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PARSIMONIOUS MODEL FOR PREDICTING MEAN CABINET DURATION ON THE BASIS OF ELECTORAL SYSTEM

Rein Taagepera and Allan Sikk

ABSTRACT

This study joins two existing logical models and tests the resulting predictions of mean cabinet duration (C). One of these models predicts C based on effective number of parties (N): $C = k/N^2$, where k is found to be around 42 years. The other predicts N on the basis of number of seats in the assembly (S) and district magnitude (M). The new combined model leads to a prediction for the mean cabinet duration in terms of these two institutional factors: $C = 42 \text{ years}/(MS)^{1/3}$. Three-quarters of the actual mean durations agree with the prediction within a factor of 2. For the purposes of institutional engineering, the model predicts that doubling the district magnitude would reduce the mean cabinet duration by 21 percent *ceteris paribus*.

KEY WORDS ■ causal model ■ electoral systems ■ party government ■ party systems

Why do most governmental cabinets tend to last long in some countries, while in some others they tend to change almost every year? This study is the first, to the best of our knowledge, to propose an explicit functional connection between the mean duration of cabinets and some institutional variables – a link that enables us to make specific quantitative predictions and to test them so as to establish the range of error within which they hold. The part of the variation that remains unexplained leaves room for other explanatory factors.

When new democracies decide on their institutions, typical duration of cabinets often is on the minds of decision-makers, among other concerns, foremost in the form of what they do *not* want: ‘Let us avoid short-lived cabinets like they have in . . .’. Excessively short-lived cabinets are seen as ineffective, and, more often than not, electoral systems get the blame –

namely excessive proportional representation (PR) in large multi-seat districts, as contrasted with PR in smaller multi-seat districts or use of single-seat districts (SSD). So the typical cabinet duration matters to the practitioners of politics, as does its presumed connection to institutions such as electoral system.

But what electoral system should we recommend to statesmen who consider two-year cabinets too short? They may not like to have 15-year cabinets either, because these may not only go stale but also exclude many politicians for too long. We are not aware of any existing theory or model that could offer a specific recommendation, supported by theoretical considerations and empirical confirmation. Here we do offer such a model.

Following the example of scholars such as Lijphart (1984, 1999), we deal with the arithmetic mean duration. If, over 30 years, the individual durations should vary appreciably, as it often does, say, 8, 1, 2, 8, 1, 1 and 9 years, then the mean (4.3 years) may be felt to reflect what is of concern to political practitioners more than the median (2 years) or the geometric mean (2.7 years).

Our aim is to establish an average relationship between this mean duration and some institutional factors that legislators could alter, in principle. This means we do not have to take a stand on whether excessively short or long durations affect government performance, a topic on which scholars disagree. A widespread view (e.g. Warwick, 1994: 139) is that short-lived cabinets are unlikely to provide effective policymaking and may over the longer run put regime survival in danger. Dogan (1989) disagrees, and Lijphart (1999: 130) puts it bluntly: 'This view is as wrong as it is widespread'. Short-lived cabinets need not be inefficient, and overly durable cabinets may go stale. Lijphart (1999: 131–9) sees mean cabinet duration as a useful indicator of something else, namely executive dominance, and he proceeds to measure it in various ways. To the extent that we can explain why some countries tend to experience shorter durations than others, we might also understand better what the mean cabinet durations signify and imply.

We do not address the 'micro' issue of why, within the same country with stable institutions, some cabinets last longer than others, taking a 'macro' view on the effect of electoral system on cabinet durability. A rich separate literature exists on this issue, focusing on bargaining models based on rational choice. Regarding the mean duration, however, the bargaining models have offered no specific quantitative predictions.¹ However, the parsimonious institutional model presented in this article is useful for other studies as it provides baseline expectations for analysing the effect of other factors – such as volatility, cabinet size, party system polarization, rules for appointing and dismissing of governments – on the mean cabinet duration. It should also be valuable for analysing the durability of individual cabinets in 'micro' level studies, where even factors like critical events could be taken into account.²

We may conjecture that certain features affect the mean cabinet duration. The number of parties visibly matters: coalition cabinets, prevalent in multi-party systems, tend to be more short-lived than one-party majority cabinets. This was extensively documented by Lijphart (1984: 124–6), using the effective number of legislative parties, $N = 1/\sum s_i^2$, where s_i is the fractional seat share of the i -th party. The number of parties, of course, cannot be altered by legislation, but it interacts with electoral systems. Single-seat districts (SSD) with a plurality seat allocation rule tend to correspond to two-party systems according to the well-known Duverger's law, while proportional representation (PR) in multi-seat districts tends to go with multiparty systems.³ In the case of PR, the number of parties is affected by district magnitude (the number of seats allocated within a district, M). The larger the district magnitude, the more parties the system can accommodate. At the same mean magnitude, the number of districts in the given country also could matter, since more districts can offer niches to more parties. This means that the total assembly size (S) could affect the number of parties.⁴

On this basis, we could go ahead and test a *directional* prediction: mean cabinet duration (C) decreases as the number of parties (N) increases, or, more briefly, $dC/dN < 0$. Also, the number of parties increases with increasing district magnitude (i.e. $dN/dM > 0$) and with increasing total number of seats in the assembly (i.e. $dN/dS > 0$). It follows that we expect mean cabinet duration to decrease as either district magnitude or assembly size increases: $dC/dM < 0$ and $dC/dS < 0$.

Testing such directional predictions is usual in political science. However, even when successful, it would only enable us to tell the political practitioners that, in order to increase an existing mean duration, they would have to reduce district magnitude and/or assembly size. But by *how much* should they reduce them? Would it be a politically feasible or unthinkable reduction? Here the directional model reaches its limit. We need the full functional relationship $C = f(M, S)$, not just the signs of its differentials (dC/dM and dC/dS). It would be reckless to assume as a matter of faith that this decreasing function of M and S is linear ($C = a - bM - cS$), because most basic equations in mature sciences are not linear (Colomer, 2007; Crease, 2004; McGregor, 1993). The format of the functional relationship must be established on logical grounds (Coleman, 2007). It often involves a numerical parameter to be determined empirically, and establishing such constants is an important part of testing the model (Hedström, 2004; Sørensen, 1998).

Can we build on these qualitative considerations presented and establish a model $C = f(M, S)$ that would enable us to make specific predictions about the expected mean cabinet duration, with a specified range of likely error? Actually, pieces for such a model have been around for some time. The present study links them, so as to establish a logical and testable chain extending from assembly size (S) and district magnitude (M) to cabinet duration (C). The resulting purely institutionally based predictions will be tested. The intermediary stages involve the fractional seat share of the largest party (s_1)

and the effective number of parliamentary parties (N). Predictions for cabinet duration at these intermediary stages will also be tested, so as to see where the random error range expands. A brief preview of this study, based on an earlier draft, has been reported in Taagepera (2007). Here the full data set is presented and used for thorough testing of the various stages of the logical model proposed.

What is random error from the viewpoint of a model in M and S includes the systematic effect of various other political – such as the frequency of majority governments and presence/absence of investiture requirements – and cultural factors. No logical quantitative model seems yet to be available for the impact of such other factors, important as they may be. This is why we focus on institutions. Once the institutional impact of M and S is factored out, the resulting residues may offer a clearer basis for studying the complementary impact of other factors. Indeed, we eagerly subscribe to Michael Laver's contention that:

... clearly needed ... is a well-grounded and coherent theoretical model generating observable implications that can be tested ... in stark contrast to almost all reported empirical work on government termination, which has tended to assemble a portfolio of independent variables gleaned from previous work and the author's own ideas, each given a brief ad hoc 'theoretical' justification in its own terms. (2003: 30)

The Institutional Model

A logical connection between the mean cabinet duration and the effective number of parties is outlined in Taagepera and Shugart (1989: 99–101). It is based on the number of communication channels, which can also become conflict channels. More parties mean more potential conflict channels that can undo a cabinet. The outcome is an inverse square law: $C = k/N^2$, where N is the effective number of legislative parties and k is a constant that comes in units of time (e.g. years or months) and is determined empirically.⁵ Using Lijphart's (1984) data, Taagepera and Shugart (1989) found that $C = 400 \text{ months}/N^2 = 33 \text{ years}/N^2$ predicts mean cabinet duration for stable democracies within a factor of 2. It will be seen that our analysis of more extensive data (Lijphart, 1999) puts the best fit at $C = 42 \text{ years}/N^2$. What this model claims is that the actual duration has an equal probability of being above or below $42 \text{ years}/N^2$. It can be considered 'deterministic' only in this limited sense.

All this applies when cabinet duration is measured according to the criteria devised by Dodd (1976). Deciding when a cabinet is terminated is a difficult matter, discussed in length by Laver (2003). As measured by the Dodd (1976) method, cabinet duration is designated as 'Average Cabinet Life' in Lijphart (1999: 132–3), who observes mean durations ranging from 1.3 years to 31 years. A cabinet is considered to last *as long as its partisan composition does*

not change. A more stringent measure ('Average Cabinet Life II' in Lijphart [1999: 132–3]) also considers a cabinet terminated upon an election, a change of prime minister, or a shift in cabinet type (oversized, minimal winning, or minority coalition). The observed mean Cabinet Life II ranges from 1.0 to 4.8 years. Frequency of elections is the cause of the severe upper cut-off. As a result, this more restrictive measure is less clearly correlated with the effective number of parties. It does not distinguish between the almost single-party Botswana, where the same party has been in power for 40 years, and multiparty Costa Rica, where the partisan composition of the cabinet has changed every 4.7 years, on average. This is why we focus on Life I, as defined by Dodd (1976). That also frees us from a constant concern of studies using Life II, namely whether a cabinet termination is 'natural' or rather due to elections (see Laver, 2003: 26, 31).

In turn, the effective number of legislative parties has been connected to the seat share of the largest party, which itself is connected to the product of district magnitude and assembly size (Taagepera, 2001). The equations are the following:⁶

$$N = 1/s_1^{3/2}$$

$$s_1 = 1/(MS)^{1/8},$$

so that

$$N = (MS)^{3/16}$$

In the case of multi-seat districts, the model presumes a PR allocation rule rather than plurality (which is rare in stable democracies).

The present study uses a slightly corrected relationship for N in terms of s_1 (see Taagepera, 2007: 160–2):

$$N = 1/s_1^{4/3},$$

so that

$$N = (MS)^{1/6}$$

It is also the first to observe that, in conjunction with $C = 42\text{yrs}/N^2$, these equations lead to:

$$C = 42\text{yrs}(s_1^{8/3})$$

and

$$C = 42\text{yrs}/(MS)^{1/3}$$

Once more, this is not a rigidly deterministic prediction but rather means that the actual duration has an equal probability of being above or below $42\text{yrs}/(MS)^{1/3}$. While the previous equations connect mean cabinet duration to other variables that cannot be stipulated legislatively, the last equation connects C to institutional factors. This is what renders it of interest for institutional engineering.

While $C = 42\text{yrs}/N^2$ is existing knowledge and $N = (MS)^{1/3}$ involves only a minor correction, connecting them is novel, even though it may look obvious in retrospect. And only this connection enables us to test the functional relationship between mean cabinet duration and factors amenable to institutional engineering.

We can now test the prediction for mean cabinet duration at three separate levels of the presumed causal chain $MS \rightarrow s_1 \rightarrow N \rightarrow C$. In this presumed chain, the mean cabinet duration is directly connected to the effective number of parties. This part – and only this part – has been tested previously. Cabinet duration is connected to the largest seat share only indirectly, through N . Hence we would expect s_1 to have less success than N in predicting C . Cabinet duration is connected even more indirectly to district magnitude and assembly size, through N and s_1 . Hence we would expect the product MS to be even less precise than s_1 in predicting C . But this is the relationship that matters for institutional engineering.

For individual countries, the impact of MS might be swamped by the impact of political culture, path-dependent developments, and institutional factors other than M and S , all specific to the given country. The worldwide median (i.e. the median of a large number of countries) can fit the model to the extent that many country-specific factors other than M and S may cancel out. Whether it does, remains to be discovered. If this is the case, then we can factor out the impact of district magnitude and assembly size when studying the impact of other factors on the mean cabinet duration. We would have narrowed down the problem for the study of cultural determinants and political processes.

At what stages do political mechanisms and processes enter in determining the mean duration of cabinets? Along with many other factors, they enter in setting assembly size and electoral rules in the first place. If these rules allocate all the seats within districts of roughly equal magnitude, then district magnitude and assembly size set restrictions on the size of the largest party and the effective number of parties, and the most probable outcomes can be calculated. This is a mechanical consequence of the political acts of choosing assembly size and district magnitude. Politics (and other factors) may enter to modify the probabilistic outcome. Even if it does not do so for the worldwide average, it may enter for individual countries.

The number of parties, in turn, has mechanical consequences for the number of potential conflict channels, which affect cabinet duration. However, some political cultures are more adept than others at managing conflict, and this may be connected to features like corruption. Thus, compared to the observed world median, some countries have double the duration of cabinets and some others have only one-half, at the same number of parties. And the number of parties itself, while depending on assembly size and district magnitude, may also be modified by political culture.⁷

The Main Prediction and Result

Rather than keep the reader in suspense, we will leave discussion of data and analysis of intervening stages to the next section, and immediately give the main result. The total range of mean cabinet durations for countries observed extends from 1.3 to 40 years – a ratio of 1 to 30. Thus, without any model, one can already say that all countries have a mean cabinet duration of 7.2 years, within a factor of 5.5 (which means multiplying or dividing by 5.5). Introducing the effective number of legislative parties (Taagepera and Shugart, 1989) narrows this range down to the point of predicting the mean cabinet duration in a stable democracy within a factor of 2 (rather than 5.5). The more distant connection of mean duration to the product of average magnitude and number of assembly seats (MS) cannot be expected to do any better than this direct connection.

Hence our prediction is that mean cabinet durations in most but not all countries will be within a factor of 2 (i.e. multiplying or dividing by at most 2) of the value given by the model $C = 42\text{yrs}/(MS)^{1/3}$. Taking the logarithms turns this nonlinear equation into a linear one. When C is in years, the expected zone is:

$$\log C = \log 42 - (1/3)\log(MS) \pm \log 2,$$

and the data can be properly subjected to linear regression analysis. Figure 1 expresses this expectation visually. Here the mean duration is graphed against the product MS , both on logarithmic scales, so that the predicted zone becomes a zone along a straight line.

Without knowing anything about the data for M and S , the theoretical model allows us to predict that, for given product MS , average cabinet durations will yield data points that lie in the zone shown, corresponding to $42\text{yrs}/(MS)^{1/3}$ multiplied or divided by 2 – if the model holds. Figure 1 highlights the fact that this rather narrow zone is a *theoretical prediction that precedes any look at data*. Our experience is that without such highlighting, some readers may mistake the theoretical prediction for a postdiction based on analysis of the data themselves.

This prediction is eminently falsifiable, because it is much more specific than a vague ‘If MS is up, then C is down’ (meaning $dC/d(MS) < 0$). Even when this statement fits and a satisfactory correlation between C and MS is found, the data points may turn out to be located all above or all below the predicted zone, or the best-fit line may turn out to have a slope vastly different from the predicted $1/3$ (on the log scale). Conversely, if most of the actual points are within the predicted zone, then the prediction holds regardless of a low R-square.

We now superimpose data to the blank format of Figure 1. This is done in Figure 2. Out of the 25 countries analysed, 19 (i.e. 76 percent) lie within the predicted zone, while three are above and three are below this zone. Linear regression of logarithms corresponds to $C = 21.8(MS)^{0.233}$. This line has a

Figure 1. Theoretical prediction of the effect of product MS on mean cabinet duration (C)

Notes:

Solid line: $C = 42 \text{ years}/(MS)^{1/3}$

Dashed lines: one-half and double the expected value.

Figure 2. Product MS and mean cabinet duration C : regression between logarithms and theoretical prediction

Notes:

Thin straight line: best fit between logarithms.

Bold straight line: theoretically based prediction ($C = 42 \text{ years}/(MS)^{1/3}$).

Dashed lines: one-half and double the expected value.

shallower slope (0.233) than the predicted $1/3 = 0.333$. What matters for testing the prediction is that it crosses the theoretical line at the centre of the data cloud and, throughout the actual range of MS , stays within the predicted zone. For the best fit line, $R^2 = 0.30$. It drops to 0.24 for the predicted line.⁸

In the presumed causal chain $MS \rightarrow s_1 \rightarrow N \rightarrow C$, the mean cabinet duration is far removed from the product MS . Yet we still have 76 percent of the data points within a factor of 2 from the predicted value. To expect more precise agreement would require blind faith in institutional design, excluding any other cultural and political inputs. Recall that, in the absence of the model, we could only expect the mean durations to be within a factor of 5.5 from the overall median of 7.2 years. We have narrowed it down to within a factor of 2, for most countries. Thus, the model represents a marked advance in precision of prediction.

It is now time to discuss data selection and, most important, examine the intermediary stages in the causal chain. This enables us to locate the links which most contribute to scatter and deviation from the theoretical model.

Detailed Testing of the Model

Lijphart (1999: 76–7, 132–3) tabulates the mean values of the effective number of legislative parties and cabinet duration for all 36 countries that by 1996 had been democratic for more than 20 years. We accepted his values of average N , which refer to the period 1945–96 or a shorter period in the case of more recent democracies. We used the same data that went into Cabinet Life I in Lijphart (1999) – graciously supplied by Lijphart. We introduced, however, an adjustment that leads to slightly longer mean durations.⁹ These relatively minor differences are readily visible when comparing C in our Table 1 and Cabinet Life I in Lijphart (1999: 132–3). We excluded Switzerland, which stands more than 2 standard deviations apart from the other countries.¹⁰ This leaves us with 35 countries.¹¹

For these 35 countries, we obtained the seat shares of the largest party in any given election, mainly from Mackie and Rose (1991, 1997), Nohlen (1993) and Nohlen et al. (2001).¹² From this, we calculated the mean largest share. We further determined the arithmetic mean S and M in those 25 cases where all seats are allocated in districts and the legal threshold is 1 percent at most.¹³ In the other 10 countries seat allocation either continues on a supra-district level or is restricted by a legal threshold larger than 1 percent, so that a meaningful effective district magnitude is hard to define.¹⁴

Table 1 first lists the 35 countries, the time periods used and the seat allocation rule used. For the 25 countries with definable district magnitudes, it further lists the mean district magnitudes and assembly sizes during these periods, and the products MS . These countries are listed in the order of increasing values of MS (which should correspond to decreasing mean cabinet duration, according to the model). Next, the largest seat shares, the

Table 1. Seat allocation rule, mean district magnitude (M), assembly size (S), product (MS), largest seat share (s_1), effective number of legislative parties (N) and mean cabinet duration (C)

	<i>Rule^a</i>	<i>M</i>	<i>S</i>	<i>MS^b</i>	<i>s₁</i>	<i>N</i>	<i>C (years)</i>
Barbados 1966–94	P	1	26	26	0.7	1.76	9.50
Trinidad 1961–2001	P	1	36	36	0.746	1.82	10.00
Botswana 1965–2004	P	1	37	37	0.749	1.35	39.60+
Bahamas 1972–2002	P	1	42	42	0.732	1.68	14.90
Jamaica 1962–89	P	1	55	55	0.755	1.62	9.20
Mauritius 1976–97	P	1	68	68	0.624	2.71	2.10
New Zealand 1946–96	P	1	85	85	0.569	1.96	6.30
Papua-NG 1977–97	P	1	108	108	0.397	5.98	1.65
Australia 1946–96	AV	1	128	128	0.507	2.22	9.90
Canada 1945–93	P	1	270	270	0.555	2.37	8.00
United States 1947–2001	P	1	435	435	0.619	2.4	7.70
France 1959–2002	TR ^c	1	508	508	0.444	3.43	3.10
India 1977–96	P	1	542	542	0.55	4.11	2.40
United Kingdom 1945–97	P	1	635	635	0.534	2.11	8.60
<i>Median for 14 M = 1 systems</i>							8.30
Malta 1966–87	STV	5	59	294	0.529	1.99	10.60
Costa Rica 1953–98	L	7.8	55	426	0.524	2.41	4.90
Ireland 1948–97	STV	3.5	154	538	0.482	2.84	3.80
Luxembourg 1945–99	L	14.2	57 ^d	809	0.411	3.36	6.00
Norway 1945–97	L	7.7	154	1190	0.466	3.35	4.30
Japan 1946–96	SNTV	4	486	1940	0.54	3.71	3.90
Spain 1977–2004	L	6.7	350	2330	0.501	2.76	9.00
Portugal 1976–2002	L	11.3	249	2810	0.43	3.33	3.20
Finland 1945–2003	L	14	200	2940	0.268	5.03	1.50
Israel 1949–96	L	120	120	14,400	0.379	4.55	1.75
Netherlands 1946–2002	L	140	140 ^e	20,000	0.344	4.65	3.30
<i>Median for 11 M > 1 systems</i>							3.90
Greece 1974–2004					0.556	2.2	4.90
Colombia 1958–96 ^f					0.526	3.32	4.70
Venezuela 1959–99					0.461	3.38	3.10
Austria 1945–2000					0.494	2.48	9.00
Sweden 1948–94					0.465	3.33	5.10
Italy 1946–94					0.412	4.91	1.33
Germany 1949–98					0.418 ^g	2.93	3.80
Denmark 1945–2001					0.368	4.51	2.80
Iceland 1946–95					0.343	3.72	2.90
Belgium 1946–2003					0.342	4.32	2.40
<i>Median for 10 systems with M undefined</i>							3.45

a: Seat allocation rules: AV = alternative vote; L = list PR; P = plurality; SNTV = single non-transferable vote; STV = single transferable vote; TR = two rounds majority.

b: When M and/or S vary, mean MS may slightly differ from $(\text{mean } M)(\text{mean } S)$ and mean $[s_1(MS)^{1/8}]$ may slightly differ from $(\text{mean } s_1)(\text{mean } MS)^{1/8}$.

c: France: list PR in 1986.

d: Luxembourg: S ranges from 25 (partial elections) to 64.

e: Netherlands: $S = 100$ in 1946–52; $S = 150$ in 1956–94.

f: Columbia: including two elections boycotted by PL, resulting in $s_1 = 1.00$.

g: Germany: CSU not included in CDU.

effective numbers of legislative parties and the mean cabinet durations are shown for all 35 countries. Here the 10 countries with undefined M are listed in the order of decreasing s_1 (which should correspond to decreasing mean cabinet duration, according to the model).

It follows from the model $C = k/N^2$ that $k = CN^2$. The value of the constant, $k = 42$ years, was determined empirically by feeding into this form the actual values of C and N in Table 1. Once we proceed beyond N , to the largest seat share and the product MS , $k = 42$ years is already part of the predictive model, because it is not affected by the actual values of s_1 and MS .

Table 2 repeats the actual values of mean cabinet duration (C) and also shows the values predicted on the basis of the various stages of the model, $C_1 = 42\text{yrs}/N^2$, $C_2 = 42\text{yrs}(s_1)^{8/3}$ and (when possible) $C_3 = 42\text{yrs}/(MS)^{1/3}$. The ratios of these predicted values to the actual values are also shown: $C/C_1 = CN^2/42\text{yrs}$, $C/C_2 = C/42\text{yrs}(s_1)^{18/3}$, and $C/C_3 = C(MS)^{1/3}/42\text{ yrs}$. These ratios should be close to 1, if the models hold.

Consider first the *predictions based on the effective number of legislative parties*, $C_1 = 42\text{yrs}/N^2$. The overall median ratio C/C_1 (0.97) is close to 1.00 by definition, given the way $k = 42$ years was chosen.¹⁵ For individual countries, the ratio ranges from 0.37 (Mauritius) to 1.72 (Botswana). The median ratio is slightly lower (0.94) for the 14 SSD systems, exactly 1.00 for the 11 PR systems with well-defined district magnitudes, and very slightly higher (1.01) for the 10 systems with indefinable M . The ratios C/C_1 indicate that $C_1 = 42\text{yrs}/N^2$ predicts the mean cabinet duration in individual countries essentially within a factor of 2.

Figure 3 shows the mean duration graphed against the mean effective number of parties, both on logarithmic scales. It highlights the exceptional case of Mauritius, the only country that deviates by more than a factor of 2. Linear regression of logarithms corresponds to $C = 31.3\text{yrs}/N^{1.757}$, with $R^2 = 0.79$. The theoretically predicted $C_1 = 42\text{yrs}/N^2$ has $R^2 = 0.77$ – only slightly less than the best fit. The two lines are extremely close to each other throughout the range of occurrence of effective numbers of parties and cross at the centre of the data cloud. This part – and only this part – of our study re-checks existing work. What follows is new.

Next, consider the *predictions based on the largest seat share*, $C_2 = 42\text{yrs}(s_1)^{8/3}$. As expected, the results are more scattered. For individual countries, the ratio C/C_2 ranges from 0.18 (Mauritius) to 2.04 (Botswana). The overall median ratio is 0.72, that is 28 percent below the expected 1.00. The median ratio is lowest for the SSD systems (0.65), higher for the PR systems with well-defined magnitude (0.78) and highest (0.93) for the systems with indefinable M . The ratios C/C_2 show that $C_2 = 42\text{yrs}(s_1)^{8/3}$ still predicts the mean cabinet duration in 28 out of the 35 individual countries (i.e. 80 percent) within a factor of 2.

Figure 4 shows mean duration graphed against the mean seat share of the largest party, both on logarithmic scales. Only one country lies above the predicted zone, while as many as six are below. Linear regression of

Table 2. Predicted values of mean cabinet duration (C), based on effective number of legislative parties (N), on largest party seat share (s_1) and on the product of district magnitude (M) and assembly size (S). Here, '42' is abbreviated notation for '42 years'

	<i>Actual C</i> (years)	$C_1 =$ 42/ N^2	C/C_1	$C_2 =$ 42 $s_1^{8/3}$	C/C_2	$C_3 =$ 42/(MS) ^{1/3}	C/C_3
Barbados	9.5	13.56	0.70	16.22	0.59	14.18	0.67
Trinidad	10	12.68	0.79	19.23	0.52	12.72	0.79
Botswana	39.6+	23.05	1.72+	19.43	2.04+	12.60	3.14
Bahamas	14.9	14.88	1.00	18.28	0.82	12.08	1.23
Jamaica	9.2	16.00	0.57	19.85	0.46	11.04	0.83
Mauritius	2.1	5.72	0.37	11.94	0.18	10.29	0.20
New Zealand	6.3	10.93	0.58	9.34	0.67	9.55	0.66
Papua-NG	1.65	1.17	1.40	3.58	0.46	8.82	0.19
Australia	9.9	8.52	1.16	6.86	1.44	8.33	1.19
Canada	8	7.48	1.07	8.74	0.92	6.50	1.23
United States	7.7	7.29	1.06	11.69	0.66	5.54	1.39
France	3.1	3.57	0.87	4.82	0.64	5.26	0.59
India	2.4	2.49	0.97	8.53	0.28	5.15	0.47
United Kingdom	8.6	9.43	0.91	7.88	1.09	4.89	1.76
<i>Median for 14 M = 1 systems</i>	8.3	8.98	0.94	10.51	0.65	9.19	0.81
Malta	10.6	10.61	1.00	7.69	1.38	6.32	1.68
Costa Rica	4.9	7.23	0.68	7.50	0.65	5.58	0.88
Ireland	3.8	5.21	0.73	6.00	0.63	5.16	0.74
Luxembourg	6	3.72	1.61	3.92	1.53	4.51	1.33
Norway	4.3	3.74	1.15	5.48	0.78	3.96	1.08
Japan	3.9	3.05	1.28	8.12	0.48	3.37	1.16
Spain	9	5.51	1.63	6.65	1.35	3.17	2.84
Portugal	3.2	3.79	0.84	4.42	0.72	2.98	1.08
Finland	1.5	1.66	0.90	1.25	1.20	2.93	0.51
Israel	1.75	2.03	0.86	3.16	0.55	1.73	1.01
Netherlands	3.3	1.94	1.70	2.44	1.35	1.55	2.13
<i>Median for 11 M > 1 systems</i>	3.9	3.74	1.00	5.48	0.78	3.37	1.08
Greece	4.9	8.68	0.56	8.78	0.56		
Colombia	4.7	3.81	1.23	7.57	0.62		
Venezuela	3.1	3.68	0.84	5.33	0.58		
Austria	9	6.83	1.32	6.41	1.41		
Sweden	5.1	3.79	1.35	5.45	0.94		
Italy	1.33	1.74	0.76	3.95	0.34		
Germany	3.8	4.89	0.78	4.10	0.93		
Denmark	2.8	2.06	1.36	2.92	0.96		
Iceland	2.9	3.04	0.96	2.42	1.20		
Belgium	2.4	2.25	1.07	2.40	1.00		
<i>Median for 10 systems with M undefined</i>	3.45	3.73	1.01	4.71	0.93		
<i>Overall medians</i>	4.70	4.89	0.97	6.65	0.72	5.54	1.08

Figure 3. Effective number of parties (N) and mean cabinet duration (C): regression between logarithms and theoretical prediction

Notes:

Thin solid line: best fit between logarithms.

Bold solid line: theoretically based prediction ($C = 42 \text{ years}/N^2$).

Dashed lines: one-half and double the expected value.

logarithms corresponds to $C = 21.9\text{yrs}(s_1)^{2.14}$, with $R^2 = 0.53$. The theoretically predicted $C_2 = 42\text{yrs}(s_1)^{8/3}$ has an appreciably lower $R^2 = 0.35$. Still, the best-fit line is contained within the predicted zone throughout the actual range of s_1 .

Finally, consider the *predictions based on the combination of assembly size and district magnitude*, $C_3 = 42\text{yrs}/(MS)^{1/3}$. In contrast to the previous two predictions, this one is a purely institutional prediction. The number of cases is reduced to 25, because in 10 countries district magnitudes cannot be defined. We are now several logical steps removed from the effective number of parties and can expect appreciable scatter. Indeed, for individual countries, the ratio C/C_3 ranges from 0.19 (Papua-NG) to 2.84 (Spain). The overall median ratio is 1.08, that is, only 8 percent above the expected 1.00. The median ratio for the SSD systems (0.81) is below the expectation, while it is higher (1.08) for the PR systems with well-defined district magnitudes. The ratios C/C_3 show that $C_3 = 42\text{yrs}/(MS)^{1/3}$ still predicts the mean cabinet duration within a factor of 2 in 19 out of the 25 countries. The corresponding graph has already been given (Figure 2) and discussed earlier.

It remains to compare the degrees of agreement with the model at the three stages of the presumed causal chain $MS \rightarrow s_1 \rightarrow N \rightarrow C$. Table 3 shows some comparisons. As expected, the degree of agreement decreases at each

Figure 4. Largest party's seat share (s_1) and mean cabinet duration (C): regression between logarithms and theoretical prediction

Notes:

Thin straight line: best fit between logarithms.

Bold straight line: theoretically based prediction ($C = 42 \text{ years}(s_1^{8/3})$).

Dashed lines: one-half and double the expected value.

stage. Each stage compounds random scatter and hence reduces correlation with the best-fit line (on a logarithmic scale). The shift from N to s_1 increases the scatter more than the shift from s_1 to MS . This contrast is even more marked for the theoretically predicted line, where correlation decreases at a steeper rate.

The percentage of points within a factor of 2 of the prediction also decreases steeply as we shift from N to s_1 and only mildly as we shift from s_1 to MS .

Table 3. Degrees of agreement with the model, at various stages of the causal chain $MS \rightarrow s_1 \rightarrow N \rightarrow C$

Stage	$N \rightarrow C$	$s_1 \rightarrow C$	$MS \rightarrow C$
R^2 for best-fit line	0.79	0.53	0.30
R^2 for predicted line	0.77	0.35	0.24
Difference	0.02	0.18	0.06
Percentage of points within a factor of 2 of the prediction	97%	80%	76%
Expected slope	-2.00	+2.67	-0.333
Best-fit slope	-1.76	+2.14	-0.233
Difference	12%	20%	30%

The absolute value of the slope of the best-fit line falls short of the expected, and it does so increasingly at each further stage. This is the pattern expected when each stage involves further accumulation of random scatter.

Figure 5 shows the frequency distributions of the *ratios* $C/C_1 = CN^2/42$, $C/C_2 = C/42s_1^{8/3}$ and $C/C_3 = C(MS)^{1/3}/42$, using logarithmic intervals. As one successively uses effective number of parties, the largest seat share, and the product MS as predictors of mean cabinet duration, the distributions widen, as expected. According to the model, the distributions should be centred at 1.00. This is largely the case for C/C_1 and C/C_3 . Surprisingly, a shift occurs for the intermediary stage C/C_2 – here cabinet durations tend to fall short of the expected. Even here, $\log(C/C_2)$ has a mean of -0.12 (instead of the expected 0), which is small compared to the standard deviation (0.23). Once this mean shifts away from 1.00, one would expect the mean of $\log(C/C_3)$ to maintain this shift. It is puzzling (though pleasant!) that the mean C/C_3 actually returns to 1.00.

How robust are these results against omission of some countries? The median values of the ratios C/C_1 , C/C_2 and C/C_3 are little affected by the removal of any single country. Suppose an extreme case, such as Botswana, is removed. The proportion of the countries within a factor of 2 of the expectation would actually increase slightly. The best-fit lines in Figures 2 to 4 would be tilted away from the predicted lines, but they would still be largely located within the predicted zone.

Figure 5. Frequency distributions of the ratios of actual cabinet duration (C) to values predicted by effective number of parties ($C_1 = 42/N^2$), largest seat share ($C_2 = 42s_1^{8/3}$) and the product of district magnitude and assembly size ($C_3 = 42/(MS)^{1/3}$), respectively

Implications for Institutional Engineering

For the purposes of institutional engineering, *the model predicts that doubling the district magnitude would reduce the mean cabinet duration by about 20 percent*, if all other factors remain the same. Indeed, if we replace M in $C = k/(MS)^{1/3}$ by $2M$, the ratio of the new duration to the previous one is $1/2^{1/3} = 0.79$, meaning a reduction by 21 percent. Note that this result is independent of the time constant k . Its value in a particular country may differ from the worldwide average, due to political culture, but as long as it can be presumed to remain constant in the given country, it does not affect the outcome. Only for new democracies would we have to depend on the worldwide average value of $k = 42$ years, with an obviously widened range of error.

Now consider a more drastic change. Suppose a country changes its mean district magnitude or assembly size or both, so that the product MS increases more than 100-fold. This was the case in 1996 for New Zealand. It shifted from single-seat districts ($M = 1$) and a mean assembly size (1946–96) of 85 to nationwide PR for 120 seats. How much change in cabinet duration could it expect? The change in MS was from 85 to 14,400, with the ratio of 169, and $(169)^{1/3} = 5.5$. Thus, if no legal threshold had been introduced, the mean duration of cabinets could be expected to fall from the previous 6.3 years to $6.3/5.5 = 1.1$ years.

However, New Zealand *did* introduce a 5 percent legal threshold. According to the approximate conversion formula $T = 75\%/(M + 1)$ reported in Lijphart (1999: 153), this would correspond to an effective magnitude of 14 rather than 120. If so, then the change in cabinet duration would be $(120 \times 14/85)^{1/3} = 2.7$, leading to an expected mean duration of $6.3/2.7 = 2.3$ years under the new set-up.

The actual mean duration of cabinets between 1996 and 2002 decreased even more drastically, from 6.3 years to 1.4 years – a striking example of institutional impact. However, future cabinets may well last somewhat longer, once voters and politicians learn to handle the altered institutional framework. Thus, a mean duration of about 2.3 years remains our prediction for New Zealand in the long run, provided that the post-1996 electoral system survives. It would take at least 20 years past 1996 to establish a meaningful empirical average and check the prediction. We shall see.

Conclusions

The study of cabinet durations involves two aspects: the central tendency for a country with stable institutions, and the dispersion of individual cabinet durations around this central tendency. Both are important but require different approaches. The present study focuses on the first issue: the mean durations (as operationalized by Dodd [1976]) over a long time span. It extends our knowledge in four respects.

First, at the level of theory, it joins two existing models so as to connect the mean cabinet duration, for the first time, to two purely institutional features – district magnitude and assembly size. *Second*, at the empirical level, this quantitatively predictive model is tested and found to predict the mean duration mostly within a factor of 2. Once political circumstances and processes have determined assembly size and district magnitude, the journey towards the mean cabinet duration is pretty much set on autopilot. *Third*, by controlling for two major factors, this study supplies an improved starting point for analysing the impact of other political and cultural factors on the mean cabinet duration. *Fourth*, it contributes to institutional engineering by making it possible to estimate *to what degree* institutional changes might alter the mean duration of cabinets – not merely the *direction* of such impact.

It may be asked how this model compares with other models for mean cabinet duration. To the best of our knowledge, no other model has been offered. Yes, there are models for variation of individual cabinet lengths *within* a country. There are *directionally* predictive models of effects of various variables x on the mean cabinet durations, that is, for whether dC/dx should be positive or negative. But we are not aware of any other functional models, $C = f(x)$, that would *quantitatively* predict the value of C for given x .

Within the framework of the quantitatively predictive model $C = k/(MS)^{1/3}$, we have also tested the predictions for an intermediary stage – the seat share of the largest party. We also have retested the connection of mean duration to the effective number of parties, leading to a revised value of the average time constant involved, from $k = 33$ years to $k = 42$ years.

For three-quarters of the stable democracies tested, the two institutional factors enable us to predict the mean duration of cabinets within a factor of 2. The next stage would be to analyse the residuals, meaning the discrepancies between institutional model and actual mean durations, from a cultural-political viewpoint. Corruption, in particular, might account for some of these residuals. With such impacts added, the prediction of average cabinet durations in stable democracies is highly likely to improve.

Notes

- 1 Laver and Shepsle (1996) is a prominent example of the public choice approach to cabinet duration. The chapter on 'Party Systems and Cabinet Stability' (1996: 195–222) offers theory, simulations, and discussion of specific past cases. The 'two basic conclusions' are that certain bargaining constellations are 'substantially more stable' than certain others, and that 'the model can be used to understand why governments might change tack between elections' (Laver and Shepsle, 1996: 215). Illuminating arguments are made, but no quantitative predictions are offered about how much duration could be expected, under some specified conditions, with a 50–50 probability.

- 2 Grofman and van Roozendaal (1997) give an excellent overview of factors that have been discussed in cabinet durability and termination studies.
- 3 Causality may go in both directions. The number of parties at the time the electoral rules are chosen influences the choice (cf. Colomer, 2005). Later on, electoral rules affect the number of parties.
- 4 The size of assembly, in turn, depends heavily on the size of population represented. A direct impact of population on the number of parties, bypassing assembly size, is conceivable but remains to be demonstrated.
- 5 Surprisingly, we have encountered opinions that $C = k/N^2$ is not a truly theoretical equation, because the constant k is not theoretically defined but is induced from observations. In contrast, Sørensen (1998) and Hedström (2004) consider establishing such constants as an important part of testing the predictive model. Indeed, if it were not so, then even the law of gravitation, $F = GMm/r^2$, would not qualify as a theoretical equation, given that the numerical value of the universal constant of gravitation (G) is not theoretically derived. Like k in the Taagepera and Shugart (1989) model, G is induced from observation. Both have a metric; in particular, k comes in units of time – years or months. What makes $F = GMm/r^2$ and $C = k/N^2$ theoretical is that the *functional form* is logically deduced.
- 6 A detailed model is available on request from the authors.
- 7 The impact of cultural and political factors at various stages may well be interconnected. For example, polities adept at managing conflict even at a large number of parties may also be more likely to choose large district magnitudes in the first place, because both are aspects of a consensual political philosophy (Lijphart, 1999). The model presented expresses the mechanical consequences of some institutional decisions. It does not address the various other potential causal links.
- 8 Fitting to M and S separately could increase the fit, since we could juggle two exponents separately, as compared to a single exponent for the product MS . But this would be curve fitting without any theoretical basis. No theoretical model for the impact of M and S separately seems to exist.
- 9 The values in Lijphart (1999) are calculated as if all cabinets collapsed at the cut-off point of data collection (30 June 1996). In contrast, we used the actual durations of cabinets that started prior to 30 June 1996 and collapsed prior to 1 September 2004 (data from *Keesing's Record of World Events*). Hence our means are larger than Lijphart's. The gap is the most pronounced for the Bahamas, where Lijphart (1999) lists the mean of two prime ministerships – $(19.9 + 3.9)/2 = 11.9$ years – although the latter continued for 6.0 years beyond June 1996. We list $(19.9 + 9.9)/2 = 14.9$ years. In Botswana the same party was in power from the beginning of independence (1965) to 1996 and beyond, winning elections again in October 2004. Here we show the duration as 39+ years, although obviously the prime minister has changed during this long spell.
- 10 The inverse square law implies that the product N^2C is conserved. The actual distribution of N^2C around the mean of 42 years is roughly normal (cf. later Figure 5), except that Switzerland stands more than 2 standard deviations apart from the rest. Such a deviation indicates that it is somehow different from all the other countries, justifying its exclusion from the set. Switzerland is the only non-presidential country where the executive, once empowered by parliament, does not depend on legislative confidence (Lijphart, 1999: 119–20). Thus a key assumption of the inverse square model does not apply.

- 11 Mauritius is problematic. Jugnauth was prime minister for 13.6 years (1982–95), which would lead to a mean duration of 9.2 years, except that the party composition of his cabinet changed seven times in quite minor ways. Thus Dodd's (1976) counting rules force us to slice Jugnauth's tenure into eight technically separate cabinets, which leads to a mean duration of only 2.1 years. Although 9.2 years seems to express better the realities on the ground, we adhere to 2.1 years, because the rules should not be changed in the middle of the game. However, the cases of Mauritius and also Botswana (see note 9) suggest that some modification of Dodd's rules might be desirable. Maybe we should give partial credit to cabinets with changed party membership but the same prime minister, and subtract some credit from same-party cabinets when prime ministers change.
- 12 The start of the time period is the one used by Lijphart (1999). In line with observations in note 9, the end point corresponds to the end of the cabinet that was in power on 30 June 1996.
- 13 As convincingly documented by Monroe and Rose (2002), the political impact at the same mean magnitude can be significantly different when individual districts have grossly divergent magnitudes. If anything, non-optimal input data would be likely to worsen the fit to a predictive model, not improve it.
- 14 Is this number of countries sufficient to test the model? More cases would be nice, but what we have used is essentially the entire set of post-World War II stable democracies available, if we accept Lijphart's (1999) criteria. Since 1999, very few democracies could be added to the list, given that those in Central East Europe have lasted as yet less than 20 years.
- 15 Below, we use the median because it does not prejudice the shape of the distribution. Using the arithmetic mean presumes a normal distribution, or at least a symmetrical one. Using the geometric mean presumes a lognormal distribution or at least a 'log-symmetrical' one (where logarithms are distributed symmetrically). The median is neutral in this respect.

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