Predicting Mean Cabinet Duration on the Basis of Electoral System

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Abstract
We join 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: $C=42\ years/(MS)^{1/3}$. 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*.

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 our best 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 excessively 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 2-year cabinets to 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 do 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: 131-139) 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 have shorter durations than others, we might also understand better what the mean cabinet durations signifies and implies.

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]

We may conjecture that certain features affect the mean cabinet duration. The number of parties visibly matters: Coalition cabinets, prevalent in multiparty systems, tend to be more short-lived than one-party majority cabinets. This was extensively documented by Lijphart (1984: 124-126), using the effective number of legislative parties, $N=1/\Sigma s_i^2$, 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 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 multi-party systems.[2] 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

[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] 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.

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.[3]

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 (McGregor, 1993; Crease, 2004; Colomer, 2007). 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 (Sørensen, 1998; Hedström, 2004).

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. 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.

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

[3] 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.

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. (Laver, 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. (See Appendix for details.) 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.[4] Using Lijphart's (1984) data, Taagepera and Shugart (1989) found that C=400 months/N^2=33 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 years/N^2. What this model of claims is that the actual duration has an equal probability of being above or below 42 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 Michael Laver (2003). As measured by the Dodd (1976) method, cabinet duration is designated as 'Average cabinet life I' in Lijphart (1999: 132-133), 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: 132-133) 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 year to 4.8 years at most. Frequency of elections is the causes of the severe upper cutoff. 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 multi-party Costa Rica, where the partisan composition of the cabinet has changed every 4.7 years, on the average.

This is why we focus on Life I, as defined by Dodd (1976). Using Life I also frees us from a constant concern of many studies using Life II, namely whether a cabinet termination is 'natural' or rather due to previous or imminent 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

[4] 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 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.

assembly size (Taagepera, 2001). The equations are the following (see Appendix for details of the model):

$$N=1/s_1^{3/2}$$
$$s_1=1/(MS)^{1/8},$$

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 Appendix):

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

so that

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

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

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

and

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

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 $42yrs./(MS)^{1/3}$. While the previous equations connect mean cabinet duration to other variables that cannot be stipulated legislatively, the latter equation connects C to institutional factors. This is what renders it of interest for institutional engineering.

While $C=42yrs./N^2$ 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 --> s_1 --> N --> 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, is to be found. 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 settling 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.[5]

The Main Prediction and Result

Rather than keep the reader in suspense, we'll leave discussion of data and analysis of intervening stages to the next section, and we will 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.

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[5] 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.

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 $C=42yrs./(MS)^{1/3}$. 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.[6] 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.

FIGURE 1 HERE

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 $42yrs./(MS)^{1/3}$ 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 the latter 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 analyzed, 19 (i.e., 76 percent) lie within the predicted zone, while 3 are above and 3 are below this zone. Linear regression of logarithms corresponds to $C=21.8(MS)^{0.233}$. This line has a 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 center 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.[7]

FIGURE 2 HERE

In the presumed causal chain MS --> s_1 --> N --> 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

[6] Why focus on factor of 2, rather than 1.5, 3 or e=2.73? It's a simple way to make the degrees of deviation from the model visible. Since the deviations from $C=42yrs./N^2$ fit within this range, but barely, the wider deviations from $C=42yrs.(s_1^{8/3})$ and $C= 42yrs./(MS)^{1/3}$ will be readily visible in this format. On the logarithmic graph, the zone within a factor of 2 appears as the narrows zone at ±log2=±0.3 from the expected value.

[7] 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.

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: 76-77 and 132-133) 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 of 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) and were graciously supplied by Lijphart. We introduced, however, an adjustment that leads to slightly longer mean durations.[8] These relatively minor differences are readily visible when comparing C in our Table 1 and Cabinet Life I in Lijphart (1999: 132-133). We excluded Switzerland, which stands more than 2 standard deviations apart from the other countries.[9] This leaves us with 35 countries.[10]

For these 35 countries, we obtained the seat shares of the largest party in any given election, mainly from Mackie and Rose (1991 and 1997), Nohlen (1993) and Nohlen et

[8] The values in Lijphart (1999) are calculated as if all cabinets collapsed at the cutoff point of data collection (30 June 1996). Thus Netanyahu, who became prime minister of Israel on 29 May 1996, contributes a cabinet lasting only 33 days, much short of its actual duration. This procedure artificially reduces Israel's mean cabinet duration. 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 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 the prime minister obviously has changed during this long spell.

[9] 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: 119-120). Thus a key assumption of the inverse square model does not apply.

[10] Among the 35 countries, Mauritius is problematic. Jugnauth was prime minister for 13.6 years (1982-1995), which would lead to a mean duration of 9.6 years, except that the party composition of his cabinet changed 7 times in quite minor ways. Thus Dodd's (1976) counting rules force us to slice Jugnauth's tenure into 8 technically separate cabinets, which leads to a mean duration of only 2.1 years for Mauritius. 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 8) 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.

al. (2001).[11] 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 legal threshold is 1 percent at most.[12] 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.[13]

TABLE 1 HERE

Table 1 first lists the 35 countries and the time periods used, and the seat allocation rule used. For the 25 countries with definable district magnitude, 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 effective numbers of legislative parties, the mean inter-elections periods 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).[14]

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.[15] 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 HERE

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, $C_1=42yrs./N^2$, $C_2=42yrs.(s_1)^{8/3}$ and (when possible) $C_3=42yrs./(MS)^{1/3}$. The ratios of these predicted values to the actual are also shown: $C/C_1=CN^2/42yrs.$, $C/C_2=C/[42yrs.s_1^{8/3}]$, and $C/C_3=C(MS)^{1/3}/42\ yrs.$ These ratios should be close to 1, if the models hold.

Consider first the *predictions based on the effective number of legislative parties*, $C_1=42yrs./N^2$. The overall median ratio C/C_1 (0.97) is close to 1.00 by definition, given

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[11] The start of the time period is the one used by Lijphart (1999). In line with observations in Note 8, the end point corresponds to the end of the cabinet that was in power on 30 June 1996.

[12] 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.

[13] 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-WWII 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.

[14] During the periods considered, the values of all variables listed have varied. Using the mean values may blur the relationships and hence worsen the fit of the predictive model, but it cannot possibly improve it.

[15] The value of k should be such that the average value of the expression C/C_1 is 1, meaning that average CN^2/k=1. The various measures of central tendencies yield somewhat different values of k. The arithmetic mean yields k=41.1, the geometric mean yields k=42.5, and the median yields k=43.3 years. We choose the rounded-off k=42 years as a compromise among these.

the way k=42 years was chosen.[16] 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 magnitude, and very slightly higher (1.01) for the 10 systems with indefinable M. The ratios C/C_1 indicate that $C_1=42yrs./N^2$ predicts the mean cabinet duration in individual countries essentially within a factor of 2.

FIGURE 3 HERE

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 $C=31.3yrs./N^{1.757}$, with R^2=0.79. The theoretically predicted $C_1=42yrs./N^2$ 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 center of the data cloud. This part -- and only his part -- of our study re-checks existing work. What follows is new.

Next, consider the *predictions based on the largest seat share*, $C_2=42yrs.(s_1)^{8/3}$. 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, i.e., 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 $C_2=42yrs.(s_1)^{8/3}$ 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 HERE

Figure 4 shows mean duration graphed against the mean seat share of the largest party, both on logarithmic scales. Only 1 country lies above the predicted zone, while as many as 6 are below. Linear regression of logarithms corresponds to $C=21.9yrs.(s_1)^{2.14}$, with R^2=0.53. The theoretically predicted $C_2=42yrs.(s_1)^{8/3}$ 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,* $C_3=42yrs./(MS)^{1/3}$. In contrast to the previous two predictions, this one is a purely institutional prediction.[17] We are reduced to 25 cases, because in 10 countries district magnitude 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, i.e., only 8 percent above the expected 1.00. The median ratio for

[16] In the following analysis, we prefer to use the median because it does not prejudge 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.

[17] No claim of unidirectional causality is involved here. The effective number of parties or the seat share of the largest party at the moment of choosing the electoral rules may well affect the choice of district magnitude -- cf. Note 2. All we say is that M and S are institutionally stipulated, while N and s_1 are not.

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 $C_3=42yrs./(MS)^{0.333}$ 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.

TABLE 3 HERE

It remains to compare the degrees of agreement with the model at the three stages of the presumed causal chain MS --> s_1 --> N --> C. Table 3 shows some comparisons. As expected, the degree of agreement decreases at each stage. Each stage compounds random scatter and hence reduces correlation with the best-fit line (on 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. 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 centered 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.

FIGURE 5 HERE

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.

The Impact of Other Factors

From the viewpoint of the predictive models tested, the deviations from expectation are random noise. From a broader viewpoint, such deviations include the impact of various other institutional, political and cultural factors. The ratios C/C_1 etc. in Table 2 represent residuals unexplained by the expected effects of N, s_1 and MS, respectively. It is now time to consider briefly the possible factors that might explain part of these residuals.

The mean inter-elections period (shown in Table 1) is another institutional factor that might interact with mean cabinet duration in both causal directions. More frequent elections offer more opportunities for cabinet reorganization. Conversely, cabinet collapse sometimes triggers early elections. Either way, once the mean cabinet duration is controlled for the impact of N, s_1 or MS, the residuals (C/C_1, C/C_2, C/C_3) could be expected to have some correlation with the mean inter-elections period, which ranges from 2.0 to 7.0 years. However, no relationship is found (R^2<.03 with all residuals).

Indications are that cultural factors such as Perceived Corruption Index (Transparency International 2004) could explain an appreciable part of the residuals. Inglehart's (1997) survival/self-expression scores and GDP per capita are also candidates. Such variables may affect the mean cabinet duration directly or indirectly, by influencing the largest seat share or the effective number of parties. Thus the entrance points of such factors need to be clarified. This tasks remains to be completed. Such a project, however, could not be carried out without making use of the present results, because otherwise the cultural inputs would be submerged among the institutional. By showing not only how but also *why* institutional factors have a specific impact on the mean cabinet duration, the present study sorts out a major institutional impact and makes the study of other factors easier to tackle. The choice of appropriate indicators would be greatly helped, if we could develop a logical model of how cultural factors may influence cabinet duration.

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 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 1x85=85 to 120x120=14,400. The ratio is 14,400/85=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 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 $(120\text{x}14/85)^{1/3}$=2.7, leading to an expected mean duration of 6.3/2.7=2.3 years under the new setup.

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 toward 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 analyzing 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 our best 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, i.e., 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 party, 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 two. The next stage would be to analyze 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.

APPENDIX. Derivation of $C= k/(MS)^{1/3}$

This derivation joins the ones in Taagepera and Shugart (1989) and Taagepera (2001), but both parts are presented in a more streamlined way, and a significant correction is introduced.

Consider *the number of parties* (p') *that could win seats in a district* of magnitude M. As a minimum, 1 party could win all the seats. As a maximum, M parties could win one seat each. The actual number could be anything from 1 to M, if nothing else is known but M. Our ignorance may seem complete. Yet we do know something very important, namely the lower and higher limits of what is possible: the number of seat-winning parties cannot be smaller than 1 nor larger than M. If nothing else is known, the best guess for median p' is the one that equalizes the possible error upwards and downwards. This means that the factor by which the upper limit exceeds p' should equal the factor by which p' exceeds the lower limit: $M/p'=p'/1$. Hence $p'=M^{1/2}$, the geometric mean of the limits.

As an alternative approach, consider the number of *seats per party* (m'). This number, too, could range from 1 (when M parties win 1 seat each) to M (when one party wins all the seats). The previous reasoning leads to expect $m'=M^{1/2}$. Multiplying the expected number of parties by the expected number of seats per party yields $p'm'=M$, as it should. Thus the two approaches are congruent.

Such congruence is not to be taken for granted. Suppose someone argued that the likeliest number of parties is the *arithmetic* mean of 1 and M: $p'=(M+1)/2$. By the same reasoning, the likeliest number of seats per party would be $m'=(M+1)/2$. But then $p'm'=(M+1)^2/4$, which exceeds M whenever $M>1$.

As an illustrative example, the Netherlands 1918-1952 had 9 elections with the entire country as one district of M=100. The model predicts that $100^{1/2}$=10 parties would win seats. The actual range of the number of seat-winning parties was 8 to 17, with a geometric mean of 10.4, arithmetic mean 10.8, and median 10. In contrast, if arithmetic mean of the extremes were used, it would yield p'=50.5 parties and m'=50.5 seats per party, which would lead to $p'm'$=638 seats, way above the actual 100. Thus, the arithmetic mean would lead to conflicting results.

Next, consider *the number parties* (p) *that could win seats in an assembly* of S members elected in districts of magnitude M. Our best guess at the number of seat-winning parties in each district was $p'=M^{1/2}$. Nationwide, it is a likely lower limit, because different parties may win seats in different districts. Hence the expectation is $p>p'=M^{1/2}$. If the entire country were made a single district of magnitude S, we would expect $p=S^{1/2}$. This would be a likely higher limit on p. Over many occurrences in such a country, the outcomes are likely to spread out mainly between these values (although the absolute limits are 1 and S). If nothing else is known besides M and S, the best guess for median p is the one that equalizes the possible upward and downward errors between $M^{1/2}$ and $S^{1/2}$. Hence $p=(M^{1/2}S^{1/2})^{1/2}=(MS)^{1/4}$ would be expected.

As an illustrative example, Malta 1947-1955 had 5 elections with S=40 and M=5 in all districts. The model predicts that $(5x40)^{1/4}$=3.76 parties would win seats. The actual range of the number of seat-winning parties, nationwide, was 2 to 6, with a geometric mean of 3.73, arithmetic mean 4.0 and median 4.

Now consider the *number of seats* (S_1) *going to the largest* among these $p=(MS)^{1/4}$ parties. If all parties have equal shares, $S_1=S/p$. If all other parties have only one seat each, then $S_1=S-p+1$. This can be approximated as $S_1=S$ when S is sufficiently high. If nothing else is known, the best guess for median S_1 is the one that equalizes the possible upward and downward errors between the limits S/p and S. Hence $S_1=S/p^{1/2}$ is the best guess.[18] The *fractional seat share* (s_1) *of the largest party* is $s_1=S_1/S=1/p^{1/2}$. Since $p=(MS)^{1/4}$, the median largest share is $s_1=(MS)^{-1/8}$.

Taagepera (2001) saw *effective number of parties*, $N=1/\Sigma s_i^2$ as subject to the following limits. When all seat-winning parties have equal shares, $N=1/s_1$. When all other parties are infinite in number and infinitesimally small, then $N \rightarrow 1/s_1^2$. In the absence of any other information but s_1, the best guess is the one that equalizes the possible upward and downward errors between these limits: $N=1/s_1^{1.5}=(MS)^{3/16}$.

However, this estimate involves a mistake. The parties cannot be infinite in number, because the number of seat-winning parties at given largest seat share is limited by $s_1=1/p^{1/2}$ (as established above) to $p=1/s_1^2$ parties. The resulting calculations are complex and will be published separately. A simple approximation to the complex result replaces 3/16=0.1785 by 1/6=0.1667 in the estimate above. Hence the new formula is $N=(MS)^{1/6}$.

Finally, the mean *duration of cabinets* (C) is affected by frequency of conflicts. As a first approximation, we assume that this frequency is constant when the number and size of parties remains the same. This is in line with the constant hazard rate assumption made by King et al. (1990) and confirmed by Diermeier and Stevenson (1999). As the frequency (f) of conflicts doubles, cabinet duration is likely to be halved. Thus $C=k'/f$, where k' is a constant. The frequency of conflicts may be assumed to be proportional to the number of potential conflict channels (c) within the system: $f=k''c$, where k'' is a constant. With n equal-sized parties, the number of such channels among them would be $c=n(n-1)/2$. When an intra-party conflict channel per party is added, the number would be $c=n(n+1)/2$. The mean of these two estimates is $c=n^2/2$. For parties of varying sizes, the effective number of parties will be used: $c=N^2/2$. Combining these links results in $C=k/N^2$, where k is a constant, $k=2k'/k''$. The model predicts an inverse square relationship, leaving k to be determined empirically. Combining $C=k/N^2$ with $N=(MS)^{1/6}$ yields $C=k/(MS)^{1/3}$. Like C itself, k has the units of time. Empirically, k=42 years, so that $C=42yrs./(MS)^{1/3}$.

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[18] The overestimate introduced by the approximation $S_1=S-p+1\approx S$ does not exceed 5 percent as long as S remains above 100 when $M=S$ and above 10 when M=1.

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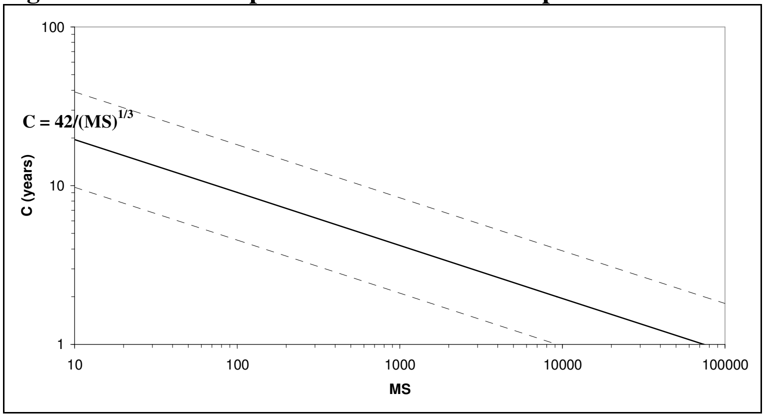
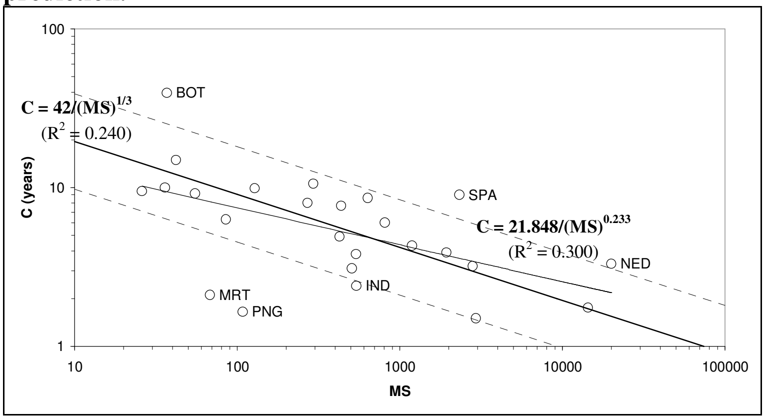
Figure 1. Theoretical prediction of the effect of product *MS* on mean cabinet duration (*C*).



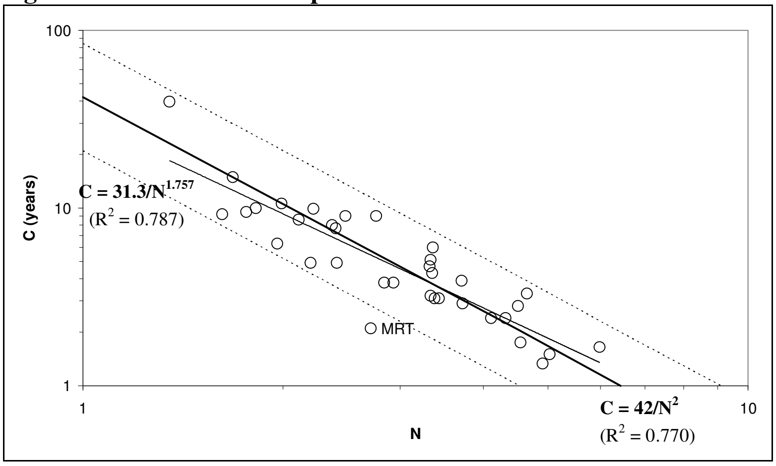
Notes:
Solid line: $C=42\ years/(MS)^{1/3}$
Dashed lines: one-half and double the expected value.

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



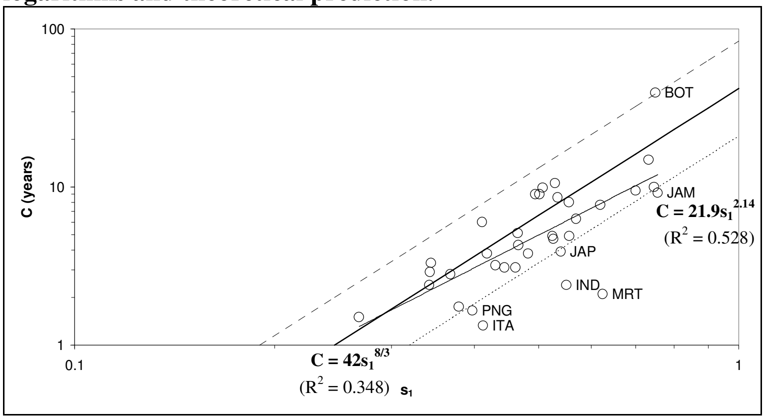
Notes:
Thin straight line: best fit between logarithms.
Bold straight line: theoretically based prediction [$C=42\ years/(MS)^{1/3}$].
Dashed lines: one-half and double the expected value.

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 [$C=42\ years/N^2$].
Dashed lines: one-half and double the expected value.

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 [$C=42\ years(s_1^{8/3})$].
Dashed lines: one-half and double the expected value.

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.

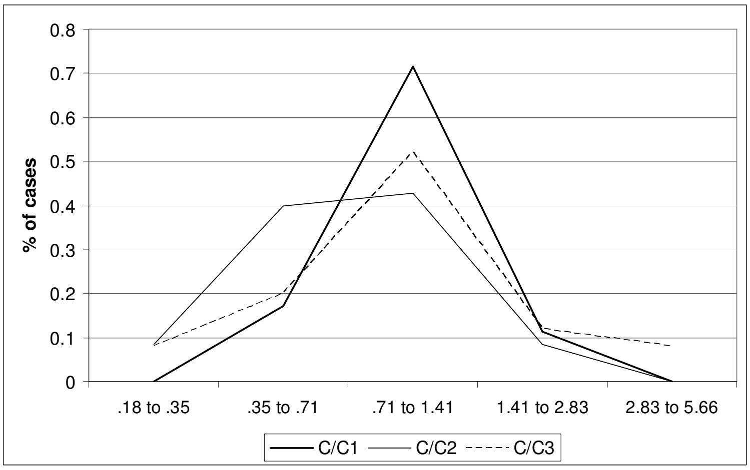


Table 1. Seat allocation rule, mean district magnitude (*M*), assembly size (*S*), product *MS*, largest seat share (s_1), effective number of legislative parties (*N*), mean inter-elections period (*I*), and mean cabinet duration (*C*).

	Rule[a]	M	S	MS[b]	s_1	N	I (years)	C (years)
Barbados 1966-94	P	1	26	26	0.7	1.76	4.7	9.50
Trinidad 1961-2001	P	1	36	36	0.746	1.82	4.9	10.00
Botswana 1965-2004	P	1	37	37	0.749	1.35	5.2	39.60+
Bahamas 1972-2002	P	1	42	42	0.732	1.68	5.0	14.90
Jamaica 1962-89	P	1	55	55	0.755	1.62	4.6	9.20
Mauritius 1976-97	P	1	68	68	0.624	2.71	3.8	2.10
New Zealand 1946-96	P	1	85	85	0.569	1.96	2.9	6.30
Papua-NG 1977-97	P	1	108	108	0.397	5.98	5.7	1.65
Australia 1946-96	STV	1	128	128	0.507	2.22	2.4	9.90
Canada 1945-93	P	1	270	270	0.555	2.37	3.2	8.00
United States 1947-2001	P	1	435	435	0.619	2.4	2.0	7.70
France 1959-2002	TR[c]	1	508	508	0.444	3.43	3.8	3.10
India 1977-96	P	1	542	542	0.55	4.11	3.8	2.40
United Kingdom 1945-97	P	1	635	635	0.534	2.11	3.1	8.60
Median for 14 M=1 systems								*8.30*
Malta 1966-87	STV	5	59	294	0.529	1.99	5.3	10.60
Costa Rica 1953-98	L	7.8	55	426	0.524	2.41	4.1	4.90
Ireland 1948-97	STV	3.5	154	538	0.482	2.84	3.1	3.80
Luxembourg 1945-99	L	14.2	57[d]	809	0.411	3.36	7.0	6.00
Norway 1945-97	L	7.7	154	1190	0.466	3.35	3.8	4.30
Japan 1946-96	SNTV	4	486	1940	0.54	3.71	2.5	3.90
Spain 1977-2004	L	6.7	350	2330	0.501	2.76	3.1	9.00
Portugal 1976-2002	L	11.3	249	2810	0.43	3.33	2.4	3.20
Finland 1945-2003	L	14	200	2940	0.268	5.03	3.6	1.50
Israel 1949-96	L	120	120	14400	0.379	4.55	3.6	1.75
Netherlands 1946-2002	L	140	140[e]	20000	0.344	4.65	3.4	3.30
Median for 11 M>1 systems								*3.90*
Greece 1974-2004					0.556	2.2	2.7	4.90
Colombia 1958-96[f]					0.526	3.32	2.9	4.70
Venezuela 1959-99					0.461	3.38	5.0	3.10
Austria 1945-2000					0.494	2.48	3.2	9.00
Sweden 1948-94					0.465	3.33	3.1	5.10
Italy 1946-94					0.412	4.91	4.2	1.33
Germany 1949-98					0.418[g]	2.93	3.7	3.80
Denmark 1945-2001					0.368	4.51	2.4	2.80
Iceland 1946-95					0.343	3.72	3.3	2.90
Belgium 1946-2003					0.342	4.32	3.1	2.40
Median for 10 systems with M undefined								*3.45*

a: Seat allocation rules: P = plurality; STV = single transferable vote; TR = two rounds majority; L = list PR; SNTV = single non-transferable vote.
b: When *M* and/or *S* vary, mean *MS* may slightly differ from (mean *M*)(mean *S*) and mean $[s_1(MS)^{1/8}]$ may slightly differ from (mean s_1)(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 2 elections boycotted by PL, resulting in s_1=1.00.
g: Germany: CSU not included in CDU.

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".

	Actual C (years)	C_1=42/N^2	C/C_1	C_2=42$s_1^{8/3}$	C/C_2	$C_3=\frac{42}{(MS)^{1/3}}$	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
Median for 14 M=1 systems	*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
Median for 11 M>1 systems	*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		
Median for 10 systems with M undefined	*3.45*	*3.73*	*1.01*	*4.71*	*0.93*		
Overall medians	*4.70*	*4.89*	*0.97*	*6.65*	*0.72*	*5.54*	*1.08*

Table 3. Degrees of agreement with the model, at various stages of the causal chain *MS* --> s_1 --> *N* --> *C*.

Stage	*N* --> *C*	s_1 --> *C*	*MS* --> *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 %**