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Sunsets and federal lawmaking: Evidence from the 110th Congress

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ABSTRACT

We test the hypothesis that the choice to include a sunset provision increases the likelihood that a bill becomes law. We develop a model where the legislator's knowledge of the increase in passage probability from including a sunset provision influences the legislator's choice to do so. Because legislators may either include a sunset provision to increase passage probability, or observe low passage probability and respond with a sunset provision, the choice to include a sunset provision is endogenous. Consequently, the causal effect of temporary enactment is identified by using the legislator's number of offspring as a source of exogenous variation in the choice to include a sunset provision. Employing recursive bivariate probit, we find that the average causal effect of including a sunset provision is sixty percent. We also find that the average causal effect of including sunset provisions in bills that already include them is about twenty percent.

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1. Introduction

Legislators can pass laws temporarily by including a sunset clause. A sunset clause automatically invalidates a statute, or a portion of a statute, on a specified date without further legislative action. While legislators can pass new legislation to extend a temporary law, temporary laws expire by default. By contrast, a law without a sunset clause governs in perpetuity by default. Law and economics scholarship first examined sunset clauses in the early 1980s. Noting a general increase in legislation volume beginning in the Progressive Era, Calabresi argued that sunset legislation was an inadequate legal device for updating obsolete statutes and sought to develop a judicial response (Calabresi, 1982). Thereafter, the scholarship remained relatively silent until the American legislature began passing a large number of tax cuts temporarily during the early 2000's. A number of articles appeared either condemning or supporting the practice (Garrett, 2004; Kysar, 2006; Yin, 2009) and public economics scholarship contributed theories that described the optimal length of a budget window for minimizing the socially suboptimal use of sunsets (Auerbach, 2006; Dharmapala, 2006). In addition to the scholarship on temporary taxation, several general articles have developed normative theories for choosing between temporary and permanent legislation according to how each type reveals information (Parisi et al., 2004; Gersen, 2007; Recordás, 2014).[1] Because lawmakers may update their preference for legislation length after information has been revealed post-enactment, a key assumption for all of these theories is that temporary and permanent legislation structure transaction costs differently over time.

For example, Gersen (2007) separates enactment costs from maintenance costs, where enactment costs are incurred during passage periods and maintenance costs are incurred during all other periods. The model assumes that the initial enactment costs of temporary legislation are less than the initial enactment costs of permanent legislation holding the substance of the legislation constant. On the one hand, temporary and permanent legislation face the same procedural requirements for passage specified by the Constitution or internal House and Senate rules. Still, variation in initial enactment costs is more plausible. Permanent legislation allocates enactment costs entirely to an initial period; temporary legislation allocates only a portion of enactment costs to an initial period and spreads the remaining costs over additional enactment periods.

We develop a model to test the enactment cost variation assumption. Our proxy for initial enactment costs is passage probability, where higher passage probabilities are associated with lower initial enactment costs and vice-versa. If initial enactment

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[1] Similarly, Levmore (2010) develops a normative theory for choosing between incremental versus non-incremental rules according to how each type impacts rent-seeking behavior.

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costs of temporary legislation are less than initial enactment costs of permanent legislation, then we should expect temporal restrictions to increase the likelihood that a bill becomes law. Our empirical results support the validity of the variation assumption. We demonstrate an average (positive) causal effect of including a statutory sunset on passage probability. Thus, initial enactment costs of temporary legislation are less than initial enactment costs of permanent legislation by proxy.

In order to test our hypothesis, we collect data from the 110th Congress on passage and various bill and sponsor's attributes. Employing a recursive bivariate probit model that takes into account the endogenous nature of temporary lawmaking, we find that the average causal effect of including a sunset provision is about sixty percent. We also find that that average causal effect of including a sunset provision for bills that already include temporal restrictions is about twenty percent.

Section 2.1 details our empirical strategy, Section 2.2 presents a summary of data, Section 2.3 discusses the regression results, and Section 3 concludes.

2. Empirical strategy

2.1. Recursive bivariate probit model

The outcome of interest is the likelihood that a bill becomes law. We assume that bill passage Y_i is determined by the latent index

$$Y_i = 1[X_{1i'}\beta_1 + X_{2i'}\beta_2 + \beta_0 T_i > \varepsilon_i] \tag{1}$$

where X_{ji} is the set of covariates, T_i is the legislator's use of temporal restriction, ε_i is the error term, and 1[.] is the indicator function taking the value of 1 if the statement in the brackets is true and 0 otherwise.

Legislators can "treat" their legislative proposals by including temporal restrictions. The treatment equation is given by the following:

$$T_i = 1[X_{1i'}\gamma_1 + \gamma_0 Z_i > \upsilon_i] \tag{2}$$

where X_{ji} is the set of covariates, Z_i is the instrumental variable, and υ_i is the error term.

We assume that the latent errors ε_i and υ_i have a bivariate standard joint normal distribution with correlation ρ. If $\rho = 0$, Eq. (1) can be estimated by a simple probit that yields the treatment effect, i.e., the effect of temporal restriction on passage. If $\rho \neq 0$, that is, the unobserved random determinants of passage are correlated with the unobserved random determinants of temporal restriction, then temporal restriction is said to be endogenous because legislators either (i) include temporal restrictions in their proposals to increase passage probability or (ii) observe low passage probability and then respond with temporal restrictions. In this case, joint estimation of Eqs. (1) and (2) is required.

The correlation of latent errors ε_i and υ_i may stem from the possibility that (i) there exists a causal relation due to the influence of temporal restriction on passage through the parameter β_1; (ii) passage and temporal restriction may depend on correlated observed covariates (X); and (iii) passage and temporal restriction may depend on unobserved components (error terms).

The model is identified by assuming that the instrumental variable Z_i is independent of ε_i, υ_i and X. Given normality, the model above can be estimated via maximum likelihood which yields consistent and asymptotically efficient estimates. The coefficient of T_i does not provide information about the size of the causal effect of including temporal restriction on passage however. Thus we calculate the average causal effect of temporal restriction, known as the average treatment effect (ATE). ATE shows the expected effect of temporal restriction on passage probability for a randomly drawn bill from the population:

$$\begin{aligned}\text{ATE} &= E[Y_{1i} - Y_{0i}] = E(Y_{1i} \mid T = 1) - E(Y_{1i} \mid T = 0)\\ &= E\{1[X_{1i'}\beta_1 + X_{2i'}\beta_2 + \beta_0 > \varepsilon_i] - 1[X_{1i'}\beta_1 + X_{2i'}\beta_2 > \varepsilon_i]\}\end{aligned} \tag{3}$$

where $Y_{1i} - Y_{0i}$ denotes the difference in outcomes due to treatment.

The other measure of interest is the average treatment effect on the treated (ATT). ATT shows the expected effect of temporal restriction for a randomly drawn bill only from those bills that include temporal restrictions:

$$\begin{aligned}\text{ATT} &= E[Y_{1i} - Y_{0i} \mid T = 1] = E(Y_{1i} \mid T = 1) - E(Y_{0i} \mid T = 1)\\ &= E\{1[X_{1i'}\beta_1 + X_{2i'}\beta_2 + \beta_0 > \varepsilon_i]\\ &\quad - 1[X_{1i'}\beta_1 + X_{2i'}\beta_2 > \varepsilon_i] \mid X_{1i'}\gamma_1 + \gamma_0 Z_i > \upsilon_i\}\end{aligned} \tag{4}$$

Another approach, advocated by Angrist and Pischke, is to employ an IV estimation on Eq. (1) disregarding the binary nature of the outcome and simply use Z_i to instrument T_i (Angrist and Pischke, 2009). As pointed out by Imbens and Angrist however, linear IV methods capture local average treatment effects (LATE) independent of whether the outcome variable is binary, non-negative, or continuous, but do not guarantee an accurate measurement of ATE (Imbens and Angrist, 1994). LATE can be interpreted as the average treatment effect of bills which could be introduced as temporary by changing the value of the instrumental variable, Z. For any two values, Z_0 and Z_1, of the instrument, the corresponding LATE is:

$$\text{LATE} = E[Y_{1i} - Y_{0i} \mid T_i(Z_1) = 1, T_i(Z_0) = 0] \tag{5}$$

where $T_i(Z_1)$ is the potential treatment status when the instrument takes the value of Z_1 and $T_i(Z_0)$ is the potential treatment status when the instrument takes the value of Z_0. Each distinct instrument yields a different LATE. The linear IV is a consistent estimator of LATE only. However, LATE is biased relative to ATE when the absolute value of the correlation ρ is large, and the probability of treatment or passage is far from 1/2 (Chiburis et al., 2012). Another difference between IV and bivariate probit estimates arises when the sample size is below 5000 as it is here (see Chiburis et al. (2012) for a discussion).

Our choice of instrument Z_i for temporal restriction is the legislative sponsor's number of offspring. The intuition behind using offspring as an instrument is that there exists a relationship between a legislative sponsor's desire to include temporal restrictions and the number of offspring that legislator has bred. There is reason to expect that the relationship is negative, that is, increasing offspring causes decreasing inclusions of temporal restrictions because legislators may desire to leave a statutory legacy to their offspring through sponsoring permanent legislation (cf. Auerbach (2006), which constructs a model where two political groups care fully about their heirs and therefore compete for long-term legislative resources). As a result, they are less likely to sponsor temporary legislation as the number of their offspring increases.

Legislator's offspring have been used in econometric studies on other occasions. For example, Washington (2008) finds that a legislator's offspring sex mix determines the propensity to vote in favor of abortion rights. Specifically, as the legislator has more females, the legislator is more likely to vote in favor of abortion rights. Similarly, Conley and McCabe (2008) use offspring sex mix as an instrument to address whether political contributions by interest groups to legislators are used as rewards or as punishments.

The legislator's number of offspring must be unrelated to unmeasured determinants of passage in order to be valid. This essentially means that number of offspring cannot have any relationship with passage probability that the specification does not