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How Time Affects EU Decision-Making

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In a recent article in this journal, Golub demonstrated that, in order to understand the determinants of legislative decision-making speed in the European Union and their theoretical implications, it is essential for analysts to fit survival models that do not make assumptions about the shape of the baseline hazard rate and that account for variables whose value and effect may change over time (Golub, 2007). To improve upon earlier findings that ignored these issues, he fit a Cox model with time-interaction terms that allow non-proportional effects to data on the 1669 proposed directives made between 1968 and 1998. The inclusion of these time-interaction terms was based on the well-known tests of proportionality (Box-Steffensmeier and Zorn, 2001). Although correct, the interpretation of the results of such a model is less straightforward than he recognized. In this paper we explain how to make sense of the estimates from survival models that contain time-interaction terms, and then investigate how a more precise interpretation affects Golub's findings. Overall, we draw three main conclusions: the effect that formal qualified majority voting (QMV) rules have on speeding up decision-making is even larger and more consistent than originally claimed; the trade-off between efficiency and legitimacy is more complicated than first thought; and the effects of some key variables do not just wear off but actually reverse direction once proposals survive long periods of time.

Survival models with time-interaction terms

The assumption that the effect of covariates remains constant over time is often unjustified, and violations of the assumption can render a model's estimates and inferences meaningless. Violations can arise naturally in all covariates when their relationship with the hazard is affected by time, but by definition they arise when we have time-varying covariates (TVCs), that is, covariates where the initial coding assigned to a case changes during its lifetime (Box-Steffensmeier and Jones, 2003; Golub, 2007). The solution is to fit a model with time-interactive terms that accommodate these non-proportional effects (Box-Steffensmeier and Zorn, 2001; Box-Steffensmeier and Jones, 2004), as Golub (2007) did in his paper on decision-making speed. However, the interpretation of such a model is not straightforward, since the impact of a variable on the hazard rate is now a combination of its time-independent and time-dependent coefficients. This can be shown as follows.

Imagine that we fit a Cox model that included one TVC and that its effect changed over time. The hazard rate would then be

$$h_i(t) = \exp(b_1 x(t)_1 + b_2 \ln(t) x(t)_1 + b_3 x_2 + \ldots b_n x_{n-1}) h_0(t) \tag{1}$$

with $x(t)_1$ as the TVC. Note that, to deal with a non-proportional effect of this covariate, time can be transformed in a way that fits better to the underlying structure in the data (in this instance, by taking the natural logarithm, as is done in Golub's study). We would then use the parameter estimates from this equation to calculate the effect of a unit change in x_1 on the hazard ratio for a case that did not undergo state changes. To get the effect of a unit change in a covariate we construct a hazard ratio where the denominator is the baseline hazard rate (i.e. hypothetical case j with all covariates set to zero). So we have

$$\frac{\hat{h}_i(t)}{\hat{h}_j(t)} = \frac{\exp(\hat{b}_1 x_1 + \hat{b}_2 \ln(t) x_1 + 0)\hat{h}_0(t)}{\exp(0)\hat{h}_0(t)} = \frac{\exp(\hat{b}_1 x_1 + \hat{b}_2 \ln(t) x_1)}{1} \tag{2}$$

where $\hat{b}_1$ is a time-independent coefficient and $\hat{b}_2 \ln(t)$ is a time-dependent coefficient. The numerator at the right-hand side of this equation can be rewritten as:

$$\exp(\hat{b}_1 x_1 + \hat{b}_2 \ln(t) x_1) = \exp([\hat{b}_1 + \hat{b}_2 \ln(t)] x_1). \tag{3}$$

In other words, the impact of x_1 on the hazard rate is now the result of $\hat{b}_1$ and $\hat{b}_2 \ln(t)$. The combined term $[\hat{b}_1 + \hat{b}_2 \ln(t)]$ is the actual coefficient of the co-variate, which depends not only on the values of $\hat{b}_1$ and $\hat{b}_2$ but also on t.

Golub recognized this much, and he calculated the magnitude of the effects of certain variables at different values of t, the point after which the effects of each variable 'wore off', and the point at which the effects of one variable outweighed those of another. He did this by examining the exact value of the combined term and the point at which it equaled exactly zero. But this approach ignores the standard error of the combined term, and herein lies the problem (Steunenberg and Kaeding, 2007). This is a common mistake of those who try to interpret interactive terms and it yields misleading inferences.

The standard error of the new, combined coefficient is actually a function of the standard errors of the original coefficients, the covariance between both terms, and time. We know from the more familiar setting of interactive terms in linear regression, logit and probit models (Friedrich, 1982; Brambor et al., 2006) that the standard error of the sum ($\hat{b}_1 + \hat{b}_2 x$) is calculated as follows:

$$\text{s.e.}_{(b_1+b_2x)} = \sqrt{\text{var}(\hat{b}_1) + x^2\,\text{var}(\hat{b}_2) + 2x\,\text{cov}(\hat{b}_1, \hat{b}_2)}\ . \tag{4}$$

In the context of a non-proportional hazards model, our 'x' is some function of t, in our case $\ln(t)$. So the standard error of the sum ($b_1 + b_2\ln(t)$) is

$$\text{s.e.}_{(b_1+b_2\ln(t))} = \sqrt{\text{var}(\hat{b}_1) + (\ln(t))^2\,\text{var}(\hat{b}_2) + 2\ln(t)\,\text{cov}(\hat{b}_1, \hat{b}_2)}\ . \tag{5}$$

The value of this term, as indicated by t, also depends on time. Consequently, the size, sign and significance of the combined coefficient may vary with time. The crucial point is that the effect of a covariate is zero not when the value of the combined term [$\hat{b}_1 + \hat{b}_2\ln(t)$] is exactly 0, but when it is indistinguishable from 0, and not when the hazard ratio in equation (2) is exactly 1, but when it is indistinguishable from 1.

There are two equivalent ways in which the standard error of the combined term allows us to determine the values of t for which these situations obtain, and thus to draw more precise inferences about the effects of covariates. First, we can compute the Wald statistic, which is defined as follows:

$$W = \left(\frac{(\hat{b}_1 + \hat{b}_2\ln(t))}{\text{s.e.}_{(b_1+b_2\ln(t))}}\right)^2 = \left(\frac{(\hat{b}_1 + \hat{b}_2\ln(t))}{\sqrt{\text{var}(\hat{b}_1) + (\ln(t))^2\,\text{var}(\hat{b}_2) + 2\ln(t)\,\text{cov}(\hat{b}_1, \hat{b}_2)}}\right)^2 \tag{6}$$

for all values of t. This statistic follows a chi-square distribution with one degree of freedom and thus provides a formal test of whether the combined term is distinguishable from 0. Alternatively, for all values of t we can construct a confidence interval around the point prediction for the hazard ratio in equation (2). This hazard ratio is indistinguishable from 1 whenever the confidence interval contains the value 1.[1] Moreover, we can estimate the hazard ratio and construct confidence intervals for any linear combination of covariates, which allows us to compare the effects of one covariate with another at any point in time. Their respective effects are indistinguishable whenever their respective confidence intervals overlap.

Time interaction and EU decision-making

Having shown how the combined coefficients and standard errors of covariates in non-proportional hazards models vary with time, in this section we re-analyze Golub's data and provide a more careful interpretation of his results. Table 1 reports the original estimates from Golub's Cox model with TVCs and non-proportional effects for six covariates.

We begin by assessing the impact of these six covariates at different values of survival time for legislative proposals. Table 2 presents the results.

Table 1 Estimates for Golub's Cox model of EU decision-making speed

Variable	*b*	*s.e.*
QMV	3.122****	0.478
QMV after SEA	2.110****	0.513
QMV after Maastricht	0.413**	0.166
Cooperation procedure	–6.041****	0.614
Codecision procedure	–5.001****	0.876
EU9	0.496**	0.198
EU10	0.457*	0.243
EU12	0.659**	0.257
EU15	0.571**	0.263
Thatcher (as prime minister)	–1.716****	0.379
Expanded legislative agenda	0.177	0.191
Legislative backlog	0.026****	0.007
QMV × ln(t)	–0.428****	0.079
QMV after SEA × ln(t)	–0.224***	0.085
Cooperation × ln(t)	0.890****	0.099
Codecision × ln(t)	0.725****	0.134
Thatcher × ln(t)	0.282****	0.061
Legislative backlog × ln(t)	–0.004***	0.0009
Likelihood ratio	311.27***	
Number of cases	1669	

Notes: See Golub (2007) for a further explanation of the variables. Data are right-censored on 17 December 1999.
* $p < .10$; ** $p < .05$; *** $p < .01$; **** $p < .001$

We chose values located up to nearly one standard deviation below and one-and-a-half standard deviations above the mean survival time of 1100 days, because decision-making time is highly positively skewed, with a median of 646 days. This range, from 80 to 2900, covers 87% of the observations in Golub's data set. For each covariate we also present the combined coefficient, its standard error and the significance level based on the Wald statistic.

The table shows how the impact of our covariates changes over time. While some covariates may have a positive impact on the hazard rate at some moment in time, leading to the speeding-up of decision-making, this impact may change into a negative one later on. In addition, the combined coefficient for each covariate is insignificant for some period in time, which indicates that during this period it temporarily has no effect on the hazard. Decision-making speed is, in that period, affected only by other covariates.

We now compare our new findings with Golub's original claims for each covariate.

QMV before the Single European Act (SEA)

> Original claims: the hazard rate for proposals formally subject to QMV, as opposed to unanimity, was 146% higher after six months and 82% higher after one year, and the effect of QMV wore off after four years.

Our results confirm the first two claims, since the combined coefficient for QMV is positive and highly significant for $t = 180$ and $t = 360$ and neither of the respective confidence intervals around the hazard ratios contains one.[2] The combined coefficient is negative and statistically significant until $t = 882$, so for the 1968–87 period the positive effects of QMV on speed wore off after 29 months (when $t = 882$), not four years. Moreover, the impact of QMV reverses sign and once again becomes significant for very long decision-making processes (more than 10 years). Apart from only five exceptional cases that survived well over two standard deviations beyond the mean, this goes beyond the period for which we have observations and thus could easily be an artifact of our analysis. For this reason, we will disregard this effect.

QMV after the Single European Act

> Original claims: the effect of QMV in the period after the SEA was the same as in the pre-SEA period, and for the first five months the pre-SEA effect was actually larger than the post-SEA effect.

Our results confirm the first but not the second claim. Judging only from the combined coefficient, as Golub (2007: 168) did, the coefficients presented in

Table 2 The impact of time on covariate effects

	Time[a]										
	–0.85	–0.75	–0.5	–0.25	Mean	+0.25	+0.5	+0.75	+1.0	+1.5	st.dev.
	80	200	500	800	1100	1400	1700	2000	2300	2900	Days
QMV	1.25**	0.85**	0.46**	0.26*	0.12	0.02	–0.06	–0.13	–0.19	–0.29	$\hat{b}^t$
	0.15	0.10	0.09	0.11	0.12	0.14	0.15	0.16	0.17	0.18	s.e.
QMV after SEA	1.13**	0.92**	0.72**	0.61**	0.54**	0.49**	0.44*	0.41*	0.38	0.32	$\hat{b}^t$
	0.13	0.13	0.12	0.14	0.15	0.16	0.18	0.19	0.20	0.21	s.e.
Cooperation	–2.14**	–1.32**	–0.51**	–0.09	0.20	0.41**	0.58**	0.73**	0.85**	1.06**	$\hat{b}^t$
	0.20	0.13	0.11	0.12	0.14	0.16	0.17	0.18	0.19	0.21	s.e.
Codecision	–1.82**	–1.16**	–0.49**	–0.15	0.08	0.25	0.39*	0.51*	0.61**	0.78**	$\hat{b}^t$
	0.31	0.21	0.15	0.15	0.16	0.18	0.20	0.21	0.23	0.25	s.e.
Thatcher	–0.48**	–0.22*	0.04	0.17*	0.26**	0.33**	0.38**	0.43**	0.47**	0.53**	$\hat{b}^t$
	0.13	0.09	0.08	0.08	0.10	0.10	0.11	0.12	0.13	0.14	s.e.
Backlog	0.0086**	0.0050**	0.0013	–0.0006	–0.0018	–0.0028	–0.0036**	–0.0042**	–0.0048**	–0.0057**	$\hat{b}^t$
	0.0026	0.0019	0.0013	0.0012	0.0012	0.0012	0.0013	0.0014	0.0014	0.0016	s.e.

Note: $\hat{b}^t = \hat{b}_1 + \hat{b}_2 \ln(t)$ with $\hat{b}_1$ as the estimated time-independent coefficient for a covariate and $\hat{b}_2$ as its estimated time-dependent coefficient, and $\ln(t)$ as the natural logarithm of time; s.e. is the standard error. The significance of $\hat{b}^t$, based on the Wald statistic, which follows the chi-square distribution with 1df, is indicated by ** $p < .01$ and * $p < .05$.

[a] Survival time is first presented in terms of standard deviations away from the mean, which is about 1100 days (the median is 646 days) with a standard deviation of about 1200 days; the second row presents the corresponding time in terms of number of days.

Table 2 seem to indicate that the speeding-up effect of QMV before the SEA was greater than the effect of QMV after 1987 for at least the first 80 days. Soon after this short period the relative impact of these coefficients seems to reverse. But taking the respective standard errors into account reveals that the effects of QMV pre- and post-SEA are indistinguishable for all values of time. Figure 1 plots the respective hazard ratios and confidence intervals for QMV in the two periods, and shows that they always overlap.

Cooperation 1987–93

> Original claim: for the period 1987–93 the negative effect of the cooperation procedure on decision-making speed outweighed the positive effect of QMV after SEA during the first year of a proposal's survival time.

Our results paint a more complicated, and in one respect slightly less severe, picture of the trade-off between speed and Parliamentary involvement. The hazard ratio for QMV after the SEA and the cooperation procedure is significantly less than 1 until $t = 249$, then indistinguishable from 1 until $t = 497$, then significantly greater than 1 thereafter. So in the six years following the SEA, the drag from cooperation overwhelms QMV for first 8 months (not 12),

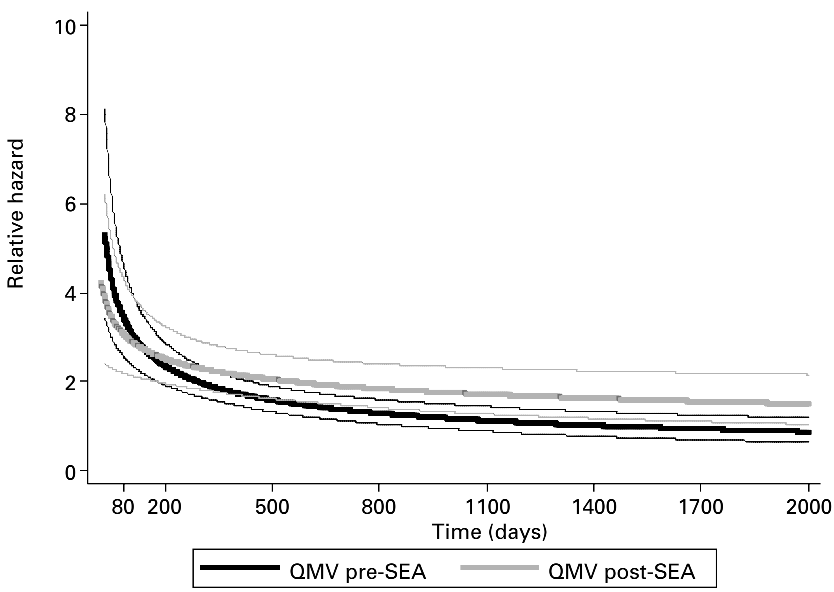

Figure 1 Effect of QMV before and after the Single European Act.
Note: The thick lines plot the relative hazard for QMV before and after the Single European Act. The thin lines are 95% confidence intervals.

then neutralizes it for the next 8 months, but the QMV effect outweighs the cooperation effect after 16 months. We can also report two new findings: the drag from cooperation wears off only after 688 days, and, more surprisingly, after 1218 days the effect of cooperation is reversed and significantly speeds decision-making. This might be a statistical artifact, though, since, although there are 67 cases in the data set that were proposed under cooperation and survived more than 1218 days, all but three of them shifted to codecision at some point. The relevant issue, then, is the effect of codecision on EU decision-making speed, which we discuss below.

QMV after the Treaty on European Union (TEU)

> Original claims: ever since the Maastricht Treaty, QMV exerted a still significant but somewhat reduced effect, there was no evidence of a unanimity norm, and for the first 18 months the effect of QMV was larger during 1968–87 than during 1994–9.

Our results confirm a significant but temporarily reduced effect of QMV compared with previous periods, show that the unanimity norm is even weaker than originally claimed, and provide a more optimistic prediction about the impact of prospective reforms to extend the scope of QMV. The effect of QMV after the Maastricht treaty (post-TEU) is not time dependent, so does not appear in Table 2. For all values of t, the hazard ratio is 1.51, with a confidence interval from 1.09 to 2.09. Thus for the entire 1994–9 period QMV was consistently faster than unanimity, so there is no sign of a unanimity norm. To compare the effects of QMV in the post-Maastricht period with its effects in earlier periods we then examined the three respective confidence intervals. These comparisons reveal that, for the first 4 months, the effect of QMV was greater during the period before Maastricht than after Maastricht, and that for only the first 5 months (not 18) the effect of QMV was greater during the period 1968–87 than during the period 1994–9. After 5 months all three confidence intervals overlap. In other words, the full effects of QMV return much more quickly than originally reported. So the unanimity norm is even weaker than Golub thought, and prospective QMV reforms would likely ease inertia more than he originally suggested.

Codecision and cooperation after Maastricht

> Original claim: the drag exerted by each of these procedures outweighs the effects of QMV for 18 months.

Again our results confirm a slightly more complicated trade-off between speed and democratic inclusiveness. The hazard ratio for proposals subject to QMV and codecision is significantly less than 1 until $t = 343$, then indistinguishable from 1 for the period between 344 and 866 days, then significantly more than 1 for $t > 866$. Thus, during the period 1994–9, the drag from codecision outweighed the effects of QMV for 11 months (not 18). The two effects balance out for the next 17 months, and QMV more than offsets the delays from codecision only after 28 months.

We can also report two important new findings. First, inspection of their respective confidence intervals shows that the effects of cooperation and codecision are indistinguishable for all values of t, so the European Parliament's growth from agenda-setter to veto player has not slowed EU decision-making. Second, the effect of codecision wears off after 671 days, then after $t = 1674$ it is reversed and significantly speeds up decisions. Overall, our results suggest that broadening participation in the legislative process by empowering the European Parliament adds nearly two extra years to a decision, but helps resolve disagreement on the especially contentious proposals that experience more than four-and-a-half years of negotiations. This result is hardly what proponents of deliberative democracy and informal norms have in mind when they claim that adding more voices to the discussion expedites the EU legislative process.

Thatcher effect

> Original claims: when Thatcher was prime minister the hazard was 22% lower after 6 months, and the 'Thatcher effect' wore off after 14 months.

Results in Table 2 show that the negative 'Thatcher effect' was slightly less than this, and, more interestingly, that it eventually reverses direction and significantly speeds the adoption of legislation. The combined coefficient for Thatcher is negative and significant until $t = 246$, indistinguishable from 0 for $t = 247$ through 793, then positive and significant for $t > 794$. Thus the hazard rate when Thatcher was prime minister was 22% lower after 6 months, as Golub originally claimed (with a confidence interval from 7% to 35%), but it wore off after 8 months (not 14), and its effect reversed after 26 months.

This curious finding deserves further study, but one possible explanation is that the reversal reflects a landmark change in Thatcher's attitude towards the EU. Thatcher became prime minister in 1979 and was highly antagonistic for nearly five years until, following an agreement at the June 1984 European Council summit at Fontainebleau, 'the British budgetary dispute was over' (Dinan, 1999: 92) and she 'got her money back'. It is certainly possible then that pre-Fontainebleau proposals were bogged down by Thatcher for much longer than 8 months, which drags down the hazard ratio, and were the ones more likely to survive 26 months and experience Thatcher's new views. So we would expect no Thatcher effect on proposals that were made after June 1984, and for those made earlier the effect should vanish if they survived past this date. This would account for the figures in Table 2, and a four-year period of intransigence would also accord well with Golub's other findings that do not derive from survival analysis – that Eurosclerosis appeared only upon Thatcher's arrival with a decline in Council output; that proposals made in 1979 and 1980 survived an unusually long time compared with those in the years immediately preceding and following; and that the volume of Council adoptions rose sharply in 1984 (Golub, 1999).

In future work, the most effective way to investigate this matter would be to construct a TVC that codes state changes for proposals under consideration before or after the Fontainebleau summit, or, better still, that reflects shifts in Thatcher's attitude. The latter could be folded into the broader TVC Golub's article suggests to capture shifts in Council heterogeneity.

Backlog

> Original claim: mounting backlog expedited Council decision-making on new proposals.

A proper interpretation of time-interaction effects yields more precise and interesting results. Table 2 shows that the combined coefficient for legislative backlog (with backlog set to its average value of 169) is positive and significant until $t = 295$, then indistinguishable from zero for $t = 296$ through 1421, then significant and negative for $t > 1421$. So, an average-sized backlog expedites decisions, in line with Golub's original claim, but only for the first 10 months. It has no discernible effect for the next three years, and then slows decision-making for proposals that have survived over 46 months. This reversal of sign shows that a large backlog spurs the Council to dispose of new proposals but does not expedite passage of the most controversial pieces of legislation.

Conclusions

Survival analysis often requires the use of models that contain time-interaction terms to deal with the non-proportional effects of certain covariates, but it is easy to misinterpret the results of such models. We have shown that the key to correct interpretation is calculating the magnitude, standard error and statistical significance of the term formed by combining a covariate's time-dependent and time-independent coefficients. In his recent study of decision-making speed in the European Union published in this journal, Golub (2007) overlooked this issue, and by addressing it we obtained more precise findings.

From our re-analysis of his data we draw three main conclusions. First, the effect that formal QMV rules have on speeding up decision-making is even larger and more consistent than originally claimed. Not only did QMV have a substantial effect as far back as the 1960s, this effect did not increase following the Single European Act. Following the Maastricht Treaty, speed under QMV was consistently faster than under unanimity, and the full effects of QMV returned much more quickly than Golub first thought. There is little sign, as often purported, that the EU legislative process operates by a unanimous consensus norm that slows down decision-making. Second, the trade-off between decision-making efficiency and legitimacy is more complicated than Golub recognized. The growth of the European Parliament's powers beyond mere consultation did significantly slow decision-making, but its evolution from an agenda-setter to a veto player added no discernible extra delay. Third, the effects of some key variables do not just wear off but actually reverse direction once proposals survive long periods of time. We suspect that some of these reversals are statistical artifacts, or, as in the case of the 'Thatcher effect', will vanish once we develop a more sophisticated TVC to capture shifts in Council heterogeneity. We attribute others, such as those for legislative backlog and codecision, with more substantive meaning.

Notes

1 The only example we know of where analysts have done this is Box-Steffensmeier and Zorn (2001: 984), and we are grateful to them for sharing their elegant Stata routine, which makes the calculations of standard errors and confidence intervals less cumbersome.

2 At $t = 180$ the hazard ratio is 2.46 with a 95% confidence interval from 2.00 to 3.03. At $t = 360$ the hazard ratio is 1.83 with a 95% confidence interval from 1.53 to 2.17.

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