King, Alt, Burns, and Laver (1990) proposed and estimated a unified model in which cabinet durations depended on seven explanatory variables reflecting features of the cabinets and the bargaining environments in which they formed, along with a stochastic component in which the risk of a cabinet falling was treated as a constant across its tenure. Two recent research reports take issue with one aspect of this model. Warwick and Easton replicate the earlier findings for explanatory variables but claim that the stochastic risk should be seen as *rising*, and at a rate which varies, across the life of a cabinet. Bienen and van de Walle, using data on the duration of leaders, allege that random risk is *falling*. We continue in our goal of unifying this literature by providing further estimates with both cabinet and leader duration data that confirm the original explanatory variables' effects, showing that leaders' durations are affected by many of the same factors that affect the durability of the cabinets they lead, demonstrating that cabinets have stochastic risk of ending that is indeed *constant* across the theoretically most interesting range of durations, and suggesting that stochastic risk for leaders in countries with cabinet government is, if not constant, more likely to rise than fall.

TRANSFERS OF GOVERNMENTAL POWER
The Meaning of Time Dependence

JAMES E. ALT
GARY KING
Harvard University

1. INTRODUCTION

The past few years have seen a surge of interest in the question of predicting and explaining transfers of governmental power. If one considers the question broadly to include any study of the frequency of government

turnovers or power transfers, whether by constitutional or nonconstitutional means, then serious arguments have appeared twice as formal "controversies" (Browne, Frendreis, & Gleiber, 1988; Jackman, O'Kane, Johnson, McGowan, & Slater, 1986; Strom, 1988) and informally in arguments among scholars in related articles (Bienen & van de Walle, 1989; Johnson, Slater, & McGowan, 1984; Strom, 1985). Other studies question whether rapid turnover might be the cause (as well as possibly the consequence) of poor economic performance, either in the form of government deficits (Roubini & Sachs, 1989) or low growth rates of national income (Londregan & Poole, 1990). Within this growing literature, disagreements abound over theory, measurement, and data.

Our present goal is similar to that of King, Alt, Burns, and Laver (1990). In that article, the coauthors unified many of the then existing models and methods in the literature, demonstrating that the important substantive theories in these literatures led to methodological disagreements that were more apparent than real. In this article, we focus on resolving some of the remaining disagreements over statistical specification, especially those involving substantive arguments about "hazard rates" and related parts of the stochastic specification of these "unified" models. We discuss the state of the scholarly consensus in this field in Section 2. The remaining disagreements outlined at the end of that section provide the structure for the remainder of this article.

2. STATE OF THE CONSENSUS

PREVIOUS WORK

King et al. (1990) introduced a unified statistical model that combined the previously distinct approaches of "attributes" theorists, who seek to explain durations as a fixed function of measured explanatory variables, and "events process" theorists, who model durations as a product of purely stochastic processes. The unified model includes a systematic component that lets the expected duration of a cabinet vary as an exponential-linear function of seven explanatory variables, and a stochastic component based on the exponential distribution. An implication of this distribution is that the hazard rate (the rate at which governments end at duration t, given that they have survived until t) is constant. We return to this point in Section 3.

The theoretical core of the King et al. (1990) analysis is that duration reflects complexity in the bargaining environment of, or the likelihood of challenges and availability of alternatives to, any one government after taking into account certain constitutional and electoral characteristics of the govern-

ment. Some of these variables characterize the situation in which a cabinet formed, whereas others describe the cabinet itself. Our specific variables explain cabinet durations with information available only at the time each government forms, thus providing *ex ante* predictions of the durability of new governments. We evaluated the adequacy of these variables by adding a set of dummy variables for the countries in the analysis. Because these indicator variables made no substantively significant contribution to explaining government durations over and above our seven substantive variables, we could exclude the indicator variables. We are therefore in the relatively unusual situation of being able to follow Przeworski and Teune's (1970) dictum to get rid of proper names, and we retain only our seven substantive variables in what follows.

Our seven substantive variables include the following:

1. Fractionalization (Rae, 1971), an index characterizing the number and size of parties in parliament such that higher values indicate dispersion into a larger number of relatively smaller blocs. This increases possible party combinations and alternative governments and thus higher values predict shorter expected durations.
2. Polarization (Powell, 1982), a measure of support for extremist parties (like fractionalization, it was averaged over decades within countries but ideally would be recalculated for each transition). It indicates willingness to create crisis through challenge from fringe parties, so more polarization should predict shorter expected durations.
3. Formation attempts, the number of attempts to form a government during the crisis. Taking several attempts to form means that there must have been (at least nearly) viable alternatives, shortening expected duration.
4. Investiture, the existence of a legal requirement for legislative investiture, a hurdle which diminished average duration by causing some governments to fail very quickly.
5. Numerical status, a dummy variable distinguishing between majority (coded 1) and minority (coded 0) governments, with majority cabinets less likely to be defeated in parliament.
6. Postelection formation, a dummy variable, which governments have longer before elections than those forming in midterm.
7. Caretaker governments, a small (5%) subset of governments specifically established only pending new elections.

POINTS OF AGREEMENT

A strong consensus now exists among several independent research teams publishing throughout this literature on many, although not all, details of this model. We have seen no theoretical objection to the unification of the attributes and events process models. There is a broad consensus about the

exponential-linear functional form, and even our specific list of variables and their coding is widely accepted. Moreover, estimates of our model have now been replicated by two other sets of researchers (Beck & Katz, 1992; Warwick & Easton, 1992), which by itself represents a significant scientific advance for this literature. Scholars also generally understand the problem that these data are often censored at or near the constitutional interelection period (CIEP).

POINTS OF DISAGREEMENT

The scholarly literature contains at least three remaining disagreements. First, scholars differ on the specific stochastic component and its corresponding hazard rate. Stochastic components and their hazard rates have potentially important implications for substantive conclusions in this area. Almost everyone seems to agree that some version of the exponential or compound-exponential distribution is appropriate. King et al. (1990) proposed an exponential distribution, which implied a constant hazard rate. Similar results were obtained by Londregan and Poole (1991). Warwick and Easton (1992) propose a distribution that implies a polynomial hazard, which they interpret as mostly increasing. Bienen and van de Walle (1991) estimate a model in different data that suggests a decreasing hazard rate. In Section 3, we demonstrate that the new evidence introduced by Warwick and Easton (1992) is largely consistent with the emerging consensus on cabinet durations having a conditionally constant hazard rate.

The second area of possible disagreement is the applicability of different models for cabinet and leader duration. Most of the literature focuses on cabinets, but Bienen and van de Walle (1991) introduce different models and methods to study leadership duration, as do Londregan and Poole (1991). We discuss the differences between these two areas and a possible consensus in Section 4.

Finally, the literature had begun to move from our original focus on ex ante prediction to other questions requiring time-varying explanatory variables. This has led to different models, new questions, and some contradictions with existing research, which we discuss in the final section.

3. DISAGREEMENTS OVER HAZARD RATES

What, in fact, does finding time dependence or a nonconstant hazard rate mean? Finding positive time dependence is like observing rust on a car: the older it gets, the more there is, and the likelier it is to fall apart. One could

think of leaders as rusting: indeed, Warwick and Easton's (1992) description of rising hazards echoes Mueller's (1970) old idea of the "coalition of minorities" in the American presidential approval literature—the longer you serve, the more people you offend. Even a politician who always sides with a (nonidentical) majority on each issue will eventually have offended a large majority of the public. Bienen and van de Walle's (1991) hazard rate is the opposite: the longer their leaders serve, the more immune to catastrophe they become. The theory here is that leaders gain expertise in problem solving and experienced leaders are better able to fend off attacks. By contrast, in King et al. (1990), the hazard rate is constant after taking into account the explanatory variables. This means that cabinets may rust or gain expertise over time, but these features of cabinets are assumed to be captured by our explanatory variables.

The hazard rate represents the rate at which cabinets end at some duration, given that they reach that duration. It is interpreted in an analogous fashion to the height of a continuous probability density function, so it is not quite a probability, but it is close. Mathematically, the hazard rate is a conditional probability density:

$$h(t) = \frac{f(t)}{1 - F(t)}$$

$$= \frac{f(t)}{Pr(T \geq t)}$$

(1)

which is the ratio of the density generating observed durations to the survivor function. King et al. (1990) focus on modeling the expected duration, which is

$$E(t) = \int_0^\infty tf(t)dt$$

$$= \int_0^\infty th(t)F(t)dt$$

(2)

Whether to study hazard rates unconditionally in the original data, or conditional on a set of explanatory variables, is a very important statistical question. In duration-type models, the analyst imposes a partition between the systematic and stochastic components by choosing explanatory variables. A hazard rate calculated from the original data will surely differ from that calculated after taking into account the effects of relevant explanatory variables.
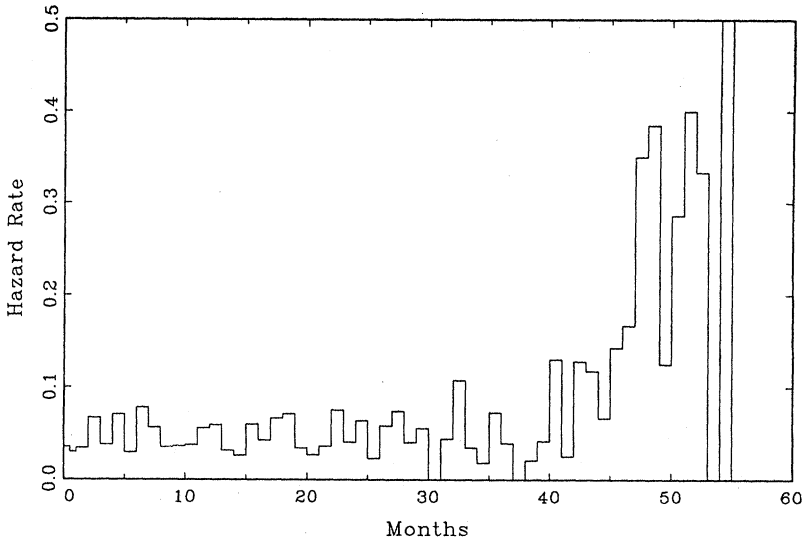
Figure 1.   **Cabinet duration without censoring corrected.**

A POLYNOMIAL HAZARD RATE?

Warwick and Easton (1992) use both methods of studying whether the hazard rate of a cabinet falling is constant. First, they estimate and plot the observed or unconditional hazard rate—that is, not controlling for the effects of any explanatory variables. They also calculate a hazard rate conditional on our seven explanatory variables, assuming a (cubic) polynomial hazard function. They find statistical evidence that the cabinets data have varying hazard rates. However, we are more interested in, and therefore will explore, the exact substantive consequences of these specific empirical results.

Figure 1 presents the hazard rates for all durations observed in the cabinet data. The hazard rate is clearly not constant (as the exponential model assumes it should be), but rather slopes up as durations become longer. This indicates that the risk of a government coming to an end does not have the Markov independence property that Browne, Frendreis, and Gleiber (1986) described. In fact, the probability of a government falling increases sharply with the passage of time.

However, many of this large group of cabinets falling toward the right end of Figure 1 did not break apart due to some combination of critical essentially
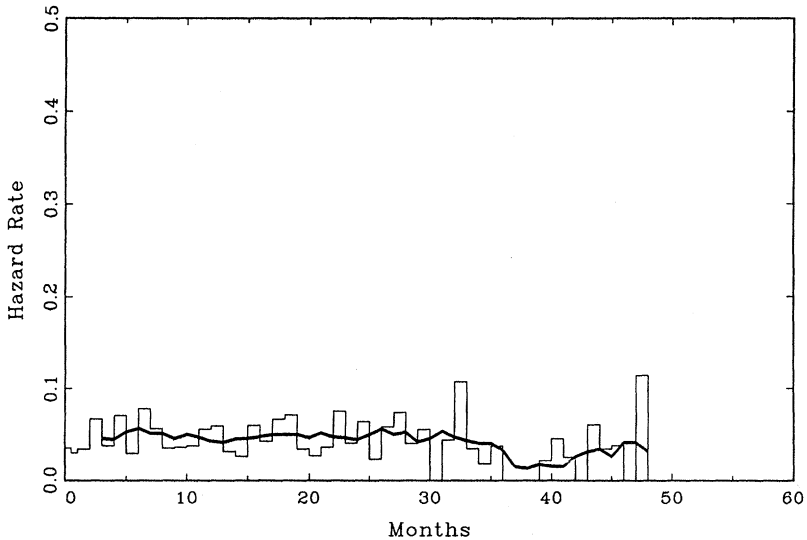
Figure 2.   **Cabinet duration with censoring corrected.**

random events, but rather are those that came near the CIEP maximum. Modeling this mechanism is a problem. King et al. (1990) treated the CIEP as having censored this second group of durations. Were it not near the CIEP, this latter group of governments would probably have lasted longer.

Figure 2 presents a similar analysis of the hazard rate for the censored cabinet durations. The solid line represents the proportion of governments that fell in that period (of those surviving to that month), but removes all of those governments that ended within a year of the CIEP from the calculation. The heavy line in this figure is an estimate of the unconditional hazard rate using a nonparametric smoothing procedure similar to that of Warwick and Easton.[1] As can be seen, this hazard rate is not constant: It increases, decreases, and then increases again. However, if substance is of concern here, one should focus on the scale of the vertical axis of this graph. This indicates that their estimated hazard rate only varies from 0.027 to 0.033, a very small change of about plus or minus 10% in the hazard rate around its mean value of .03. This small change is more visible in the vastly exploded scale in Warwick and Easton (1992, p. 136, Figure 3).

1. For convenience, we use a moving average rather than their Nelson-Aalen procedure; the two give similar results in this case.

To get a feel for the size of such narrow variability in the hazard rate, think of another rate, say an economic growth rate. Imagine that from period to period, mean economic growth was 3%, but that recessions featured growth of 2.7% whereas booms featured growth of 3.3%. We think any economist would regard such growth as effectively constant. In the same spirit, by any substantive measure, a duration that had a hazard rate averaging .03 and varying by a maximum of only ±0.003 is effectively constant. One might also consider how small this change is by comparing to Figure 1, which also portrays a clearly increasing hazard rate, the range of which is from just above 0 to about 0.5.

Warwick and Easton (1992) also reestimated our final model with the addition of a polynomial hazard rate function. They find that (with slight modifications to our 314 observations and slightly different assumptions about the effects of the CIEP) the parameters of their polynomial are statistically significant at conventional levels. They took this as evidence against our model's constant hazard rate, which of course it is—from a statistical perspective.[2] However, we can also interpret these results from a substantive perspective.

To do this, we present a plot of the time-dependence component of their estimated hazard rate in Figure 3. This was calculated directly from the estimates reported for variables representing powers of time in their Table 2 (Warwick & Easton, 1992, p. 138). The vertical axis of this figure is only proportional to the hazard rate because the actual hazard rate is also multiplied by the exponential-linear function of our seven explanatory variables.

Two features of Figure 3 are worthy of note. First, across most of the graph, the polynomial hazard rate is indeed constant. It is constant in exactly the area where most of the field's substantive interest lies, where durations of about 1 to 3 years are observed. There is no question that something unusual is happening near the CIEP, and we agree that there is no one perfect way to deal with this censoring problem. The period just following cabinet formation is similarly open to controversy. It may very well be that our explanatory variables are so successful in predicting short-duration cabinets that the residual stochastic risk of falling, systematic effects apart, is actually lower at very short durations. But their results show that the hazard rate from 12 to 36 months is very close to constant, thus confirming a significant part of our results.

Finally, it is worth noting that polynomials have a well-known tendency to estimate ends (in this case, of the hazard rate) poorly. This is because only

2. Estimates revealing positive time dependence of government duration in the case of Italy were also independently reported by Merlo (1991).
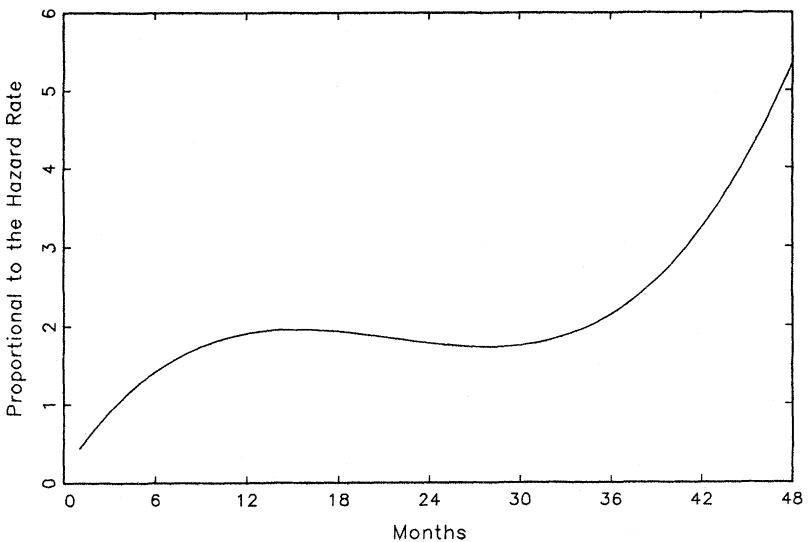
Figure 3.   **Cubic hazard rate.**

a few odd observations that happened to be at the end of a series (in this case, near the CIEP) will cause the polynomial to shoot upward or downward without limit. Polynomials were used more frequently earlier on in distributed lag models, but have been largely (although not completely) abandoned. Polynomial hazard rates are also not frequently used for this same statistical reason.

REANALYZING THE CABINET DATA

Given the statistical problem with polynomial hazard rate specifications, it pays to explore further the possibility of substantively important changing hazard rates. We therefore reestimate King et al.'s "final model" (1990, Model 2.4) with a variety of techniques allowing different sorts of time dependence in the hazard rates and describe any time dependence we find, as well as the behavior of the parameter estimates. King et al. (pp. 855-856) reported analyses of time dependence using the Pareto and gamma distributions as generalizations of the final exponential model. We expand these in this article to include the Gompertz, Weibull, and proportional hazards models. These results appear in Table 1.

The shape of the hazard function and all the claims about time independence and dependence are conditional on specifying and estimating the

Table 1
Survival Models of Cabinet Durations

Explanatory Variable	Model			
	1	2	3	4
Fractionalization	−.0012	−.0008	−.0011	−.0013
	(.0011)	(.0014)	(.0007)	(.0009)
Polarization	−.016	−.014	−.015	−.019
	(.008)	(.006)	(.005)	(.006)
Formation attempts	−.091	−.144	−.087	−.107
	(.058)	(.057)	(.039)	(.046)
Investiture	−.369	−.324	−.330	−.423
	(.165)	(.146)	(.107)	(.140)
Numerical status	0.515	0.566	0.464	0.571
	(.160)	(.154)	(.107)	(.132)
Postelection	0.723	0.800	0.664	0.818
	(.173)	(.181)	(.113)	(.141)
Caretaker	−1.30	−1.84	−1.32	−1.53
	(.353)	(.161)	(.217)	(.280)
Constant	3.724	3.570	3.696	
	(.742)	(1.08)	(.514)	
σ		34.9	0.770	
		(3.55)	(.043)	
Log likelihood	−425	−1171	−413	−1299
Model	Exponential	Gompertz	Weibull	Prop. Hazards

Note. Observations are 314 cabinet durations in months in 15 countries, 1946-1987. Standard errors are in parentheses. The Gompertz model does not treat any observations as censored by the constitutional interelection period. The log likelihoods from columns 2 and 4 are not comparable to those in other columns. Parameter estimates are given with signs to predict duration.

systematic part of the model. Table 1 gives a series of relevant results. Model 1 replicates the exponential Model 2.4 in King et al. (1990) with minor data corrections. Models 2 (Gompertz), 3 (Weibull), and 4 (proportional hazards) introduce varying forms of time dependence. The Gompertz model makes the log of the hazard rate a linear function of time; the Weibull model makes it a logarithmic function of time. The proportional hazards model is nonparametric so the hazard rate is only specified up to a proportionality factor that is an unknown function of time; it does however make the nontrivial assumption that hazards are proportional. Thus when Models 2 and 3 are estimated, the fitted hazard rates are obtained as follows:

$$\text{Model 2 (Gompertz) } h(t) = pe^{\lambda t}$$

$$\text{Model 3 (Weibull) } h(t) = \lambda p(\lambda t)^{p} - 1_{e}^{\lambda t}$$

where $\lambda = \exp(x\beta)$ is the parameterization[3] given in Table 1 and $p = 1/\sigma$.

The first thing we notice is that the parameter estimates are extremely robust across different stochastic specifications, just as reported in King et al. (1990). The role of a constitutional investiture requirement, which reduces duration, is a little smaller in Models 2 and 3, but increased in Model 4, and much the same could be said of the rest. All of the signs are the same; fractionalization, always the most marginally significant, remains so; but any qualitative judgment about effects of these variables would remain largely unaltered by the choice of specification.

The log likelihoods of the Weibull and exponential models can be compared directly, and the Weibull, which adds only one parameter, σ, is clearly a statistically significant, although substantively modest, improvement. (Recall that in the parameterization given above, $\sigma = 1/p$, where p indicates how rapidly the hazard function is rising or falling.) The Weibull model produces a monotonically rising hazard rate, steepest at first and gradually flattening out. We feel that this result is picking up the tendency for coalition partners to enjoy a brief honeymoon period immediately after forming. In other words, once the effects of caretaker status and investiture requirements have been taken into account, for a few months the residual randomly induced chance of falling is smaller. Nevertheless, we continue to believe that the stability of the parameter estimates for the explanatory variables across the wide variety of specifications of time dependence should not be overlooked.

We are therefore in the happy situation of the evidence reported by different researchers again being in agreement on the same important problem.

4. DO THE SAME PROCESSES
DRIVE LEADER AND CABINET DURATIONS?

Bienen and van de Walle (1991, p. 61), analyze 2,256 leader durations in 167 countries since 1801 and find that the risk of a leader's losing power decreases over time, independent of a small number of significant effects of

3. Throughout this article, parameter estimates for the explanatory variables in the tables are presented so as to predict durations. The signs of these estimates must be reversed when predicting hazard rates. Alt and King (1992) discuss the hazard rate for the gamma model.

measurable explanatory variables. They find the risk of losing office to be lower (that is, expected durations to be longer) for leaders drawn from the military and in countries with higher incomes and larger populations, but the opposite to be true for leaders entering nonconstitutionally and for older leaders. They find much shorter durations (higher risk) in Latin America and particularly North America, Europe, and Australasia, which they attribute to the shorter tenures of leaders in parliamentary regimes, and an increase in risk around Year 4, which they attribute to electoral terms in Latin American countries. Although these leaders do not last as long, the temporal pattern of risk that they face is the same as elsewhere. The same patterns of effects (except for income and population) describe the subsample of leader durations beginning after 1945 (p. 87). In this shorter period, independent of age, region, and type of entry, Bienen and van de Walle find no significant systematic predictors of leaders' risk of losing office.[4]

Bienen and van de Walle's (1991) results thus pose two important challenges. First, other research (at least the attributes approach) amassed evidence of effects on cabinet durations of both enduring structural differences between countries and more transitory characteristics of the strategic environment in which cabinets formed. Attribute variables that explain duration predict the ability of governments and/or their leaders to forge support and avert overthrow once in office. But when looking at leaders rather than cabinets, Bienen and van de Walle find few significant effects of attributes, whether structural or strategic. If leaders' durations are not explained by the same sort of attributes, then leaders' personal characteristics or circumstances (or good fortunes) dominate structural and behavioral variables in explaining duration. Part of the explanation is that many explanatory variables are difficult or impossible to gather in the massive collection of leaders data.

At the same time, Bienen and van de Walle's (1991) results strongly indicate a pattern of decreasing risk over time: the longer a leader has been in power, the longer still he or she is likely to remain. As they point out, this could result from "unobserved heterogeneity" (pp. 50-52) or the effects of factors omitted from the model. Another way of saying this is that the explanatory variables they chose to include partition the existing variation in durations in a different way than most of the rest of the literature (because of problems in collecting data on the explanatory variables for most of their

4. To extend their data to over 100 countries and nearly 200 years, Bienen and van de Walle were unable to collect information on some explanatory variables that others found significant. However, they also include median leader duration by country as an explanatory variable. Because a lower risk of losing office corresponds to longer expected durations in office, this is circular, and probably creates bias in the estimates of the effects of other explanatory variables on risk.

countries). Even so, if a falling hazard rate survives further effort to measure these factors, this clearly contradicts the Markov-like constant conditional probability assumptions of the events process theorists, which King et al. (1990) endorsed, even more with the possibility of rising hazard rates at the beginning and end of governments, as discussed in the previous section. It also appears to be inconsistent with the Londregan and Poole (1991) finding that past coups make future coups more likely, although this represents a substantially different specification of time dependence. Finally, of course, if explanations of cabinet durations also apply to leaders, it is inconsistent with the Warwick and Easton (1992) estimates of rising hazard rates.

COMPARING RESULTS FOR CABINETS AND LEADERS

Unfortunately, all of the King et al. (1990) explanatory variables do not exist for most of the leaders data. However, we can use data on those leaders who appear as a subset of the cabinet data. We will reexamine the subset of Bienen and van de Walle's (1991) leaders' durations data, which corresponds to those cabinet durations—that is, leaders of the 15 European countries since World War II—to see whether the same substantive variables predict leadership durations as well as cabinet durations.

Thus, to compare results, we extracted from the Bienen-van de Walle data the duration in office—coded in years—of all of the postwar leaders in those European parliamentary democracies previously analyzed by King et al. (1990). There were 200 such cases. We merged these durations with the data used in King et al., obtained from Strom's (1985) analysis of cabinet durability and extended to include the period until the end of 1987. This latter comprised a set of 314 governments in 15 countries for which durations were recorded in months and there was also complete data on a series of country, party-system, and cabinet attributes (Strom, 1985, p. 747). The differences between the sets (apart from the use of units of months in one and years in the other) is that a change of leader occurs whenever there is a change of prime minister, but a change of cabinet also occurs with a change in the party composition of the cabinet, a formal resignation, or an election.

We proceed in two steps. First, we reanalyze a subset of the cabinet data for which leaders data is available with methods sensitive to time dependence. Then, having satisfactorily replicated other results and chosen a final model, we substitute leaders data and repeat Bienen and van de Walle's (1991) analysis for an important subset of cases, but also with explanatory variables different from theirs. What might we expect to find? A number of results are possible.

1. Everything has been done correctly: conditional on these explanatory variables, leaders' durations are time dependent but cabinet transitions are not. A change of leader is only one of the factors in the definition of a cabinet change, and our systematic factors predict durations of cabinets but are erased from the durations of leaders whose careers may span many cabinets.[5]
2. There is a grouping-induced measurement error problem with the dependent variable, with some months-level effects lost in the yearly aggregates of the leaders data.
3. King et al. (1990) are wrong: There is a conditional time dependence in the cabinets data that makes a substantive difference in causal and predictive estimates. (This turns out to be false, although there is some time dependence.) Moreover, it could even be that, as claimed by Bienen and van de Walle (1989) for the case of African leaders, once time dependence is correctly modeled, other substantive effects disappear. (This, however, does not happen.)
4. King et al. (1990) are right about the systematic effects on cabinet durations and also about the leader durations that correspond to the times and countries covered in their model. (This turns out to be true.) In this case, a significant subset—about 10%—of the cases in Bienen and van de Walle's model may be misspecified, and more variables belong in their model to explain postwar power transitions in 15 European countries.

And of course, given Result 4, it could also be true that the apparent finding of time dependence of other leader durations in Bienen and van de Walle's (1989) work may turn out to be the results of differing specifications of explanatory variables. Dealing with this requires analysts to use different models for different subsets of data (that is, to model interactions between patterns of time dependence and other explanatory variables).

AN ALTERNATIVE SPECIFICATION OF THE LEADERS MODEL

If the length of time a leader remains in power is actually a result of his or her skills, ability, personality, or other individual characteristics, the sort of explanatory variables used in the analysis of cabinet durations will not explain leader durations. That is, what will count is not, for instance, party polarization in parliament, but whether or not the system happens (or is likely) to select a leader with particular attributes who can systematically overcome the obstacles polarization presents to extended tenure. Like Bienen and van de Walle (1991), we lack measures of leader attributes with which to model leader durations directly. Instead, determining the extent to which the systematically predictable component of leader durations is, in fact, the durabil-

5. It is clear at points in Bienen and van de Walle's discussion that this is what they have tried to do, to isolate the factors affecting leaders' duration from those affecting the political system in which they operated, but only a complete specification of system-level characteristics would allow this to be done satisfactorily.

ity of cabinets they lead will indirectly reveal whether system structure dominates individual attributes in producing duration outcomes. We expect to find that most of the explanatory variables, particularly those that indicate complexity and instability in the bargaining environment and the probability of confrontation, challenge, and defeat will significantly predict leader durability. If they do, we will have identified a source of unmeasured heterogeneity in their results.

We test this expectation in three steps. First, we extract the subset of cases of leader durations for which there are data on cabinets and show that it is a representative subset of that data. Then we calculate life tables for these leaders' durations and show that they reveal a hazard rate constant out to 8 years or more. Finally, we experiment with alternative parameterizations that show that time dependence is still present but increasing throughout the subset of leaders' data, whereas the parameter estimates for the effects of explanatory variables, with a few predictable exceptions, remain quite stable.

King et al.'s (1990) postwar sample of 15 countries contained 200 postwar leader durations, which appeared in Bienen and van de Walle's data. Bienen and van de Walle treat Finland as led by its president, so the durations do not correspond in that case, and the French Fifth Republic as well had to be excluded for lack of cabinets data. There are also a few scattered leaders whose cabinets data was lacking, but we ended up with 173 cases of leader durations. In each case, to keep the information set as purely ex ante predictive as possible, we used data from the first cabinet that a multicabinet leader led as the basis of prediction, unless the first cabinet was of less than 6 months and was followed by a term of over a year (that is, a durable leader initially took over as a caretaker).[6]

First, as before, we construct the life table and display its hazard rates in Figure 4. We see that there is no evidence of decreasing hazard rates nor of high hazard rates only in the first few years. The hazard rates vary only a little across the first 7 years, although there is clearer evidence of increase in Year 8, after which the number of cases is so small that none of the hazard rates move more than their standard errors by very much. Although no conclusions are appropriate without first controlling for systematic effects, there is clearly no evidence of declining hazard rates in our subset of the leaders data, nor is there any evidence of a surge in the hazard rate around Year 4, to which Bienen and van de Walle called attention.

Table 2 gives the statistical results when the variables used in King et al. (1990) predict leader durations. Model 5 replicates Model 3 from Table 1 on

6. We also corrected four miscoded leader durations, most of which were coded substantially longer than the actual durations.
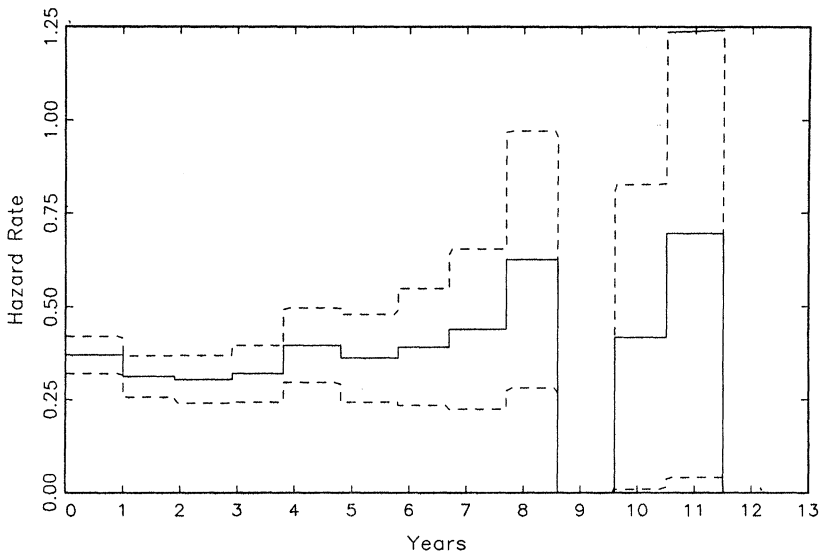
Figure 4. **Unconditional hazard rate and standard errors.**

the subset of 173 cases out of the original 314: Data are in months and censoring is as before. The standard errors of the parameter estimates all go up because the number of cases is now smaller, but few of the estimates themselves change by as much as 10%. On the whole, we feel that the subset preserves the contours of the earlier analysis of cabinet duration very well.

Model 6 replicates the exponential Model 1 from Table 1, but now with the leaders' durations rather than the underlying cabinet durations. Of course the constant term changes dramatically; it would be reduced by about $\ln(12) =$ 2.5, simply by the change from months to years, although in fact, the decline is a little larger.[7] Some of the parameter estimates remain hardly altered (conditional on the change in scale picked up by the constant term, there is no reason that this change in the measurement of duration should affect any parameter estimates except for the constant term). The effects of investiture requirements, polarization, and even numerical status are evident. The effects that are the most reduced are those relating to timing of accession. Clearly, many caretakers succeeded themselves, and many durable leaders took over

7. Suppose m is durations measured in months, and we specify the usual exponential functional form, $E(m) = \exp(\beta_0 + \beta x)$. This equation may be rewritten as $E(m) = \exp(\beta_0)\exp(\beta x)$. Then, if we estimate the same equation in years, the factor $\exp(\beta_0)$ adjusts by 12, which implies that β_0 decreases by 2.5.

Table 2
Survival Models of Leaders Durations

Explanatory Variable	Model			
	5	6	7	8
Fractionalization	−.0006	−.0006	−.0005	−.0004
	(.0009)	(.0012)	(.0006)	(.0010)
Polarization	−.017	−.015	−.014	−.015
	(.006)	(.010)	(.005)	(.008)
Formation attempts	−.050	−.091	−.092	−.087
	(.053)	(.115)	(.060)	(.066)
Investiture	−.310	−.345	−.397	−.354
	(.138)	(.231)	(.122)	(.189)
Numerical status	0.559	0.376	0.413	0.379
	(.140)	(.239)	(.126)	(.184)
Postelection	0.618	0.172	0.111	0.206
	(.140)	(.234)	(.130)	(.172)
Caretaker	−1.16	−.455	−.486	−.829
	(.282)	(1.05)	(.568)	(.421)
Constant	3.447	0.850	1.008	
	(.623)	(1.08)	(.452)	
Sigma	0.692		0.706	
	(.052)		(.053)	
Log likelihood	−211	−214	−198	−699
Model	Weibull	Exponential	Weibull	Prop. Hazards

Note. Observations are 173 leader durations, measured in years, in 14 countries, 1946-1987, except in Model 5 where the data are cabinet durations measured in months. Standard errors are in parentheses. Models 6-8 treat leaders dying of natural causes or still in office as censored. The log likelihood from Model 8 is not comparable to those in other columns. Parameter estimates are given with signs to predict duration.

in midterm. Although the cabinet they lead often changes after the first election they face, they are not sufficiently likely to lose that election for timing of accession to predict their durability. Other than this, enough of the earlier structure shows through the change in dependent variable to suggest strongly that leaders' durations can reasonably be modeled at least partly in terms of the determinants of their underlying cabinet durability, where it is appropriate.

Model 7 confirms that the Weibull model's pattern of increasing hazard rates is not confined to the first few years within cabinets, but relates to the expected durations of leaders in parliamentary systems generally. The systematic effects are much as we just described in the exponential model, and in fact, the standard errors of our estimates are smaller. (Model 8, using Bienen and van de Walle's preferred proportional hazards model, confirms

these systematic parameter estimates yet again.) Within post-1945 parliamentary systems, hazard rates appear to increase for leaders as well as cabinets.

However, knowing only whether the hazard rate is increasing, constant, or decreasing is often largely uninformative. Just as we wish to know the sign and magnitude of regression coefficients, we should also pay attention to how steeply increasing or decreasing a hazard rate is. The models in Table 2 produce dramatically differing hazard rates, several of which we plot out in Figure 5. With the explanatory variables set to their average values, Figure 5 shows the difference between the Weibull model's (Model 7) increasing hazard rate and the constant rate of the exponential model (Model 6), which overstates risk at short durations and understates it at longer durations.[8] But the differences between these functional forms pale in comparison to the differences made by the explanatory variables. Setting the main explanatory variables a standard deviation away from their means to predict shorter durations (for this example, imagine a minority government forming after several attempts in a country having an investiture requirement with fairly high fractionalization like that observed in Israel and polarization typical of Belgium or Iceland) produces an explosively upward hazard rate. By contrast, setting the main explanatory variables to produce long durations (a majority government forming at the first attempt with Sweden's polarization and the UK's fractionalization) produces the other hazard rate in Figure 5, which slopes up so gradually that for longer intervals the constant hazard rate is by no means a bad approximation. To conclude, when explanatory variables are included in the analysis and make such substantive difference, it makes less sense to make summary comments about the overall hazard rate because there are different hazard functions for every different combination of values of the explanatory variables.

Results from the smaller but overlapping sample of countries appear to differ significantly from the results of Bienen and van de Walle's (1991) analysis in two ways. First, there are significant covariates at the political system and party system levels, which predict leader duration in an important subset of countries but were omitted from their analysis. Second, conditional on these effects, hazard rates appear to increase, not decrease, among leaders of European parliamentary regimes, that is, the longer a leader is in office, the greater his or her risk of losing office. Other subsets of their data also reveal different systematic effects and different conditional patterns of time dependence, as Londregan and Poole's (1990) results for coups showed. Bienen and van de Walle's finding of an apparently declining hazard rate for

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8. In the case of binary variables like majority status, the plotted hazard rate is based on the proportion of cases coded as majorities in the data. Recall that the Weibull model's hazard rate is given by $h(t) = \lambda p(\lambda t)^{p-1} e^{\lambda t}$ where $\lambda = \exp(x\beta)$, while the exponential model's is just λ.
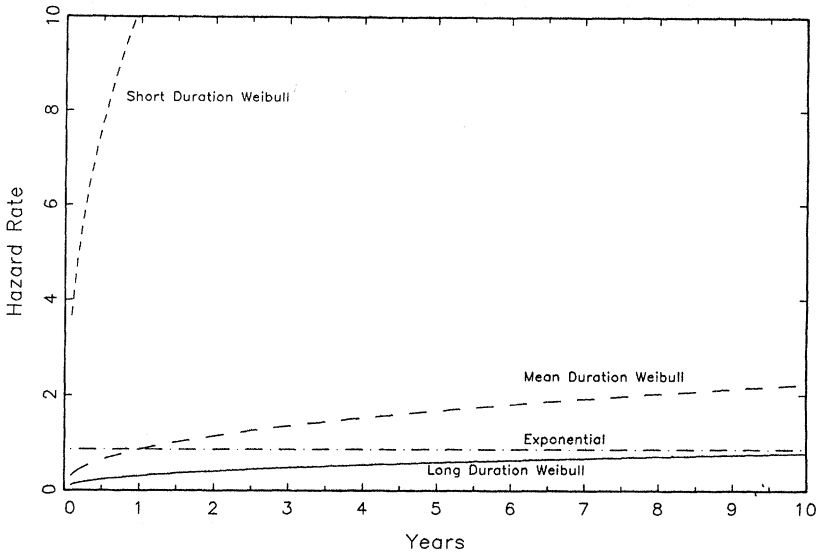
Figure 5. **Hazard rates for different values of explanatory variables.**

leaders could combine our high, rising rate for parliamentary leaders (pre-dominantly of shorter duration) with the lower, flat (and of longer duration) rate for those entering in coups, and other uncontrolled sources of heteroge-neity in the data.

5. FUTURE DIRECTIONS

This article demonstrates that the literature on transfers of governmental power, whether dealing with cabinet durations or indeed coups, provides models that can be compared. It produces models in which stochastic and systematic elements appear, with the latter containing the effects of variables relating to structural and strategic characteristics. Equivalent social-level data on polarization and fractionalization (and even crisis duration) could predict leader durations in nonparliamentary settings too.

The natural next step for the field is to generalize these models to include time-varying covariates, such as economic conditions, poll results, and so on. Londregan and Poole (1990), Merlo (1991), and others have made progress in developing these models. However, the precise goal of the analysis will greatly affect the specification of the explanatory variables.

For example, the goal of King et al. (1990), Warwick and Easton (1992), and many others was ex ante predictions of, and estimates of the causal effects of, characteristics of coalitions and their environment at the time of formation on government duration. To pursue these particular goals, one should not control for variables that change over the course of a government's tenure because these variables are in part a consequence of the characteristics of the government at formation. Thus models incorporating time-varying covariates are irrelevant to the goal of creating ex ante forecasts and causal estimates (see King, 1991); indeed, including them will destroy the desirable properties of all of the ex ante causal effects.

However, suppose the goal were to evaluate the causal effect of time-varying covariates such as monthly economic conditions. In this case, controlling for King et al.'s (1990) ex ante variables is essential to avoid omitted variable bias. Thus it would make the most sense for future researchers to include both specifications—one with only ex ante variables, presumably replicating existing results, and one with ex ante variables and time-varying covariates. The coefficients on the ex ante variables should only be interpreted from the former and the time-varying covariates from the latter.[9]

We close with a further conjecture, that parliamentary leaders face increasing hazard rates because popularity and elections can still be lost even after a leader has managed to secure two or more turns in office. Then our model predicting increasing hazard rates for European parliamentary leaders should fit durations of parliamentary regime leaders on all continents. Other kinds of leaders who can consolidate power (perhaps after an initial period of very high risk) face decreasing hazard rates. This would be a straightforward example of a valuable further step in modeling durations, which is to allow the dispersion parameter (which describes the slope of time dependence in the hazard function) itself to be a function of some independent variables. Indeed, perhaps even the difference between presidential and parliamentary systems itself predicts hazard functions of different shapes.

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9. To facilitate this development, we have elsewhere (Alt & King, 1992) begun linking together a variety of statistical models of government transfers data, so that conclusions can be compared even where methods of data collection and the level of analysis at which the model is estimated differ.

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James E. Alt is professor of government at Harvard University and director of the Harvard-MIT Research Training Group in Positive Political Economy.

Gary King is professor of government at Harvard University and director of the Harvard Data Center.