

$$FT_{it} = 1\{MV_{it} \geq 0\} \tag{4.57}$$

and $1\{\cdot\}$ is the indicator function. So in municipalities above the cutoff citizens are represented by a first-term mayor who can run for reelection, while municipalities below the cutoff have reelected the incumbent and received a second-term mayor who does not have reelection incentives.

The control function $f(MV_{it})$ is a polynomial of order n on each side of the cutoff. I consider different functional forms for $f(\cdot)$ in order to allow for nonlinearities and for different relationships between the running variable MV_{it} and the outcome above and below the cutoff point. Not modelling the relationship between MV_{it} and y_{it}, restricting such relationship to be the same above and below the cutoff, or assuming linearity when the relationship is nonlinear could result in biased estimates. The optimal bandwidth $\hat{h}$ is computed by bandwidth selectors developed in Calonico et al. (2018, 2019, 2020).

4.4.2 Validity checks

The internal validity of the regression discontinuity design relies on the assumption that potential outcomes are a continuous function of the running variable at the threshold, which is satisfied if individuals are unable to precisely manipulate the assignment variable. Under this assumption, individuals just above and just below the cutoff will have the same probability of receiving the treatment, even if they have imprecise control over the running variable (Lee and Lemieux, 2009). As a result, variation in treatment around the cutoff is as random, so units just above the cutoff are not systematically different from those just below the threshold.

To check for the presence of bunching, below I test if the margin of victory is continuous at the cut off. Figure 4.6 shows the histogram of the running variable and figure 4.7 presents the formal density test from McCrary (2008), which tests for a jump at the threshold.

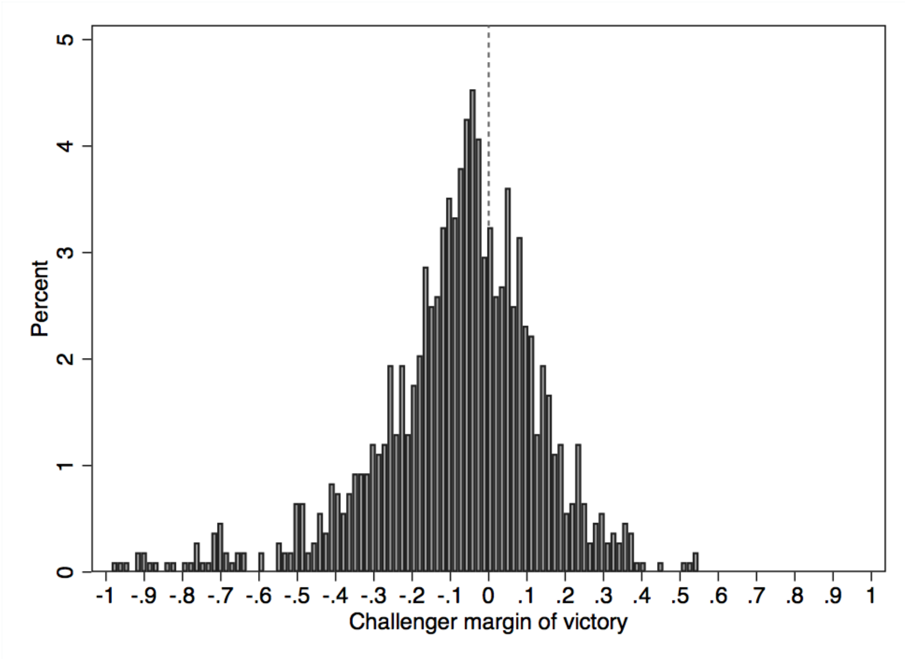


Figure 4.6: Histogram - Global

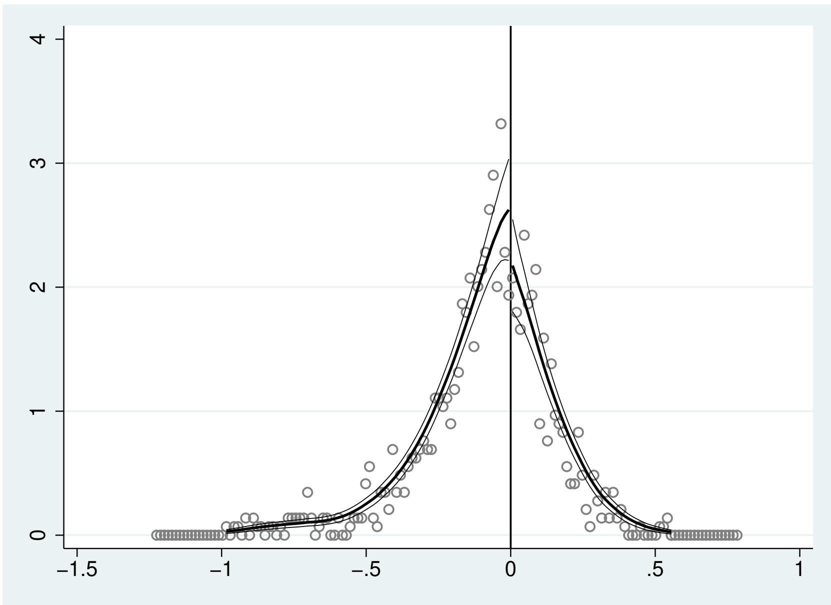


Figure 4.7: Density test (McCrary, 2008): Discontinuity estimate= -.175; S.E.= .124; bin size = .013; bandwidth = .235

Figure 4.6 is consistent with the presence of a type of incumbency advantage, since challengers win with an advantage of at most 55 percentage points, while some incumbents win with almost all votes. However, precise manipulation of election results is not likely and the formal McCrary (2008) test does not reject the null hypothesis that the density is continuous at the cutoff.

In figure 4.8 I address the concern that other changes at the cutoff affect the outcome and that such effects are erroneously attributed to the treatment. I inspect whether there are significant jumps in the covariates used in the regressions and other pre-treatment variables at the cutoff.

The bottom left panel shows that the average frequency of dry years just below the threshold is similar to just above. Municipalities with first and second-term mayors are not systematically different regarding their pre-election exposure to droughts: the frequency of dry years during the term prior to the election was similar for both groups and, in the election year, treatment and control groups had similar shares of cities with precipitation was below average.

I also find that treatment and control groups are similar with respect to the share of the population reliant on water wells and to the share with access to the main water networks. I do not find population and per capita income discontinuities at the cutoff. Mayor-level covariates such as age, gender, and education also do not show significant differences around the cutoff.

The RDD design does not control for mayor-level endogeneity, so bias could arise if mayors select into rerunning according to their types. While I cannot directly address unobserved mayor heterogeneity, Figure 4.8 does not show any discontinuities on observable mayor characteristics around the cutoff. Moreover, Tables B.1 and B.2 in Appendix B show no correlation between treatment status and the provision of water wells as predicted by observable mayor characteristics. So I find no indication that the local treatment effects are driven by politician's selection into treatment based on observables.

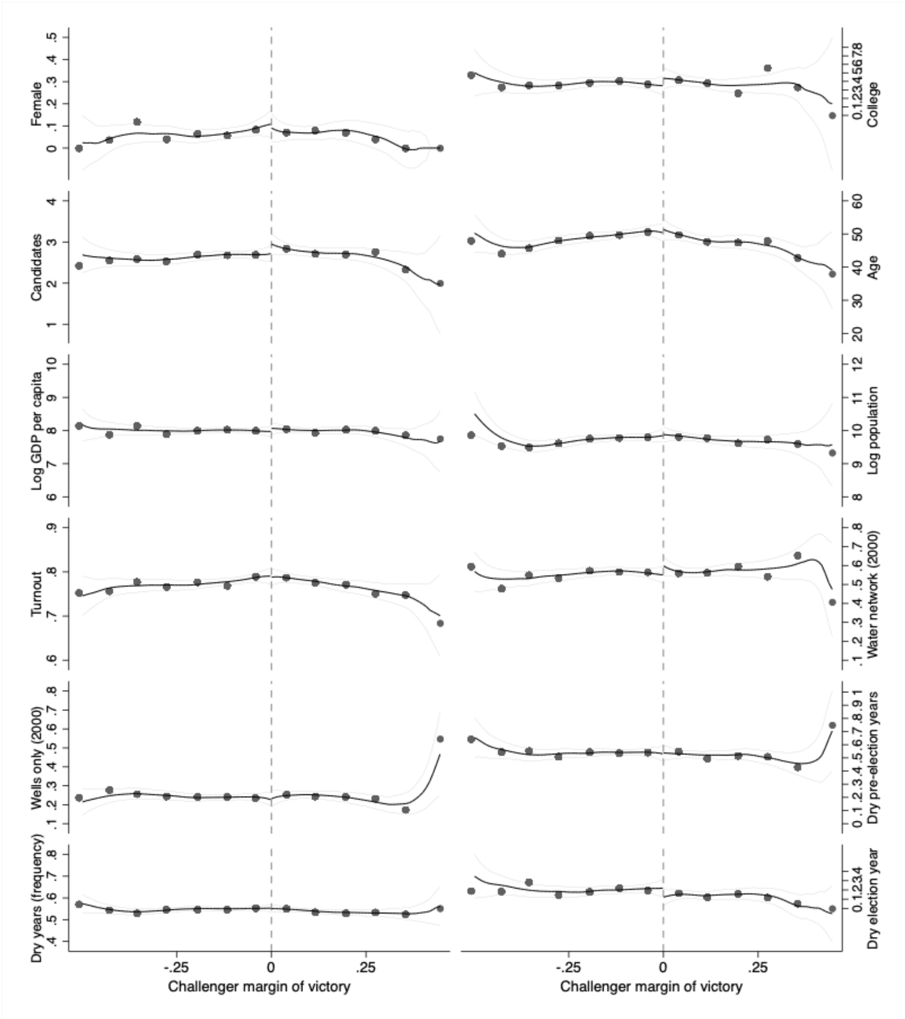


Figure 4.8: Balance in covariates

4.5 Results

4.5.1 The heterogeneous effects of reelection incentives

The predictions of the model in Section 4.2 imply that the effect of electing a first-term mayor on the provision of water wells is heterogeneous, depending on the frequency of dry years. To test this prediction, the specification below allows for an interaction term between having a first-term mayor (FT_it) and the frequency of dry years (DR_it):

$$y_{it} = \alpha + \beta_0 FT_{it} + \beta_1 DR_{it} + \beta_2 (FT_{it} \times DR_{it}) + \boldsymbol{X}_i' \gamma + \boldsymbol{Z}_t + f(MV_{it}) + \varepsilon_{it}. \quad (4.58)$$

Tables 4.3 and 4.4 present the main sets of results. Municipality-level controls are population and log GDP per capita, while the mayor-level controls include a indicator for gender, for having at least a college degree, and party affiliation.

Columns 1 an 2 in Table 4.3 show the raw OLS estimates for the coefficients of interest, based on an cross-section of all municipalities in dry states of northeast Brazil for which data on the outcome variable is available. I find evidence that having a first-term mayor is associated to higher provision of water wells, as long as the frequency of dry years is sufficiently high. The Global linear specification, which restricts the sample to municipalities with mixed incumbent-challenger elections corroborates the baseline OLS results.

Table 4.3: Interactions between Reelection Incentives and Precipitation Profile

	(1) OLS	(2) OLS	(3) Global	(4) Global
First term	-0.937***	-1.275***	-0.987***	-1.316***
	(-12.50)	(-13.74)	(-6.84)	(-22.61)
First term=1 × Frequency of dry years	1.672***	2.224***	1.689***	2.240***
	(10.71)	(10.32)	(8.59)	(17.86)
Term dummies	No	Yes	No	Yes
Mun. controls	No	Yes	No	Yes
Mayor controls	No	Yes	No	Yes
Observations	1081	1081	1081	1081
R^2	0.006	0.229	0.006	0.230

t statistics in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

These estimates, however, may be biased. For example, in cities where citizens are more risk averse they may be more likely to elect a second-term mayor and to put higher weights on the availability of water wells as an alternative or redundant source of water, which could bias downwards the effect of the having a first-term mayor. Similarly, better administrative capacity at the city-level could be associated to higher rates of incumbent reelection and to higher levels of drilling. In other words, risk preferences of citizens and other city-level unobservables could affect both voting decisions and and the provision of water wells.

The RDD specifications in Table 4.4 control for such municipality-level confounders. In columns 1, 3, and 5 I present the estimates without controlling for extra covariates, while the specifications in Columns 2, 4, and 6 include term dummies, municipality, and mayor controls. Including covariates significantly increases the R^2 but does not change the estimates substantially, which is consistent with the interpretation of treatment assignment being as good as random near the cutoff.

In all specifications, on average, the provision of water wells by first-term mayors is larger as long as the frequency of dry years is sufficiently high, with the largest effects of reelection incentives arising from the driest municipalities, where the distribution of annual precipitation is positively skewed. Low frequency of dry years, on the other hand, has a negative local treatment effect, which is consistent with first-term mayors allocating effort into the provision of other goods with higher reelection returns if the risk of a drought is low.

Figures 4.9 and 4.10 present the marginal RD effects as a function of the frequency of dry years. Within the optimal bandwidth (Figure 4.9), the RD treatment effect at the cutoff is positive if the frequency of dry years is higher than 0.46 and increases steeply with the frequency of dry years in both specifications, with and without controls (panels (a) and (b)). Interestingly, the drilling of wells by first term mayors is lower than by second term mayors, on average, if the frequency of dry years is low. This pattern is consistent with mayors eligible for reelection prioritizing other types of public goods where drought risk is less salient. The relative impact of reelection incentives on the provision of public goods unrelated to drought risk shall be assessed in future work.

With a bandwidth twice as large (Figure 4.10), the local treatment effect is positive as long as the frequency of dry years is larger than .53 (panels (a) and (b)) and increases steeply as with the frequency of dry years as in the optimal-bandwidth model.

These findings relate to a broader literature on the relationship between term limits and fiscal performance. Overall, if voters also care about fiscal austerity, the

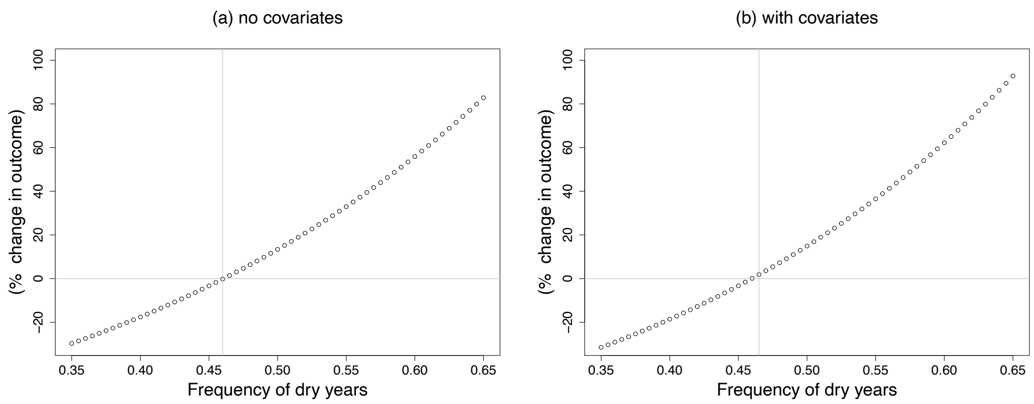


Figure 4.9: Local treatment effect at the cutoff - h

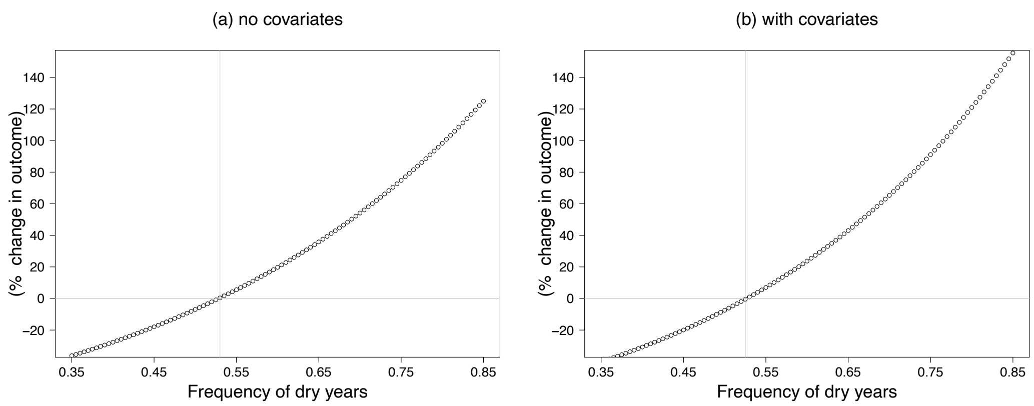


Figure 4.10: Local treatment effect at the cutoff - 2h

Table 4.4: Interactions between Reelection Incentives and Precipitation Profile

	(1) Linear, h	(2) Linear, h	(3) Linear, 2h	(4) Linear, 2h	(5) Linear, 0.5h	(6) Linear, 0.5h
First term	-1.467*	-1.585***	-1.333**	-1.528***	-0.889*	-0.858*
	(-2.49)	(-9.66)	(-3.98)	(-13.71)	(-2.63)	(-2.56)
First term=1 $\times$ Frequency of dry years	3.185*	3.448***	2.522**	2.901***	1.678	1.564
	(2.92)	(13.68)	(3.62)	(8.11)	(1.64)	(1.57)
Term dummies	No	Yes	No	Yes	No	Yes
Mun. controls	No	Yes	No	Yes	No	Yes
Mayor controls	No	Yes	No	Yes	No	Yes
Observations	547	547	834	834	288	288
R^2	0.020	0.354	0.011	0.238	0.017	0.427

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

effect of term limits on fiscal performance is theoretically ambiguous (Johnson and Crain, 2004; Reed et al., 1998): on the one hand, term-limited incumbents (second-term mayors in this application) do not have electoral incentives, but on the other hand they have more experience and may have lower austerity reputation-building concerns, which could result is higher levels of overall spending and provision of public goods, even if in deviation of voter's preferences. Similarly, politicians eligible for reelection may respond to electoral motives, including being more concerned about building a reputation of austerity, while being less familiar with the state bureaucracy. That is, the net effect of electoral incentives on the provision of public goods would depend, among other factors, on how incumbents perceive voter's weights on austerity versus the provision of public goods, and the extent to which experience reduces the cost of effort.

I find that first-term mayors outperform term-limited incumbents. In this deprived setting, it is reasonable to assume that voters value access to water and other basic public goods more than austerity, so that first-term mayors are not responding to austerity reputation-building concerns. Moreover, the returns to experience for the provisions of water wells could be limited, meaning that voter's preferences over the public good consumption could be the key driver of the incumbent's performance. Finally, my findings are consistent with term-limited incumbents being more likely to deviate from voter's preferences over the provision of water wells.

In Table 4.5 I show that the results above are robust to using quadratic and cubic control functions. Column (1) shows the estimates from a model linear in the running variable (as shown above in Column (2) of Table 4.4), and Columns (2) and (3) indicate that including polynomials of higher order does not change the magnitude nor the significance of the coefficients substantially.

In sum, the results obtained from different specifications of Equation4.56 suggest that the raw estimates from a simple OLS are biased and that controlling for observable and unobservable municipality level heterogeneity by adding covariates and using a RD design results in positive and statistically significant local treatment effects in most cases. Additionally, allowing for an interaction between first-term

Table 4.5: Polynomial control functions

	(1) Linear, h	(2) Quadratic, h	(3) Cubic, h
First term	-1.585***	-1.821***	-1.809***
	(-9.66)	(-13.28)	(-12.69)
First term=1 × Frequency of dry years	3.448***	3.361***	3.302***
	(13.68)	(9.47)	(6.83)
Term dummies	Yes	Yes	Yes
Mun. controls	Yes	Yes	Yes
Mayor controls	Yes	Yes	Yes
Observations	547	547	547
R^2	0.354	0.361	0.362
AIC	1529.7	1523.4	1522.7

t statistics in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

and the frequency of dry years indicates that the RD treatment effect of reelection incentives on the provision of water wells is driven by areas where dry years are more frequent, and it is significantly larger where the distribution of precipitation is positively skewed.

It is important to note that, while the magnitude of the effects in the local linear regressions is sensitive to the choice of bandwidth, all specifications suggest a positive treatment effect as long as the frequency of dry years is sufficiently high. Unfortunately, within narrow bandwidths the number of observations is low, but, overall, increasing the bandwidth reduces the magnitude of the local treatment effect, which is consistent with the hypothesis of unobserved city heterogeneity biasing down the estimates that include observations far from the cutoff.

Additionally, the RD design does not control for unobserved mayor heterogeneity. Nevertheless, the top panels in Figure 4.8 do not indicate significant discontinuities in observable mayor characteristics at the cutoff. Finally, the main results are obtained under the assumption that the relationship between the outcome variable and the running variable does not vary with the frequency of dry years.

The next section addresses this issue and presents additional evidence of a political cycle in the provision of water wells in the Brazilian semiarid.

4.5.2 Additional Results

4.5.2.1 Control Function Specification: Robustness Checks

The main results above assume that the relationship between the outcome and the running variables does not vary with the frequency of dry years. This issue has been overlooked in the literature but is formally raised in Carril et al. (2019), where the authors consider a model in which treatment status is interacted with a binary indicator of subgroup membership. In Table 4.6 I present the results from specifications that use the optimal bandwidth and include interactions between the running variable and the frequency of dry years, linear, and quadratic.

In general, allowing for an interaction between the margin of victory and the frequency of dry years does not reduce the *AIC* significantly, so based on this criteria there is no strong evidence against the simpler specification.

Additionally, qualitatively the relationship between reelection incentives, frequency of dry years, and log wells is the same even though the magnitude and significance of the coefficients decline as the additional interactions are allowed, as shown in Table 4.7.

Table 4.6: Additional RD specifications

	(1) Linear, h	(2) Lin. w/ full interaction, h	(3) Quadratic, h	(4) Quad. w/ full interaction, h
First term	-1.585***	-0.492	-1.821***	-0.913*
	(-9.66)	(-2.06)	(-13.28)	(-2.49)
First term=1 × Frequency of dry years	3.448***	1.456*	3.361***	1.686
	(13.68)	(3.11)	(9.47)	(2.01)
Term dummies	Yes	Yes	Yes	Yes
Mun. controls	Yes	Yes	Yes	Yes
Mayor controls	Yes	Yes	Yes	Yes
Observations	547	547	547	547
R^2	0.354	0.355	0.361	0.364
AIC	1529.7	1528.7	1523.4	1520.6

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4.7: Linear and full interaction

	(1) Lin. w/ full int., h	(2) Lin. w/ full int., 2h	(3) Lin. w/ full int., 0.5h
First term	-0.492	-1.102**	-1.491
	(-2.06)	(-4.06)	(-1.38)
First term=1 $\times$ Frequency of dry years	1.456*	2.127**	2.728
	(3.11)	(3.71)	(1.77)
Term dummies	Yes	Yes	Yes
Mun. controls	Yes	Yes	Yes
Mayor controls	Yes	Yes	Yes
Observations	547	834	288
R^2	0.355	0.238	0.427

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

4.5.2.2 Time to Election and Maintenance of Water Wells

Next I use data on the share of the public water wells in use to investigate whether reelection motives affect the maintenance of water wells in dry areas. Given the cross-sectional nature of the Water Wells Census, the following regression is used to estimate a relationship between time to election and the share of public water wells in use:

$$E_i = \beta_0 + \beta_1 DR_i + \beta_2 TE_i + \beta_3 (DR_i \times TE_i) + Z_i \upsilon + \varepsilon_i, \tag{4.59}$$

where E_i is the share of public water wells in use for municipality i, DR_i is the frequency of dry years (defined as precipitation below the climate normals), TE_i is the number of months to the next election in municipality i at the time of measurement of E_i, and the vector Z_i is a set of municipal characteristics that affect the provision of water wells.

Table 4.8 consolidates the regression results from different versions of our main Equation. Column 1 shows that the higher the frequency of dry years, the higher the share of public water wells in use and that this relationship is statistically significant. Columns 2 and 3 report the unconditional relationship between the number of months to the next municipal election and the share of public water wells in good working conditions. Even if controlling for the municipal per capita income (in log) the point estimates on the political cycle for public water wells are statistically indistinguishable from zero. That is, without additional controls there is no evidence of a political cycle affecting the provision of water wells. The specifications presented in Columns 4 and 5 control for the frequency of dry years and include an interaction term to allow for the presence of a political cycle conditional on the drought risk profile of each unit.

Besides being statistically significant at the one percent level, the estimates are economic relevant. The results indicate that, as long as the drought risk is sufficiently high, the share of public water wells in use is lower the longer the remaining time to the next election. Under the specification in Column 5, only in municipalities where the frequency of dry years is higher than 63% incumbent mayors increase the provision of public water wells as the election approaches. This finding is consistent with the idea that only in municipalities with a high drought risk profiles water wells are perceived as potentially valuable public goods and, therefore, used as part of a political cycle.

Table 4.8: The Political Cycle and Water Wells Maintenance: Linear Model

Dependent variable	Share of public water wells in use				
	OLS	OLS	OLS	OLS	OLS
	(1)	(2)	(3)	(4)	(5)
Months to election		-0.0014	.0009	0.0360***	0.0377***
		(0.001)	(0.001)	(0.012)	(0.012)
Drought risk	0.6529***			1.6045***	1.5591***
	(0.122)			(0.3383)	(0.334)
Months to election ×				-0.0607***	-0.0597***
Drought risk				(0.020)	(0.019)
R^2	0.054	0.001	0.02	0.04	0.06
Observations	1,096	1,096	1,096	1,096	1,096
Municipal characteristics	Yes	No	Yes	No	Yes

Notes: t statistics in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

This is the case of 532 municipalities, which corresponds to about 48% of the sample being considered. Among the municipalities where a political cycle is observed, the timing of elections (in months) is associated to up to 0.023 percentage point decrease in the share of public water wells in good working conditions. Where the frequency of dry years is the highest, in 12 months this would represent a decrease of 8.7 percentage points in the share of wells in use, a 13 percent decline from the average level (0.623). The lower bound of these estimates, however, are much smaller. In municipalities where the frequency of dry years is just high enough to induce a political cycle, the timing of elections is associated to 0.0002 percentage point decrease in the outcome variable, which corresponds to a decline of 0.2 percentage points in a 12-month period. That is, the OLS estimates also suggest that once the frequency of dry years is high enough, the magnitude of the political cycle increases steeply with drought risk.

Table 4.9: The Political Cycle and Water Wells Maintenance: Fractional Logit

Dependent variable	Share of public water wells in use				
	Fractional logit	Fractional logit	Fractional logit	Fractional logit	Fractional logit
	(1)	(2)	(3)	(4)	(5)
Months to election		-0.0061	0.0045	0.1568***	0.1657***
		(0.004)	(0.005)	(0.055)	(0.055)
Drought risk	2.7884***			6.9474***	6.7835***
	(0.519)			(1.475)	(1.455)
Months to election ×				-0.2650***	-0.2627***
Drought risk				(0.087)	(0.085)
Observations	1,096	1,096	1,096	1,096	1,096
Municipal characteristics	No	Yes	No	Yes	No

Notes: t statistics in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Another issue that arises from using a simple OLS to model the share of working water wells is that the domain of the dependent variable is the unit interval (closed) and the linear model produces fitted values outside this range. One approach to modeling proportions is to use the quasi-maximum likelihood (QMLE) fractional logit estimator in Papke and Wooldridge (1996). The results obtained from this specification are presented in Table 4.9. Column 4 of Table 4.9 contains the results of estimating the complete version of the main equation by QMLE, with the variables of interest highly statistically significant and the same directions of effects as in the linear model.

The level of the QMLE marginal effects are slightly lower than those obtained by OLS, but qualitatively the results are similar to those obtained in the linear model: it takes a certain level of drought risk to make water wells part of a political cycle, and among the municipalities where a political cycle is observed, the magnitude of the cycle also increases steeply with the frequency of dry years. At the average per capita income and an annual frequency of dry years of 65%, 12 months prior to the election the share of public water wells in use is 1.4 percentage lower than in the election month. At the highest level of drought risk in the sample (75%), the decline amounts to 7.4 percentage points. Moreover, based on the QMLE estimates the drought risk turning point is only slightly higher. At a frequency of dry years of 63%, 12 months prior to an election the share of wells in good working condition is only 0.01 percentage point lower, so both the OLS and QMLE models produce essentially the same results.

Finally, given that the measurement of the dependent variable was carried out by a national-level institution in coordination with state governments, the timing of measurement could be capturing state effects along with the remaining time to municipal elections. Table 4.10 presents the QMLE point estimates for an specification of Equation 4.59 that includes state dummies.

The results indicate that, at an annual probability of a dry year of 75%, the share of water wells in use in *Piauí* 12 months before municipal elections is 3.8 percentage points lower than immediately after election (about 6% below the state average), while in *Paraíba* the magnitude of the cycle is higher, with a 5.9 percentage point decline (approximately 8.8% lower than the state average).That is, allowing for different intercepts indicates that the magnitude of the cycle can be significantly different across states.

Table 4.10: The Political Cycle and Drought Risk Management: State Effects

Dependent variable	Share of public water wells in use	
	Fractional logit	Fractional logit
	(1)	(2)
Months to election	0.1657***	0.1970***
	(0.055)	(0.057)
Drought risk	6.7835***	6.4908***
	(1.455)	(1.564)
Months to election ×	-0.2627***	-0.2928***
Drought risk	(0.085)	(0.881)
Municipal characteristics	Yes	Yes
State dummies	No	Yes
Bahia	-	0.2724
		(0.192)
Ceará	-	0.1301
		(0.175)
Minas Gerais	-	0.4332**
		(0.204)
Paraíba	-	0.3706**
		(0.182)
Pernambuco	-	0.4983***
		(0.184)
Piauí	-	0.8977***
		(0.172)
Rio Grande do Norte	-	0.5399***
		(0.184)
Observations	1,096	1,096

Notes: t statistics in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

4.6 Concluding Remarks

Incumbents with political career concerns can use policy to signal their types to voters and distinguish themselves from challengers. However, whether specific policies are affected by reelection concerns of incumbents should depend on how voters value the policy in question.

This paper studies the relationship between reelection incentives and the provision of water wells in dry areas of the Northeast and Southeast Brazil. Using a regression discontinuity design, I show that reelection incentives increase the drilling of water wells as long as the frequency of dry years is sufficiently high and that this effect is significantly larger in the driest areas. In order to allow for heterogeneous effects, I interact treatment status with a measure of the frequency of dry years. I also show a negative relationship between the maintenance of public water wells and time to election in municipalities where drought frequency is sufficiently high.

The above relationship is explained in the context of a Rogoff-type model (Rogoff, 1990) of political budget cycles, which I modify to allow for the provision of a public good from which voters derive utility only under certain states of nature (state-dependent utility).

These findings are in line with a body of literature that studies the effects of electoral institutions on the behaviour of politicians and voters, finding relevant impacts on policy and welfare outcomes.

Chapter 5

General Conclusions

The design of political institution may significantly affect the behaviour of voters and politicians, with major implications for economic and social welfare outcomes.

In this paper I studied the role of political parties in explaining important outcomes of legislators in office and the effects of reelection incentive on policy outcomes of major importance to citizens exposed to weather related shocks.

Firstly, I find that parties play an important distributive role in the competition for a type of public funding that can be used for pork-barrel and, as such, increase the probability of reelection and the political career progression of incumbent legislators in Brazil. More specifically, I find that the variance in access to budgetary amendments is mostly explained by differences in individual heterogeneity and that negative sorting of politicians with lower influence into parties with higher party premiums contributes to limit the observed inequalities in access to funding.

This relationship is particularly relevant for the outcomes of female legislators. More specifically, I document a gender gap in access to funding, which I decompose into person and party related components. I find that the gender gap is mainly explained by differences in individual heterogeneity and that the allocation of females into parties with higher fixed effects contributes to partially offset the disadvantage of females. I also find evidence that this pattern of negative sorting is associated to ideology, as left-wing parties tend to have higher party premiums but be composed by politicians with lower individual influence. As females tend to segregate in left-wing parties, this benefit is particularly important for their outcomes.

Regarding the effects of reelection incentives, I exploit term limits on Brazilian mayors and, using a regression discontinuity design, I find evidence of heterogeneous effects of reelection concerns. More specifically, in a context of recurrent droughts in the remote Brazilian Semiarid, I find that first-term mayors, who are eligible for reelection, provide more water wells to citizens than second-term mayors, as long as the frequency of dry years in the municipalities is sufficiently high. These results are robust to different specifications. I also show a negative relationship between time to the next election and the share of water wells in good working condition.

In sum, the studies in this project contribute to a large body of literature that investigates the impacts of political institutions on policy and welfare outcomes. I focus on the provision of public goods particularly valuable to vulnerable populations and on gender gaps in access to sources of funding that have the potential to hinder the political careers of women in office. As increasing female representation in politics is an important channel to reduce inequalities, Chapter 2 also brings a discussion of importance to economic development and inequality reduction.

Finally, Chapter 3 extends the gender gap analysis to a broader variance decomposition and investigates in detail the application of a two-way fixed effect model, which is a framework widely used in fields such as Labour and Education Economics, to questions pertaining to Political Economy and Gender Economics, and to the Political Sciences discipline.

Appendix A

Evolution of Brazilian Parties

Brazil has a large number of active parties and over the period considered we observe the creation of new parties, merging parties, and parties changing names. For the purposes of adequately computing party-level variables, we track the changes between 2006 and 2018, which are summarized in Table A.1.

Table A.1: Evolution of Brazilian Parties

party_sigla	party_number_18	party_sigla_18	party_name_18	comments
AVANTE	70	AVANTE	AVANTE	No changes
DC	27	DC	DEMOCRACIA CRISTÃ	No changes
DEM	25	DEM	DEMOCRATAS	No changes
MDB	15	MDB	MOVIMENTO DEMOCRÁTICO BRASILEIRO	No changes
NOVO	30	NOVO	PARTIDO NOVO	No changes
PATRI	51	PATRI	PATRIOTA	No changes
PCB	21	PCB	PARTIDO COMUNISTA BRASILEIRO	No changes
PCdoB	65	PCdoB	PARTIDO COMUNISTA DO BRASIL	No changes
PCO	29	PCO	PARTIDO DA CAUSA OPERÁRIA	No changes
PDT	12	PDT	PARTIDO DEMOCRÁTICO TRABALHISTA	No changes
PEN	51	PATRI	PATRIOTA	Name changed in 2016
PFL	25	DEM	DEMOCRATAS	Name changed in 2007
PHS	31	PHS	PARTIDO HUMANISTA DA SOLIDARIEDADE	No changes
PL	22	PR	PARTIDO DA REPÚBLICA	PL and PRONA merged into PR in 2006
PMB	35	PMB	PARTIDO DA MULHER BRASILEIRA	No changes
PMDB	15	MDB	MOVIMENTO DEMOCRÁTICO BRASILEIRO	Name changed in 2016
PMN	33	PMN	PARTIDO DA MOBILIZAÇÃO NACIONAL	No changes
PMR	10	PRB	PARTIDO REPUBLICANO BRASILEIRO	Name changed in 2006
PODE	19	PODE	PODEMOS	No changes
PP	11	PP	PROGRESSISTAS	No changes
PPL	54	PPL	PARTIDO PÁTRIA LIVRE	No changes
PPS	23	PPS	PARTIDO POPULAR SOCIALISTA	No changes
PR	22	PR	PARTIDO DA REPÚBLICA	No changes
PRB	10	PRB	PARTIDO REPUBLICANO BRASILEIRO	No changes
PRONA	22	PR	PARTIDO DA REPÚBLICA	PL and PRONA merged into PR in 2006
PROS	90	PROS	PARTIDO REPUBLICANO DA ORDEM SOCIAL	No changes
PRP	44	PRP	PARTIDO REPUBLICANO PROGRESSISTA	No changes
PRTB	28	PRTB	PARTIDO RENOVADOR TRABALHISTA BRASILEIRO	No changes
PSB	40	PSB	PARTIDO SOCIALISTA BRASILEIRO	No changes
PSC	20	PSC	PARTIDO SOCIAL CRISTÃO	No changes
PSD	55	PSD	PARTIDO SOCIAL DEMOCRÁTICO	No changes
PSDB	45	PSDB	PARTIDO DA SOCIAL DEMOCRACIA BRASILEIRA	No changes
PSL	17	PSL	PARTIDO SOCIAL LIBERAL	No changes
PSOL	50	PSOL	PARTIDO SOCIALISMO E LIBERDADE	No changes
PSTU	16	PSTU	PARTIDO SOCIALISTA DOS TRABALHADORES UNIFICADO	No changes
PT	13	PT	PARTIDO DOS TRABALHADORES	No changes
PTB	14	PTB	PARTIDO TRABALHISTA BRASILEIRO	No changes
PTC	36	PTC	PARTIDO TRABALHISTA CRISTÃO	No changes
PTdoB	70	AVANTE	AVANTE	Name changed in 2016
PTN	19	PODE	PODEMOS	Name changed in 2016
PV	43	PV	PARTIDO VERDE	No changes
REDE	18	REDE	REDE SUSTENTABILIDADE	No changes
SOLIDARIEDADE	77	SOLIDARIEDADE	SOLIDARIEDADE	No changes
S/Partido	111	S/Partido	S/Partido	No changes
PPB	11	PP	PROGRESSISTAS	Name changed in 2003
PPR	11	PP	PROGRESSISTAS	PP and PPR merged into PPB in 1995, changed name to PP in 2003
PDC	11	PP	PROGRESSISTAS	PDC and PDS merged into PPR in 1993, PPR merged into PPB in 1995, then changed name to PP in 2003
PAN	14	PTB	PARTIDO TRABALHISTA BRASILEIRO	Merged into PTB in 2007
PSDC	27	DC	DEMOCRACIA CRISTÃ	Name changed in 2017
PST	22	PR	PARTIDO DA REPÚBLICA	PST (second branch) and PGT merged into PL in 2003, then PL and PRONA merged into PR in 2006
SD	77	SOLIDARIEDADE	SOLIDARIEDADE	To follow sigla in the TSE file

Appendix B

Selection concerns

Table B.1: Predicted values of log wells and First-term

	(1) Observed	(2) Predicted	(3) Observed	(4) Predicted
First term		0.00246		-0.0187
		(0.28)		(-1.18)
College	-0.309***		-0.240**	
	(-3.91)		(-2.12)	
Female	-0.0953		-0.00575	
	(-0.56)		(-0.03)	
Party dummies	Yes	No	Yes	No
Observations	1081	1081	547	547
R^2	0.047	0.000	0.060	0.001
AIC	3512.6	213.8	1774.5	227.2

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B.2: Predicted values of log wells and First term

	(1) Observed	(2) Predicted	(3) Observed	(4) Predicted
College	-0.309***		-0.240**	
	(-3.91)		(-2.12)	
Female	-0.0953		-0.00575	
	(-0.56)		(-0.03)	
First term		-0.113		-0.201
		(-1.49)		(-1.23)
Frequency of dry years		-0.0544		-0.107
		(-0.45)		(-0.85)
First term × Frequency of dry years		0.212		0.333
		(1.51)		(1.14)
Party dummies	Yes	No	Yes	No
Observations	1081	1081	547	547
R^2	0.047	0.001	0.060	0.003
AIC	3512.6	214.9	1774.5	228.2

t statistics in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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