A Copula Approach to the Problem of Selection Bias in Models of Government Survival

Daina Chiba
Department of Government, University of Essex
dchiba@essex.ac.uk

Lanny W. Martin
Department of Political Science, Rice University
lmartin@rice.edu

Randolph T. Stevenson
Department of Political Science, Rice University
stevenso@rice.edu

Abstract

Theories of coalition politics in parliamentary democracies have suggested that government formation and survival are jointly determined outcomes. An important empirical implication of these theories is that the sample of *observed* governments analyzed in studies of government survival may be nonrandomly selected from the population of *potential* governments. This can lead to serious inferential problems. Unfortunately, current empirical models of government survival are unable to account for the possible biases arising from nonrandom selection. In this study, we use a copula-based framework to assess, and correct for, the dependence between the processes of government formation and survival. Our results suggest that existing studies of government survival, by ignoring the selection problem, overstate the substantive importance of several covariates commonly included in empirical models.

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Scholars of coalition politics in parliamentary democracies have often suggested that government formation and survival are jointly determined outcomes of multiparty bargaining (see, e.g., Laver and Schofield, 1990; Laver and Shepsle, 1996; Diermeier, Eraslan and Merlo, 2003). That is, when legislative parties negotiate over the composition of the government, they presumably do so with an eye to the *expected stability* of that government. For example, De Swaan (1973) argued long ago that when faced with numerous alternative governments that are more or less equally desirable in all other respects, parties have incentives to choose the government they anticipate will last the longest. In contrast, Diermeier, Eraslan and Merlo (2003) have recently developed a bargaining model that predicts that parties will sometimes prefer coalitions with shorter expected durations.[1] Whether parties prefer to form coalitions that they expect to be more stable or less stable, both possibilities suggest that the processes of government formation and survival are almost certainly linked in significant (perhaps complicated) ways. An important implication is that current empirical studies of government survival—all of which use a single-equation framework to model the effects of various cabinet-level and system-level attributes on the durability of observed governments—may need to be fundamentally reconsidered. As Diermeier (2006, 175) points out, if "expectations about cabinet duration influence the choice of initial cabinets...this implies that we cannot treat cabinet-specific features as proper independent variables in a regression model. Rather, we are facing a selection problem." By assuming away this problem, current research runs the risk of producing inaccurate parameter estimates, resulting in mistaken inferences about the characteristics of cabinets and legislatures that lead to more or less stability.

Of course, the theoretical possibility of such a risk does not mean that nonrandom selection is necessarily damaging for any particular empirical model. Its impact depends on, among other things, the extent to which there are *unmeasured* factors that impact both government formation and survival. If such relevant unmeasured factors do not exist, the two processes are stochastically independent, and their joint likelihood can be safely factored into separate formation and duration

[1] This can happen in their model if larger majority coalitions are more stable. In this case, longer lasting coalitions must divide the spoils of office among more parties, and so it may sometimes be in the interest of the party leading negotiations (the *formateur*) to choose smaller coalitions even if they are expected to be shorter-lived.

components. Certainly, one could argue that after decades of theoretical and empirical modeling, scholars have identified most of the systematic factors explaining government formation and survival and have been able to measure these factors with ever smaller amounts of error. It could be the case, then, that the logical possibility of a selection problem is not empirically consequential because extant models are sufficiently well-specified, and the relevant variables sufficiently well-measured, that there exists little systematic covariation in the error terms of the two processes.

Unfortunately, the methodological toolkit for assessing whether this is a problem is currently inadequate to the task. In this study, we present a relatively new empirical approach, based in copula theory, that allows us to model the processes of government formation and survival *jointly*. Our approach allows us both to evaluate whether current single-equation models of government survival suffer from problems of nonrandom selection and to correct for such problems if they exist. In the next section, we describe in more detail the nature of the potential selection problem in models of government survival. We then present the details of our copula approach and demonstrate, through Monte Carlo simulations, its effectiveness in overcoming the problems of selection bias. Finally, using new data on government coalitions in 17 parliamentary democracies, we compare our joint model with single-equation survival models. Our findings indicate that current research on government survival, by ignoring the selection problem, significantly overstates the substantive importance of several covariates commonly included in empirical models.

Selection Effects in Models of Government Survival

The generic problem of nonrandom selection is well understood. In the context of our particular question, if there are unmeasured (and hence excluded) variables that impact which alternative coalitions (or "potential governments") get selected into the sample of observed governments *and* how long these observed governments last, then our inferences about the effects of the included variables on government survival may be biased. As Achen (1986) has demonstrated, such bias will be more severe as selection becomes less deterministic (i.e., as fewer of the important factors that determine government formation are included in the model of government survival) and as the correlation between the stochastic processes of formation and survival (conditional on the included factors) becomes larger. Thus, if we concede that our empirical models of government survival are incomplete in terms of the variables included that affect both formation and survival, then we must acknowledge the possibility of selection effects. It becomes an empirical question as to whether this generic kind of selection problem has any substantive impact on our inferences.

It is interesting to note that even though most empirical researchers of government survival have not explicitly modeled the relationship between survival and formation, their concept of what a government is comes directly from the formation literature. That is, many studies of survival implicitly rely on the notion that a government represents an *equilibrium allocation* of office and policy payoffs among legislative parties, in that no party necessary for supporting the government can unilaterally improve its payoff by supporting an alternative. These payoffs are realized (at least in expectation) at the time a government forms and can only change when the factors underlying the equilibrium change. Thus, as Laver and Shepsle (1996, 196) point out, "a general model of cabinet stability must identify the key parameters [supporting an equilibrium government] and specify the types of change in these that are likely to destabilize the government." That is, if scholars want to identify the factors that predict government duration, they need to look for factors that can either "shock" the equilibrium payoffs to parties or make the current equilibrium robust to such shocks.

Much of the empirical literature on government survival reflects this logic. For example, Browne, Frendreis and Gleiber (1984), King et al. (1990), Warwick (1992), and Diermeier and Stevenson (1999, 2000) have all modeled government termination as a random variable in which "critical events" can bring the government down. Considerable work has gone into exploring the nature of the distribution of critical events, which equates to specifying the form of the hazard rate for failure times. Some have argued that these events follow a Poisson process, some have argued for more complicated stochastic processes, and still others (in the "competing risks" camp) have advocated the idea that different stochastic processes may govern different kinds of terminations.

Whatever one's view of the nature of the underlying process generating critical events, all of the recent empirical work in this area recognizes that the "government-as-equilibrium" theoretical approach implies that the impact of critical events will not be the same for all governments—characteristics of governments and the bargaining environments in which they exist can make them more or less sensitive to shocks of different magnitudes (Lupia and Strom, 1995; Laver and Shepsle, 1996; Diermeier and Stevenson, 2000). For example, no shock to the distribution of party preferences is likely to bring down a single-party majority government (at least where there is strong party discipline), while one can imagine any number of shocks that would bring down an ideologically diverse coalition controlling less than a majority of legislative seats. This literature has characterized the institutional and political context in which an equilibrium government exists (as well as the corresponding collection of government attributes) as "bargaining environment complexity" (see also Laver and Schofield, 1990) and has suggested that, in general, governments that exist in more complex bargaining environments will be more vulnerable to shocks to party preferences and expected seat distributions. The issue of which exact indicators constitute a more or less complex bargaining environment depends on the specific theoretical model one adopts to produce predictions of equilibrium governments. However, the empirical literature has generally thought of more complex bargaining environments as those in which there is high *legislative fragmentation* (which is usually measured by the effective number of legislative parties). This bargaining complexity characteristic tends to produce equilibrium governments that are less robust to shocks (Laver and Shepsle, 1996, 98–106). In addition, some have argued that high *legislative polarization*—generally thought of as a greater presence of anti-establishment parties, which tend to use destabilizing parliamentary tactics to attract protest votes (Powell, 1982)—should also make equilibrium governments more sensitive to shocks. In short, the key prediction is that governments that exist in complex bargaining environments should possess characteristics that are less robust to shocks, thereby making the governments less durable. Empirically, this indeed appears to be the case (see, e.g., King, Alt, Burns and Laver, 1990).

Given this understanding of governments as equilibria that can potentially be shocked by random events, along with the corresponding understanding that observable features of both the government and the larger legislative environment regulate the impact of these events, it seems clear that strategic politicians could use this information to estimate the "survivability" of any potential government in a particular coalition bargaining situation (or "formation opportunity") and then incorporate this estimate into their choice among potential governments. When will this be a problem for estimating the impact of various characteristics of governments and the legislative bargaining environment on government duration? In general, it will be a problem when (1) politicians choose governments based on their expected duration and (2) their expectations depend on factors other than the measurable features of the governments and bargaining environments included in empirical models of survival. Of course, it is possible that neither of these conditions hold. In the next section, we propose an empirical strategy that allows us to assess whether selection bias is a problem in current work on government survival and to correct for such bias if it exists.

Methods

To obtain accurate parameter estimates of the determinants of government survival in the face of nonrandom government selection, we adopt an empirical strategy to model both processes simultaneously. Our proposed solution is to estimate a (competing risks) duration model while conditioning on the selection process.[2] In this sense, it shares many features of the approach developed by Heckman (1976) for linear models, Dubin and Rivers (1989) for logit and probit models, and Boehmke, Morey and Shannon (2006) for duration models. Our estimator extends previous work on sample selection and continuous-time duration models in two directions. First, to model the selection process, we use a polychotomous conditional logit model (McFadden, 1973), which has been used in numerous studies of government formation in parliamentary democracies (e.g., Martin and Stevenson, 2001, 2010; Bäck and Dumont, 2007; Indridason, 2008; Warwick, 2005, 2006). To our knowledge, this is the first survival analysis to model the selection process as a multinomial

[2] Competing risks duration models are the prevailing standard in the government survival literature (see, e.g., Diermeier and Stevenson, 1999, 2000; Gordon, 2002). Our empirical models will distinguish between two types of government terminations: those due to the dissolution of parliament and the calling of early elections (*dissolution* terminations), and those due to the direct replacement of the government, with no intervening election, by an alternative administration (*replacement* terminations).

choice problem. Second, in testing and correcting for the sample selection bias that may result from the stochastic dependence of the two stages, we use a flexible distribution that accommodates an unrestricted range of correlation between the selection and duration processes. This overcomes one of the limitations in previous survival models with nonrandom selection (Boehmke, Morey and Shannon, 2006).[3]

These two innovations are made possible through the use of copula functions, which have recently gained momentum in applied econometrics (Trivedi and Zimmer, 2005). A copula is a function that parameterizes the dependence between univariate marginal distributions to form a joint distribution function. Consider two random variables y_1 and y_2 with associated univariate distribution functions $F_1(y_1)$ and $F_2(y_2)$. Sklar's (1959) theorem establishes that there exists a copula $C(\cdot, \cdot; \theta)$ such that a bivariate joint distribution is defined for all y_1 and y_2 in the extended real line as,

$$F(y_1, y_2) = C\left\{F_1(y_1), F_2(y_2); \theta\right\}, \tag{1}$$

where the association between the two marginal distributions is represented by the association parameter, θ. It is important to note that the functional form of a copula does not depend on the functional form of the univariate marginals. This is useful because we can construct a new bivariate distribution based on univariate marginal distributions that may be from different families. As long as the univariate marginal distributions are known, an appropriate choice of copula function C in (1) enables us to represent the unknown bivariate distribution. Taking advantage of this feature, we construct a joint model of government formation and duration where the government duration is modeled by the Weibull distribution and the government selection process is described by the conditional logit model. The flexible copula framework allows us to specify each marginal distribution as a function of a set of covariates.

Our goal is to make an inference about the duration of a government j that was chosen in a formation opportunity i. The selection process leads us to observe the actual duration of the

[3] For a model without such limitations, see the work of Prieger (2002).

formed government, while the duration is censored for all other possible coalitions in the formation opportunity. If some unobserved variables influence both the formation and duration of a government, then the observed sample of governments across formation opportunities constitutes a biased sample on which to base an inference about government duration. Modeling the duration process thus requires us to account for the selection process by writing down a probability model for government selection.

In a given formation opportunity $i = 1, \cdots, n$, only one coalition j can in fact form out of M_i possible coalitions. Let us denote a government formation outcome by Y_i, such that $Y_i = j$ when coalition j is formed in formation opportunity i. For convenience we also define a binary censoring variable c_{ij} that takes on the value of 1 for the coalition that is formed, and 0 for all the other possible coalitions that are not formed. Then, for each i, one and only one of the c_{ij}'s is 1. Once a particular government is formed, then we observe a government duration. Let us denote a continuous duration outcome by T_i. In addition to the prior selection process of a coalition, we allow for right-censoring by calculating the probability that an observation has a duration greater than the right-censoring point, t_i^0. With a right-censoring indicator, r_i, that takes on the value of 1 when the observation is right-censored and 0 otherwise, we have the following likelihood function for the observed duration of a government:

$$\mathcal{L} = \prod_{i=1}^{n} \prod_{j=1}^{M_i} \left\{ \Pr(T_i > t_i^0,\ Y_i = j)^{r_i}\ \Pr(T_i = t_i,\ Y_i = j)^{(1-r_i)} \right\}^{c_{ij}} . \tag{2}$$

Notice that the censoring indicator c_{ij} never takes on the value of 0 for the observed government. Nevertheless, we need a joint probability representation to describe how information from the non-formed coalitions contribute to the total likelihood function.

The next step is to calculate each of the two joint probabilities by first specifying the univariate marginal distributions for T_i and Y_i and then binding them together with a copula function. We start with the specification of the government formation process by describing the univariate marginal distribution for Y_i. Following the multinomial conditional logit model (McFadden, 1973), a latent utility for each potential coalition is given as a function of a set of covariates, z_{ij}, a vector of regression coefficients, γ, and a stochastic component,

$$u_{ij} = z_{ij}\gamma + \eta_{ij}, \tag{3}$$

where the stochastic parts of the utility functions, $\eta_{ij},\ j = 1, \cdots, M_i$, are assumed to be independent and identically Gumbel-distributed. That is, we are assuming a standard random utility framework in which the utilities across alternatives in a given formation opportunity are uncorrelated with one another. The probability that a coalition j forms a government in a given formation opportunity i is,

$$\begin{aligned} \Pr(Y_i = j) &= \Pr\left\{\max_{j=1,\cdots,M_i}(u_{ij}) = u_{ij}\right\} \\ &= \frac{\exp(z_{ij}\gamma)}{\sum_{j=1}^{M_i}\exp(z_{ij}\gamma)} = G(z_{ij}\gamma). \end{aligned} \tag{4}$$

We then specify the univariate marginal distribution for the government duration. We allow the duration of the government to be conditioned on a set of covariates by specifying the hazard rate, λ_i, as a function of a vector of covariates x_i, such that $\lambda_i = \exp(-x_i\boldsymbol{\beta})$. Using the Weibull specification, the univariate density function, $f(t)$, the survivor function, $S(t)$, and the distribution function, $F(t)$, are given as,

$$\begin{aligned} f(t) &\equiv \Pr(T = t) = \lambda p\,(\lambda t)^{p-1}\exp\left(-\,(\lambda t)^p\right) \\ S(t) &\equiv \Pr(T > t) = \exp\left(-\,(\lambda t)^p\right) \\ F(t) &\equiv \Pr(T \leq t) = 1 - S(t), \end{aligned} \tag{5}$$

where p is the shape parameter that determines whether the hazard is increasing, decreasing, or constant over time.[4]

[4] We use the Weibull specification here for illustrative purposes. In practice, we can use other distributions such as Exponential, Log-logistic, Gompertz, or Generalized Gamma, etc., to model the duration process. In the empirical application that follows, we estimate both Weibull and Log-logistic models and choose the model that

From the univariate marginal distribution functions for duration and formation, $G(\cdot)$ and $F(\cdot)$, we calculate the joint probabilities for the duration given that the observation is selected. This is done by using a copula function to represent the joint cumulative distribution function. The first component of (2), the probability that an observation is selected and has a duration greater than the right-censoring point t_i^0 is,

$$\begin{aligned}\Pr(T_i > t_i^0,\ Y_i = j) &= \Pr\left\{T_i > t_i^0 \ \cap \ \max_{j=1,\cdots,M_j}(u_{ij}) = u_{ij}\right\} \\ &= \Pr\left\{\max_{j=1,\cdots,M_j}(u_{ij}) = u_{ij}\right\} - \Pr\left\{T_i \leq t_i^0 \ \cap \ \max_{j=1,\cdots,M_j}(u_{ij}) = u_{ij}\right\} \\ &= G(z_{ij}\gamma) - C\left\{F(t_i^0), G(z_{ij}\gamma); \theta\right\}. \end{aligned} \tag{6}$$

The second component of (2), the probability that an observation is selected and has a duration equal to t_i, is obtained by using the density function,

$$\begin{aligned}\Pr(T_i = t_i,\ Y_i = j) &= \Pr(Y_i = 1 \mid T_i = t_i) \times \Pr(T_i = t_i) \\ &= \Pr\left\{\max_{j=1,\cdots,M_j}(u_{ij}) = u_{ij} \mid T_i = t_i\right\} \times f(t_i) \\ &= \frac{\partial C\left\{F(t_i), G(z_{ij}\gamma); \theta\right\}}{\partial F(t_i)} \times f(t_i). \end{aligned} \tag{7}$$

To complete the derivation, the last step is to choose a particular copula function for $C(\cdot, \cdot; \theta)$. There are a number of different copula functions that can be used to construct a multivariate distribution from univariate marginals (Trivedi and Zimmer, 2005), but some copulas are more flexible than others in that they can accommodate a greater range of dependence between the marginals. In this application, we use the Gaussian copula, one of the most flexible copula functions that can

better fits the data.

accommodate both positive and negative dependence. It has the following form,

$$
\begin{aligned}
C(u, v; \theta) &= \int_{-\infty}^{\Phi^{-1}(u)} \int_{-\infty}^{\Phi^{-1}(v)} \frac{1}{2\pi(1-\theta^2)^{1/2}} \exp\left\{ \frac{-(s^2 - 2\theta st + t^2)}{2(1-\theta^2)} \right\} ds dt \\
&= \Phi_2 \left\{ \Phi^{-1}(u), \Phi^{-1}(v); \theta \right\}, \qquad (8)
\end{aligned}
$$

where Φ^{-1} is the inverse Gaussian function, Φ_2 is the bivariate Gaussian distribution function, $-1 < \theta < 1$, and $u = F_1(y_1)$ and $v = F_2(y_2)$ for random variables y_1 and y_2. This implies that the conditional probability in equation (7) is given as,

$$
\frac{\partial C\,(u, v; \theta)}{\partial u} = \Phi \left\{ \frac{\Phi^{-1}(v) - \Phi^{-1}(u)\theta}{\sqrt{1-\theta^2}} \right\}, \qquad (9)
$$

where Φ is the standard Gaussian distribution function.

The Gaussian copula has a number of desirable characteristics. First, it allows for independence as a special case ($\theta = 0$). We can thus test the existence of selection bias by testing whether θ is different from 0. Second, the Gaussian copula is *comprehensive* in that as θ approaches the lower (upper) bound of its permissible range, the copula approaches the theoretical lower (upper) bound.[5] This is not true with other copulas that have been utilized to address selection bias in political science applications. For example, the copula function proposed by Sartori (2003) forces one to *assume* that one or the other of the theoretical bounds represents the true data generating process. The consequence of this is not only that we are unable to test the existence of selection bias but also that, depending on the assumption made about the direction of the dependence, we make completely opposite inferences about the effect of explanatory variables on outcomes. The copula function utilized in Boehmke, Morey and Shannon (2006) can accommodate both positive and negative dependence, and allows for testing the direction of dependence, but the permissible range is limited to $\theta \in [-0.25, 0.25]$.

It is worth noting several (potentially limiting) features of the assumed correlation structure in

[5] The upper and lower theoretical bounds of a joint distribution, called Fréchet bounds, F^- and F^+, are defined as $F^-(u, v) = \max(0, u + v - 1)$ and $F^+(u, v) = \min(u, v)$.

our model. First, for each potential coalition we assume a pairwise non-zero correlation between the error term in the formation process and the error term in the duration process, but we assume that the error terms in the formation process are not correlated across potential coalitions. We also assume that the correlation is constant among pairs of formation and duration error terms across different potential coalitions j and different formation opportunities i. While we believe these are reasonable assumptions, analysts may want to relax them in future research.[6]

Monte Carlo Simulation

We now perform a Monte Carlo analysis to assess the consequences of nonrandom selection for the conventional (single-equation) survival model and to demonstrate the effectiveness of our joint estimation approach. The analysis reveals that the conventional survival model produces biased estimates of the model parameters. It also shows that, when the processes underlying selection and survival are nontrivially correlated, our joint approach significantly outperforms the conventional model in terms of root mean-squared error.

A single round of our Monte Carlo simulation consists of 200 hypothetical coalition formation opportunities. Each formation opportunity has 4 political parties, giving rise to $\binom{4}{1} + \binom{4}{2} + \binom{4}{3} + \binom{4}{4} = 4+6+4+1 = 15$ possible combinations of political parties per each formation opportunity. We thus have a total of $200 \times 15 = 3000$ potential governments. For all of these potential governments, we generate data on latent utilities and coalition duration. The latent utilities are generated according to equation (3). We use these latent utilities to determine which government is formed out of 15 possible governments in each of the 200 formation opportunities. We then generate data on duration according to a Weibull duration model. We use the Weibull model because it is one of the simplest duration models that can capture positive and negative duration dependence.[7] Recall that researchers are able to observe the actual duration of government only for those governments

[6] In the Supplementary Materials for this article, we discuss one of the potentially limiting features of the correlation structure in our model, which arises from the fact that we use an independent competing risks specification for observed durations.

[7] In an empirical application that follows, we also estimate other parametric duration models as a robustness check.

that formed. We thus retain those 200 duration observations and discard the remaining duration data in conducting the statistical analysis. Nevertheless, our joint estimation approach allows us to utilize information from the government selection process to correct for bias in estimation.

We assume that two independent variables influence the government formation process, one of which also influences the duration of the chosen government. The purpose of our Monte Carlo simulation is to test whether our approach and standard models can recover the true effects of these independent variables on government formation and duration. We generate 3000 values of two independent variables, x_1 and x_2, each according to an independent uniform distribution over the interval $(-2, 2)$. We hold these variables constant throughout the simulations.

Each round of simulation begins by generating two correlated random variables (v_1, v_2) for the 3000 observations according to a bivariate normal distribution with a given correlation coefficient θ. We then transform v_1 into a Gumbel random variate, η, using the inverse transformation method. We then generate the latent utilities for coalition $g = 1, \cdots, 3000$ as follows:

$$u_g = -1.5 * x_1 + x_2 + \eta_g. \tag{10}$$

To generate government durations, we transform v_2 into an exponential variate using the inverse transformation method, such that $\epsilon_g = -\log(1 - \Phi(v_2))$, where Φ is the standard Gaussian distribution function. Then, with a duration dependence parameter, p, we generate a government duration, t_g, such that,

$$t_g = (\exp(\alpha + \beta x_1) * \epsilon_g)^{1/p}, \tag{11}$$

where α and β are the regression coefficients of interest.[8]

To compare the performance of our proposed model to that of conventional duration models, we vary the correlation parameter from -0.9 (very high negative correlation) to 0.9 (very high

[8] Note that this is the accelerated failure time (AFT) interpretation—positive coefficients correspond to longer expected duration times.

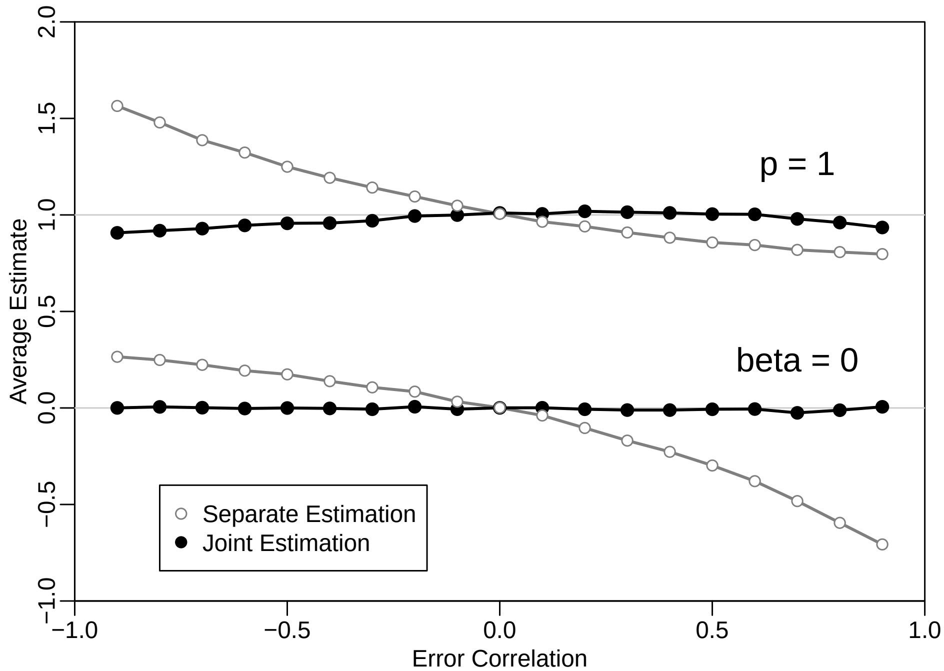

Figure 1: *Summary of Monte Carlo Simulation Results*

Note: Black solid circles show the average estimates from our joint model, whereas gray hollow circles are the average estimates from a naive Weibull model that ignores sample selection.

positive correlation) while fixing the other parameters. We set $\alpha = 0$, $\beta = 0$, and $p = 1$. The assumption $p = 1$ means that there is no duration dependence.[9] For a single round of simulation, we generate 500 draws of the errors for a given value of the correlation parameter, calculate t_g, then estimate our model and a "naive" Weibull model that ignores selection. We save the resulting coefficient estimates and standard errors and then repeat the process for different values of the correlation parameter.

Figure 1 summarizes our Monte Carlo results graphically. The plotted values provide a simple illustration of the bias that is introduced by ignoring sample selection. Solid circles in black show the estimates from our joint model, and hollow circles in gray are the estimates from a naive Weibull model that ignores sample selection. The two connected lines at the bottom show the average estimates of the slope parameter, $\beta = 0$, and the two connected lines at the top show the

[9] Therefore, our Weibull model reduces to an exponential model. We nevertheless use a Weibull specification in the estimation to test if we can recover this true relationship.

Table 1: *Monte Carlo Simulation Results for Slope Parameter ($\beta = 0$)*

	Joint Model			Weibull				
θ	**Mean**	**SD**	**RMSE**	**Mean**	**SD**	**RMSE**	**RMSE Ratio**	**t-ratio**
-0.9	0.00	0.06	0.06	0.26	0.06	0.27	4.69	72.82
-0.8	0.00	0.07	0.07	0.25	0.06	0.25	3.79	60.22
-0.7	0.00	0.07	0.07	0.22	0.06	0.23	3.26	51.69
-0.6	0.00	0.08	0.08	0.20	0.07	0.21	2.61	41.77
-0.5	0.00	0.09	0.09	0.17	0.08	0.19	2.17	33.72
-0.4	-0.01	0.10	0.10	0.14	0.08	0.16	1.66	26.68
-0.3	-0.01	0.10	0.10	0.11	0.08	0.14	1.42	20.14
-0.2	0.00	0.10	0.10	0.08	0.09	0.12	1.15	12.96
-0.1	0.00	0.12	0.12	0.04	0.09	0.10	0.86	6.18
0	0.00	0.12	0.12	-0.01	0.10	0.10	0.81	-0.33
0.1	0.00	0.13	0.13	-0.04	0.11	0.12	0.90	-5.89
0.2	-0.01	0.13	0.13	-0.11	0.11	0.15	1.13	-13.15
0.3	-0.01	0.14	0.14	-0.16	0.11	0.20	1.43	-19.31
0.4	-0.01	0.15	0.15	-0.23	0.12	0.26	1.78	-26.08
0.5	0.00	0.15	0.15	-0.30	0.12	0.32	2.16	-34.46
0.6	0.00	0.16	0.16	-0.38	0.12	0.40	2.58	-42.71
0.7	-0.01	0.16	0.16	-0.48	0.13	0.50	3.08	-49.99
0.8	-0.01	0.16	0.16	-0.59	0.13	0.60	3.81	-62.48
0.9	0.00	0.16	0.16	-0.71	0.14	0.72	4.61	-75.17

average estimates of the duration dependence parameter, $p = 1$. We can see that our joint model consistently recovers values very close to the true value of the slope parameter, β, whereas the common Weibull duration model that ignores selection generates positive (negative) bias when the errors are negatively (positively) correlated. The figure also displays the average estimates of the duration dependence parameter: values of p greater than 1 indicate positive duration dependence, while values of p less than 1 suggest negative duration dependence.[10] Recall that we generated duration data by assuming that there is no duration dependence. However, the naive Weibull model produces positive (negative) duration dependence when the errors are negatively (positively) correlated. These results demonstrate that ignoring selection can lead to biased inferences.

Table 1 presents the detailed results of the average estimates for β. We report average estimates

[10] Since p has to be greater than zero, we estimate $\log(p)$ instead of p.

Table 2: *Monte Carlo Simulation Results for Auxiliary Parameters(θ and $\log(p)$)*

	Joint Model $\log(p)$				$\hat{\theta}$		**Weibull** $\log(p)$				
θ	**Mean**	**SD**	**RMSE**	**t-value**	**Mean**	**SD**	**Mean**	**SD**	**RMSE**	**t-value**	**RMSE Ratio**
-0.9	-0.10	0.08	0.13	-1.12	-0.82	0.14	0.45	0.06	0.45	8.24	3.56
-0.8	-0.08	0.10	0.12	-0.84	-0.74	0.15	0.39	0.06	0.40	7.17	3.18
-0.7	-0.07	0.09	0.12	-0.77	-0.66	0.13	0.33	0.05	0.33	6.02	2.88
-0.6	-0.06	0.10	0.11	-0.58	-0.58	0.14	0.28	0.05	0.28	5.10	2.52
-0.5	-0.04	0.10	0.11	-0.44	-0.48	0.14	0.23	0.06	0.23	4.11	2.15
-0.4	-0.04	0.10	0.11	-0.40	-0.40	0.14	0.18	0.06	0.18	3.19	1.72
-0.3	-0.03	0.10	0.11	-0.28	-0.31	0.15	0.13	0.05	0.14	2.44	1.37
-0.2	-0.01	0.10	0.10	-0.09	-0.20	0.14	0.09	0.06	0.11	1.68	1.12
-0.1	0.00	0.09	0.09	-0.04	-0.10	0.15	0.05	0.06	0.07	0.88	0.81
0	0.00	0.09	0.09	0.05	0.00	0.15	0.01	0.06	0.06	0.14	0.64
0.1	0.00	0.08	0.08	0.06	0.10	0.15	-0.03	0.06	0.06	-0.57	0.77
0.2	0.02	0.07	0.07	0.20	0.20	0.16	-0.06	0.05	0.08	-1.09	1.09
0.3	0.01	0.07	0.07	0.18	0.30	0.16	-0.09	0.06	0.11	-1.68	1.53
0.4	0.01	0.06	0.06	0.15	0.40	0.15	-0.13	0.06	0.14	-2.28	2.19
0.5	0.01	0.06	0.06	0.12	0.50	0.16	-0.15	0.06	0.16	-2.74	2.76
0.6	0.00	0.06	0.06	0.05	0.60	0.16	-0.17	0.06	0.18	-3.08	3.23
0.7	-0.02	0.05	0.06	-0.34	0.68	0.17	-0.20	0.05	0.20	-3.58	3.61
0.8	-0.04	0.06	0.07	-0.69	0.77	0.17	-0.21	0.06	0.22	-3.90	3.25
0.9	-0.07	0.06	0.09	-1.17	0.85	0.17	-0.23	0.05	0.23	-4.17	2.58

and their standard deviations, and the Root Mean Squared Error (RMSE), which is the square root of the sum of variance and squared bias. As we can see from the table, our joint model almost always produces smaller RMSE than the naive Weibull does, except for the cases where the error correlation is close to 0. The "RMSE ratio" column shows how big the RMSE from the Weibull model is relative to the RMSE from the joint model. Values greater than 1 suggest that RMSE is bigger in the Weibull than in the joint model. Finally, the "t-ratio" column shows the result of t-test of difference between the estimates from the Weibull and the joint model. We can see that the difference between the estimates are highly statistically significant except for the case where the error correlation is 0.

Table 2 reports the results for the auxiliary parameters, $\log(p)$ and θ. For the duration dependence parameter, we report the mean, standard deviation, RMSE, and average t-value of the significance test of $\log(p)$ (the estimated value of $\log(p)$ divided by the standard error). Since the true value of $\log(p)$ is 0, the t-value should not be large enough (in absolute values) to reject the null hypothesis. We can see that our joint model never produces a t-value greater than 1.64, which is the critical value to reject the null hypothesis at the 90% confidence level. However, the naive Weibull model frequently generates t-values that are much greater than the critical value in absolute terms. In short, the Monte Carlo simulations clearly show that, when estimating models of government survival, ignoring the selection process of government formation can result in seriously mistaken inferences about the key parameters of interest.

Data and Analysis

We now turn to assessing whether such inferential problems are present in existing models of government survival. To do so, we use a recently updated version of the Martin and Stevenson (2010) data set on parliamentary governments.[11] The unit of analysis in our study is a coalition bargaining situation (or "formation opportunity"). Each formation opportunity consists of observations on the $2^p - 1$ "potential governments" that could form in that formation opportunity (where p refers to the number of legislative parties). The dependent variable in the government selection component of our joint model is a dichotomous indicator that takes a value of 1 for the potential government that formed in a formation opportunity and a value of 0 for all other potential governments in that formation opportunity. As we discussed earlier, we will use a conditional logit specification to examine government formation.

It is worth noting that the conditional logit model makes the independence of irrelevant alternatives (IIA) assumption, which means that the underlying "utilities" for different potential govern-

[11] Our data include governments in the following countries and time periods: Austria (1949–2006), Belgium (1946–2007), Denmark (1945–2011), Finland (1945–2007), France (1959–2011), Germany (1949–2011), Greece (1977–2004), Iceland (1946–2007), Ireland (1951–2011), Italy (1946–2008), Luxembourg (1945–2004), the Netherlands (1946–2006), Norway (1945–2005), Portugal (1976–2011), Spain (1977–2011), Sweden (1948–2011), and the United Kingdom (1945–2011).

ments are independent, conditional on the included covariates.[12] Given the complexity and novelty of our joint model, choosing the conditional logit specification seems to be a prudent decision (as opposed to using, for example, a mixed logit in the selection component), but it is worth clarifying a few points about that decision. First, researchers have found (through Monte Carlo experiments) that selection bias corrections based on multinomial choice models are quite good even when the IIA assumption is violated (Bourguignon and Fournier, 2007). Second, while much ink has been spilled about the IIA assumption in studies of government formation, its imposition in our model is unlikely to lead to serious problems. To understand why, one must remember a basic fact that seems to have been forgotten by the few consistently strident critics of the conditional logit model: IIA is assumed to hold *conditional on the covariates included in the model*. To the extent that one has identified and measured the most important variables in the utilities for different potential coalitions—as 50 years of work on the problem has surely done—then whatever is left in the errors should be essentially white noise.[13]

We examine the processes of government formation and survival for 432 formation opportunities in our set of democracies, encompassing a total of $95,576$ potential governments. Because only one potential government can form in a given formation opportunity, the sample consists of 432 governments with an observed (non-zero) duration time. This duration time (measured in days) is the dependent variable in the single-equation duration model and in the government survival component of our joint model.[14] Following Diermeier and Stevenson (1999, 2000), we distinguish between two types of government terminations. Specifically, some governments in our

[12] As we assume the utilities of a potential cabinet are independent, we estimate a pairwise correlation between the utility of a potential cabinet and its duration risk. As we mention in the Supplementary Materials (see footnote 6), this pairwise correlation assumption can become less innocuous when we have an independent competing risks model, as the assumption implies that each of the competing duration risks is independent from one another while they are correlated with the utility.

[13] Furthermore, the status of the published tests of the IIA assumption in this literature that have purported to show violations is quite unclear. We discuss this issue in more detail in the Supplementary Materials for this article.

[14] In the models estimated in the next section, we always treat a certain set of government duration times as right-censored. Namely, an observation is right-censored if the government was still in office as of December 31, 2011 (the cutoff date for our sample) or if it terminated for any of the following reasons: the occurrence of a regularly scheduled election, a technical resignation required for constitutional reasons, or the death of the prime minister. Had these (technical) events not occurred, the government probably would have lasted longer, which is explicitly taken into account by the censoring component of the likelihood function for an observed duration (King et al., 1990). Of our 432 governments, 89 are right-censored for these reasons.

sample ended due to the dissolution of parliament and the calling of early elections (*dissolution* terminations), while other governments ended due to their being directly replaced, with no intervening election, by an alternative administration (*replacement* terminations). As Diermeier and Stevenson argue, the effects of bargaining environment and government characteristics on survival (as well as the underlying stochastic process) may differ across the two modes of termination. For example, factors that make it more likely that a government resigns and calls new elections may have no effect (or even the reverse effect) on the likelihood that a government is replaced without elections by an alternative coalition (see also Lupia and Strom, 1995, for a theoretical rationale). To account for this possibility, Diermeier and Stevenson (1999) use a competing risks framework, in which each mode of termination is examined separately. As they found dramatic differences in the effects of the covariates and the shape of the baseline hazard, we also adopt this approach in our study. Of our 432 governments, 112 terminated due to parliamentary dissolution, and 231 terminated due to replacement without elections.[15]

Over the long history of empirical work on government survival, scholars have introduced a wide array of covariates into their models. However, following recent theoretical work, we focus only on variables that we (and other scholars) believe should have an impact on the ability of governments to survive random shocks. Thus, as independent variables in the competing risks single-equation duration model, and in the government survival component of our competing risks joint model, we include several attributes of governments and of the bargaining environment in which they exist (e.g., King et al., 1990; Warwick, 1994; Diermeier and Stevenson, 1999). The attributes of the government that scholars have commonly focused on in previous work include the government's numerical status, its ideological diversity, and the "returnability" of the parties that comprise it. The central attributes of the bargaining environment in these studies relate to the fragmentation and polarization of the legislature.

[15] The remaining governments fell due to technical reasons (see previous footnote). In our analysis of survival by termination type, all observations that experienced the other termination type are treated as right-censored. This approach presumes that the competing risks are stochastically independent. Fortunately, as Gordon (2002) demonstrates in his reanalysis of the Diermeier and Stevenson (1999) model, this assumption does not appear to be problematic.

Beginning with the government attributes, numerical status simply indicates whether the government controls a majority of legislative seats. To remain in office, governments must maintain the confidence of a parliamentary majority, which should be easier to do if they do not have to reach out for support to parties in the opposition (which have no office-related benefits at stake if the government falls). We measure numerical status with the dichotomous variable, *Minority Government*, which takes a value of 1 when government parties do *not* collectively control a parliamentary majority (making it less likely they will survive), and a value of 0 when they do. Governments should also be less likely to survive the higher the level of policy disagreement between coalition partners (Warwick, 1979, 1992, 1994). Greater policy differences imply greater compromise (for at least one party in the coalition), which can generate significant outside pressure on party leaders from voters, activists, and backbenchers to withdraw from the government. We capture the degree of government policy disagreement with a variable from Martin and Stevenson (2010), *Ideological Divisions in Coalition*, which is the range between the most distant parties in the government on the left-right socioeconomic dimension (as measured in the CMP data). The "returnability" of government parties captures the idea that premature government termination should be more likely in cases in which coalition members believe they will be able to immediately form a new government in which they are included, i.e., cases in which the opportunity costs of withdrawing from the coalition are relatively low. Following Warwick (1992, 1994), we measure this as the proportion of parties in the current government that were part of the previous government.

Turning now to the bargaining environment attributes, a fragmented legislature is one in which legislative seats are dispersed across a large number of parties. As discussed earlier, high legislative fragmentation signifies a complex bargaining environment in which there are numerous alternative governments that may be viable replacements for the incumbent. The measure of fragmentation we (and most other studies of government survival) use is the *Effective Number of Legislative Parties* index created by Laakso and Taagepera (1979). Similarly, legislative polarization, characterized by a large anti-establishment presence in the legislature, also signifies a complex bargaining environment and should thus lead to lower rates of government survival. Our *Polarization Index* variable differs somewhat from the previous measure of this concept (from Powell, 1982), which was the share of legislative seats controlled by parties assumed to be "anti-establishment" (such as communist parties or far-right nationalist parties). Our measure, rather than assuming that some parties are anti-establishment while others are not, makes use of the CMP data on anti-establishment views.[16] Specifically, following Martin and Stevenson (2001, 2010), we first calculate the anti-system presence within each potential coalition as the maximum anti-establishment saliency score for parties within the coalition and then weight this by the share of legislative seats controlled by the potential coalition. We then sum the seat-weighted anti-system scores across all potential coalitions in the bargaining situation to create a system-level measure. Finally, we include a control variable in the analysis, *Time Remaining in CIEP (Logged)*, to account for the possibility that governments that form early in the constitutional inter-election period (CIEP)—the period of time between mandatory parliamentary elections—may last longer than governments that form late in the period simply because they have a longer possible tenure.

To model the government selection component of our joint model, we use the full set of independent variables from the Martin and Stevenson (2001) and Martin and Stevenson (2010) studies of government formation. These include several variables measuring the size and ideological characteristics of potential coalitions, several variables measuring incumbency (interacted with the context in which incumbent governments previously terminated), and several variables capturing constraints on coalition bargaining. The construction these variables is described in greater detail in Martin and Stevenson (2001) and Martin and Stevenson (2010).[17]

[16] See Laver and Budge (1992) for a factor analysis of the CMP categories corresponding to the anti-establishment dimension.

[17] Before we turn to our findings, it is important to consider the issue of identification. Similar to other models of selection, our model is technically identified even when the covariates in the selection and outcome models are identical; however, this kind of identification depends only on nonlinearities in the model and not on variation in the covariates, making the results quite fragile and "model dependent." For example, when one replicates the simulations described above with the same variables in both marginal models, one can obtain estimates, but convergence is slower, and estimates of both the substantive parameters and the correlation parameter have a greater amount of error. More robust identification comes from finding variables that are good predictors of selection but that do not have an independent impact on outcomes, and vice-versa. Fortunately, in our case there are several variables that affect government formation that should not have an impact on duration. For example, while it has been argued theoretically, and shown empirically, that the party of the previous prime minister, and incumbent parties as a whole, enjoy procedural advantages that make them more likely to participate in the next government, there is little reason to think that a government comprising these parties will last a longer

Table 3: *Competing Risks Analysis of Government Survival: Models without Selection versus Models with Selection*

Explanatory Variables	**Replacement Terminations**		**Dissolution Terminations**	
	Without Selection	*With Selection*	*Without Selection*	*With Selection*
Minority Government	-0.271***	-0.201**	-0.362***	-0.325**
	(0.091)	(0.091)	(0.137)	(0.139)
Ideological Divisions in Coalition	-0.005***	-0.002	0.003	0.005
	(0.002)	(0.002)	(0.004)	(0.004)
Returnability	-0.201**	-0.359***	-0.015	-0.078
	(0.100)	(0.111)	(0.140)	(0.150)
Effective Number of Legislative Parties	-0.063**	-0.006	0.074	0.107
	(0.031)	(0.035)	(0.058)	(0.067)
Polarization Index	-0.032*	-0.022	-0.066**	-0.064**
	(0.020)	(0.020)	(0.027)	(0.028)
Time Remaining in CIEP (Logged)	0.894***	0.895***	0.752***	0.753***
	(0.065)	(0.066)	(0.117)	(0.151)
Intercept	1.304***	1.334***	2.058**	2.018*
	(0.494)	(0.505)	(0.891)	(1.155)
Duration Dependence (Logged)	0.540***	0.683***	0.488***	0.543***
	(0.057)	(0.060)	(0.082)	(0.092)
Error Correlation ($\tanh^{-1}(\theta)$)		0.310***		0.112
		(0.073)		(0.093)
Log-likelihood	-2655.72	-2646.22	-1803.16	-1802.42

Note: Cell entries are coefficient estimates (with standard errors in parentheses) expressed in the accelerated failure-time metric. All models assume a Weibull parameterization of the baseline hazard rate. Total number of government terminations: 432. Number of terminations resulting in non-electoral replacement: 231. Number of terminations resulting in early elections: 112. Number of potential governments in selection models: 95,576 (output from selection component of models with selection shown in Appendix Table 1 in the Supplementary Materials for this article). Significance levels : $*$: 10% $**$: 5% $***$: 1%.

In Table 3, we present the results from our competing risks analysis of government survival.[18] As discussed earlier, we examine separately those governments that ended in a replacement with no intervening election and those that ended in a parliamentary dissolution and the calling of early elections. For each termination mode, we present the results from the conventional single-equation duration model and the results from our joint duration model, which explicitly takes into account the process of government selection.[19] We begin by examining those governments that were replaced by an alternative government without early elections. The estimated effects from the naive survival model (i.e., the model that does not account for selection) are very much in line with the expectations and findings from previous research.[20] For example, we see that minority governments are expected to be replaced earlier than majority governments, and that governments that are more ideologically diverse are expected to be replaced earlier than governments with ideologically compatible parties. Further, when the government consists of a larger proportion of parties from the previous government, which suggests that turnover in coalition parties tends to be relatively low from one government to the next, then premature government replacement is more likely. In short, the set of government attributes highlighted in previous work also appear to be important factors in our (single-equation) analysis. We note the same pattern for the two variables measuring the complexity of the bargaining environment. A fragmented legislature (with a large effective number of parties) is more likely to lead to early cabinet replacements, as is a legislature containing numerous parties with pronounced anti-establishment views. Thus, our findings with

time in office. Meanwhile, some factors that affect government duration—such as legislative fragmentation and polarization—do not vary across potential governments in a formation opportunity, and so they cannot directly impact which potential government is selected to form.

[18] To conserve space, we show the results for the selection component of the joint models in Appendix Table 1 in the Supplementary Materials for this article. A comparison of those results to those of the government formation model of Martin and Stevenson (2010) reveals no major differences in the effects of the independent variables.

[19] The estimates are based on a Weibull parametrization of the baseline hazard. The Weibull is a simple and flexible specification that allows the baseline hazard rate to increase, decrease, or stay constant. One limitation of the Weibull specification, however, is that non-monotonic change in the hazard rate is not permitted. Therefore, we also estimate models based on a log-logistic parametrization that allows for non-monotonicity in the hazard rate. Across all four models shown in Table 3, a Weibull specification produces a better model fit than a log-logistic specification.

[20] The coefficients are expressed in the accelerated failure-time metric; thus, a negative coefficient indicates that an increase in the corresponding variable leads to a shorter expected duration while a positive coefficient indicates that an increase in the corresponding variable leads to a longer expected duration.

respect to both cabinet-level and system-level attributes should come as no surprise to researchers of government survival. We now assess whether these findings continue to hold once we correct for the problem of nonrandom government selection.[21]

Proceeding with our examination of governments that terminated due to non-electoral replacement, we first note that the estimate of the correlation between government formation and survival is positive and statistically significant.[22] This implies that the governments that parties choose to form are those that have a *lower probability of ending in early replacement* based on factors that are unaccounted for by the *observed* set of variables associated with government survival. We also see that, once the selection process is taken into account, several of the factors previously believed to shorten the life of the government no longer have a discernible effect. For example, conditional on the selection of the government, the ideological diversity of the government is no longer a significant predictor of early government replacement. Nor are the two factors associated with the complexity of the bargaining environment. The coefficient on the polarization measure falls in magnitude by almost 35% from its value in the single-equation model, while the effect of legislative fragmentation falls almost to zero. The only theoretical variables continuing to exert an impact on the likelihood of premature government replacement are the numerical status of the cabinet and the "returnability" of cabinet parties. In short, we conclude that there are inferential costs for models of government survival—at least for those examining governments that end in replacement—when they ignore the process of government selection.

For governments that terminate due to early parliamentary elections, the story is different. The estimate of the correlation between government formation and survival, while still positive, is considerably smaller than in the case of replacement terminations, and is not statistically different

21 The conditional logit model of government formation shown in Appendix Table 1 (in the Supplementary Materials for this article) is estimated first as a stand-alone model, and then as a component in our joint model for each of the competing risks.

22 The error correlation estimate shown in Table 3 is actually the inverse hyperbolic tangent of the error correlation parameter, θ—i.e., $\tanh^{-1}(\theta)$—rather than θ itself. Using this function is a standard technique to facilitate estimation of correlation parameters that are bounded by -1 (perfect negative correlation) and $+1$ (perfect positive correlation). Thus, the value of θ in the replacement terminations joint model is 0.301; in the dissolution terminations joint model, θ is 0.112.

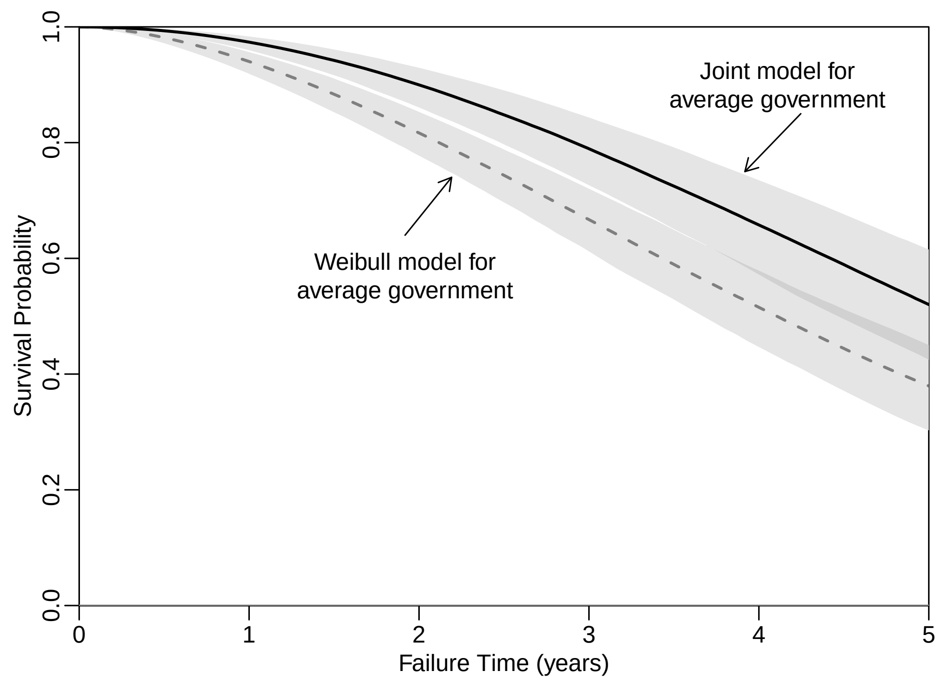

Figure 2: *Survivor Functions for Naive and Joint Models of Government Survival (Replacement Terminations)*

from zero.[23] Thus, to the extent that parties choose governments based on factors expected to lower their chances of collapsing with early parliamentary elections, these factors seem to be adequately captured by the observed variables in the survival model. Specifically, the results for both the naive and joint survival models show that minority governments are more likely than majority governments to experience early dissolution terminations, just as they are more likely to experience early replacement terminations. Further, the results suggest that governments existing in more polarized bargaining environments are more likely to terminate prematurely with a subsequent early election than governments in less polarized environments. The effects of all variables are very similar across the models with and without selection, implying that there is very little inferential cost to estimating a single-equation survival model for these types of government terminations (conditional, of course, on the set of covariates included in the model).

The overwhelming majority of governments, however, are terminated by being replaced with

[23] A likelihood ratio test between the naive and joint models for dissolution terminations reveals no statistically discernible difference ($p > 0.25$), in contrast to the case of replacement terminations, where the likelihood ratio test shows that the joint model significantly improves model fit ($p < 0.01$).

an alternative government with no intervening elections (approximately 67% of the non-technical terminations in the sample). Thus, most of the time, the inferences drawn from single-equation survival models are mistaken ones. To illustrate the severity of the bias in these models, we display in Figure 2 the survivor functions (with 95% confidence bounds) from the naive Weibull model and our joint model (for governments ending in replacement), holding all covariates at their average sample values (except for *Minority Status*, which is set at its modal value of 0, and *Time Remaining in CIEP (Logged)*, which is set at $ln(1825)$, corresponding to a maximum 5-year term).

The survivor functions indicate the probability that a government will last beyond a particular time. A comparison of the survivor functions clearly shows that an average government is expected to last longer under the joint model than under the naive model. For example, the naive model predicts that an average (majority) government facing a five-year maximum term has roughly a 65% chance of surviving past three years, whereas our joint model estimates the probability to be almost 80%. The divergence between the expected survival rates is due primarily to the attenuation that occurs in the effects of ideological divisions, legislative fragmentation, and polarization after we account for government selection. These results suggest that party leaders, when they are making the decision about which government to form, base their choice partly on unobserved characteristics that increase government survival, thereby reducing the impact of several of the destabilizing factors highlighted in previous research.

Conclusion

In this study, we have presented a new estimator, based in copula theory, that allows researchers to model the processes of government formation and survival jointly. The results clearly indicate that current models of government survival, especially those focusing on governments that end with non-electoral replacement, suffer from significant selection bias when the formation process is not taken into account. Conventional approaches significantly overstate the substantive importance of several covariates commonly included in empirical models. Ideally, as the government survival literature moves forward both theoretically and empirically, scholars will be able to identify (and measure) the various factors affecting the duration of governments that politicians take into account in coalition bargaining. In the meantime, we hope that researchers in this area will use our approach to assess, and correct for, selection bias in models of government survival.

Our estimator extends important new work in political methodology on the problem of nonrandom selection in event history analysis (e.g., Boehmke, Morey and Shannon, 2006). It incorporates a polychotomous selection component and overcomes certain limitations in previous survival models with nonrandom selection. We believe our copula-based approach could prove useful for other substantive questions beyond those explored here. For example, there have been several notable studies of the US Congress examining how legislator tenure on committees depends on factors such as party control of assignments, legislator loyalty on roll call votes, and the electoral incentives of committee members (see, e.g., Katz and Sala, 1996; Heberlig, 2003). However, whether legislators remain (or are retained) on a particular committee may be related to some of the same (unobserved) factors that got them selected onto the committee in the first place. In recent work, Cann (2008) explores the committee assignment decision using a conditional logit model. Thus, using our joint approach, one could easily model both committee tenure and committee assignment simultaneously and investigate whether nonrandom selection is a problem in this area. Another possible application is in the area of international relations, such as research on the duration of civil wars. Some scholars suggest that the length of civil wars may depend, to some extent, on the alignment pattern among multiple rebel groups (Cunningham, 2011). For example, multiple rebel groups may fight against a common enemy (i.e., the government) and, in doing so, they can choose one form of alignment (i.e., none, all allied, various combinations allied) from among many potential alignment patterns. It is possible that the choice made by rebel coalitions to continue fighting, and thus extend the war, is related to some of the same (unmeasured) factors that led them to come together. Our approach would allow researchers in this area to take such a possibility explicitly into account, and potentially overcome the associated problems of nonrandom selection.

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