E C O N O M I C S   B U L L E T I N

A note on death penalty executions and business cycles in U.S. federal states: Is there any nexus?

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Abstract

Only legal facts should affect the application of the death penalty. This maxim is manifested in the U.S. constitution and stipulated by the Supreme Court. With regard to extra−legal economic factors, it has been applied to practice from 1950 to 1990 with the exception of two states: For Louisiana and Utah a significant countercyclical enforcement is found. A marginal increase of the economic conditions in these two states significantly reduces the probability of the performance of executions in a certain year.

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1. Introduction

When prisoners on death row in the U.S. have exhausted all their appeals, their only chance of survival rests on having their death sentences commuted. If justice in such cases is truly fair, then only the facts of the crime should affect an inmate's chances of receiving clemency.[1] Following this argumentation, we would expect no systematic relationship between the number of performed executions and economic activity across states over time, i.e. *pre-* and *post-Furman*.[2] In contrast to studies on gubernatorial election cycles and the timing of executions like Pridemore (2000) and Kubik and Moran (2003), the present note isolates and examines this hypothesis. A timing of executions independent from any budgetary politics (cf. Poterba, 1994)[3] implies no significant relationship between the cyclical component of real p.c. state income and the annual number of performed executions. The task of the present study is to scrutinize this relationship for the 39 states that imposed and enforced the death penalty during the 1950-90 period.

2. Data and Methodology

For the 1950-90 period the number of performed executions per state and year is compiled from the *ESPY* file provided by DPIC (2003). To capture the transitory economic conditions of a state in a certain year the business cycle component of personal disposable real income (p.c.) for each of the 39 states is chosen. In the context of quantifying the role of gubernatorial election year politics in the decision to execute, similar series are used as regressors; see Pridemore (2000) and Kubik and

[1] However, a recent strand of applied literature raises concerns about a disparate treatment of defendants based on racial, political, and other extra-legal factors; see Langbein (1999), Pridemore (2000), and Kubik and Moran (2003).

[2] In *Furman v. Georgia*, 1972, all executions were halted and hundreds of inmates had their death sentences lifted by a Supreme Court decision. By 1976 the Court ruled that states could again impose the death penalty provided that its application was neither arbitrary nor discriminatory.

[3] Figures of direct and indirect costs of crime are given in Levitt (1997) and NASBO (2003).

Moran (2003).[4] The state income series are taken from the BEA *Regional Economic Accounts*. As is common practice, the national CPI taken from the IMF *International Financial Statistics* has been used to deflate them (base year: 1990). Contrary to existing studies, the assessment of the relationship between executions and transitory economic conditions follows the methodology of applied business cycle research that has been established since the seminal study by Kydland and Prescott (1990). The cornerstones of this descriptive approach are (i) HP filtering to extract business cycle components (see Hodrick and Prescott, 1997) along with (ii) a comprehensive correlation analysis, investigating the lead-lag relationships of the series at stake. With regard to (i), an HP filter with a smoothing weight $\lambda = 6.25$, as recently suggested for series of annual frequency by Ravn and Uhlig (2003), is used to extract the state business cycle components (*SBC*) from the raw series.

3. Findings

3.1 Correlation Analysis. In 95% of all reported correlations of executions (*EXC*) and *SBC* the respective coefficient is insignificant (Tab. 1). However, for LA, UT, and WA a significant countercyclical relationship is found.[5] The strongest correlation in terms of magnitude and significance is found for LA. UT and WA show only a weakly significant contemporaneous correlation of *EXC* and *SBC*. For LA and UT a significant correlation is also found at first as well as second lead and at first lag, respectively. Furthermore, LA and UT differ from the case of WA, in so far as in the state of WA *EXC* were performed up to 1964 only, while in LA and UT the application of the death penalty continued in the *post-Furman* period. Up to 1964 (*pre-Furman*), however, the contemporaneous correlation coefficient for WA is significant, negative and equals 64.13%. For UT the contemporaneous correlation for the subperiods 1950-71 and 1976-90 is insignificant. Nevertheless, its correlation coefficient for the first lag in the *SBC* series remains high in magnitude and significance terms equalling

[4] Similarly, these authors consider idiosyncratic trend and cycle components across states. However, the following analysis abstracts from seeking to explain data of events on the sentencing stage or the duration on death row before the transition to either a commutation or an execution. The latter seem to be of more relevance in the context of election cycles.

[5] The weakly significant scattered correlation coefficients for IL and IN (NV and NC) at first and second lag (lead) of the respective *SBC* seem to be a product of chance rather than a robust result.

0.624* for 1950-71 and 0.300*** for 1976-90, respectively. The same holds for LA and its contemporaneous correlation of *EXC* and *SBC* equalling -0.527^{*} for 1950-71 and -0.281^{**} for 1976-90. The correlation analysis revealed that for the 39 U.S. states that enforced the death penalty from 1950 to 1990 the respective *SBC* is not associated with the timing of *EXC* apart from two notable exceptions: LA and UT.

3.2 The Case of Louisiana and Utah. This section raises the question whether and by how much a marginal increase of the transitory state-wide economic conditions (as measured by the *SBC*) reduces the probability of the performance of one or more executions in the states of LA and UT for a certain year. An answer can be given by estimating a qualitative response model. For this purpose the respective *EXC* series for LA and UT is transformed into a binary variable by encoding each observation ≥ 1 with a value of 1. The time series of the two states are pooled and state specific fixed-effects estimated: The resulting sample contains $N = 82$ pooled annual observations; 23 show a positive number of executions with 15 cases (i.e. 9 *pre-* and 6 *post-Furman*) referring to LA and 8 (i.e. 5 *pre-* and 3 *post-Furman*) to UT. The specification of choice is a standard binary choice model:

$$EXC_{i,t} = \begin{cases} 1 \text{ for } V_{i,t}^{*} > 0 \\ 0 \text{ for } V_{i,t}^{*} \leq 0 \end{cases}, \text{ where } V_{i,t}^{*} = \beta' X_{i,t} = \beta_{0,i} + \beta_1 SBC_{i,t} + \varepsilon_{i,t}, \tag{1}$$

where $V_{i,t}^{*}$ represents the latent, $\varepsilon_{i,t}$ an i.i.d. normal error, and the likelihood function is given by

$$L = \Pi_{i=1,t=1}^{N,2T} F\left(\beta' X_{i,t}\right)^{EXC_{i,t}} \left[1 - F\left(\beta' X_{i,t}\right)\right]^{1-EXC_{i,t}}.$$

For the following estimates a cumulated normal distribution for F, i.e. a Probit model framework, is chosen.[6]

Marginal effects are calculated, using the mean of the fitted value of the index $\widehat{\beta}' X_{i,t}$, according to

$$\frac{\partial E\left(EXC_{i,t} \left| \beta, X_{i,t}\right.\right)}{\partial X_{i,t}} = f\left(-\widehat{\beta}' X_{i,t}\right) \widehat{\beta}.$$

Estimates of this model are reported in the first row (S1) of Tab. 2. Normalized *SBC* refers to the estimate for which the exogenous series was transformed to have a zero

[6]Results remain unchanged or change only marginally for a Logit specification.

mean and a standard deviation of one. All estimated effects are significant, at least, at the 10% level of significance. Accordingly, an increase of the *SBC* by 1000 US $ (in prices of 1990) reduces the probability of one or more executions in a certain year by about 4% in either of the two considered states - a one unit increase of the normalized *SBC* by an even higher 10%. Specification S2 considering the *SBC* series as a regressor as well at its first lag confirms these figures.[7] Additionally, it lends support to the view that the enactment of *EXC* is following a peaking of the *SBC* with a lag of one year. Astonishingly, a *post-Furman* dummy in neither specification showed a significant impact on the probability of annual *EXC*-performances conditional on the *SBC* series. In the case of LA and UT, this calls the compliance with the 1976 Supreme Court's ruling in favor of a timing of executions across states solely based on legal factors into question.

4. Conclusion

This study examined how the performance of executions in U.S. states relates and responds to transitory changes in economic conditions as measured by the business cycle component in real p.c. state income. The question raised in its title can not unequivocally be answered with 'no:' For LA and UT, a significant countercyclical enforcement is revealed. Probit model estimates imply that a marginal increase of the transitory economic conditions in these two states reduces the probability of the enactment of one or more executions in a certain year by 3.5-10%. However, as warned by Ehrlich and Brower (1987), these findings remain empirical associations in the context of social sciences and should be interpreted with the adequate caution.

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[7] Although, the negative marginal effect of a one thousand $ increase in *SBC* slightly diminishes to 3.4%.

Tab. 1. Correlation analysis: Executions and state business cycle components

State: $EXC_t \rightarrow$	SBC_t	SBC_{t-1}	SBC_{t-2}	SBC_{t+1}	SBC_{t+2}
Alaska (AK)	$2e^{-18}$ (0.02)	−.001 (0.02)	−.003 (0.01)	$2e^{-18}$ (0.01)	$2e^{-18}$ (0.01)
Alabama (AL)	0.081 (0.08)	−.127 (0.11)	−.189 (0.08)	0.018 (0.17)	0.069 (0.14)
Arkansas (AR)	−.062 (0.08)	0.001 (0.06)	−.228 (0.20)	−.051 (0.07)	0.104 (0.14)
Arizona (AZ)	−.004 (0.13)	0.011 (0.18)	−.029 (0.10)	−.057 (0.11)	−.137 (0.11)
California (CA)	0.076 (0.16)	−.007 (0.10)	−.085 (0.12)	0.049 (0.14)	0.092 (0.15)
Colorado (CO)	−.184 (0.47)	0.324 (0.36)	0.406 (0.58)	−.374 (0.37)	−.051 (0.24)
Conneticut (CT)	0.302 (0.25)	0.164 (0.23)	−.209 (0.15)	−.157 (0.11)	−.099 (0.20)
Washington D.C. (DC)	−.196 (0.12)	−.199 (0.14)	0.095 (0.12)	0.029 (0.12)	0.445 (0.35)
Florida (FL)	0.092 (0.11)	0.038 (0.15)	−.097 (0.12)	−.087 (0.11)	0.233 (0.18)
Georgia (GA)	−.124 (0.13)	0.018 (0.10)	0.076 (0.16)	0.048 (0.14)	0.099 (0.10)
Iowa (IA)	0.052 (0.11)	0.099 (0.15)	−.067 (0.08)	−.023 (0.03)	0.180 (0.14)
Idaho (ID)	−.217 (0.15)	−.048 (0.06)	0.144 (0.15)	−.174 (0.12)	0.039 (0.08)
Illinois (IL)	−.066 (0.09)	−.126 (0.11)	−.104* (0.06)	−.011 (0.08)	0.210 (0.23)
Indiana (IN)	−.109 (0.09)	−.170* (0.09)	0.123 (0.12)	0.219 (0.24)	−.066 (0.12)
Kansas (KS)	0.179 (0.21)	0.001 (0.12)	−.007 (0.09)	0.091 (0.11)	−.070 (0.12)
Kentucky (KY)	0.165 (0.19)	0.290 (0.29)	0.058 (0.10)	−.174 (0.14)	−.139 (0.22)
Louisiana (LA)	−.36*** (0.11)	0.325 (0.24)	0.175 (0.12)	−.196** (0.09)	0.249** (0.09)
Maryland (MD)	0.246 (0.20)	−.169 (0.17)	−.094 (0.11)	0.045 (0.18)	0.089 (0.13)
Missouri (MO)	−.082 (0.05)	−.063 (0.10)	−.001 (0.13)	0.047 (0.14)	0.163 (0.20)
Mississippi (MS)	0.050 (0.15)	0.309 (0.22)	0.123 (0.17)	−.194 (0.16)	−.294 (0.20)

Tab. 1 (continued). Correlation analysis: Executions and state business cycles

State: $EXC_t \rightarrow$	SBC_t	SBC_{t-1}	SBC_{t-2}	SBC_{t+1}	SBC_{t+2}
North Carolina (NC)	−.040 (0.09)	−.060 (0.07)	−.070 (0.07)	−.010 (0.14)	0.250* (0.14)
Nebraska (NE)	0.036 (0.13)	−.209 (0.15)	−.154 (0.15)	−.116 (0.10)	0.306 (0.24)
New Jersey (NJ)	0.296 (0.28)	0.278 (0.25)	−.027 (0.04)	−.119 (0.09)	−.145 (0.22)
New Mexico (NM)	0.116 (0.21)	0.170 (0.14)	−.049 (0.12)	0.213 (0.22)	−.151 (0.16)
Nevada (NV)	0.053 (0.17)	−.131 (0.09)	−.146 (0.12)	0.404** (0.18)	−.032 (0.07)
New York (NY)	0.051 (0.18)	0.079 (0.13)	−.070 (0.12)	0.093 (0.18)	−.028 (0.14)
Ohio (OH)	−.149 (0.13)	−.073 (0.11)	−.078 (0.09)	0.142 (0.16)	0.029 (0.17)
Oklahoma (OK)	−.271 (0.16)	0.116 (0.14)	0.024 (0.16)	−.045 (0.10)	0.142 (0.13)
Oregon (OR)	0.020 (0.05)	−.064 (0.11)	−.163 (0.13)	0.274 (0.26)	0.346 (0.30)
Pennsylvania (PA)	0.307 (0.25)	0.134 (0.18)	−.120 (0.07)	−.038 (0.13)	−.107 (0.15)
South Carolina (SC)	0.190 (0.14)	0.048 (0.17)	0.114 (0.14)	0.046 (0.11)	−.092 (0.14)
Tennessee (TN)	0.155 (0.20)	0.157 (0.20)	0.014 (0.09)	−.206 (0.16)	−.105 (0.16)
Texas (TX)	0.112 (0.13)	0.085 (0.07)	−.045 (0.09)	0.044 (0.10)	−.019 (0.09)
Utah (UT)	−.205* (0.11)	0.43*** (0.14)	0.199 (0.14)	−.209 (0.15)	−.251 (0.16)
Virginia (VA)	−.067 (0.09)	−.126 (0.09)	−.013 (0.04)	0.071 (0.10)	−.050 (0.05)
Vermont (VT)	0.274 (0.27)	−.029 (0.03)	−.102 (0.10)	0.296 (0.29)	−.135 (0.13)
Washington (WA)	−.262* (0.14)	0.084 (0.10)	0.096 (0.16)	0.036 (0.16)	0.210 (0.17)
West Virginia (WV)	−.024 (0.05)	−.038 (0.04)	−.009 (0.03)	0.049 (0.09)	−.100 (0.10)
Wyoming (WY)	0.008 (0.02)	0.035 (0.05)	−.040 (0.04)	−.093 (0.09)	0.041 (0.05)

Note: *, **, *** denotes significance at 10, 5, 1% level of significance;

Newey-West HAC standard errors given in parentheses.

Tab. 2. Two-state pooled Probit model estimation: Dependent is *EXC* (binary)

Specification	State-effects	Marginal effects	
		Non-normalized *SBC*	Normalized *SBC*
S1	LA: -0.360^{***}	SBC_t: -0.039^{*}	SBC_t: -0.100^{*}
	UT: -0.898^{***}		
		Non-normalized *SBC*	Normalized *SBC*
S2	LA: -0.405^{***}	SBC_t: -0.034^{*}	SBC_t: -0.103^{*}
	UT: -0.957^{***}	SBC_{t-1}: 0.032^{**}	SBC_{t-1}: 0.095^{**}

Fit	S1	S2
McFadden R^2	0.066	0.115
log L	−45.4	−41.9
LR statistic	6.458	10.90

Note: *, **, *** denote effects corresponding to coefficient estimates significant at the 10, 5, 1% level of significance, according to standard errors based on GLM robust covariances.

For S2 the lag length is chosen on the base of the Schwarz Criterion.

The non-normalized *SBC*-series are given in thousands of US 1990 $.

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