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Legislative organization and government spending: cross-country evidence

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Abstract

The ''Law of $1/n$'' posits that an increase in the number of elected representatives fuels excessive government spending. Despite its wide acceptance as a stylized fact, the Law of $1/n$ has received only limited empirical scrutiny, and the existing evidence for the American States provides mixed support for the thesis. This paper examines the Law of $1/n$ in bicameral and unicameral legislative structures using a cross-section of democratic countries. The results indicate that legislative size matters under both legislature structures, but bicameralism dampens the $1/n$ effect relative to unicameralism.

JEL classification: D72; H11; H50

Keywords: ; Law of 1/n; Logrolling; National legislatures; Government spending; Bicameralism

1. Introduction

Beginning with Gordon Tullock (1959), political economists have modeled fiscal policy in democratic regimes as a common pool problem. In this perspective voters or their elected representatives view the tax base as a common pool from which to finance constituent-specific projects, leading to the familiar problem of overutilization. Constituent groups internalize the benefits of the expenditures their

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legislators propose, while they internalize only a fraction of the requisite costs imposed on the whole economy.[1] Weingast et al. (1981, hereinafter, W-S-J) extended this insight and formalized the ''Law of $1/n$,'' which posits that fiscal inefficiency in the form of excessive spending increases with the number of legislative districts.[2] With the exception of a few studies little empirical evidence stands behind this conceptual proposition despite its broad acceptance as a stylized fact.

Gilligan and Matsusaka (forthcoming, 1995) examine the relationship between legislature size and spending using American State data in the first half and the second half of the 20th Century. They find in both time periods that spending increases with the size of state upper chambers, but not with the size of state lower chambers, or in effect, partial support for the Law of $1/n$. DelRossi and Inman (1999) examines spending decisions of the US Congress and finds that legislator support for public expenditures is negatively related to the constituent-specific cost burden. Internationally, the W-S-J hypothesis has been tested only indirectly. This literature refers to 'size fragmentation' and uses the number of participants in the fiscal policy process as analytically equivalent to the number of geographically-defined legislative districts.[3]

This paper extends the theory and evidence underlying the fiscal commons literature in two ways. First, we bring the role of bicameralism into the analysis of the Law of $1/n$, a framework originally developed within the context of a unicameral legislative structure. Bicameralism is widespread among democratic countries, yet analysis regarding this institution is minimal. Most studies of legislative organization ignore the potential effects of dividing the legislature into two distinct chambers. This has been particularly true in the theoretical and empirical examinations of the Law of $1/n$. Because both chambers must consent under a bicameral structure, chamber size may exert a different effect in a bicameral legislature than in a unicameral legislature. Though studies do examine

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[1]See Buchanan and Tullock (1962) and the surveys of the seminal public choice contributions by Dennis Mueller (1985); Ingberman and Inman (1988) and Tollison (1988). Inman and Fitts (1990) provides a useful framework for organizing the numerous studies of fiscal institutions. For further examinations of the relevant literature with regard to the American States, see Poterba (1996a,b) and Crain and Muris (1995). For literature summaries with regard to international fiscal institutions, see von Hagen (1992); von Hagen and Harden (1994); Alesina and Perotti (1996); Poterba and von Hagen (1999).

[2]See Stockman (1975) for a nonformal precursor of the distributive politics framework, and Stigler (1976) for an early exploratory analysis of legislature size. Stratmann (1997) addresses the role of logrolling in the pork barrel process and summarizes the theoretical and empirical literature.

[3]Kontopoulos and Perotti (1999) examines the fiscal effects of the number of spending ministers and the fragmentation of party coalitions. Both proxies for size fragmentation are positively correlated with government expenditures in OECD countries. Stein et al. (1999) examines a sample of Latin American countries and finds that government spending is positively related to legislative fractionalization which they proxy by the number of political parties and the number of seats held by parties. Other empirical studies include Hallerberg and von Hagen (1999), and Jones et al. (1999).

the Law of $1/n$ in bicameral chambers, no study accounts for the interaction between chambers. Instead, prior studies treat the chambers as if they are separate unicameral legislatures rather than treating them as separate veto players. The second extension of the paper is to examine the relationship between legislature size, bicameralism and government expenditures internationally. This comparative institutional analysis is not possible using American State data because only Nebraska has a unicameral system.

Section 2 extends the Law of $1/n$ analytical model to incorporate the role of bicameralism on fiscal policy decisions. Section 3 specifies the empirical model and presents the results for a sample of 24 democracies with bicameral legislatures for the 1970s and 1980s. Then, we expand the sample to include 14 democracies with unicameral systems and compare the magnitude of the $1/n$ effect under these two forms of legislative organization. Section 4 concludes by discussing the impact of dual-chamber legislative bodies on fiscal inefficiencies.

2. The Law of 1/*n*

2.1. The traditional model

W-S-J base their model of fiscal inefficiency on the incentives of elected representatives to return projects that benefit their geographically-based constituents. Formally, let $b_i(x)$ stand for the benefits of spending x dollars in district i to the constituents of legislator i, and let $c(x)$ stand for the costs the associated revenues impose on the whole economy. The efficient level of spending in district i is the x that solves $b_i'(x) = c_i'(x)$. Given n districts and an equal distribution of the tax burden across districts, the constituents of legislator i bear only $(1/n)$th of the cost of spending in district i. The legislator thus seeks spending for his or her district up to the point where $b_i'(x) = (1/n)\ c'(x)$. Assuming the norm of universalism where legislators logroll and defer to each other regarding district-targeted expenditures, total spending increases with the number of districts, n.

2.2. Bicameralism and the law of 1/n

Roughly two-thirds of all democratic countries have a unicameral legislative branch. However, most developed democracies have adopted a bicameral structure.[4]

Montesquieu, whose ideas heavily influenced the US constitutional creators, was the first person to notice the virtue of separating the legislature to differentiate bases of representation in bicameral chambers to limit majority exploitation. James

[4]Tsebelis and Rasch (1995) and Tsebelis and Money (1997) provide extensive discussions of the history and logic of bicameralism.

Madison, who was well aware of Montesquieu's reasons for favoring a divided legislature, states in *The Federalist*, Number 63, ''the danger [that representatives will betray the interests of constituents] will be evidently greater where the whole legislative trust is lodged in the hands of one body of men, than where the concurrence of separate and dissimilar bodies is required in every public act (Hamilton et al., 1992).'' Until recently, scholars have engaged in little discussion beyond these initial observations.

Along with the heuristic praise lauded on bicameral legislatures as a solution to social class conflicts, economists and political scientists provide theoretical support for the positive qualities of this institution. Tullock (1959) argues that two-chamber legislatures with differing electoral bases reflect a constitutional response to mitigate the common-pool problem. Buchanan and Tullock (1962) finds the most important feature of bicameral chambers is the difference in constituent preferences between chambers. The difference in inter-chamber constituent preferences ranges from identical fiscal policy preferences (meaning the chambers act de facto as a large unicameral chamber) to perfectly divergent policy preferences. In most cases, the true outcome lies between the extremes, meaning that intermediate outcomes result from ''the overlapping of the interest-group coalitions in each house.'' Stigler (1976) and Hayek (1979) make similar arguments, namely if the chambers were composed of the identical constituencies they would produce legislation no different than if it were one body. Tsebelis and Money (1997) highlight the importance of differing preferences across chambers stating, ''[T]he *political* dimension of bicameralism recognizes that different interests or preferences may be expressed in the two legislative bodies. Where preferences differ, conflict between the two houses arises over the legislative outcome.'' In sum, the general presumption echoed by these scholars is that bicameralism serves a purposeful function in democracy only if the chambers are composed of differing constituencies.

Of course, it is not obvious that spreading interests into separate chambers holds any advantage over the agglomeration of many diverse interests into a single unicameral legislature. However, an important element of bicameralism is that both chambers have veto power over the other. Persson et al. (1997) demonstrate formally the importance of 'unanimity' in the bicameral separation of powers in limiting budgetary outcomes. Each chamber in a bicameral structure must approve outcomes, and the interaction across chambers differs from logrolling within a chamber.[5] In essence, a bicameral legislature constitutes a bilateral monopoly in legislative power. Any legislative trade to which both chambers agree must result from the dual consent of a monopoly buyer and seller, unlike trade within each

[5]Buchanan and Tullock (1962) demonstrates the difference in outcomes that result from different decision rules. Intra-cameral trades are subject to majority rule, which makes trade less costly than the unanimity rule required from agreement with inter-cameral trade. McCormick and Tollison (1981) investigates the effect of bicameralism on interest group activities. They argue that more equal chamber sizes promote redistributional activity associated with pressure group politics.

chamber where many alternative opportunities for trade exist among numerous legislators.[6] Differences in the two chambers' constituencies add to the costs of such bilateral agreements. As Gilligan et al. (1989) note, ''[W]hen different interests dominate different houses, each interest, in effect, holds a veto over legislation.''

The traditional Law of $1/n$ thesis applies specifically to unicameral legislatures. In this sense the theoretical analysis in W-S-J of the common pool problem does not fully describe the budget process in many countries, or in all but one of the American States. Including the effect of bicameralism in the analysis of the common pool problem adds a key institutional element to empirical studies of the legislative process. Although Gilligan and Matsusaka examine the Law of $1/n$ in bicameral legislatures, because the American States utilize bicameral chambers they are unable to modify the W-S-J framework to account for bicameral interaction. Instead, their empirical analysis treats state legislatures as having two independent unicameral chambers. The presence of two interdependent chambers alters the analysis for two reasons, asymmetries in legislative power and differences in the representation across chambers. First, if one chamber has more political power than the other, outcomes may be asymmetrically favorable to the stronger house. In this case, the policy preference of one chamber is more relevant than that of the other chamber.[7]

We incorporate the asymmetry in power into the traditional model of the Law of $1/n$. Stated simply, if one chamber has less bargaining power than the other chamber, the policy preference of the weaker chamber has a smaller effect on final outcomes than the preference of the stronger chamber. And because a chamber's policy preferences are in part determined by its size, the size of the weaker chamber will have a smaller effect on final outcomes than the size of the more powerful chamber.

The institutional design of most national legislatures gives a greater amount of power to the lower chamber (Tsebelis and Money, 1997). Such asymmetries in power result from clearly identifiable constitutional provisions that range from allowing the lower house to resolve conflicts (e.g., Austria, Belgium, France), having a greater number of members when conflicts are decided in a joint session (e.g., Australia, Bolivia, India), and the power to propose budgets (e.g., the United States).[8] If this presumption is correct, we expect lower chambers to be more

[6]Diermeir and Myerson (1999) make a related point, that bicameral legislatures ''are more like monopolistic producers of complementary goods than like duopolistic producers of a common good.''

[7]See Binmore et al. (1986) for a game-theoretic model of bargaining that demonstrates that asymmetric power and expectations about agreement can influence bargaining outcomes.

[8]Gilligan and Matsusaka (forthcoming) suggest that an asymmetric delegation of power to upper chambers in American States may explain why the sizes of lower chambers have an insignificant effect on government spending. This favorable bias towards the upper chambers may derive from state constitutional provisions. Another possibility is that the smaller size of upper chambers allows members to solve collective action problems better than the lower chamber.

important than upper chambers in determining policy outcomes, thereby lessening the $1/n$ effect in upper chambers relative to lower chambers.

A second reason that bicameralism alters the traditional fiscal commons analysis relates to the diversity of constituents across the two chambers. This is the classic virtue of bicameralism reviewed above.[9] If the composition of chambers leads to differing equilibrium policy preferences, these differences can limit the potential legislative choice set. For example, similarity in the median legislators from the respective chambers implies that the chambers will be more likely to agree on the level and composition of parochial spending drawn from the common pool of government funds. Because both chambers must agree on fiscal outcomes, bicameral chambers will be more likely to agree on which constituents to tax or subsidize as the constituencies across chambers become more similar. In other words, similarity among the two chambers facilitates inter-cameral agreements.[10]

Tsebelis and Money (1997) further stresses that the rationality of representatives in bicameralism chambers alters legislator behavior. Consequently, the application of a single-chamber framework to the study of a bicameral institution induces bias ''because of the very assumption that legislators are rational.'' Rational legislators who desire parochial spending projects realize that constituent-specific spending projects that would receive majority approval within their legislative chamber are subject to the preferences of the second chamber. Taking this into account they demand a level of spending on constituent-specific projects consistent with the total amount of spending that will receive majority support in both chambers. Where the chamber constituencies differ, each legislator will rationally propose a lower amount of parochial spending than he or she would propose if one chamber were the monopoly supplier of legislation.

A final effect of differences in constituencies across chambers relates to its impact on the stability of legislative outcomes over time. Riker (1992) argues that legislative cycling between different majorities leads to unrest in legislative decisions that can lead to increased spending in a unicameral body. If a given majority feels that its policy may be overturned by cycling in the existing legislature or future legislatures, the majority may seek excessive spending to compensate for the uncertainty of future outcomes.[11] Adding a second chamber increases the chance that the median outcome will be a Condorcet winner, thereby

[9]See Hammond and Miller (1987); Cox and McKelvey (1984); Tsebelis and Rasch (1995) and Brennan and Hamlin (1992) for a discussion of the difficulty of reaching agreement across chambers. Also see Bradbury and Crain (2000) for a discussion of the importance of the median legislator in each chamber.

[10]See Crain (1979); Bradbury and Crain (1999); Bradbury (2000); Zywicki (1994) and Zywicki (1997) for further discussion of constituent similarity in bicameral chambers.

[11]Besley and Coate (1998) develops a model where the optimal fiscal policy in the present can lead to undesirable outcomes in the future by altering constituent preferences, thereby forcing legislatures to select the non-optimal, second-best policy that will minimize the likelihood that the less preferred policy will result in the future.

reducing the swings of cycles that bring about inefficient excessive spending with election cycles.[12]

In sum, countries typically design bicameral chambers to represent different constituencies, and this creates differences in policy preferences between the chambers. Different constituent preferences means that chambers will be less likely to agree on which groups to tax and subsidize in the legislative process, and power asymmetries affect the ability of each chamber to institute its own legislative preference. This constrains the level of constituent-specific spending below the amount each chamber would prefer individually. Bicameralism predictably muffles the Law of $1/n$ effect from adding an additional seat to bicameral legislatures compared to the effect in unicameral legislatures. This implies that the effect on spending from an additional chamber seat will be larger in unicameral legislatures than in bicameral legislatures.

3. Empirical model and results

3.1. Specification issues, variable definitions, and sample

Initially we apply the Gilligan and Matsusaka (1995, and forthcoming) empirical strategy to a cross-section of countries. The model is specified in Eq. (1).

$$G_{it} = \Psi \boldsymbol{N}_{it} + \Phi X_{it} + \alpha + \tau_t + \varepsilon_{it}. \tag{1}$$

The data sample pools time-series and cross-sectional data, and the variable subscript i denotes an observation on an individual country, and the subscript t denotes an observation in a particular year.[13] α is a constant term, and $\boldsymbol{\tau}_t$ represents a set of time dummy variables, one for each year in the sample. ε_{it} is the error term.

G_{it} is the level of government spending which we measure in two ways, as a percent of GDP (in Models 1–5 in Table 2), and in per capita terms (in Models 6–8 in Table 2). The vector $\boldsymbol{N}_{\mathrm{it}}$ includes two variables, one measuring the size of a country's upper chamber and the second measuring the size of its lower chamber. On average, the size of the upper chamber is 121 members with a standard

[12] In a related paper, Stratmann (1996) finds no evidence of cycling in the US Congress, which in part may result from its bicameral structure. Shepsle and Weingast (1981) argues that the costs of majoritarian instability encourage the proliferation of institutions, such as bicameralism, that induce stability.

[13] We selected countries using the Gastil Index of political freedom, including countries with an average Gastil rating of five or less. Gastil ratings greater than five mean that general elections with relatively wide suffrage are not the primary means of determining political leadership.

deviation of 97. The number of seats ranges from 20 (Iceland) to the maximum of 319 (France). The average size of the lower chamber is 281 members with a standard deviation of 172 ranging from 40 (Iceland) to 630 (Italy). The vector X_{it} includes a set of four control variables: log of population, population growth, log of real GDP per capita, and 'openness.'[14] These controls mirror those used by Gilligan and Matsusaka with the addition of the openness variable, which Rodrik (1999) finds to be a robust predictor of government spending.[15]

We estimate Eq. (1) using Weighted Least Squares (WLS), to control for detected heteroskedasticity for a sample of 24 democratic countries with bicameral legislatures using the period 1971 through 1989.[16] Table 1 lists the countries included, relevant legislative characteristics, and summary statistics. Fiscal data were obtained from the Penn-World Tables (Summers and Heston, 1995). Institutional data were obtained from yearly editions of the *Statesman's Yearbook,* and verified with the *CIA World Factbook,* 1995 and *Constitutions of the World* (Maddex, 1995).

Table 2 presents the results for three specifications, representing slight variations of Eq. (1). Models 1 and 2 serve as baseline estimates against which we compare the other models. Model 1 includes only the control variables (log of population, population growth, real GDP per capita, and openness). Model 2 includes only the legislative structure variables. Models 3 and 6 include the all variables without fixed-country effects. Models 4 and 7 exclude the openness variable. For comparison, Models 5 and 8 present the results using fixed-effects, although as noted, the slight variation in legislative size within countries limits the reliability of these parameter estimates.

In all the models reported in Table 2, the size of the lower chamber is positively related to government expenditures. The relationship is statistically significant in all regression specifications except Model 2 that excludes the controls. The size of the upper chamber is negatively related to spending except when the controls are dropped or when we include fixed-country effects. The coefficient estimates

[14] Of course, other factors such as electoral and party systems, climatic conditions or historical–cultural preferences may cause spending to differ among the nations. Thus, we examine fixed-effects dummy variables to control for such country-specific factors in some specifications. Although the number of seats across countries exhibits some variation over the sample time period, this variation is insufficient to render much confidence in a fixed-effects specification. For example, the correlation coefficient between the first year observations and the last year observations is 0.99 for upper chambers and 0.94 for lower chambers.

[15] *Openness* measures imports plus exports as a fraction of GDP, and we use a one-year lag in this variable. Rodrik (1999) finds that the log of openness is positively associated with government expenditures. He postulates that open economies need large government expenditures to offset volatility associated with external economic shocks.

[16] The following observations are excluded either because data were missing for certain years, or because the country was not a democracy at some point during the sample period: Bolivia (1971–80), Ecuador (1971–9, 1983–4), Greece (1971–4), Honduras (1971–80), India (1971–72), Jamaica (1971–2), Panama (1971–78), Portugal (1974–5), Spain (1971–7), and Thailand (1971–78).

Table 1
Summary statistics

Bicameral countries

Countries	Mean legislature size		Government spending as a fraction of GDP				
	Upper	Lower	Mean	Median	S.D.	Min	Max
Argentina	46	208	10.43	1.66	10.70	5.30	13.50
Australia	64	128	13.78	0.46	13.80	12.90	14.40
Austria	59	183	12.05	0.76	12.20	10.40	12.90
Belgium	175	212	12.40	0.52	12.60	11.20	13.10
Bolivia	27	117	18.98	3.07	18.50	14.70	24.30
Canada	103	274	12.81	0.70	13.00	11.60	13.80
Columbia	111	200	11.56	0.84	11.40	10.40	12.80
Dominican Republic	27	120	10.26	2.07	10.10	6.30	13.30
France	307	504	14.73	0.61	14.60	13.60	15.50
Iceland	20	40	15.37	1.22	15.50	12.70	17.50
India	243	538	27.39	2.36	27.10	24.10	31.50
Italy	315	630	11.96	0.34	12.00	11.20	12.60
Jamaica	35	58	15.14	1.47	15.60	12.40	16.80
Japan	252	500	8.81	0.39	8.90	7.90	9.20
Mexico	63	330	8.93	1.01	8.50	7.40	10.60
Netherlands	75	150	12.28	0.33	12.30	11.60	12.90
Norway	39	116	15.88	0.72	15.70	14.90	17.10
Panama	58	505	27.87	1.31	27.90	25.00	29.20
Spain	248	350	11.51	0.99	11.70	9.80	12.70
Switzerland	44	200	8.82	0.50	9.00	7.90	9.30
Thailand	270	352	17.29	1.37	17.80	14.50	19.10
Turkey	184	450	11.68	1.06	11.00	10.60	13.40
United States	100	435	13.49	0.52	13.40	12.70	14.70
All	122	281	13.61	12.60	4.70	5.30	31.50

Unicameral Countries

Country	Legislature size	Government spending as a fraction of GDP				
		Mean	Median	S.D.	Min	Max
Botswana	36	27.35	25.80	6.60	15.50	43.10
Denmark	179	20.63	20.80	1.67	17.80	23.50
Ecuador	71	17.14	16.45	1.53	15.40	19.10
Finland	200	14.89	15.50	1.23	12.60	16.20
Greece	300	13.60	13.80	0.64	12.70	14.30
Honduras	104	17.73	17.50	0.89	16.80	19.10
Israel	120	35.55	35.00	4.37	28.70	44.10
Mauritius	71	16.67	16.95	1.38	13.70	18.40
New Zealand	90	14.44	14.50	0.81	13.00	16.10
Panama	505	26.41	27.00	1.43	24.80	28.50
Portugal	223	17.48	18.20	2.27	13.40	20.20
Sri Lanka	165	17.96	18.60	2.42	13.90	22.10
Sweden	339	21.99	22.30	1.26	20.10	23.60
Turkey	412	11.84	11.65	0.64	11.20	12.90
All	182	19.96	18.25	6.95	11.20	44.10

Table 2
The effect of chamber size on government spending in bicameral legislatures[a]

Independent	Percent of GDP					Per Capita		
variables	1	2	3	4	5	6	7	8
Upper size		0.004964	−0.01134	−0.00559	0.012620	−0.00048	−0.00027	0.000545
		(2.90)**	(4.52)**	(2.39)**	(2.12)*	(2.72)**	(1.60)	(1.37)
Lower size		0.000517	0.016946	0.012337	0.003627	0.000885	0.000694	0.000395
		(0.47)	(8.65)**	(6.65)**	(2.51)**	(7.13)**	(5.83)**	(3.29)**
Log of population	−0.540324		−0.79287	−1.05727	3.510723	−0.05844	−0.07445	0.290762
	(5.57) **		(5.21)	(7.99)**	(3.98)**	(5.83)**	(8.50)**	(4.65)**
Population growth	−13.74258		38.93611	−27.8935	0.873937	1.90037	−2.69833	0.013426
	(0.85)		(1.51)	(1.17)	(0.10)	(1.03)	(1.69)	(0.02)
Log of real GDP	−1.979701		−1.87419	−1.68975	−1.935639	0.892647	0.892357	0.840671
	(9.00) **		(6.83)**	(6.56)**	(3.54)**	(52.59)**	(58.10)**	(22.23)**
Log of openness	0.518642		1.65187		−1.540253	0.098488		−0.10718
(lagged 1 year)	(2.27)*		(6.31)**		(4.68)**	(5.04)**		(4.59)**
Constant	33.43479	11.69562	27.3522	35.68703		−1.13361	−0.53819	
	(11.85)**	(39.23)**	(9.22)**	(13.40)**		(5.80)**	(3.45)**	
Country effects	NO	NO	NO	NO	YES	NO	NO	YES
Year effects	YES	YES	YES	YES	YES	YES	YES	YES
Adjusted *R*-squared	0.16	0.02	0.30	0.30	0.99	0.87	0.87	0.99
F-statistic	334	322	208	126	3595	13760	12510	202828
Panel observations	437	409	392	392	392	392	392	392

[a] *T*-statistics (absolute value) in parentheses, where ** indicates statistical significance at the one percent level, * indicates statistical significance at the five percent level.

indicate that a one-percent increase in the size of lower chambers leads to an increase in government expenditures between 0.24 and 0.35 percent (evaluated at the sample mean of expenditures = 13.61 percent). A one-percent increase in the size of upper chambers leads to a reduction in spending by between 0.05 and 0.1 percent. However, when we include fixed-country effects, in Models 5 and 8, the estimated coefficients for the size of the upper and lower chambers are positive. These parameters indicate that a one-percent increase in the size of an upper chamber results in a 0.11 percent increase in government spending as a percent of GDP. A one-percent increase in the size of the lower chamber results in a 0.08 percent increase in spending as a percent of GDP. Interestingly, with fixed-country effects, the estimated elasticity of spending with respect to chamber size across countries conforms quite closely to the elasticity for upper chambers in American States presented in Gilligan and Matsusaka (1995 and forthcoming). Their elasticity estimates for the state upper chambers range from 0.08 to 0.17.[17]

In sum, the cross-country results provide systematic support for the Law of $1/n$ in lower chambers in a bicameral system, but results are ambiguous with respect to upper chambers. Importantly, the results empirically support the perceived constitutional bias towards power in lower chambers. The operation of the Law of $1/n$ in one chamber, and not the other is consistent with the findings of Gilligan and Matsusaka (1995 and forthcoming). Power asymmetries that give the lower chamber more power over legislation may explain why the Law of $1/n$ is robust only in this chamber. If one chamber holds more power than the other, then adding members to the weaker chamber may not result in the hypothesized increase in spending.

We expect the fiscal impact of marginal changes in legislature size to be greater in unicameral than in bicameral systems. However, this econometric strategy does not explicitly treat the effects of bicameral interaction as the conceptual analysis above suggests. That is, Eq. (1) measures the fiscal commons effect of legislature size in the separate chambers on government spending assuming no bicameral effect. We explore the importance of chamber interaction empirically using a specification that allows us to identify the dampening effect of bicameral chambers.

We account for the bicameral effect by modifying Eq. (1), separating $\boldsymbol{N}$ into chamber-specific components N^L and N^U, which stand for lower and upper chamber sizes. The ideal amount of logrolled spending occurs where $G^L = zN^L$ for the lower chamber, and where $G^U = zN^U$ for the upper chamber. The outcome of G is determined by interaction between chambers influenced by asymmetric power and constituent differences. The power asymmetries influence the bargaining outcome so that $G = aG^L + (1 - a)G^U$, where a parameterizes the strength of the lower chamber relative to the upper chamber. Second, constituent heterogeneity

[17] Gilligan and Matsusaka find different elasticities in different time periods. The effect is 0.17 for the last period in their sample (1960–1990) and 0.08 for the first period (1902–1942). We reiterate that Gilligan and Matsusaka find insignificant effects for the size of lower chambers in American States.

should dampen the proclivity of legislatures to logroll spending. To capture this effect we add a bargaining parameter θ, such that $G = \theta(aG^L + (1-a)G^U)$. If $\theta = 1$, bargaining has no effect other than forcing a compromise between chambers. If $\theta < 1$, bargaining has a dampening effect. Thus, the spending equation modified to account for bicameral interaction is $G = \theta(aG^L + (1-a)G^U) = \theta azN^L + \theta(1-a)zN^U$. In a regression, $G = \alpha + d_1N^L + d_2N^U +$ controls, the parameters correspond to $d_1 = \theta az$ and $d_2 = \theta(1-a)z$. The coefficients impound the asymmetric power effect, the constituent homogeneity effect, and the $1/n$ effect. By adding unicameral countries to the regression it is possible to recover the bicameral parameters of interest.

We use Eq. (2) to estimate the relevant parameters.

$$G_{it} = b_1 N_{it}^{UNI} + d_1N_{it}^L + d_2N_{it}^U + \Phi X_{it} + \alpha + \tau_t + \varepsilon_{it}. \tag{2}$$

Like Eq. (1), Eq. (2) pools time-series and cross-sectional data. Here the sample includes an enlarged sample of 35 countries, 14 with unicameral and 24 with bicameral legislatures.[18] Vector $\boldsymbol{X}_{it}$ represents the same set of control variables used in Eq. (1).

To separate the effect of the marginal member in the two legislative structures Eq. (2) includes three variables measuring chamber size for unicameral, upper, and lower chambers. N_{it}^{UNI} is the number of seats in unicameral legislatures (and equal to zero for observations with bicameral legislatures), N_{it}^L and N_{it}^U are the number of upper chamber and lower chamber seats in bicameral legislatures (set equal to zero for unicameral chambers). By including all the coefficients in a single equation we can recover the three hypothesized effects, because N_{it}^{UNI} captures the $1/n$ effect absent bicameral interaction. We recover the asymmetric bargaining coefficient by noting that $d_1/d_2 = a/(1-a)$, so $a = d_1/(d_1 + d_2)$. We recover θ using b_1, d_1, and d_2, here noting that $b_1 = z$ and $d_1 = \theta az$, and thus $\theta = (d_1 + d_2)/b_1$. The coefficient b_1 does not include the bicameral interaction and therefore is a pure measure of the fiscal commons effect in unicameral legislatures.

Table 3 lists the regression results for Eq. (2), again using spending as a share of GDP and spending per capita as alternative dependent variables. As predicted, unicameral size is positively related to government spending. A one-percent increase in unicameral legislature size is associated with a 0.17 percent increase in government spending as a percent of GDP (evaluated at the unicameral sample

[18] Table 1 also lists these 14 unicameral countries. Three countries changed from bicameral to unicameral or vice versa during the sample period; thus these three countries are sometimes included as bicameral and sometimes unicameral countries. The countries and the years they were bicameral are: Panama (1979–89), Sri Lanka (1971–2, listed only as unicameral because of its brief history as a bicameral country), and Turkey (1982–9). See Note 16 for a list of excluded observations. The average unicameral legislature contains 182 members with a standard deviation of 118. The size ranges from 36 (Botswana) to 505 (Panama).

Table 3
The effect of chamber size on government spending in unicameral versus bicameral legislatures[a]

Independent	Percent of GDP			Per Capita		
variables	1	2	3	4	5	6
Upper size	−0.00919	−0.00159	−0.00057	−0.0004	−7.32E−05	0.000137
	(3.96)**	(0.85)	(0.16)	(2.40)*	(0.54)	(0.66)
Lower size	0.01249	0.007432	0.005134	0.00056	0.000353	0.000507
	(7.62)**	(5.57)**	(3.36)**	(5.34)**	(3.94)**	(4.95)**
Unicameral size	0.019286	0.020288	0.005933	0.00113	0.001201	0.000658
	(14.97)**	(17.34)**	(3.27)**	(15.01)**	(18.21)**	(6.25)**
Log of population	−0.60981	−0.87644	2.84494	−0.04391	−0.06003	0.218156
	(5.81)	(8.74)**	(3.60)**	(6.50)**	(11.05)**	(4.38)**
Population growth	73.02741	50.95634	1.253346	2.602726	1.304295	0.126179
	(4.97)**	(3.65)**	(0.15)	(3.41)**	(1.99)*	(0.23)
Log of real GDP	−1.4788	−1.75553	−1.62532	0.90144	0.883557	0.823023
	(7.92)**	(9.73)**	(3.18)**	(86.52)**	(93.33)**	(25.54)**
Log of openness	2.625156		−1.79334	0.126594		−0.12455
	(10.88)**		(5.95)**	(7.35)**		(6.28)**
Constant	18.83146	34.96437		−1.39948	−0.56274	
	(7.46)**	(19.02)**		(8.87)**	(5.84)**	
Country effects	NO	NO	YES	NO	NO	YES
Year effects	YES	YES	YES	YES	YES	YES
Adjusted R-squared	0.26	0.22	0.92	0.99	0.85	0.99
F-statistic	273	929	4099	21564	55553	333808
Panel observations	606	606	606	606	606	606

[a] T-statistics (absolute value) in parentheses, where ** indicates statistical significance at the one percent level, * indicates statistical significance at the five percent level.

average spending of 20 percent). The results in Table 3 also give support to the positive relationship between lower chamber size and government spending. A one-percent increase in the size of the lower chamber is associated with a 0.26 percent increase in government spending (using the estimated coefficients in Model 1). Again, upper chamber size is ambiguously associated with government expenditures; only in Model 1 and Model 4 is upper chamber size statistically significant. Based on the results in Model 1 a one-percent increase in upper chamber size leads to a 0.08 percent decrease in government spending. Table 4 lists the recovered bicameral parameter values estimated from the different specifications. These parameters allow us to determine whether the differences in chamber influence result from bicameral interaction. Models 1 and 4 indicate that the asymmetric power parameter is 3.8 and 3.4, meaning that lower chambers are approximately 3.5 times more powerful than upper chambers in determining fiscal policy. However, we can reject the hypothesis that $a=1$ at the 10 percent level for Model 1, and we cannot reject the hypothesis for Model 4. The other estimates

Table 4
Parameter estimates of bicameralism

Model from Table 3	*a* Asymmetric power	*θ* Dampening effect	*z* Fiscal commons
Spending as a percent of GDP			
Model 1.	3.78[a]	0.17[b]	0.0193[b]
Model 2.	1.27[b]	0.28[b]	0.0202[b]
Log of spending per capita			
Model 4.	3.44	0.14[b]	0.00113[b]
Model 5.	1.26[b]	0.23[b]	0.001201[b]

[a] Indicates statistical significance at the ten percent level.
[b] Indicates statistical significance at the one percent level.

(Models 2 and 5) indicate a more modest power asymmetry, but these values are significantly different from 1 at the one-percent confidence level. Based on these values, the lower chamber appears 1.26 to 1.27 times more powerful than upper chambers. This finding is consistent with a constitutional bias among countries towards strong lower chambers.

The bargaining parameter indicates a bicameral dampening effect ranging from 0.14 to 0.28. Note that as θ approaches one, the bicameral result grows closer to the unicameral spending outcome. Again, this indicates that bicameral interaction dampens the marginal impact of the Law of $1/n$ relative to unicameral chambers by up to one-third. In all of the models the bargaining parameter is statistically significant at the 1 percent level,

In sum, we document two empirical regularities in bicameral chambers across countries. First, the influence of lower chambers on fiscal policy outcomes exceeds the influence of upper chambers by a factor of about 3.5. Second, a bicameral structure reduces the $1/n$ effect by roughly one-third relative to a unicameral structure.

4. Concluding remarks

The cross-country evidence indicates that the partition of legislative power into two chambers exerts a considerable impact on fiscal policy. This dimension was absent from the previous empirical literature analyzing legislative institutions. The results further support the positive relationship between legislative size and spending across countries, but the effect is far greater in unicameral legislatures than in bicameral legislatures. The policy implication is straightforward: splitting the legislative branch into two chambers mitigates the fiscal commons problem. Because different constituency preferences in the two chambers limit the level of spending of each chamber, future research might address in detail the degree to which the bases of representation differ between the chambers. Different bases of representation likely cause differences in inter-chamber policy preferences and thereby influence fiscal outcomes. The results also strongly support the general supposition that lower chambers hold more legislative power than upper chambers across countries. The power asymmetry likely results from explicit constitutional provisions that favor the lower chamber. This may explain why past studies of the Law of $1/n$ find support for the hypothesized effect in only one chamber.

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