Are Rules-based Government Programs Shielded from Special-Interest Politics? Evidence from Revenue-Sharing Transfers in Brazil

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Are Rules-based Government Programs Shielded from Special-Interest Politics? Evidence from Revenue-Sharing Transfers in Brazil*

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Abstract

Manipulation of government finances for the benefit of narrowly defined groups is usually thought to be limited to the part of the budget over which politicans exercise discretion in the short run, such as earmarks. Analyzing a national revenue-sharing program in Brazil, I find that in spite of an allocation rule based on local population, funds ended up being channeled to political allies as well as to communities likely to be swayed by economic benefits, exactly as theory would predict for discretionary transfers. Specifically, I find that over the decade of the 1990s, revenue-sharing transfers were targeted at local governments run by right-wing parties as well as municipalities that were both right-leaning and electorally fragmented. These findings suggest that the exclusive focus on discretionary transfers in the extant empirical literature may considerably understate the true scope of tactical redistribution that is going on under programmatic guise.

Keywords: Bureaucracy, institutions, redistributive politics, electoral competition

JEL: H77, D72, D73

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1 Introduction

Manipulation of government finances for the benefit of narrowly defined groups—special-interest politics for short—is usually thought to be limited to the part of the budget over which politicians exercise discretion in the short run, such as earmarks. Examples of such tactical redistribution include regulatory or fiscal favors to special interests, such as particular industries or particular districts that receive public construction projects and government jobs. In contrast, rules-based or programmatic redistribution—carried out using income taxes and the social welfare system—is considered to be relatively stable over time and driven by general-interest politics, which pits the economic interests of large groups of voters against each other (Dixit and Londregan 1996; Persson and Tabellini 2000). Whether in practice the scope of tactical redistribution is really limited to discretionary parts of the budget is an important question from both theory and policy perspectives. Perhaps surprisingly, however, little is known about this issue because the voluminous empirical literature on redistributive politics has focused almost exclusively on discretionary government spending, implicitly assuming that rules-based programs are implemented without regard to special interests.[1]

In this paper, I examine whether a rules-based transfer program in Brazil, the Fundo de Participação dos Municípios (FPM), which supposedly makes payments to municipal governments uniquely based on population size, was manipulated to favor special interests. The design of the revenue-sharing mechanism considered here is similar to the General Revenue Sharing program in the US from 1972 to 1986, which is common in many other federations around the world today.[2] These programs bypass the annual budget process and redistribute a substantial part of national tax revenues to local governments based on objective criteria, such as population size. While the explicit goals of such revenue-sharing mechanisms are many, an important common feature is that they aim to redistribute income from rich to poor communities, irrespective of political charac-

[1] Ames (1995a, 1995b, 2001), Levitt and Snyder (1995), Schady (2000), Case (2001), Dahlberg and Johansson (2002), Finan (2003), Ansolabehere and Snyder (2006), Khemani (2007), Solé-Ollé and Sorribas-Navarro (2007), Arulampalam, Dasgupta, Dhillon, and Dutta (2009).

[2] Other major federations include Canada, Germany and India (Boadway and Shah 2007).

teristics of the community. Ideological alignment with the party in control of the central government should play no role in the allocation of resources under such programs.

The first result of the paper is that the population estimates which went into the transfer allocation formula for the year 1991 were manipulated, as evidenced by their discontinuous distribution around several thresholds that determine the amount of transfers received by the municipality. This is in stark contrast to the distribution of *actual* 1991 municipality population—known from the census—and to official estimates from prior years, which are all smooth around the same thresholds. The 1991 manipulation substantively increased the number of municipalities which received higher amounts of transfers than warranted by their actual population and resulted in economically important funding differentials. Municipalities that located above the various population cutoffs in 1991 received additional transfers of about US$ 3.6 million on average over the entire decade of the 1990s because the 1991 allocations were subsequently grandfathered.[3] For small local governments the annual transfer differential amounted to about 15% of the public budget.

In the second step of the analysis I evaluate which, if any, of several theories about special-interests politics outlined in section 3 below are consistent with the observed program manipulation. In particular, I test whether political determinants, such as political alignment with the conservative central government, support for right-wing candidates for Congress, and the extent of interparty fragmentation at the municipality level predict official population estimates. The results suggest that center-local alignment matters—in particular—, resources were targeted at local governments run by right-wing parties. For the kind of transfers considered here, given as general budget support and therefore difficult for voters to attribute to the central government, this finding is consistent with recent theoretical models of special-interest politics (Arulampalam, Dasgupta, Dhillon and Dutta, ADDD for short, 2008; Khemani 2007; Solé-Ollé and Sorribas-Navarro 2007). The finding that resources were targeted at aligned local governments is also consistent

[3] The cumulative difference in FPM transfers over the period from 1991 to 1999 was about R$ 5 million in 2008 prices. The Real/$ purchasing power parity exchange rate in 2005 was about 1.36 (World Bank 2008).

with a model that assumes a more uncertain electoral response to economic favors among opposition or uncommitted groups relative to core support groups when politicians are risk averse (Cox and McCubbins 1986).

I also provide evidence that among municipalities governed by right-wing mayors, the main beneficiaries were those that were both right-leaning and electorally fragmented. Under the assumption that high political fragmentation at the local level proxies for swing constituencies—those that are not ideologically attached to any given party—these results are consistent with the "aligned swing" prediction of Arulampalam, Dasgupta, Dhillon and Dutta (2008). In contrast, none of these political determinants have predictive power for the 1985 allocations, suggesting that the military government, which had set up the revenue-sharing mechanism in 1965, indeed played by its own rules.[4] The evidence thus suggests that, although the grand redistribution scheme discussed here was shielded from tactical redistribution during the dictatorship, the same program became subject to special-interest politics after the transition to democracy.

The fact that the programmatic revenue-sharing scheme considered here was used for tactical redistribution is a new result in the literature. Although the findings are in line with recent work on redistributive politics in federations—where the parties in power may differ across levels of government—existing studies have generally taken for granted that programmatic redistribution is implemented without regard to special interests. Specifically, ADDD (2008) show that alignment matters for the allocation of project-specific discretionary grants in India over the period 1974-1997, but only in those states with a high proportion of close constituency elections. Khemani (2007) finds that over essentially the same time period, aligned Indian states received more general purpose discretionary grants, irrespective of the closeness of previous state legislature elections. Finally, the results in Solé-Ollé and Sorribas-Navarro (2008) suggest that over the period 1993-2003, partisan alignment had a sizeable positive effect on the amount of grants received by Spanish municipalities.

[4] Whether special-interest politics was already at play during the 1986-1990 Congress, the first under the new democratic regime, I cannot tell because electoral data from that period are not readily available.

More generally, many studies have found that politicians tend to reward their core constituents, as measured by the proportion of votes in a district that go to the party in power at the center. Levitt and Snyder (1995) show that the Democratic vote share is an important predictor of the amount of federal spending across congressional districts for the period 1975-1981, when the federal government was under control of the Democratic party, but not during the 1981-1990 period of divided government. Case (2001) provides evidence of a positive relationship between commune level voting with the central government party in a 1994 constitutional referendum and the subsequent receipt of block grants in Albania. Schady (2000) likewise shows that expenditures by the Peruvian Social Fund over the period 1991-1995 were in part targeted at communities that had helped elect the incumbent government. Using variation in party control of U.S. state governments across states and over time, Ansolabehere and Snyder (2006) also find that the distribution of intergovernmental transfers to local (county) governments is skewed towards loyal constituents.

Some studies have tested explicitly whether transfers are targeted at swing communities. Wright (1974) finds that states exhibiting higher variability in Democratic vote shares for Presidential elections received more federal spending and more work-relief jobs. Case (2001) shows that block grants were also targeted at communes that were relatively swing (close to 50% voting with the central government party on the referendum). Similarly, Schady (2002) also finds that central government funds were targeted at communities where support for the government in previous elections was close to 50%. Dahlberg and Johansson (2002) provide evidence that the central government in Sweden targeted transfers towards regions where the last center government election was close or the estimated proportion of swing voters was high. They find no evidence that core-constituents were favored. In contrast, Ansolabehere and Snyder (2006) find no evidence that parties reward municipalities where partisan vote shares are close to 50% Democratic and 50% Republican or where the volatility of the Democratic vote share in the past was high.

A number of empirical papers deal with special-interest redistributive politics in Brazil.

Ames (1995a) demonstrates that federal deputies in the 1987-1990 legislature were more likely to make amendments to the national budget in municipalities where their individual vote share in the previous election was high. He also finds that deputies target vulnerable municipalities, that is, municipalities where incumbent deputies retired, in-migration was high and interparty and intraparty fragmentation were high. Similarly, Finan (2003) investigates federal deputies' amendments to the national budget over the legislative cycle 1995-1998, and finds that they tend to reward municipalities for past electoral support. Arretche and Rodden (2004) find that those states which provided more votes in past presidential elections received more intergovernmental transfers over the period 1991-2000.

The remainder of the paper is organized as follows: Section 2 presents institutional background on the revenue-sharing mechanism between the federal and local governments in Brazil and provides evidence of program manipulation in 1991. Section 3 presents the theoretical framework as well as empirically testable hypotheses given the political and institutional environment in Brazil around 1990. Section 4 describes the data. Section 5 gives details on the estimation approach. Estimation results are presented in Section 6. The final section concludes with a discussion of limitations and extensions.

2 Institutional background

In this section, I first describe the economic importance, mechanics and origins of the federal revenue-sharing fund for municipal governments in Brazil. Next, I give details on the forecasting procedure for local population estimates in inter-censal years. I then document a manipulation of the program that occurred with the 1991 population estimates and show that this manipulation substantively increased the number of municipalities that were over-classified relative to transfer brackets warranted by actual population. I also show that the manipulation had economically significant effects on the distribution of revenue-sharing funds. Finally, I discuss why the effects of the manipulation extended over the entire decade of the 1990s to the present day.

2.1 Importance, mechanics and origins of revenue-sharing in Brazil

Intergovernmental transfers finance most of local government spending on primary education, primary health care, housing and urban infrastructure, and local public transportation in Brazil. Over the period of the 1990s, total government revenue in Brazil was about 28% of GDP, of which municipalities collected about 5%. At the same time, local governments managed about 16% of public resources. Intergovernmental transfers to local governments therefore represented about 3.08% of GDP. The most important among these transfers is the Fundo de Participação dos Municípios (FPM), a constitutionally guaranteed and largely unconditional revenue-sharing grant funded by federal income and industrial products taxes.[5] The FPM grant alone accounted for about 50% of revenue in small to medium sized local governments.

According to the national tax code (Decree 1881/81), transfer amounts depend on municipality population in a discontinuous fashion. More specifically, based on municipality population estimates, pop^e, municipalities are assigned a coefficient $k = k(pop^e)$, where $k(.)$ is the step function shown in Table 1. For municipalities with up to 10'188 inhabitants, the coefficient is 0.6; from 10'189 to 13'584 inhabitants, the coefficient is 0.8; and so forth. There is a total of 18 population brackets and although the population thresholds were supposed to evolve with population growth in Brazil, they remained unchanged since 1966, as further detailed below. The coefficient $k(pop^e_{mst})$ determines the share of FPM resources available for state s that are distributed to municipality m in year t. The amount of transfers to state s in turn depends on a percentage f_s of federal tax collection earmarked for revenue-sharing in year t, rev_t. The state shares are determined in the constitution and have remained unchanged since their introduction in 1989.[6] FPM_{mst} is the amount transferred to municipality m in state s during year t according to the

[5]Federal Constitution of Brazil, 1988, Art. 159 Ib. The one condition is that municipalities must spend 25 percent of the transfers on education (Art. 212). This constraint is usually considered non-binding, in that municipalities typically spend about 20% of their total revenue on education. It is not clear how this provision was enforced in practice, since there is no clear definition of education expenditures and accounting information provided by local governments was not systematically verified.

[6]Supplementary Law n^o 62/1989 and Decision n^o 242/1990 of the Federal Court of Accounts. The state shares f_s correspond to the shares of each state in the total population of Brazil according to the 1991 census.

following formula:

$$FPM_{mst} = \frac{k(pop^e_{mst})}{\sum\limits_{i \mid s} k^e_{ist}} f_s rev_t \tag{1}$$

Equation (1) makes it clear that local population estimates should be the only determinant of cross-municipality variation in FPM funding in a given state. Before proceeding it is worth discussing why politicians would choose to allocate resources based on objective criteria, such as population, rather than use discretion? The answer to this question lies in the political agenda of the military dictatorship which came to power in 1964. As detailed by Hagopian (1996), one of the major objectives of the military was to wrest control over resources from the traditional political elite and at the same time to depoliticize public service provision. The creation of a revenue-sharing fund for the *municípios* based on an objective criterion of need, population, was part of this greater agenda. It reflected an attempt to break with the clientelistic practice of the traditional elite, which manipulated public resources to the benefit of narrowly defined constituencies.

The reason for allocating resources by brackets, i.e. as a step function of population as in Decree 1881/81, is less clear. One explanation could be that compared to a linear schedule, for example, the bracket design mutes incentives for local officials at the interior of the bracket to tinker with their population figures or to contest the accuracy of the estimates in order to get more transfers. A related question is where the exact cutoffs come from—that is, why 10'188, 13'584, 16'980, etc.? While I was unable to trace the origin of these cutoffs precisely, I know roughly how they came about. The initial legislation from 1967 created cutoffs at multiples of 2'000 and stipulated that these should be updated proportionally with population growth in Brazil.[7] The cutoffs were thus presumably updated twice, once with the census of 1970 and then with the census of 1980, which explains the "odd" numbers. It is noteworthy that the thresholds are still equidistant from one another, the distance being 6'792 for the first 7 cutoffs (except for the second cutoff which lies exactly halfway in between the first and the third cutoffs).

[7] Supplementary Law No. 35, 1967, Art. 1, Paragraphs 2 and 4.

2.2 The forecasting procedure for local population estimates

According to equation (1), municipality population is the key determinant of this revenue-sharing mechanism. However, exact municipality population figures are only available for census years or years when a national population count is conducted. For all other years, official population estimates are produced by the national statistical agency, IBGE.[8] Prior to 1989 these estimates were updated only in years ending with the number 5. Beginning in 1989 the estimates are updated on a yearly basis. The currently used forecasting procedure is based on a top-down approach that ensures consistency of estimates for lower level units (municipalities) with the higher levels (states and the country as a whole) (IBGE, 2002).

First, IBGE produces a population estimate for Brazil, pop_t^e, based on estimated birth rates, mortality and net migration for Brazil. Individual states are then assigned their share of the national estimate, pop_{st}^e, in proportion to past state level census population numbers. municipalities within a given state are grouped by quartile of both census population levels and past population growth between census years and growing municipalities are separated from shrinking municipalities. Each of these 20 groups of municipalities is then assigned its share of the state population estimate, pop_{jst}^e, proportional to past group level census population. Finally, each municipality within each group is assigned its population estimate, pop_{mjst}^e, based on past municipality level census information. The specific formula for municipality population estimates is as follows:

$$pop_{mjst}^e = (pop_{mjs80}/pop_{js80})[a_{js}pop_{st}^e + b_{js}] \quad t > 1988 \tag{2}$$

where

$$a_{js} = \frac{pop_{js80} - pop_{js70}}{pop_{s80} - pop_{s70}} \quad j = 1, 2, ..., 20$$

$$b_{js} = pop_{js80} - a_{js}pop_{s80}$$

According to equation (2) local population forecasts are essentially a continuous function of past census information and state level population projections. This top-down proce-

[8]Supplementary Law n^o 59/1988, Art. 91, Paragraph 3.

dure ensures the consistency of estimates for lower level units (municipalities) with the higher levels (states and the country as a whole):

$$pop_t^e = \sum_s \sum_j \sum_m pop_{mjst}^e$$

Since local population estimates directly determine funding levels it is important to verify whether they are indeed derived from this forecasting procedure. Figure 1 plots 1989 official population estimates (coming directly from reports issued by the Federal Court of Accounts (TCU)) against predicted estimates calculated using the above formula. It is clear from the scatterplot that the formula predicts 1989 official population estimates quite well although there is some dispersion around the 45 degree line. The dispersion is related to the fact that the predicted estimates are not based on the same 1970 and 1980 census data that were used at the time official estimates were made in 1989. Another and probably more important reason for the dispersion is that origin municipalities ceded some population to newly created municipalities.[9] Finally, the dispersion might be related to political manipulation as further discussed below. The important point here, however, is that as a first approximation, official 1989 estimates are indeed consistent with the forecasting procedure described above.

2.3 Evidence on manipulation of population estimates

The first empirical fact established in this paper is that the tight link between formula-driven predictions and official estimates broke down over the next two years.[10] This point is best demonstrated with the use of histograms for 1989 official estimates, for the 1991 official estimates and for 1991 actual population.[11] Figures 2 and 3 show that while the distribution of 1989 official estimates is smooth at the thresholds, the distribution of 1991 official estimates exhibits gaps immediately below the thresholds determining

[9] In order to obtain forecasts for the newly created and origin counties I would need to know which counties lost territory to the newly created counties as well as access to census tract population numbers from 1980 which are not readily available.

[10] 1990 estimates already exhibit some irregularities but the 1991 manipulation is much more pronounced and produced more lasting effects as further discussed in Section 2.3 below.

[11] The bin-width in these histograms is 566, which ensures that the various cutoffs coincide with bin limits—that is, no bin counts observations from both sides of any cutoff.

transfer brackets and even more obvious spikes immediately above those cutoffs. The histogram for 1991 official estimates actually understates the discontinuity of the density around the cutoffs because the spikes occur at specific points on the support.[12] The total number of municipalities that were placed on any one of these bunching points is 1870, which represents 42% of the municipalities receiving FPM transfers at the time. Figure 4 makes it clear that these gaps (to the left) and spikes (to the right) of the thresholds do not reflect actual 1991 census population. While I was not able to confirm with IBGE what forecast model they were using in 1991, it seems clear that government officials did not rely exclusively on some variant of the forecast procedure outlined above, which is essentially a continuous function of past census information and population projections. The discontinuous distribution of population estimates is thus almost surely the result of an adjustment which went beyond the mechanical application of the forecasting procedure.

The reasons for this manipulation or adjustment of population estimates are less clear. For example, it is conceivable that bureaucrats used some administrative rule to determine which estimates to revise. Officials were likely more averse to underestimate a municipality relative to a given threshold than overestimating it because underestimated municipalities were much more likely to appeal against IBGE's preliminary population estimates. Although IBGE has the final authority to determine official estimates, i.e. there is no external review of IBGE decisions, dealing with municipality complaints involves scarce administrative resources. Bureaucrats' attempts to preempt such complaints would explain the curious gaps in the distribution of estimates just below the thresholds as well as a part of the spikes just above. One sensible administrative rule would be that all municipalities within a given distance to the next higher threshold were placed just above the threshold to take account of the uncertainty surrounding the formula based estimates. The mass of missing municipalities from the gaps to the left of each threshold is too low to account for the mass on the spikes, however. In other words, IBGE officials must have bumped up municipalities for other reasons as well.

[12] The exact bunching points are as follows: 10189, 10298, 13730, 17162, 24027, 30891, 37756, 44620, 51484, 61781, 72078, 82375, 92671, 102968, 116697, 130426, 144155, 157884.

Alternatively, administrators might have had access to evidence about actual local population levels justifying selective revision of population estimates. For example, some mayors may have presented IBGE with administrative data, such as local vital and migration statistics indicating that they were in fact eligible for higher transfers. It is also possible that IBGE used electoral data from 1988 to reclassify municipalities. If this were the case—and if the information IBGE acted upon was more reliable than the predictions from the model—one would expect that the number of correctly classified municipalities in terms of transfer brackets increased with the manipulation. Since actual populations are known ex post, I can test whether this is indeed the case by comparing the classification performance that arises using the 1991 manipulated estimates to the classification performance using the 1991 pre-manipulation or first-pass population estimates. Such a comparison holds the inherent uncertainty surrounding population estimates constant and allows a quantification of the distortion of public funds generated by the manipulation.

Since I do not observe 1991 pre-manipulation estimates I use the 1989 official estimates instead.[13] Equation (2) shows that the only information relevant for local population forecasts that changes between 1989 and 1991 are state-level population estimates. Since these changes are unlikely to be large from year to year, the resulting classification error is likely limited. I focus on the bracket error, defined as defined as the difference between the predicted transfer bracket for 1991, $k(pop_m^e)$, and the correct transfer bracket for 1991, based on actual (unknown at the time of the forecast) local population, $k(pop_m)$:

$$bracket\ error\ = 5 \times [k(pop_m^e) - k(pop_m)]$$

Table 2 shows the distribution of bracket errors under the 1989 official estimates (which proxies for 1991 pre-manipulation estimates) and the manipulated 1991 official estimates. From panels A and B it is apparent that for bunched municipalities, that is, those located on any of the bunching points, the manipulation increased the percentage of mis-classified municipalities (bracket error $\neq 0$) from about 52% to about 83%. Even more strikingly, the

[13] I also use the 1989 predicted population estimates discussed above and results are almost identical to those obtained using the 1989 official estimates.

manipulation shifted the entire bracket error distribution to the right, moving the percent over-classified (bracket error > 0) from 31% to 80%. For non-bunched municipalities, the percentage mis-classified increased only slightly from 20% (Panel C) to 21% (Panel D), while the percentage over-classified increased from 10% to 20%. Figure 5 presents the histograms corresponding to Table 2. Overall, the manipulated 1991 official estimates increased the number of mis-classified municipalities from 33% to 48% and the number of over-classified municipalities from 19% to 46%.

These results suggest that the information used to revise the formula-driven estimates was not a good predictor of actual levels of population in 1991. It is also worth noting that manipulation may not have been limited to the bunched municipalities since the percentage over-classified also increased for the non-bunched municipalities. Similarly, the 1991 manipulation may not have been an isolated incident. Even prior to 1991 there might have been more subtle manipulations of the program, which left the distribution of population estimates smooth at the cutoffs. I take up this issue in Section 3 below where I test whether conditional on non-political municipality characteristics, political conditions are able to predict official estimates, both in 1991 and in 1985, the last year of the military government.

2.4 Economic significance of the manipulation

The 1991 manipulation resulted in economically important transfer differentials. Municipalities that located above a population cutoff in 1991 received additional transfers of about US$ 3.6 million on average over the entire decade of the 90s (and beyond) because coefficients were subsequently grandfathered.[14] For small local governments the annual transfer differential amounted to about 15% of their public budgets. Figure 6 illustrates the persistence of this effect by showing cell means of cumulative FPM transfers over the period 1991-1999 against the 1991 official population estimate.[15]

[14] The cumulative difference in FPM transfers over the period from 1991 to 1999 was about R$ 5 million in 2008 prices. The Real/$ purchasing power parity exchange rate in 2005 was about 1.36 (World Bank 2008).

[15] The bin-width in this plot is 566, which ensures that the various cutoffs coincide with bin limits.

Grandfathering began in 1992 when all coefficients remained virtually unchanged, partly because census results had not been available by the end of 1991. When census population estimates were finally released in 1993, the majority of municipalities would have had their coefficients reduced because the law stipulated that the thresholds be adjusted with population growth and these municipalities had grown less than the population average for Brazil. Some municipalities would have incurred a significant loss of transfers as a result of this reclassification (Brandt 2002).

Another law was approved in April 1993, still by the same congress, which determined that both coefficients and population thresholds were to be maintained without adjustment.[16] The only exception was for municipalities that were subdivided and lost population to newly-created municipalities. The revision of coefficients for these types of municipalities was done according to the existing population thresholds using the latest census population figures. Underestimated municipalities' coefficients were updated pursuant to the publication of the census while overestimated municipalities' coefficients were not.

In 1996, there was a population count carried out by IBGE and Congress approved another supplementary law at the end of 1997. It stated that in 1998 all coefficients of the FPM were to remain the same as in 1997.[17] From 1999 onwards however, coefficients would be based on the 1996 population count and the grandfathering would be phased out over the next five years. In each year, coefficients of municipalities that had benefited from the grandfathering would be reduced by 20% of the excess coefficient, the difference between the grandfathered coefficient and that resulting from current population estimates. As a result of the 1997 law, coefficients for fiscal years from 1999 onwards were increasingly based on current population estimates. Denoting $\overline{k}_m$ as the grandfathered coefficient for municipality m, $1[.]$ as the indicator function and α_t as the percentage reduction in the excess coefficient $\overline{k}_m - k(pop_{mt})$, coefficients are currently calculated as

$$k_{mt} = 1[k(pop_{mt}) \geqslant \overline{k}_m]k(pop_{mt}) + 1[k(pop_{mt}) < \overline{k}_m][\overline{k}_m - \alpha_t(\overline{k}_m - k(pop_{mt}))]$$

[16] Supplementary Law n^o 74/1993.

[17] Supplementary Law n^o 91/1997.

In March 2001 a new supplementary law was enacted in order to postpone full adjustment to 2008.[18] The 1991 manipulation thus extends its effects to the present day.

To sum up this section, there is clear evidence that the 1991 official population estimates were somehow adjusted or manipulated. The adjustments resulted in economically important transfer differentials extending up to the present day because coefficients were grandfathered. The fact that the manipulation of municipality population estimates documented above significantly increased the number of mis-classified municipalities casts doubts on technocratic explanations. The remainder of the paper turns to political explanations of the program manipulation.

3 Theoretical framework and predictions

3.1 Theoretical framework

This sections presents a simple model of central government resource allocation across municipalities, borrowed from Arulampalam, Dasgupta, Dhillon and Dutta, henceforth ADDD (2008), and similar to models by Solé-Ollé and Sorribas-Navarro (2007) and Khemani (2007). The key prediction from these models is that a vote-maximizing central government incumbent will skew fiscal transfers in favor of aligned local governments if credit-claiming is difficult, that is, if the implementing local government gets sufficiently high partial credit for additional resources.

There are two parties, L and R, and two levels of government, center and local. The central government incumbent party R, decides on the allocation of transfers and is assumed to care about its own re-election. There is a set of municipalities S^R where the local incumbent party is R, and a set of municipalities S^L where the local incumbent party is L. Transfers from the center to each of M municipalities m finance local public services valued by voters. Actual service provision is done by the local government and imperfectly informed voters do not perceive perfectly that the R party is the source of the grants. As a result, the goodwill generated by these transfers is shared between in-

[18] Supplementary Law n^o 106/2001.

cumbent parties at both levels of government. Let $\theta \in [0,1]$ denote the share of goodwill from per capita transfers that accrues to the central incumbent. θ is known by the central government and assumed exogenous. For the kind of transfers considered here, given as general budget support, θ is likely to be low.

Within each municipality m, there is a continuum of voters of mass N_m who may differ in their ideologies. A voter j, located at X_j on the ideology spectrum $[\underline{X}, \overline{X}]$ has preference X_j for party L over party R. X_j is private information while the cumulative distribution function of X in municipality m, denoted $\Phi_m(X)$, is common knowledge. $\Phi'_m(X)$ is strictly positive and continuous for all X. For simplicity, $\underline{X} = -\overline{X}$, so that the midpoint is 0.

Voters in each municipality vote on the basis of ideology and economic benefits generated by grants. Consider a voter j in municipality $m \in S^R$ which has received per capita grant g_m from the center. Party R has received a goodwill of $U(g_m)$, with $U(0) = 0$, $U'(g_m) > 0$, $U''(g_m) < 0$ and so he votes for party R iff:

$$U(g_m) - X_j \geq 0 \tag{3}$$

and he votes for party L otherwise. In contrast, in a municipality governed by party L, goodwill is split between the two parties: party R gets $\theta U(g_m)$ while party L gets $(1-\theta)U(g_m)$. Voter j will vote for party R iff:

$$\theta U(g_m) - (1-\theta)U(g_m) - X_j \geq 0 \tag{4}$$

The inequalities (3) and (4) generate cutpoints, $X(g_m, R)$ and $X(g_m, \theta, L)$ for each municipality such that a voter located at X_j votes for party R iff $X_j \leq X(g_m, \theta, p)$ for $p = L, R$. The central incumbent uses grants in order to shift the location of these cutpoints:

$$\frac{\partial X(g_m, R)}{\partial g_m} = U'(g_m), \frac{\partial X(g_m, \theta, L)}{\partial g_m} = (2\theta - 1)U'(g_m)$$

Increasing grants to aligned local governments unambiguously improves electoral prospects of the R party. Increasing grants to non-aligned local governments improves electoral prospects of the R party only if θ is sufficiently high (above 0.5) and hurts party $R's$ prospects if goodwill leakage is large (θ below 0.5).

Tactical redistribution by the central incumbent is subject to two constraints. First, transfers must satisfy an overall budget constraint. Second, the incumbent is also interested in maximizing total welfare accruing from transfers. This aspect is captured by specifying a per capita welfare function $\gamma(g_m)$, assumed increasing and concave in g_m. If voters vote along party lines, that is, ideology of voters at the local level is the same as at the central level, it is reasonable to assume that the central incumbent maximizes its vote total across municipalities. The objective function is then:

$$\sum_{m\in S^R} N_m\Phi_m(X(g_m,R)) + \sum_{m\in S^L} N_m\Phi_m(X(g_m,\theta,L) + \sum_m N_m\gamma(g_m) \tag{5}$$

which the incumbent R government maximizes by choice of grant allocation $\{g_m\}_{m=1}^M$, subject to the budget constraint:

$$\sum_m N_m g_m = B$$

At an interior solution the first-order condition for a municipality $m \in S^R$ is:

$$\gamma'(g_m^*) + \Phi_m'(X(g_m^*,R)U'(g_m^*) = \lambda \tag{6}$$

and for a municipality $m \in S^L$ it is:

$$\gamma'(g_m^*) + \Phi_m'(X(g_m^*,\theta,L)(2\theta-1)U'(g_m^*) = \lambda \tag{7}$$

where λ denotes the Lagrange multiplier and g_m^* is the optimal allocation of grants for the central incumbent R. If the objective function (5) is concave, then the necessary conditions are also sufficient and the solution $\{g_m^*\}_{m=1}^M$is unique.

From the first order conditions (6) and (7), and the concavity assumption on $\gamma(g_m)$ it follows that, when goodwill leakage is large ($\theta < 0.5$), the central incumbent R will allocate higher per capita grants to aligned (R) municipalities than to those that are non-aligned (L).[19] The assumption of large goodwill leakage is very plausible for the kind of transfers considered here, given as general budget support and therefore difficult for voters to attribute to the central government.

In terms of earlier theoretical work on redistributive politics, goodwill leakage in non-aligned municipalities corresponds to an inefficient targeting technology relative to aligned municipalities, which leads to core-support targeting (Lindbeck and Weibull 1987, Dixit and Londregan 1996). The prediction that aligned municipalities will be favored is also consistent with a more uncertain electoral response in opposition or uncommitted groups relative to core support groups when politicians are risk averse (Cox and McCubbins 1986).

When the potential for credit claiming for the center is low, a further prediction, developed fully in ADDD (2008) and Solé-Ollé and Sorribas-Navarro (2007) is that, among aligned local governments, those that are more easily swayed by economic benefits (swing communities) will receive higher transfers. And according to the Cox and McCubbins (1986) logic, among aligned local governments, municipalities that are more supportive of the R party should receive higher transfers. These predictions are not mutually exclusive. In fact, municipalities that are *both* supportive of the R party *and* relatively swing might be favored most. In the next sub-section I discuss how I translate these predictions into empirically testable hypotheses, given the political and institutional environment in Brazil around 1990.

3.2 Testable predictions

The first prediction obtained above is that aligned municipalities, that is, those that were governed by mayors affiliated with the ruling party at the center, were more likely to obtain population estimates above a given threshold and hence receive more federal funding than non-aligned municipalities. Determining allied parties and hence center-local alignment in Brazil's fragmented party system is somewhat difficult. It is even more complicated during the presidency of Fernando Collor (PRN) from 1990 until 1992. since he did not

[19] Proposition 1 from Arulampalam, Dasgupta, Dhillon and Dutta (2008).

enter into formal coalitions with other parties until the end of his term. Observers agree, however, that he needed to rely on legislative support from right-wing parties, PDS and PFL in particular, in order to pass legislation (Ames 1995b, 2001). Other right-wing parties at the time included the PL, the PDC and the PTB.[20] In the empirical analysis below, the binary variable *right-wing mayor* indicates a municipality headed by mayors affiliated with any of the above political parties.

The second testable prediction is that among aligned municipalities, those with a higher proportion of swing voters should be favored. This is what ADDD (2008) refer to as the "aligned swing" effect. I operationalize the swing voter concept as interparty fragmentation of the municipality vote in the 1990 federal government Camara (House) elections. Specifically, *interparty fragmentation* is defined as 1-H, where H is the Herfindahl index applied to party vote shares v_{mp} at the municipality level:

$$H_m = \sum_p v_{mp}^2$$

High levels of interparty fragmentation (relatively equal vote shares) are supposed to proxy for an electorate with weak ideological preferences. While certainly not perfect, this measure has been used as a proxy for how susceptible communities are to economic favors in earlier work on determinants of fiscal transfers in Brazil (Ames 1995a).

The third prediction is that among aligned local governments, municipalities that are more supportive of the R party should receive higher transfers. To measure core-support I use the fraction of municipality votes cast for right-wing parties in the Camara 1990 elections and call it *right-wing vote share*. Finally, municipalities that are *both* right-wing *and* relatively swing might be favored most. I operationalize this prediction by including the interaction of *interparty fragmentation* with *right-wing vote share*, separately for *right-wing mayor* and *non-right-wing mayor* municipalities.

[20] Partido da Reconstrução Nacional (PRN), Partido da Frente Liberal (PFL), Partido Democrático Social (PDS), Partido Trabalhista Brasileiro (PTB), Partido Democrata Cristão (PDC), Partido Liberal (PL).

4 Data

The data used in this study are complied from several sources. Official population estimates stem from successive reports issued by the federal court of accounts (TCU). Although estimates are produced by the national statistical agency, it is the responsibility of the TCU to compute municipalities' coefficients k_{mt} in accordance with decree 1881/81. 1991 census population figures come from the national statistical agency (IBGE). Data on FPM transfers were self-reported by municipality officials and compiled into reports by the secretariat of economics and finance inside the federal ministry of finance. The FPM data are somewhat noisy as there is sometimes substantial under-reporting of transfers received from the federal government. Unfortunately, more reliable data from the ministry of finance are not available for the early nineties. The financial data were converted into 2008 currency units using the GDP deflator for Brazil. Electoral data for municipal executive (1988), national congress (1990) and presidential (1989) elections are from the Supreme Electoral Tribunal. Again these data are somewhat incomplete both in terms of available variables and observations. Table 3 gives full party names and descriptive statistics of the political determinants used in the empirical analysis below.

5 Estimation approach

The goal of this paper is to evaluate which, if any, of the predictions outlined in Section 3.2 are borne out by the data. Specifically, I test whether *right-wing mayor, interparty fragmentation* and *right-wing vote share,* as well as interactions among these variables, are able to predict the 1991 official population estimates. Controlling for municipality characteristics is important for all of these tests because revision of estimates may have been based on (local) evidence that a municipality's actual population placed it into higher transfer brackets. If these municipalities happened to favor right-wing parties in previous elections for example, the correlation between right-wing support and population estimates would be an upwardly biased measure of special-interest influence. If, however, there turns out to be a correlation between political determinants and official population estimates, *controlling* for municipality characteristics that might account for revisions of population estimates, such as 1988 electorate data and actual 1991 population, this would be indicative of political interference.

In order to address concerns about variables that might have been used to improve the municipality classification into transfer brackets, the regression specification includes second- and third-order polynomials of the 1988 electorate and of 1991 actual (census) local population, a set of indicators for the 1991 actual population classification as well as 1991 municipality characteristics such as income per capita, average years of schooling of those 25 years of age and older, the poverty rate, gini index and the urbanization rate. Denoting by Y_{ms} the 1991 official population estimate for municipality m in state s for the year 1991, $\mathbf{X}_{ms}$ the vector of controls mentioned above and a_s a state fixed effect, the estimation equation is as follows:

$$\begin{aligned} Y_{ms} = \; & \alpha_1 \text{Right-wing mayor}_{ms} \\ & + \alpha_2 \text{Right-wing mayor}_{ms} \times \text{Right-wing vote share}_{ms} \\ & + \alpha_3 \text{Right-wing mayor}_{ms} \times \text{Interparty fragmentation}_{ms} \\ & + \alpha_4 \text{Right-wing mayor}_{ms} \times \text{Right-wing vote share}_{ms} \times \text{Interparty fragmentation}_{ms} \\ & + \alpha_5 \text{Non-right-wing mayor}_{ms} \times \text{Right-wing vote share}_{ms} \\ & + \alpha_6 \text{Non-right-wing mayor}_{ms} \times \text{Interparty fragmentation}_{ms} \\ & + \alpha_7 \text{Non-right-wing mayor}_{ms} \times \text{Right-wing vote share}_{ms} \times \text{Interparty fragmentation}_{ms} \\ & + \boldsymbol{\gamma} \mathbf{X}_{ms} + a_s + u_{ms} \end{aligned} \tag{8}$$

One drawback of the continuous dependent variable, *1991 official population* is that it does not explicitly take into account the various brackets mandated by decree 1881/81, on which transfer allocations are ultimately based. As a specification check I also use the 1991 *bracket error,* $5 \times [k(pop^e_m) - k(pop_m)]$, as the dependent variable in the statistical analysis. As a final specification check, I use the binary dependent variable *positive bracket error*, equal to 1 if bracket error is positive and zero otherwise. All parameters are estimated using OLS. In order to interpret statistically and economically significant estimates of the $\alpha' s$ as evidence of political interference, the key assumption is that conditional on covariates $\mathbf{X}_{cs}$, unobserved factors in u_{ms} are uncorrelated with the political determinants.

Before presenting estimation results it is useful to briefly discuss the expected signs on the parameters in equation (8). The common prediction from the models by ADDD, (2008) Solé-Ollé and Sorribas-Navarro (2007), Khemani (2007), and Cox and McCubbins (1986) is that $\alpha_1 > 0$. According to the Cox and McCubbins core-supporter logic, $\alpha_2 > 0$. If swing communities are favored, then $\alpha_3 > 0$. And if municipalities that are *both* right-wing *and* relatively swing are favored most, one would expect $\alpha_4 > 0$. If goodwill leakage is large, or put differently, if the implementing local government gets sufficiently high partial credit for additional transfers, then it is not clear whether among municipalities run by non-right-wing mayors, those that are relatively right-wing or swing or both should be favored by the right-wing central government. In fact, ADDD (2008) suggest that swing communities governed by non-aligned parties should receive *less* transfers. While the theory has therefore no firm predictions for α_5, α_6 and α_7 individually, they should be jointly different from α_2, α_3 and α_4 if alignment matters.

There might also have been more subtle manipulations of the program prior to 1991, which left the distribution of population estimates smooth at the cutoffs. Unfortunately, electoral data for the 1987-1990 congressional session, the first under the new democratic regime, is not readily available. Instead, I use data from 1985, the last year of the military government, to run the exact same tests as discussed above. *Right-wing mayor* in this period refers to a municipality headed by mayors affiliated with the PDS (the party of the military regime) or the PTB. *Interparty fragmentation* is defined as 1-H, where H is the Herfindahl index applied to party vote shares at the municipality level in the 1982 federal government Camara (House) elections. The fraction of municipality votes cast for right-wing parties in the Camara 1982 elections is again referred to as the *right-wing vote share*.

6 Estimation results

Table 4 presents estimates of α_1 through α_7 based on equation (8) using 1991 official population as the dependent variable. The results provide clear statistical evidence that official population numbers were politically manipulated. In fact, the F-test rejects the null hypothesis of no effect of any political determinant at the 5% level across specifications.

More specifically, the results suggest that center-local alignment matters, in particular, that resources were targeted at local governments run by right-wing parties. Estimates of α_1 are all positive and become statistically significant when actual 1991 population is added as a control variable. According to the last estimate of α_1 in column 6, which also controls for 1991 municipality characteristics, in addition to 1991 actual population, municipalities governed by right-wing mayors got an expected fictional population gain of about 2'200. For the kind of transfers considered here, given as general budget support and therefore difficult for voters to attribute to the central government, this finding is consistent with recent theoretical models (ADDD 2008, Khemani 2007, Solé-Ollé and Sorribas-Navarro 2007). The finding is also consistent with a more uncertain electoral response to fiscal transfers in municipalities run by non-right-wing mayors relative to right-wing mayors (Cox and McCubbins 1986).

The table also shows that among municipalities governed by right-wing mayors, the main beneficiaries were those that were both right-leaning and electorally fragmented. Using column 6 again, the estimates of α_3 and α_4 suggest that among aligned municipalities with a high right-wing vote share of say 0.9, a 2 standard deviation difference in interparty fragmentation (0.28) led to a fictional population gain of about $-4+9\times0.9\times0.28\approx1150$. This result is consistent with the "aligned swing" prediction of ADDD (2008), although with the qualification that the effect seems to depend on a relatively right-leaning electorate.

Among municipalities governed by non-right-wing mayors there is less statistical evidence that official population numbers were politically manipulated. However, there is some evidence that the regression function among aligned municipalities is different from the regression function in non-aligned municipalities, which is as theory would predict if the central government worries about goodwill leakage. Specifically, while the F-test fails to reject the joint null hypotheses that $\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$, the last three specifications suggests that the effect of interparty fragmentation is different between aligned and non-aligned municipalities.

Table 5 reports the results from the same tests discussed above but using the 1991 bracket error as the dependent variable. The results again provide clear statistical evidence that official population numbers were politically manipulated (F-tests have zero p-values throughout the table). According to the last estimate of α_1 in column 6, municipalities governed by right-wing mayors got an expected fictional bracket error gain of about 0.5. And using again the estimates of α_3 and α_4 from column 6, among aligned municipalities with a high right-wing vote share of 0.9, a 2 standard deviation difference in interparty fragmentation (0.28) led to a fictional population gain of about $1 \times 0.9 \times 0.28 \approx 0.25$. Finally, there is clear evidence that the effects of right-wing vote share and interparty fragmentation among aligned municipalities were different from the effects in non-aligned municipalities (almost all F-tests significant at 1%).

Table 6 reports the results from specifications where the binary variable *positive 1991 bracket error* is the dependent variable. Results are again qualitatively similar to those obtained above: The null hypothesis of zero effect of any political determinant is clearly rejected in all specifications. Municipalities governed by right-wing mayors were about 26 percentage points more likely to get classified above the transfer bracket warranted by their actual population (column 6). And among aligned municipalities with a high right-wing vote share, a 2 standard deviation difference in interparty fragmentation led to an increased probability of getting over-classified of about $0.47 \times 0.28 \approx 13$ percentage points. Finally, there is again clear evidence that the effects of right-wing vote share and interparty fragmentation among aligned municipalities were different from the effects in non-aligned municipalities, a finding consistent with substantive goodwill leakage in non-aligned municipalities.

In sharp contrast to the clear evidence of special-interest politics in the 1991 official estimates, there is very little evidence of similar interference in the 1985 official estimates. Table 7 shows the results for 1985 official population as the dependent variable. Most of the individual estimates are statistically insignificant and most F-tests fail to reject the joint null hypotheses that all coefficients on political determinants are zero. Moreover, there is little evidence that the regression function among aligned municipalities is different from the regression function in non-aligned municipalities as most F-tests fail to reject the joint hypotheses. The fact that none of the political determinants are correlated with the 1985 official population estimates suggests that the military government, which had set up the revenue-sharing mechanism in 1965, indeed played by its own rules.

Tables 8 and 9 show that these results are robust to coding the dependent variable as *1985 bracket error* and as the binary *positive 1985 bracket error*, respectively. Although the estimates in the first three columns of each table might suggest that there was political interference, these correlations disappear when actual 1985 population is added as a control. Controlling for actual population, there is virtually no evidence of any interference in the 1985 official estimates, regardless of how the dependent variable is specified. None of the individual estimates are statistically insignificant and all F-tests fail to reject the joint null hypotheses that all coefficients on political determinants are zero. Moreover, there is no evidence that the regression function among aligned municipalities is different from the regression function in non-aligned municipalities as all F-tests fail to reject the joint hypotheses. The evidence thus suggests that although the grand redistribution scheme discussed here was shielded from tactical redistribution during the dictatorship, the same program became subject to special-interest politics after the transition to democracy.

7 Conclusion

This paper presents evidence that even a rule-based transfer program anchored in the constitution and in the national tax code—as opposed to programs funded through the annual budget—and based on apparently technocratic inputs is not always immune to special-interest politics. Specifically, the results suggest that over the decade of the 1990s FPM revenue-sharing transfers were targeted at local governments run by right-wing parties. For the kind of transfers considered here, given as general budget support and therefore difficult for voters to attribute to the central government, this finding is consistent with recent theoretical models of special-interest politics (Arulampalam, Dasgupta, Dhillon and Dutta 2008; Khemani 2007; Solé-Ollé and Sorribas-Navarro 2007). Among municipalities run by right-wing mayors, the main beneficiaries were those that were both right-leaning and electorally fragmented. Under the assumption that political fragmentation proxies for swing constituencies, these results are consistent with the "aligned swing" prediction of ADDD's model. Overall, these findings provide clear evidence of special-interest politics in what was supposed to be a non-partisan government program.

There are two main caveats to the analysis presented here. The first is that although the findings are consistent with the theories discussed above, alternative interpretations are also possible. For example, Ames (1995a, 1995b, 2001) argues that presidential coalition-building strategies are partly based on deputies trading votes for discretionary grants from the federal executive. It seems reasonable to speculate that alliances between deputies and local mayors are more likely to happen if they share the same political orientation. Such alliances, in turn make it more likely for a deputy to trade his vote for presidential favors, which would be an alternative mechanism through which transfers are skewed towards local governments run by right-wing parties.

The second caveat is that the reported correlations between political conditions and municipality population estimates might significantly understate the true extent of special-interest politics. For example, bureaucrats may have simply bumped up those municipalities which paid the highest bribes. This type of corruption would be exceedingly hard to detect in the data. It is also conceivable that favored municipalities were part of influential federal politicians' networks. In exchange for funds transferred under the FPM, federal politicians likely received monetary kickbacks which they used to finance their campaign spending and cultivate their personal vote. Municipalities that are in the network are not necessarily the municipalities that provided most electoral support for federal politicians, however, which makes this type of special-interest politics difficult to detect (Samuels, 2002).

Nonetheless, the results presented here do suggest that the exclusive focus on discretionary transfers in the extant literature may considerably understate the true scope of tactical redistribution that is going on under programmatic guise. Investigation of other seemingly special-interest-proof programs, including direct transfer programs to individuals, is thus an obvious avenue for future research.

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Table 1: Brackets and coefficients for the FPM transfer

Population bracket				*Coefficient*
up to	10,188			0.6
from	10,189	to	13,584	0.8
from	13,585	to	16,980	1
from	16,981	to	23,772	1.2
from	23,773	to	30,564	1.4
from	30,565	to	37,356	1.6
from	37,357	to	44,148	1.8
from	44,149	to	50,940	2
from	50,941	to	61,128	2.2
from	61,129	to	71,316	2.4
from	71,317	to	81,504	2.6
from	81,505	to	91,692	2.8
from	91,693	to	101,880	3
from	101,881	to	115,464	3.2
from	115,465	to	129,048	3.4
from	129,049	to	142,632	3.6
from	142,633	to	156,216	3.8
above	156,216			4

Source: Decree 1881/81

Table 2: Bracket error distribution

Panel A, bunched municipalities
1989 official population classification

Bracket error	Freq.	Percent	Cum.
-6	2	0.11	0.11
-5	7	0.38	0.48
-4	6	0.32	0.80
-3	14	0.75	1.56
-2	49	2.63	4.18
-1	301	16.15	20.33
0	911	48.87	69.21
1	387	20.76	89.97
2	132	7.08	97.05
3	36	1.93	98.98
4	15	0.80	99.79
5	3	0.16	99.95
6	1	0.05	100.00
7	0	0.00	100.00

Panel B, bunched municipalities
1991 official population classification

Bracket error	Freq.	Percent	Cum.
-6	0	0.00	0.00
-5	0	0.00	0.00
-4	2	0.11	0.11
-3	2	0.11	0.21
-2	5	0.27	0.48
-1	38	2.04	2.52
0	318	17.06	19.58
1	1051	56.38	75.97
2	333	17.86	93.83
3	76	4.08	97.91
4	24	1.29	99.20
5	11	0.59	99.79
6	3	0.16	99.95
7	1	0.05	100.00

Panel C, non-bunched municipalities
1989 official population classification

Bracket error	Freq.	Percent	Cum.
-7	1	0.04	0.04
-6	1	0.04	0.08
-4	3	0.12	0.20
-3	4	0.16	0.36
-2	33	1.34	1.70
-1	200	8.10	9.81
0	1976	80.06	89.87
1	212	8.59	98.46
2	29	1.18	99.64
3	6	0.24	99.88
4	2	0.08	99.96
5	0	0.00	99.96
9	1	0.04	100.00

Panel D, non-bunched municipalities
1991 official population classification

Bracket error	Freq.	Percent	Cum.
-7	0	0.00	0.00
-6	1	0.04	0.04
-5	0	0.04	0.04
-3	3	0.12	0.16
-2	9	0.36	0.53
-1	43	1.74	2.72
0	1941	78.65	80.92
1	404	16.37	97.29
2	45	1.82	99.11
3	16	0.65	99.76
4	5	0.20	99.96
5	1	0.04	100.00

Notes: Total number of municipalities in Panels A and B is 1864 and in Panels C and D the number is 2468. Bunched municipalities refers to those located on any of the bunching points identified in the main text. The tabulation excludes municipalities that were created between 1989 and 1991. Bracket error is defined as 5×[k(19XX official population)-k(1991 actual population)], where k(.) is the step function defined in decree 1881/81 and XX=89,91.

Table 3: Summary Statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
1991 actual population, census ('000)	4451	24.32	48.83	0.75	846.4
1991 population forecast error ('000)	4451	2.57	10.10	-108.5	476.9
1991 bracket error using 1991 official pop.	4451	0.57	0.89	-6	7
1991 bracket error using 1989 official pop.	4332	0.07	0.9	-7	9
1991 bunch status (0/1)	4451	0.42	0.49	0	1
1989-1992 right-wing mayor (0/1)	4276	0.53	0.49	0	1
1989-1992 PFL mayor (0/1)	4276	0.24	0.43	0	1
1989-1992 PDS mayor (0/1)	4276	0.10	0.30	0	1
1989-1992 PTB mayor (0/1)	4276	0.07	0.26	0	1
1989-1992 PDC mayor (0/1)	4276	0.05	0.22	0	1
1989-1992 PL mayor (0/1)	4276	0.05	0.22	0	1
1988 electorate ('000)	4276	18.6	118.6	0	6'057.5
1990-1994 interparty fragmentation	3761	0.67	0.14	0.06	0.98
1990-1994 right-wing vote share	3761	0.37	0.22	0	0.98
1985 actual population, interpolated ('000)	4490	29.01	172.8	0.78	9'016.8
1985 population forecast error ('000)	3942	-1.44	7.55	-115.8	69.5
1985 bracket error	3942	-0.10	0.76	-9	8
1982-1988 right-wing mayor (0/1)	3932	0.64	0.48	0	1
1982-1988 PDS mayor (0/1)	3932	0.64	0.48	0	1
1982-1988 PTB mayor (0/1)	3932	0.002	0.042	0	1
1980 actual population, census ('000)	4017	29.33	171.2	0.73	8'493.2
1982-1986 interparty fragmentation	4086	0.57	0.30	0	0.99
1982-1986 right-wing vote share	4086	0.62	0.23	0.03	1

Right-wing mayor includes mayors affiliated with the PFL, PDS, PTB, PDC and PL. Partido Frente Liberal (PFL), Partido Democratico Social (PDS), Partido Trabalhista Brasileiro (PTB), Partido Democrata Cristao (PDC), Partido Liberal (PL), Partido da Reconstrução Nacional (PRN). Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the respective Camara (House) elections.

Table 4

Dependent Variable: 1991 official population						
Right-wing mayor	0.053	0.579	0.287	2.599**	2.608**	2.283*
	(0.967)	(1.340)	(1.294)	(1.206)	(1.196)	(1.183)
Right-wing mayor× Right-wing vote share	0.736	-4.505***	-4.339***	-4.531***	-4.522***	-4.404***
	(0.720)	(1.694)	(1.660)	(1.462)	(1.461)	(1.490)
Right-wing mayor× Interparty fragmentation	2.521**	-1.249	-2.339	-4.244***	-4.231***	-4.021***
	(1.065)	(1.636)	(1.554)	(1.439)	(1.445)	(1.485)
Right-wing mayor× Right-wing vote share× Interparty fragmentation		9.179***	8.846***	9.861***	9.827***	9.137***
		(3.072)	(3.006)	(2.707)	(2.712)	(2.735)
Non-right-wing mayor× Right-wing vote share	0.430	-4.893	-5.587*	-3.137	-3.060	-3.487*
	(1.446)	(3.006)	(3.024)	(2.103)	(2.078)	(2.102)
Non-right-wing mayor× Interparty fragmentation	3.546***	0.702	-0.995	0.564	0.606	0.222
	(1.215)	(2.130)	(2.086)	(1.625)	(1.577)	(1.542)
Non-right-wing mayor× Right-wing vote share× Interparty fragmentation		9.023	10.38	6.471*	6.284*	6.386*
		(6.839)	(6.854)	(3.872)	(3.814)	(3.817)
Electorate 1988, (Electorate 1988) ^{2}	Y	Y	Y	Y	Y	Y
(Electorate 1988) ^{3}	N	N	Y	Y	Y	Y
Actual 1991 population	N	N	N	Y	Y	Y
(Actual 1991 population) ^{2}	N	N	N	Y	Y	Y
(Actual 1991 population) ^{3}	N	N	N	N	Y	Y
1991 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero	2.89	3.08	2.06	2.90	2.71	2.31
$\alpha_1=\alpha_2=\alpha_3=\alpha_4=\alpha_5=\alpha_6=\alpha_7=0$	(0.013)	(0.003)	(0.044)	(0.005)	(0.008)	(0.024)
Same regression functions	1.14	1.05	1.00	1.88	1.80	1.54
$\alpha_1=0,\alpha_2=\alpha_5,\alpha_3=\alpha_6,\alpha_4=\alpha_7$	(0.333)	(0.382)	(0.406)	(0.111)	(0.121)	(0.187)
Same effect of interparty fragmentation	0.54	0.86	0.81	3.73	3.57	3.02
$\alpha_3=\alpha_6,\alpha_4=\alpha_7$	(0.462)	(0.423)	(0.406)	(0.024)	(0.028)	(0.049)
Observations	3564	3564	3564	3564	3564	3563
R-squared	0.970	0.970	0.986	0.986	0.986	0.986

Notes: Right-wing consists of the following political parties: PFL, PDS, PTB, PDC, PL. Right-wing vote share is from the 1990 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1990 Camara elections. Other covariates (not shown) included with actual 1991 population are 1991 actual population bracket classification effects. County characteristics are 1991 income per capita, average years of schooling, poverty rate, gini index and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 5

Dependent Variable: 1991 bracket error						
Right-wing mayor	0.427***	0.662***	0.634***	0.554***	0.553***	0.512***
	(0.135)	(0.199)	(0.194)	(0.159)	(0.159)	(0.158)
Right-wing mayor× Right-wing vote share	-0.016	-0.522*	-0.506*	-0.610***	-0.611***	-0.594**
	(0.0952)	(0.282)	(0.273)	(0.224)	(0.224)	(0.231)
Right-wing mayor× Interparty fragmentation	0.376***	0.012	-0.094	-0.331*	-0.333*	-0.294
	(0.145)	(0.249)	(0.243)	(0.195)	(0.195)	(0.196)
Right-wing mayor× Right-wing vote share× Interparty fragmentation		0.884*	0.852*	1.151***	1.155***	1.071***
		(0.480)	(0.468)	(0.388)	(0.388)	(0.398)
Non-right-wing mayor× Right-wing vote share	0.075	0.175	0.107	0.085	0.076	0.023
	(0.117)	(0.392)	(0.384)	(0.330)	(0.330)	(0.329)
Non-right-wing mayor× Interparty fragmentation	0.917***	0.960***	0.794***	0.496**	0.491**	0.448**
	(0.153)	(0.233)	(0.230)	(0.202)	(0.201)	(0.200)
Non-right-wing mayor× Right-wing vote share× Interparty fragmentation		-0.161	-0.028	0.029	0.052	0.086
		(0.647)	(0.632)	(0.562)	(0.564)	(0.560)
Electorate 1988, $(\text{Electorate 1988})^2$	Y	Y	Y	Y	Y	Y
$(\text{Electorate 1988})^3$	N	N	Y	Y	Y	Y
Actual 1991 population	N	N	N	Y	Y	Y
$(\text{Actual 1991 population})^2$	N	N	N	Y	Y	Y
$(\text{Actual 1991 population})^3$	N	N	N	N	Y	Y
1991 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero	8.93	7.30	5.70	4.30	4.32	3.85
$\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_6 = \alpha_7 = 0$	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Same regression functions	3.41	3.45	3.58	3.47	3.47	3.25
$\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.016)	(0.008)	(0.006)	(0.007)	(0.007)	(0.011)
Same effect of interparty fragmentation	7.87	5.47	5.36	5.67	5.67	4.82
$\alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.005)	(0.004)	(0.005)	(0.003)	(0.003)	(0.008)
Observations	3564	3564	3564	3564	3564	3563
R-squared	0.117	0.118	0.152	0.359	0.359	0.376

Notes: Right-wing consists of the following political parties: PFL, PDS, PTB, PDC, PL. Right-wing vote share is from the 1990 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1990 Camara elections. Other covariates (not shown) included with actual 1991 population are 1991 actual population bracket classification effects. County characteristics are 1991 income per capita, average years of schooling, poverty rate, gini index and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6

Dependent Variable: positive 1991 bracket error (0/1)						
Right-wing mayor	0.187**	0.353***	0.335***	0.289***	0.285***	0.267***
	(0.0806)	(0.121)	(0.117)	(0.102)	(0.101)	(0.100)
Right-wing mayor× Right-wing vote share	-0.031	-0.193	-0.183	-0.270*	-0.274**	-0.272**
	(0.052)	(0.170)	(0.164)	(0.138)	(0.136)	(0.137)
Right-wing mayor× Interparty fragmentation	0.383***	0.267*	0.201	-0.027	-0.033	-0.028
	(0.081)	(0.144)	(0.140)	(0.122)	(0.120)	(0.120)
Right-wing mayor× Right-wing vote share× Interparty fragmentation		0.282	0.262	0.485**	0.501**	0.476**
		(0.287)	(0.278)	(0.238)	(0.234)	(0.236)
Non-right-wing mayor× Right-wing vote share	-0.019	0.314	0.272	0.204	0.167	0.141
	(0.063)	(0.229)	(0.221)	(0.192)	(0.192)	(0.194)
Non-right-wing mayor× Interparty fragmentation	0.640***	0.809***	0.706***	0.438***	0.418***	0.390***
	(0.091)	(0.139)	(0.137)	(0.122)	(0.121)	(0.120)
Non-right-wing mayor× Right-wing vote share× Interparty fragmentation		-0.559	-0.476	-0.358	-0.267	-0.250
		(0.370)	(0.358)	(0.317)	(0.316)	(0.318)
Electorate 1988, $(\text{Electorate } 1988)^2$	Y	Y	Y	Y	Y	Y
$(\text{Electorate } 1988)^3$	N	N	Y	Y	Y	Y
Actual 1991 population	N	N	N	Y	Y	Y
$(\text{Actual 1991 population})^2$	N	N	N	Y	Y	Y
$(\text{Actual 1991 population})^3$	N	N	N	N	Y	Y
1991 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero	12.71	10.08	7.69	3.99	4.11	3.69
$\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_6 = \alpha_7 = 0$	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Same regression functions	1.89	2.27	2.21	2.16	2.11	1.98
$\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.129)	(0.060)	(0.065)	(0.070)	(0.077)	(0.094)
Same effect of interparty fragmentation	5.26	4.26	4.03	3.95	3.85	3.37
$\alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.022)	(0.014)	(0.018)	(0.019)	(0.021)	(0.034)
Observations	3564	3564	3564	3564	3564	3563
R-squared	0.118	0.119	0.163	0.344	0.358	0.366

Notes: Right-wing consists of the following political parties: PFL, PDS, PTB, PDC, PL. Right-wing vote share is from the 1990 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1990 Camara elections. Other covariates (not shown) included with actual 1991 population are 1991 actual population bracket classification effects. County characteristics are 1991 income per capita, average years of schooling, poverty rate, gini index and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 7

Dependent Variable: 1985 official population						
Right-wing mayor	-4.689**	-0.210	1.228	-2.465	-3.991	-3.312
	(2.222)	(3.484)	(3.431)	(3.776)	(3.934)	(3.771)
Right-wing mayor× Right-wing vote share	-0.919	-2.032	-1.174	-1.495	-0.099	-0.310
	(1.284)	(1.964)	(1.782)	(2.637)	(2.860)	(2.632)
Right-wing mayor× Interparty fragmentation	-1.083	-2.520	-1.761	-3.279	-2.052	-2.054
	(0.864)	(2.336)	(2.255)	(2.595)	(2.623)	(2.462)
Right-wing mayor× Right-wing vote share× Interparty fragmentation		1.414	1.220	3.384*	2.908	2.955
		(2.207)	(2.199)	(2.035)	(1.951)	(1.875)
Non-right-wing mayor× Right-wing vote share	-5.138***	-0.918	1.289	-5.119*	-5.187	-4.629
	(1.741)	(2.941)	(3.025)	(2.857)	(3.159)	(3.051)
Non-right-wing mayor× Interparty fragmentation	-4.589***	-0.821	1.606	-5.562*	-5.561*	-4.808
	(1.612)	(2.875)	(3.006)	(2.876)	(3.167)	(3.059)
Non-right-wing mayor× Right-wing vote share× Interparty fragmentation		-4.295	-5.643*	6.930**	6.804**	6.119**
		(2.960)	(3.119)	(2.800)	(2.940)	(2.838)
1980 population, $(\text{1980 population})^2$	Y	Y	Y	Y	Y	Y
$(\text{1980 population})^3$	N	N	Y	Y	Y	Y
Actual 1985 population	N	N	N	Y	Y	Y
$(\text{Actual 1985 population})^2$	N	N	N	Y	Y	Y
$(\text{Actual 1985 population})^3$	N	N	N	N	Y	Y
1980 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero $\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_6 = \alpha_7 = 0$	1.93	1.50	1.39	2.18	2.12	1.64
	(0.086)	(0.163)	(0.207)	(0.033)	(0.038)	(0.119)
Same regression functions $\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	1.52	1.53	1.59	1.24	1.31	1.25
	(0.208)	(0.190)	(0.175)	(0.292)	(0.264)	(0.287)
Same effect of interparty fragmentation $\alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	4.38	3.02	3.00	0.66	0.56	0.41
	(0.036)	(0.049)	(0.050)	(0.517)	(0.570)	(0.664)
Observations	3877	3877	3877	3877	3877	3838
R-squared	0.991	0.991	0.991	0.993	0.994	0.994

Notes: Right-wing consists of the following political parties: PDS, PTB. Right-wing vote share is from the 1982 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1982 Camara elections. Other covariates (not shown) included with actual 1985 population are 1985 actual population bracket classification effects. County characteristics are 1980 income per capita, average years of schooling, poverty rate and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 8

Dependent Variable: 1985 bracket error						
Right-wing mayor	-3.048***	-3.294***	-2.833***	-0.414	-0.649	-0.633
	(1.041)	(1.097)	(1.034)	(0.703)	(0.654)	(0.639)
Right-wing mayor× Right-wing vote share	0.794**	0.229	0.430	-0.042	0.317	0.318
	(0.370)	(0.554)	(0.550)	(0.547)	(0.417)	(0.412)
Right-wing mayor× Interparty fragmentation	0.357	-0.331	-0.155	-0.243	0.067	0.067
	(0.250)	(0.572)	(0.570)	(0.515)	(0.425)	(0.421)
Right-wing mayor× Right-wing vote share× Interparty fragmentation		0.621	0.574	0.261	0.140	0.161
		(0.483)	(0.484)	(0.370)	(0.350)	(0.344)
Non-right-wing mayor× Right-wing vote share	-0.629	-2.674**	-2.234**	-0.313	-0.219	-0.240
	(0.622)	(1.078)	(1.033)	(0.550)	(0.566)	(0.571)
Non-right-wing mayor× Interparty fragmentation	-2.610***	-3.626***	-2.975***	-0.557	-0.422	-0.394
	(0.912)	(0.967)	(0.902)	(0.495)	(0.537)	(0.525)
Non-right-wing mayor× Right-wing vote share× Interparty fragmentation		2.654*	2.584*	0.191	0.157	0.212
		(1.542)	(1.501)	(0.709)	(0.735)	(0.746)
1980 population, $(\text{1980 population})^2$	Y	Y	Y	Y	Y	Y
$(\text{1980 population})^3$	N	N	Y	Y	Y	Y
Actual 1985 population	N	N	N	Y	Y	Y
$(\text{Actual 1985 population})^2$	N	N	N	Y	Y	Y
$(\text{Actual 1985 population})^3$	N	N	N	N	Y	Y
1980 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero $\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_6 = \alpha_7 = 0$	8.22	6.27	6.97	1.08	1.06	0.87
	(0.000)	(0.000)	(0.000)	(0.376)	(0.384)	(0.529)
Same regression functions $\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	4.43	3.21	2.62	0.17	0.56	0.68
	(0.004)	(0.012)	(0.049)	(0.954)	(0.694)	(0.609)
Same effect of interparty fragmentation $\alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	10.15	4.39	7.09	0.20	0.39	0.33
	(0.001)	(0.012)	(0.007)	(0.815)	(0.679)	(0.716)
Observations	3754	3754	3754	3754	3754	3723
R-squared	0.069	0.071	0.079	0.499	0.512	0.520

Notes: Right-wing consists of the following political parties: PDS, PTB. Right-wing vote share is from the 1982 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1982 Camara elections. Other covariates (not shown) included with actual 1985 population are 1985 actual population bracket classification effects. County characteristics are 1980 income per capita, average years of schooling, poverty rate and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 9

Dependent Variable: positive 1985 bracket error (0/1)						
Right-wing mayor	-0.626**	-0.624*	-0.787**	-0.202	-0.311	-0.298
	(0.291)	(0.377)	(0.398)	(0.357)	(0.331)	(0.326)
Right-wing mayor×	0.315**	0.248	0.177	0.007	0.176	0.195
Right-wing vote share	(0.154)	(0.213)	(0.223)	(0.236)	(0.185)	(0.184)
Right-wing mayor×	0.182*	0.099	0.037	-0.036	0.109	0.131
Interparty fragmentation	(0.103)	(0.204)	(0.210)	(0.211)	(0.176)	(0.173)
Right-wing mayor×		0.075	0.092	0.035	-0.021	-0.028
Right-wing vote share×		(0.160)	(0.159)	(0.141)	(0.134)	(0.129)
Interparty fragmentation						
Non-right-wing mayor×	-0.014	-0.173	-0.328	0.070	0.115	0.146
Right-wing vote share	(0.148)	(0.562)	(0.569)	(0.497)	(0.489)	(0.491)
Non-right-wing mayor×	-0.366	-0.446	-0.676*	-0.156	-0.093	-0.047
Interparty fragmentation	(0.228)	(0.330)	(0.348)	(0.293)	(0.296)	(0.292)
Non-right-wing mayor×		0.204	0.229	-0.289	-0.305	-0.337
Right-wing vote share×		(0.651)	(0.654)	(0.591)	(0.578)	(0.581)
Interparty fragmentation						
1980 population, (1980 population)^2	Y	Y	Y	Y	Y	Y
(1980 population)^3	N	N	Y	Y	Y	Y
Actual 1985 population	N	N	N	Y	Y	Y
(Actual 1985 population)^2	N	N	N	Y	Y	Y
(Actual 1985 population)^3	N	N	N	N	Y	Y
1980 municipality characteristics	N	N	N	N	N	Y
F-statistics and (p-values)						
All political determinants zero	2.67	2.43	2.61	0.71	0.84	1.15
$\alpha_1 = \alpha_2 = \alpha_3 = \alpha_4 = \alpha_5 = \alpha_6 = \alpha_7 = 0$	(0.020)	(0.017)	(0.011)	(0.663)	(0.554)	(0.331)
Same regression functions	3.86	3.21	2.96	0.62	1.08	1.30
$\alpha_1 = 0, \alpha_2 = \alpha_5, \alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.009)	(0.012)	(0.018)	(0.647)	(0.366)	(0.267)
Same effect of interparty fragmentation	4.69	1.57	2.41	0.56	0.90	0.87
$\alpha_3 = \alpha_6, \alpha_4 = \alpha_7$	(0.030)	(0.208)	(0.090)	(0.569)	(0.406)	(0.420)
Observations	3754	3754	3754	3754	3754	3723
R-squared	0.099	0.099	0.109	0.310	0.337	0.347

Notes: Right-wing consists of the following political parties: PDS, PTB. Right-wing vote share is from the 1982 elections for the Camara dos Deputados. Interparty fragmentation is defined as 1-H, where H is the Herfindahl Index applied to party vote shares at the municipality level in the 1982 Camara elections. Other covariates (not shown) included with actual 1985 population are 1985 actual population bracket classification effects. County characteristics are 1980 income per capita, average years of schooling, poverty rate and urbanization rate. State fixed effects included in all regressions. Heteroskedasticity-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Figure 1: 1989 predicted and official populations

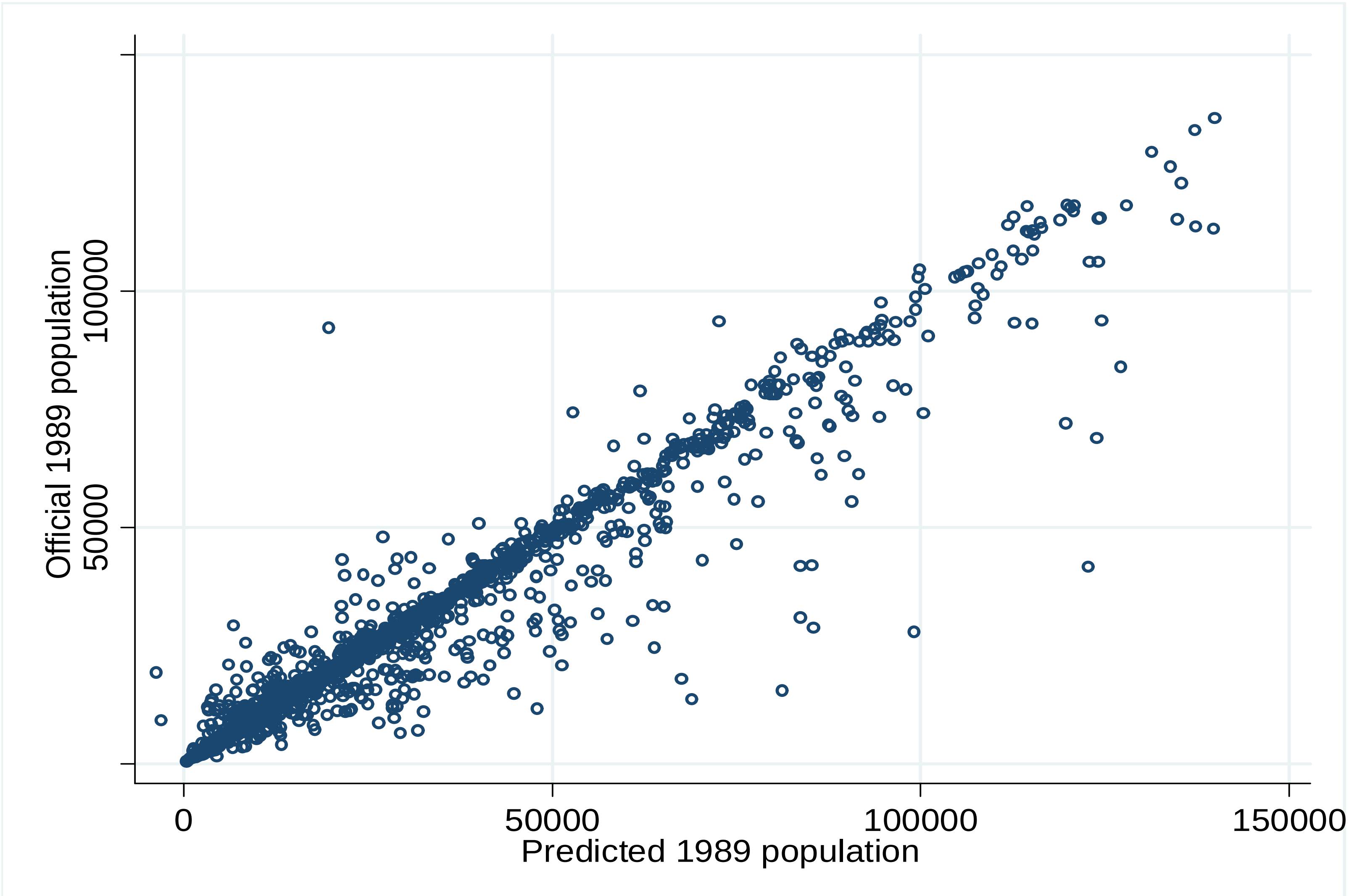


Figure 2: Histogram of 1989 official population

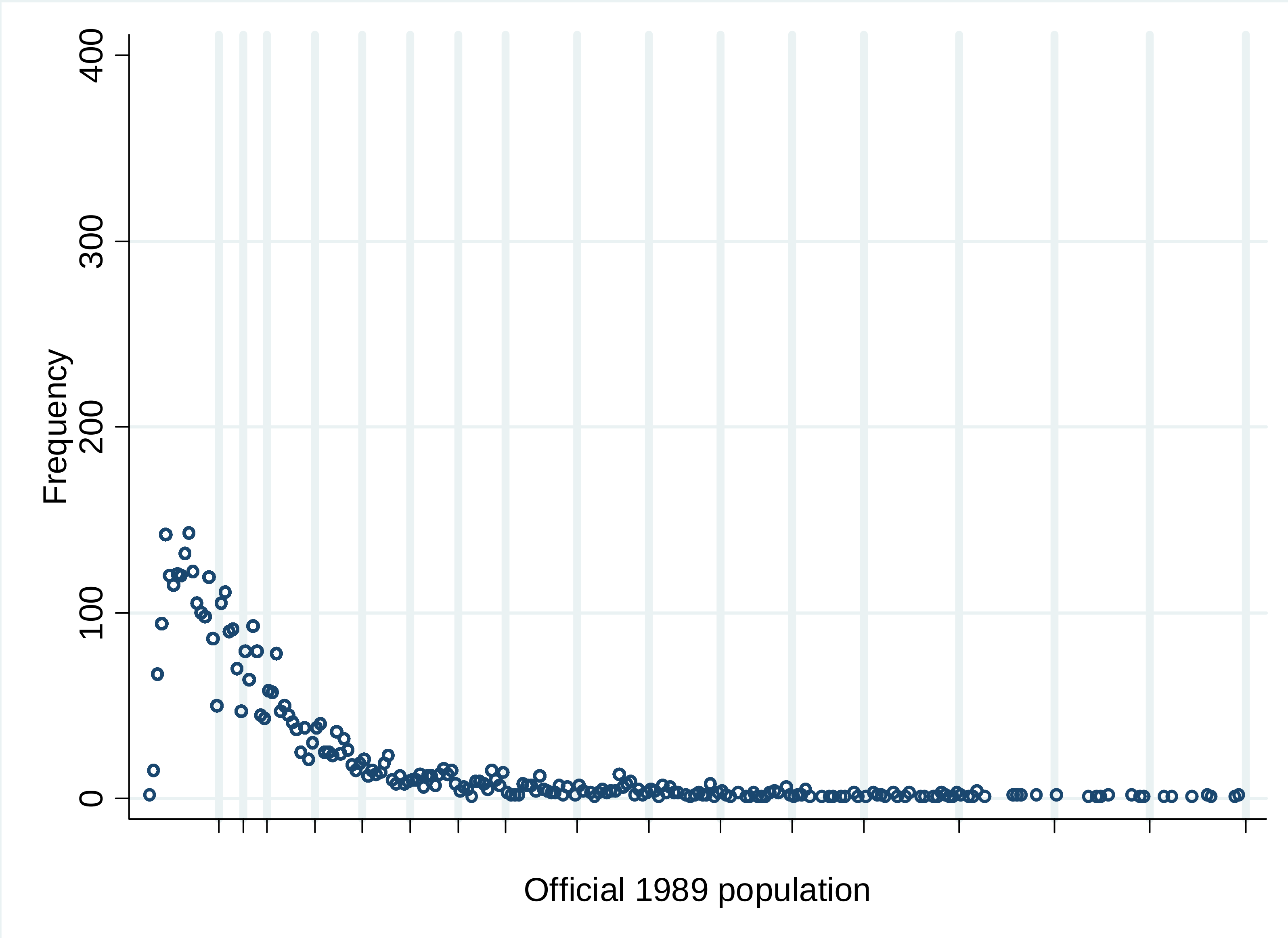


Figure 3: Histogram of 1991 official population

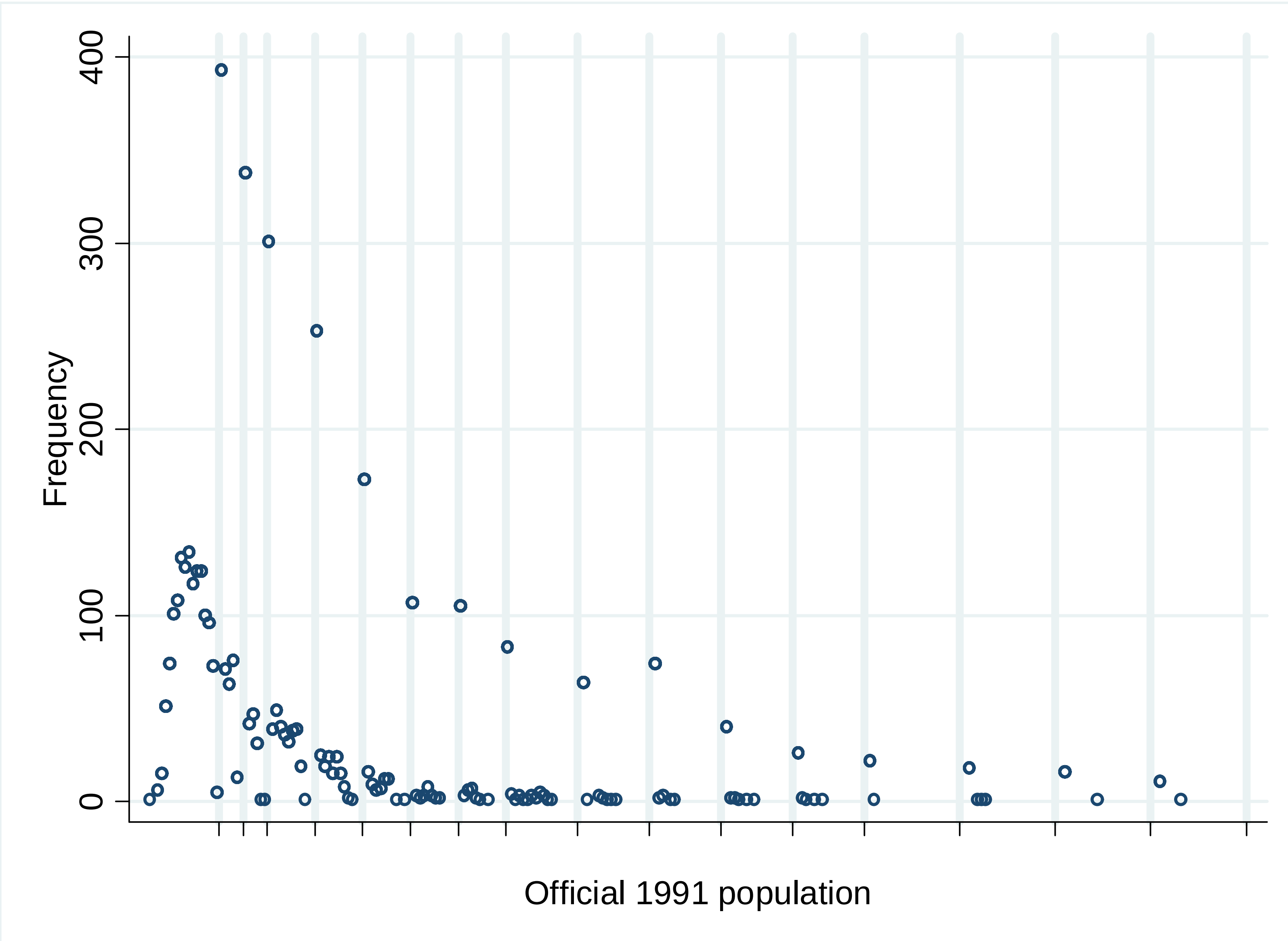


Figure 4: Histogram of 1991 actual population

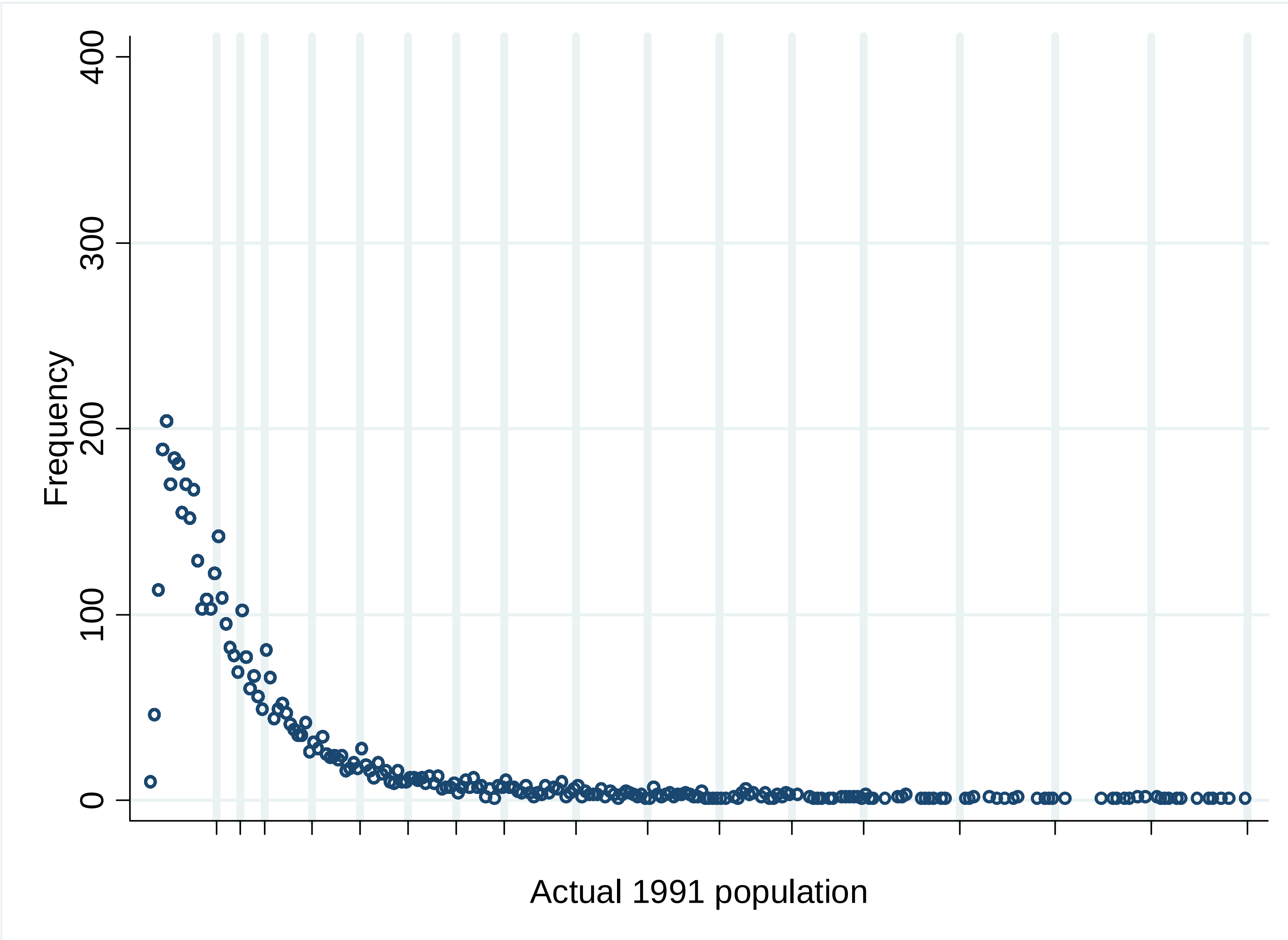
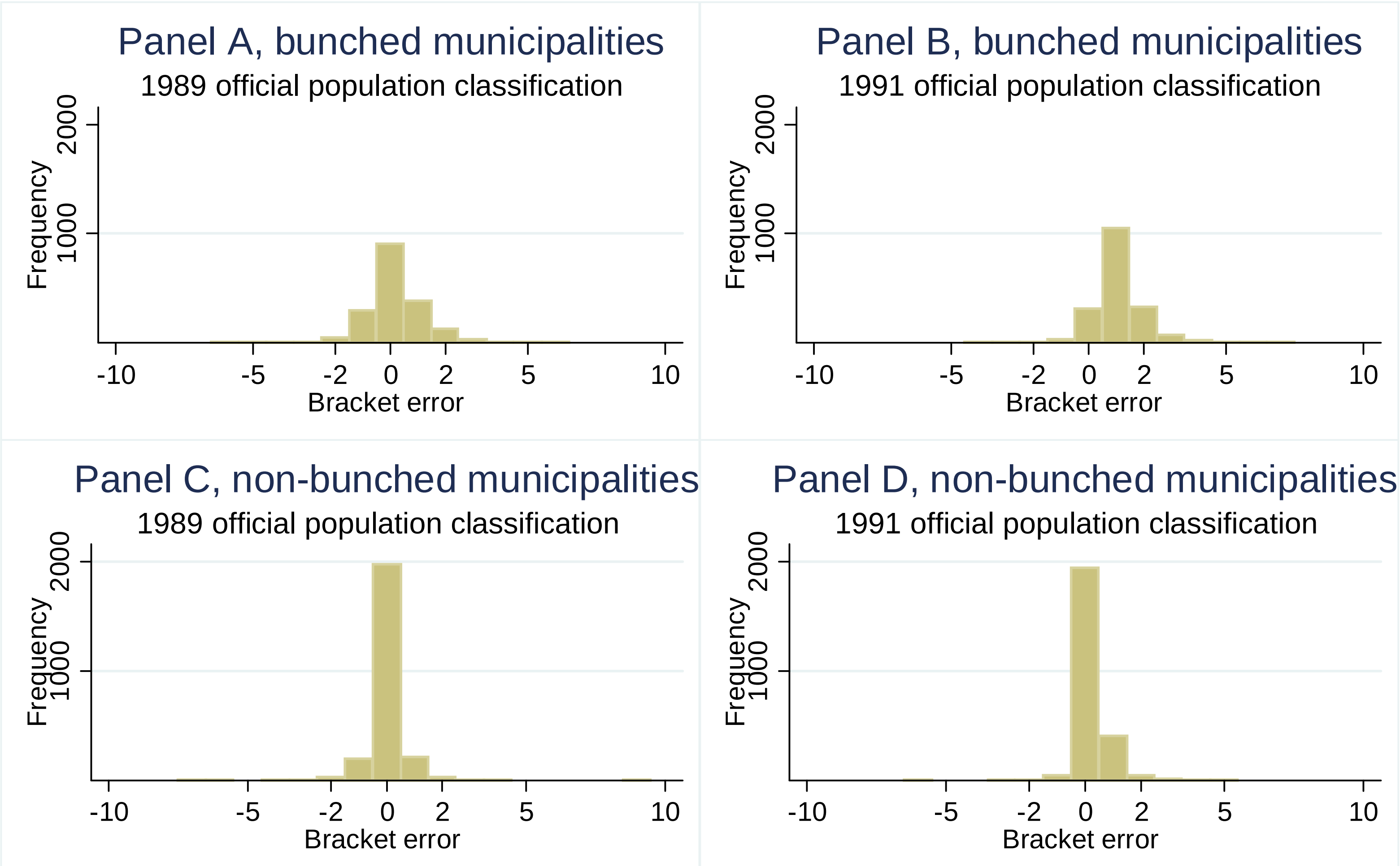


Figure 5: Bracket error distribution



Total number of municipalities in Panels A and B is 1864 and in Panels C and D the number is 2468. Bracket error is defined as 5×[k(19XX official population)-c(1991 actual population)], where k(.) is the step function defined in decree 1881/81 and XX=89,91. Bunched municipalities refers to those located on any of the bunching points identified in the main text.

Figure 6: FPM Transfers, 1991-1999 (in ´000 of 2008 Reais)