Appendix to accompany

*Rethinking the Conflict "Resource Curse":*
*How Oil Wealth Prevents Center-Seeking Civil Wars*

## Contents

# A Theoretical Appendix

Section A.1 formally defines the equilibrium concept. Section A.2 proves all the results from the article. Section A.3 solves a modification of the model presented in the article. In the modified model, $G$ sets a state-dependent military spending schedule prior to the first period of the game, and it is assumed that any future government can commit to this schedule. Comparing the modified model to the original model explains why a non-obvious source of inefficiency arises in the original model, specifically, why it is possible for low enough $R$ that $m_s^*$ has an interior solution but $\frac{d\sigma}{dO} > 0$ even if oil does not negatively impact bureaucratic capacity.

## A.1 Equilibrium Concept

The analysis below solves for the conditions under which a Markov Perfect Equilibrium (MPE) exists that does not involve fighting along the equilibrium path of play. An MPE requires players to choose best responses to each other, with strategies predicated upon the state of the world and on actions within the current period. There is one state variable that denotes whether $C$ is strong ($\mu_t = \mu^s$) or weak ($\mu_t = \mu^w$) in the current period. $G$ chooses $m_t$ and $x_t$ as a function of the state of the world. Formally, $m_t : \{\mu^s, \mu^w\} \to [0, R]$. The equilibrium strong-period armament amount is denoted $m_s^*$ whereas equilibrium weak-period spending is $m_w^*$. Additionally, $x_t : \{\mu^s, \mu^w\} \to [0, R]$. The equilibrium strong-period patronage offer is $x_s^*$ and the equilibrium weak-period offer is $x_w^*$. $C$ chooses whether or not to accept $G$'s offer as a function of the military spending amount, patronage offer, and state of the world. Formally, $\alpha_t : \{\mu^s, \mu^w\} \times [0, R]^2 \to \{0, 1\}$, and 1 denotes acceptance whereas 0 denotes fighting. The equilibrium strong-period acceptance decision is $\alpha_s^*$ and the equilibrium weak-period decision is $\alpha_w^*$. A strategy profile $(m_s^*, m_w^*, x_s^*, x_w^*, \alpha_s^*, \alpha_w^*)$ is an MPE if there is no information set at which a single deviation from the strategy profile yields a strictly higher utility than what would be obtained from following the proposed strategy profile. By definition, a peaceful equilibrium requires $\alpha_s^* = \alpha_w^* = 1$.

## A.2 Proving Results from the Article

The following analysis proves all the lemmas and propositions from the article. I also prove nine auxiliary lemmas that ease the exposition of the proofs for the main results, as well as present four additional assumptions about how revenues affect $\theta$.

Assumption A1 imposes an upper bound on $\frac{d\theta}{dR}$ that, as the proof for Lemma 1 demonstrates, is sufficient to establish that $G$'s equilibrium total strong-period expenditures if $m_s^* = 0$ is a strictly increasing function of $R$, which is consistent with the prize argument.

**Assumption A1.** $\frac{d\theta}{dR} < \frac{(1-\delta)\theta+\delta\sigma(1+\theta)}{[1-\delta(1-\sigma)]R}$

***Proof of Lemma 1.***

- *C1 always binds.* C1 from Equation 2 always binds because, for reasons discussed in the article, $G$ always optimally makes $C$ indifferent between accepting or fighting in a strong period.

- *Optimal choices when only C1 binds.* If only C1 binds, after substituting $x_s' = \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \cdot \frac{1+m_s^*}{1+m_s'}$ into the objective function, the first derivative is $-1+\frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \cdot \frac{1+m_s^*}{(1+m_t)^2}$. Because $m_t$ enters only as a strictly positive term in a denominator in this derivative, the second derivative is strictly negative. This implies the optimal solution to the program is a unique maximizer. Solving for the first order condition as well as substituting in $m_t = m_s^*$ implicitly characterizes $m_s^*$:

$$1 = \frac{\delta(R-\sigma m_s^*)}{(1+m_s^*)\big[(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)\big]}. \tag{A1}$$

  The explicit solution is $m_s^* = \frac{\sqrt{\delta[4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2]}-\delta\sigma(2+\theta)}{2(1-\delta)\theta} - 1$.

  Substituting $m_t = m_s^*$ (which is a function only of parameters) into Equation 1 yields a unique solution:

$$x_s^* = \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \tag{A2}$$

  If C4 from Equation 2 does not bind, then $R > \sigma m_s^*$. Therefore, Lemma 1's assumption that C4 does not bind implies $x_s^* > 0$, and therefore C2 from Equation 2 never binds.

- *Optimal choices if C3 also binds.* C3 binds if $R$ is below a threshold value $R_{m_s^*=0}$. After substituting $m_t = m_s^*$ into the first derivative of the objective function solved above, it is clear that this function is strictly decreasing in $m_s^*$. Therefore, if the first derivative is strictly negative even if $m_s^* = 0$, then an implicit solution characterized by Equation A1 does not exist and the program has a corner solution with $m_s^* = 0$. Substituting $m_t = m_s^* = 0$ into $-1 + \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \cdot \frac{1+m_s^*}{(1+m_t)^2}$ demonstrates that C3 binds if $\frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)} \leq 1$ and there is an interior solution if $\frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)} > 1$. Define $G(R) \equiv \frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)} - 1$. Applying the intermediate value theorem establishes the existence of at least one $R_{m_s^*=0} \in (0, \tilde{R})$ (for $\tilde{R}$ defined below) that satisfies $G(R_{m_s^*=0}) = 0$.

  ◦ $G(0) = -1 < 0$.

◦ A sufficient condition to establish the upper bound is that $G(R)$ converges monotonically to $\infty$. This is true because $\lim_{R\to\infty} G(R) = \infty$ and because Assumption A1 ensures $G'(R) = \frac{\delta}{(1-\delta)\theta+\delta\sigma(1+\theta)} - \frac{\delta[1-\delta(1-\sigma)]R}{[(1-\delta)\theta+\delta\sigma(1+\theta)]^2} \cdot \frac{d\theta}{dR}$ is strictly positive at all $R$. Therefore, a finite $\tilde{R}$ exists such that $G(\tilde{R}) > 0$.

◦ The continuity of $\theta(R)$ in $R$ implies $G(R)$ is continuous in $R$.

Because $G(R)$ strictly increases in $R$, $R_{m_s^*=0}$ is unique and creates a threshold such that C3 binds if $R < R_{m_s^*=0}$ and C3 does not bind if $R > R_{m_s^*=0}$.

- *Optimal choices a function of R.*

◦ If $R < R_{m_s^*=0}$, then $(m_s^*, x_s^*) = \left(0, \frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)}\right)$.

◦ If $R > R_{m_s^*=0}$, then $(m_s^*, x_s^*)$ is implicitly characterized by Equations A1 and A2. ■

***Proof of Lemma 2.*** If there exists an allocation $(m_s^*, x_s^*)$ that would induce peace but $G$ deviates to an allocation $(m_s', x_s')$ that $C$ would reject, this deviation yields lifetime expected utility of $\delta\left[\frac{m_s'}{1+m_s'}V^G + \frac{1}{1+m_s'}V^C\right]$ for $G$. To demonstrate any such deviation is not profitable, need to show $R - m_s^* - x_s^* + \delta V^G \geq \delta\left[\frac{m_s'}{1+m_s'}V^G + \frac{1}{1+m_s'}V^C\right]$ for any $(m_s', x_s')$ that $C$ will reject, which simplifies to $R - m_s^* - x_s^* + \frac{\delta}{1+m_s'}\left(V^G - V^C\right) \geq 0$. Because I am examining cases in which C4 does not bind, $R - m_s^* - x_s^* > 0$. Therefore, it suffices to demonstrate that $V^G \geq V^C$. $V^G - V^C = R - \sigma m_s^* - \sigma(1+\theta)x_s^*$. Substituting in $x_s^* = \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)}$ and simplifying yields $\frac{(1-\delta)(1+m_s^*)\theta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)}$, which is strictly positive. ■

The following statement disaggregates Lemma 3 into two different cases.

**Lemma 3.** If $C$ is sufficiently patient, then there exists a unique value of $\sigma$, denoted as $\overline{\sigma}$, such that $G$ will not be able to buy off $C$ in a strong period if $\sigma < \overline{\sigma}$. If $C$ is not sufficiently patient, then $G$ will able to buy off $C$ in a strong period regardless of other parameter values. Formally, there are two cases.

**Case 1.** Assume $R < R_{m_s^*=0}$. If $\delta > \underline{\delta}_1$, for $\delta > \underline{\delta}_1$ defined in the proof, then there exists a unique value $\overline{\sigma} = 1 - \frac{\theta}{\delta(1+\theta)}$ such that $m_s^* + x_s^* > R$ if $\sigma < \overline{\sigma}$, and $m_s^* + x_s^* < R$ if $\sigma > \overline{\sigma}$. If $\delta < \underline{\delta}_1$, then $m_s^* + x_s^* < R$ for all $\sigma$ and $\overline{\sigma} = 0$.

**Case 2.** Assume $R > R_{m_s^*=0}$. If $\delta > \underline{\delta}_2$, for $\delta > \underline{\delta}_2$ defined in the proof, then there exists a unique value $\overline{\sigma}$ such that $m_s^* + x_s^* > R$ if $\sigma < \overline{\sigma}$, and $m_s^* + x_s^* < R$ if $\sigma > \overline{\sigma}$. This threshold $\overline{\sigma}$ is implicitly defined as $m_s^*(\overline{\sigma}) + x_s^*(\overline{\sigma}) = R$, for $m_s^*$ and $x_s^*$ defined in Lemma 1. If $\delta < \underline{\delta}_2$, then $m_s^* + x_s^* < R$ for all $\sigma$ and $\overline{\sigma} = 0$.

***Proof.*** In Case 1, total strong-period expenditures consist solely of the patronage offer and equal $x_s^*(0) \equiv \frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)}$, whereas total strong-period expenditures in Case 2 are $m_s^* + x_s^*$. Because $\frac{d}{d\sigma}\big(x_s^*(0)\big) < 0$ and $\frac{d}{d\sigma}(m_s^* + x_s^*) < 0$ (the first is obvious because $\sigma$ enters $x_s^*(0)$ only a positive term in the denominator; Lemma A1 proves the latter claim), in both cases total strong-period expenditures reach their upper bound when $\sigma = 0$, i.e., $C$ is strong in the current period but will never be strong again in the future. In both cases I will show that if $\delta$ is sufficiently low, then $G$ can buy off $C$ even if $\sigma = 0$. That is, $C$ is impatient enough that it prefers to accept the offer and consume in the current period rather than fight, even though $C$ foregoes consumption in all future periods because it will never be strong again. Then I will establish the threshold claim for $\sigma$ for higher values of $\delta$.

*Case 1.* Substituting $\sigma = 0$ into $x_s^*(0)$ and setting less than $R$ demonstrates that $x_s^*(0)\big|_{\sigma=0} < R$ if $\delta < \underline{\delta}_1 \equiv \frac{1}{1+\theta}$. If instead $\delta > \underline{\delta}_1$, then $\overline{\sigma} = 1 - \frac{\theta}{\delta(1+\theta)}$ is unique and creates a threshold. If $\sigma > \overline{\sigma}$, then $\frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)} < R$ and fighting does not occur; if $\sigma < \overline{\sigma}$, then $\frac{\delta R}{(1-\delta)\theta+\delta\sigma(1+\theta)} > R$ and fighting occurs.

*Case 2.* Substituting $\sigma = 0$ into Equation A1, solving for $m_s^*$, substituting that term into Equation A2, and calculating $m_s^* + x_s^*$ demonstrates $G$ must spend a total of $2\sqrt{\frac{\delta R}{(1-\delta)\theta}} - 1$ to buy off $C$ when $\sigma = 0$. This amount is strictly less than $R$ if $\delta < \underline{\delta}_2 \equiv \frac{(R+1)^2}{\theta(R^2+1)+4R(1+\theta)}$. If instead $\delta > \underline{\delta}_2$, then applying the intermediate value theorem establishes there exists at least one $\overline{\sigma}$ such that $m_s^*(\overline{\sigma}) + x_s^*(\overline{\sigma}) = R$.

- By definition of $\underline{\delta}_2$, $m_s^* + x_s^* > R$ if $\sigma = 0$ and $\delta > \underline{\delta}_2$.
- The following shows $m_s^* + x_s^* = m_s^* + \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} < R$ if $\sigma = 1$ and $\delta > \underline{\delta}_2$. Substituting $\sigma = 1$ into the inequality and rearranging yields $\big[(1-\delta)(1+m_s^*) + \delta\big]\theta > 0$, a true statement.
- Because the implicit characterizations of both $x_s^*$ and $m_s^*$ are each continuous functions of $\sigma$, their sum is also continuous in $\sigma$.

Establishing the threshold claim, because $m_s^* + x_s^*$ is a strictly decreasing function of $\sigma$, $\overline{\sigma}$ is unique. Furthermore, $m_s^* + x_s^* < R$ if $\sigma > \overline{\sigma}$ and $m_s^* + x_s^* > R$ if $\sigma < \overline{\sigma}$. ■

***Proof of Proposition 1.*** For $\sigma > \overline{\sigma}$, demonstrating the stated strategy profile is *an* equilibrium requires verifying $G$ does not have a profitable deviation at either of its two information sets, nor does $C$ at any of its infinite information sets (since its strategy specifies a best-response to any possible armament/patronage pair chosen by $G$ in both types of periods). Demonstrating the stated strategy profile is the *unique* equilibrium requires demonstrating that for any other strategy profile there exists a profitable deviation at least at one information set.

1. $\alpha_w^*(m_t, x_t) = 1$, where $m_t$ and $x_t$ compose $G$'s current-period allocation, is always consistent with equilibrium behavior. For any set of equilibrium strategies that determine $V^C$, $C$'s lifetime expected utility from accepting an offer in a weak period is $\theta x_t + \delta V^C$. Because $C$ wins a center-seeking civil war with probability 0, the lifetime expected utility to rejecting any offer is $\delta V^C$. Because $x_t \geq 0$, this is not a profitable deviation.

   $C$ also does not have a profitable deviation from $\alpha_w^*(m_t, 0) = 0$, because deviating to accepting $x_t = 0$ is not strictly profitable.

2. Any equilibrium features:

$$\alpha_s^*(m_t, x_t) = \begin{cases} 1 & \text{if } \theta x_t + \delta V^C \geq \delta\left[\frac{1}{1+m_t}V^G + \frac{m_t}{1+m_t}V^C\right] \\ 0 & \text{otherwise} \end{cases}$$

   Because $C$'s expected utility to accepting is $\theta x_t + \delta V^C$ and to fighting is $\delta\left[\frac{1}{1+m_t}V^G + \frac{m_t}{1+m_t}V^C\right]$, $C$ cannot reject any offer $\theta x_t + \delta V^C \geq \delta\left[\frac{1}{1+m_t}V^G + \frac{m_t}{1+m_t}V^C\right]$ because this deviation yields either the same or strictly less utility than accepting. Similarly, $C$ cannot accept any offer $\theta x_t + \delta V^C < \delta\left[\frac{1}{1+m_t}V^G + \frac{m_t}{1+m_t}V^C\right]$ because this deviation yields strictly less utility than fighting.

3. Because $C$'s probability of winning is not affected by $m_t$ in a weak period, $m_w^* = 0$ is the unique optimal military spending choice.

   If the equilibrium features $\alpha_w^*(m_t, x_t) = 1$ for all $(m_t, x_t)$, then $x_w^* = 0$ is the unique optimal patronage offer. Because $C$ accepts any offer in a weak period, $G$'s utility strictly decreases if it deviates to any $x_t > 0$ because $C$ will still accept but $G$ consumes less than if $x_t = 0$.

   Furthermore, there does not exist an equilibrium strategy profile in which $x_w^* > 0$, which also implies that $\alpha_w^* = 1$ in any equilibrium strategy profile. To prove this claim by contradiction, suppose instead that there exists an equilibrium in which $\alpha_w^*(m_t, 0) = 0$ for at least one $m_t$. Then, to induce acceptance, $G$ optimally chooses the lowest $x_t$ that is strictly greater than 0. By the completeness of real numbers, for any $x_t > 0$, there exists $\epsilon > 0$ small enough that $x_t - \epsilon > 0$. Therefore, $G$ has a profitable deviation from any strategy profile with $x_t > 0$ in a weak period, which generates a contradiction. (An identical argument also shows why explicitly allowing $C$ to mix over acceptance/fighting in response to $x_t = 0$ in a weak period would not yield additional equilibria).

4. Given the unique $\alpha_s^*$ function in any equilibrium, $G$'s unique optimal allocation in a strong period is $(m_s^*, x_s^*)$ implicitly defined in Lemma 1. Suppose $G$ chooses $m_t' \neq m_s^*$ and/or $x_t' \neq x_s^*$. There are two cases.

   (a) $C$ optimally accepts $(m_t', x_t')$. Suppose $(m_t', x_t')$ is $G$'s most profitable deviation among the set of allocations that induce $C$ to accept. Then, given $m_t'$, $x_t'$ must make $C$ indifferent between accepting or fighting. Otherwise, this would not be $G$'s most profitable deviation because it could choose the same $m_t'$ but a lower $x_t$.

   ∘ If $R < R_{m_s^*=0}$, the Lemma 1 proof shows that deviating to any $m_t > 0$ strictly lowers $G$'s utility. Furthermore, because the posited $x_s^*$ makes $C$ indifferent between accepting or fighting at $m_t = 0$, $G$ cannot profitably deviate from offering $x_s^*$.

   ∘ If $R > R_{m_s^*=0}$ and if $x_t'$ makes $C$ indifferent between accepting and fighting, then $E\big[U_C(\text{accept } x_t')\big] = E\big[U_C(\text{fight} \mid \text{strong}, m_t')\big]$. Because $E\big[U_C(\text{accept } x_t')\big] = \theta x_t' + \delta V^C$ and $E\big[U_C(\text{fight} \mid \text{strong}, m_t')\big] = \delta\Big[\frac{1}{1+m_t'}V^G + \frac{m_t'}{1+m_t'}V^C\Big]$, setting these two terms equal and rearranging shows that $x_t' = \frac{\delta}{\theta(1+m_t')}\Big(V^G - V^C\Big)$. This simplifies to $x_t' = \frac{\delta}{(1-\delta)\theta(1+m_t')}\Big\{R - \sigma\big[m_s^* + (1+\theta)x_s^*\big]\Big\}$ after substituting in for the continuation values, and to $x_t' = \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \cdot \frac{1+m_s^*}{1+m_t'}$ after substituting in Equation A2 for $x_s^*$ and rearranging. This yields an implicit characterization of the military-spending amount consistent with the most profitable deviation: $1 = \frac{\delta(R-\sigma m_s^*)}{(1+m_s^*)\big[(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)\big]}$. Because this is identical to Equation A1, there are no profitable deviations from the stated strategy.

   (b) $C$ optimally rejects $(m_t', x_t')$. Lemma 2 establishes this is not a profitable deviation. ■

If $\sigma < \overline{\sigma}$, then Lemma 3 proves $m_s^* + x_s^* > R$, and hence a peaceful equilibrium does not exist.

To focus attention on the substantively interesting parameter ranges, the comparative statics analysis assumes that $\delta > \max\{\delta_1, \delta_2\}$ at the initial amount of revenues $\underline{R}$. Otherwise, the amount of oil does not impact fighting prospects because $C$ is sufficient impatient that a center-seeking civil war will never occur (see proof Lemma 3).

***Proof of Proposition 2, Part a.*** The proof proceeds in three steps. The first provides a preliminary result. Second, if there exists $\Delta_O' > 0$ such that $\overline{\sigma}(\underline{R}) < \overline{\sigma}(\underline{R} + \Delta_O')$, then there exists $\underline{\Delta}_O > \Delta_O'$ such that $\overline{\sigma}(\underline{R}) = \overline{\sigma}(\underline{R} + \underline{\Delta}_O)$. Third, $\underline{\Delta}_O$ is unique and creates a threshold such that $\overline{\sigma}(\underline{R}) > \overline{\sigma}(\underline{R} + \Delta_O)$ for any $\Delta_O > \underline{\Delta}_O$.

1. The following result will be used to generate a contradiction to complete a proof in the next step. Using explicit solutions it can easily be demonstrated that $\lim\limits_{\Delta_O \to \infty} \left( \frac{m_s^*(\underline{R}+\Delta_O)+x_s^*(\underline{R}+\Delta_O)}{\underline{R}+\Delta_O} \right) = 0$. This result does not require any additional functional form assumptions about $\theta(R)$ because the assumption $\theta \in (0, 1]$ for all $R$ implies $\theta$ does not diverge at the infinite limit.

2. Applying the intermediate value theorem proves the existence of at least one $\underline{\Delta}_O > \max\left\{0, \overline{\overline{R}} - \underline{R}\right\}$ such that $\overline{\sigma}(\underline{R}) = \overline{\sigma}(\underline{R} + \underline{\Delta}_O)$, under the imposed assumptions.

   - We are assuming there exists $\Delta_O' > 0$ such that $\overline{\sigma}(\underline{R}) < \overline{\sigma}(\underline{R} + \Delta_O')$. This is not true for all parameter values; note that if this assumption does not hold, then $\underline{\Delta}_O = 0$ and any-sized increase in oil revenues decreases $\overline{\sigma}$.

   - Need to show there exists an $\Delta_O'' > \Delta_O'$ such that $\overline{\sigma}(\underline{R}) > \overline{\sigma}(\underline{R} + \Delta_O'')$. The existence of a finite $\Delta_O''$ that satisfies this inequality can be established because $\overline{\sigma}(\underline{R} + \Delta_O)$ converges monotonically to 0 as $\Delta_O$ increases (for large enough $\Delta_O$). To prove $\overline{\sigma}(\underline{R} + \Delta_O)$ strictly decreases in $\Delta_O$ if $\Delta_O > \max\left\{0, \overline{\overline{R}} - \underline{R}\right\}$, the proofs below characterize the threshold $\overline{\overline{R}}$ such that $\frac{d\overline{\sigma}}{dO} < 0$ if $R = \underline{R} + \Delta_O > \overline{\overline{R}}$. To prove $\lim\limits_{\Delta_O \to \infty} \overline{\sigma}(\underline{R} + \Delta_O) = 0$ by contradiction, suppose instead that $\lim\limits_{\Delta_O \to \infty} \overline{\sigma}(\underline{R} + \Delta_O) > 0$ (although it is possible that the implicit solution for $\sigma$ in $m_s^*(\sigma) + x_s^*(\sigma) = R$ is less than 0 at $R = \underline{R} + \Delta_O$, in this circumstance $\overline{\sigma}$ attains its lower bound value of 0). Then, by the definition of $\overline{\sigma}$, there exists at least one $\sigma'$ such that $\lim\limits_{\Delta_O \to \infty} \left[ \underline{R} + \Delta_O - \left( m_s^*(\underline{R} + \Delta_O; \sigma') + x_s^*(\underline{R} + \Delta_O; \sigma') \right) \right] < 0$, which rearranges to $\lim\limits_{\Delta_O \to \infty} \left( \frac{m_s^*(\underline{R}+\Delta_O;\sigma')+x_s^*(\underline{R}+\Delta_O;\sigma')}{\underline{R}+\Delta_O} \right) > 1$. This contradicts the result established in step 1.

   - $\overline{\sigma}(\underline{R} + \Delta_O)$ is continuous in $\Delta_O$.

3. Because $\overline{\sigma}(\underline{R} + \Delta_O)$ strictly decreases in $\Delta_O$ if $\Delta_O > \max\left\{0, \overline{\overline{R}} - \underline{R}\right\}$, $\underline{\Delta}_O$ is unique. Furthermore, $\overline{\sigma}(\underline{R}) > \overline{\sigma}(\underline{R} + \Delta_O)$ for any $\Delta_O > \underline{\Delta}_O$. ■

Assumption A2 simplifies the number of cases to consider for Proposition 2, Part b by assuming the effect of oil on institutional quality has a constant sign.

**Assumption A2.** For any oil amounts $O_1$ and $O_2$, $\text{sgn}\left(\frac{d\theta}{dO}\Big|_{O=O_1}\right) = \text{sgn}\left(\frac{d\theta}{dO}\Big|_{O=O_2}\right)$.

To clarify the intuition behind Proposition 2, Part b, I split it into two parts. The first part assumes $\frac{d\theta}{dO} = 0$ (referred to below as Proposition 2, Part b.1), and the second part considers $\frac{d\theta}{dO} \neq 0$ (Proposition 2, Part b.2). The following formally states Proposition 2, Part b.1. It first characterizes $\frac{d\overline{\sigma}}{dO}$ if $R$ is small enough that $m_s^*$ hits a corner solution of 0. Intuitively, because the purpose of military spending is to decrease $C$'s desire to fight for the prize, arming to guard a small prize yields low benefits. Because $m_s^* = 0$, the game becomes strategically equivalent to the baseline scenario analyzed on pages XX and XX of the article. The statement below of Proposition 2, Part b.1 also shows that $R$ only needs to be larger than a threshold $\overline{R}$ to yield $\frac{d\overline{\sigma}}{dO} < 0$. Below, Lemma A9 proves that $\overline{R}$ is strictly less than the threshold $\overline{\overline{R}}$ (which is derived in the proof for Proposition 2, part b.2) that is stated in the article. Therefore, Proposition 2, Part b.1. is more general than Proposition 2 from the article.

**Proposition 2. Part b.1.** Assume $\frac{d\theta}{dO} = 0$.

**Case 1.** If $R < R_{m_s^*=0}$, then $\frac{d\overline{\sigma}}{dO} = 0$.

**Case 2.** If $R > \max\left\{R_{m_s^*=0}, \overline{R}\right\}$, for $\overline{R}$ defined in the proof, then $\frac{d\overline{\sigma}}{dO} < 0$.

***Proof of Case 1.*** If $R < R_{m_s^*=0}$, then $\overline{\sigma} = 1 - \frac{\theta}{\delta(1+\theta)}$ by the proof of Lemma 3. Therefore, $\frac{d\overline{\sigma}}{dO} = -\frac{\delta}{[\delta(1+\theta)]^2} \cdot \frac{d\theta}{dO}$. This term is 0 because $\frac{d\theta}{dO} = 0$ by assumption in part b.1 of Proposition 2. ■

Lemma 3 states that $\overline{\sigma}$ is defined implicitly by $R - m_s^*(\overline{\sigma}) - x_s^*(\overline{\sigma}) = 0$ if $R > R_{m_s^*=0}$. Applying the implicit function theorem yields:

$$\left.\frac{d\overline{\sigma}}{dO}\right|_{\frac{d\theta}{dO}=0} = \frac{1 - \left.\frac{d}{dO}\Big(x_s^* + m_s^*\Big)\right|_{\frac{d\theta}{dO}=0}}{\frac{d}{d\sigma}\Big(x_s^* + m_s^*\Big)}. \tag{A3}$$

Before proving Proposition 2, Part b.1, Case 2, the following lemmas characterize the signs of the different terms in Equation A3. Lemma A1 demonstrates that the denominator of Equation A3 is strictly negative. Providing intuition for the result, $C$'s expected utility to accepting is contingent not only on its patronage offer in the current period, but also on how frequently it will receive a positive patronage offer in the future. More frequent future patronage offers engendered by a higher $\sigma$ imply that $C$ will accept a lower $x_s^*$ in the current period. This also decreases the marginal effectiveness of arming at reducing the patronage offer, which lowers $m_s^*$.

**Lemma A1.** $\frac{d}{d\sigma}\Big(x_s^* + m_s^*\Big) < 0.$

***Proof.*** Because $\frac{d}{d\sigma}\Big(x_s^* + m_s^*\Big) = \frac{\partial x_s^*}{\partial \sigma} + \Big(1 + \frac{\partial x_s^*}{\partial m_t}\Big) \cdot \frac{\partial m_s^*}{\partial \sigma}$, it suffices to show $\frac{\partial x_s^*}{\partial \sigma} < 0$, $1 + \frac{\partial x_s^*}{\partial m_t} > 0$, and $\frac{\partial m_s^*}{\partial \sigma} < 0$.

- $\frac{\partial x_s^*}{\partial \sigma} = -\Big(\frac{\delta m_s^*}{(1-\delta)(1+m_s^*)\theta + \delta\sigma(1+\theta)} + \frac{\delta^2(1+\theta)(R-\sigma m_s^*)}{\big[(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)\big]^2}\Big) < 0.$
- Using explicit solutions, $1 + \frac{\partial x_s^*}{\partial m_t} = \frac{2\delta\sigma\theta}{\delta\sigma\theta + \sqrt{\delta\Big(4(1-\delta)(R+\sigma)\theta + \delta\sigma^2(2+\theta)^2\Big)}} > 0.$
- Because $m_s^*$ and $\sigma$ each appear only as negative terms in the numerator of Equation A1, and only as positive terms in the denominator, the implicit function theorem easily establishes $\frac{\partial m_s^*}{\partial \sigma} < 0$. ■

Regarding the numerator of Equation A3, $1 - \left.\frac{d}{dO}\Big(x_s^* + m_s^*\Big)\right|_{\frac{d\theta}{dO}=0} = 1 - \frac{\partial x_s^*}{\partial O} - \Big(1 + \frac{\partial x_s^*}{\partial m_t}\Big) \cdot \frac{\partial m_s^*}{\partial O}$. The sign of this term is ambiguous because the direct effect $1 - \frac{\partial x_s^*}{\partial O} > 0$ whereas the indirect effect $-\Big(1 + \frac{\partial x_s^*}{\partial m_t}\Big) \cdot \frac{\partial m_s^*}{\partial O} < 0$, as shown in Lemmas A2 and A3, respectively. The direct effect $1 - \frac{\partial x_s^*}{\partial O}$ intuitively should be positive. Because $G$ optimally raises military spending in response to a larger prize, an increase in oil wealth raises $G$'s total revenues by a greater amount than it increases the equilibrium patronage offer. If $G$'s strong-period military spending minimized lifetime strong-period expenditures subject to inducing $C$ to accept, this would be the only term in the numerator of Equation A3. However, because $x_s^* + m_s^*$ does not in fact minimize lifetime strong-period expenditures subject to inducing $C$ to accept, there is also an indirect effect—how oil affects total strong-period expenditures through its effect on military spending—that diminishes the positive direct effect oil exerts on $R - x_s^* - m_s^*$. The next subsection explains this indirect term more thoroughly.

**Lemma A2.** $1 - \frac{\partial x_s^*}{\partial O} > 0$.

***Proof.*** If a peaceful equilibrium exists, then $\frac{m_s^*+x_s^*}{R} < 1$. Therefore, to demonstrate $\frac{\partial x_s^*}{\partial O} = \frac{\delta}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} < 1$, we can tighten the upper bound of the inequality and solve $\frac{\delta}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} < \frac{m_s^*+x_s^*}{R}$. Substituting in for $x_s^* = \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)}$ and rearranging slightly yields $[(1-\delta)(1+m_s^*)\theta + \delta\sigma(1+\theta)]m_s^* + \delta(R - \sigma m_s^*) > \delta R$. This easily reduces to $(1-\delta)(1+m_s^*)\theta + \delta\sigma\theta > 0$, a true statement. ■

**Lemma A3.** $-\left(1 + \frac{\partial x_s^*}{\partial m_t}\right) \cdot \frac{\partial m_s^*}{\partial O} < 0$.

***Proof.*** It suffices to demonstrate $1 + \frac{\partial x_s^*}{\partial m_t} > 0$ and $\frac{\partial m_s^*}{\partial O} > 0$.

- The proof of Lemma A1 establishes $1 + \frac{\partial x_s^*}{\partial m_t} > 0$.
- Because $O$ only appears as a positive term in the numerator of Equation A1 (recall that Part b.1 of Proposition 2 assumes $\frac{d\theta}{dO} = 0$; also recall $R = O+N$) and because $m_s^*$ appears only as a negative term in the numerator and as a positive term in the denominator of Equation A1, the implicit function theorem easily demonstrates $\frac{\partial m_s^*}{\partial O} > 0$. Using the explicit solution, $\frac{\partial m_s^*}{\partial O} = \sqrt{\frac{\delta}{4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2}} > 0$. ■

Because the indirect effect term $-\left(1+\frac{\partial x_s^*}{\partial m_t}\right)\cdot\frac{\partial m_s^*}{\partial O}$ diminishes the numerator of Equation A3, under some parameter values there exists a range $R \in (R_{m_s^*=0}, \overline{R})$ such that the numerator of Equation A3 is negative, which implies $\frac{d\overline{\sigma}}{dO} > 0$. However, the proof of Proposition 2, Part b.1, Case 2 proves there exists a unique $\overline{R}$ such that the numerator of Equation A3 is strictly positive if $R > \overline{R}$—which in turn implies $\frac{d\overline{\sigma}}{dO} < 0$. Lemma A4 provides a preliminary result that will be used to establish this claim. The intuition behind Lemma A4 is that although more oil increases $x_s^* + m_s^*$, it does so at a diminishing rate.

**Lemma A4.** The numerator of Equation A3 is a strictly increasing function of $O$. Formally, $\frac{d}{dO}\left[1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}\right] > 0$.

*Proof.* $\frac{d}{dO}\left[1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}\right] = -\frac{d^2}{dO^2}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} =$
$-\frac{d}{dO}\left[\frac{\partial x_s^*}{\partial O} + \left(1+\frac{\partial x_s^*}{\partial m_t}\right)\cdot\frac{\partial m_s^*}{\partial O}\right] = -\frac{\partial^2 x_s^*}{\partial O^2} - \frac{\partial^2 x_s^*}{\partial O \partial m_s}\cdot\frac{\partial m_s^*}{\partial O} - \left(1+\frac{\partial x_s^*}{\partial m_t}\right)\cdot\frac{\partial^2 m_s^*}{\partial O^2} - \frac{d}{dO}\left(1+\frac{\partial x_s^*}{\partial m_t}\right)\cdot\frac{\partial m_s^*}{\partial O}$.

It suffices to demonstrate $\frac{\partial^2 x_s^*}{\partial O^2} = 0$, $\frac{\partial^2 x_s^*}{\partial O \partial m_s} < 0$, $\frac{\partial m_s^*}{\partial O} > 0$, $1 + \frac{\partial x_s^*}{\partial m_t} > 0$, $\frac{\partial^2 m_s^*}{\partial O^2} < 0$, and $\frac{d}{dO}\left(1 + \frac{\partial x_s^*}{\partial m_t}\right) < 0$.

- Using the expression for $\frac{\partial x_s^*}{\partial O}$ derived in the proof for Lemma A2, we can calculate $\frac{\partial^2 x_s^*}{\partial O^2} = 0$.
- Using the expression for $\frac{\partial x_s^*}{\partial O}$ derived in the proof for Lemma A2, we can calculate $\frac{\partial^2 x_s^*}{\partial O \partial m_s} = -\frac{\delta(1-\delta)\theta}{[(1-\delta)(1+m_s^*)\theta + \delta\sigma(1+\theta)]^2} < 0$.
- The proof for Lemma A3 establishes $\frac{\partial m_s^*}{\partial O} > 0$.
- The proof for Lemma A3 establishes $1 + \frac{\partial x_s^*}{\partial m_t} > 0$.
- The proof for Lemma A3 demonstrates $\frac{\partial m_s^*}{\partial O} = \sqrt{\frac{\delta}{4(1-\delta)(R+\sigma)\theta + \delta\sigma^2(2+\theta)^2}}$. Because $O$ only appears as a positive term in the denominator (recall that Part b.1 of Proposition 2 assumes $\frac{d\theta}{dO} = 0$; also recall $R = O + N$), $\frac{\partial^2 m_s^*}{\partial O^2} < 0$.
- The proof for Lemma A3 demonstrates $1 + \frac{\partial x_s^*}{\partial m_t} = \frac{2\delta\sigma\theta}{\delta\sigma\theta + \sqrt{\delta\left(4(1-\delta)(R+\sigma)\theta + \delta\sigma^2(2+\theta)^2\right)}}$. Because $O$ only appears as a positive term in the denominator (recall that Part b.1 of Proposition 2 assumes $\frac{d\theta}{dO} = 0$; also recall $R = O + N$), $\frac{d}{dO}\left(1 + \frac{\partial x_s^*}{\partial m_t}\right) < 0$. ■

***Proof of Proposition 2, Part b.1, Case 2.*** The proof proceeds by establishing two different results. First, if there exists $R' > R_{m_s^*=0}$ such that the numerator of Equation A3 is negative, then there exists an $\overline{R} > R'$ such that the numerator of Equation A3 is 0. Second, $\overline{R}$ is unique and creates a threshold such that the numerator of Equation A3 is strictly positive for any $R > \overline{R}$. This proof strategy incorporates the Lemma A1 result that the denominator of Equation A3 is strictly negative, which implies $\frac{d\overline{\sigma}}{dO} < 0$ for any $R$ such that the numerator of Equation A3 is strictly positive.

1. Applying the intermediate value theorem proves the existence of $\overline{R}$ such that $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0, R=\overline{R}} = 0$, under the imposed assumptions.
    - We are assuming there exists $R' > R_{m_s^*=0}$ such that the numerator of Equation A3 is negative, i.e., $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0, R=R'} < 0$. This is not true for all parameter values; note that if this assumption does not hold, then $\frac{d\overline{\sigma}}{dO} < 0$ for all $R > R_{m_s^*=0}$.

- Need to show there exists an $R'' > R'$ such that $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0, R=R''} > 0$. The existence of a finite $R''$ with these properties can be established because $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}$ converges monotonically to 1 as $O$ increases. The monotonic convergence claim follows from Lemma A4 and from the following proof establishing $\lim_{O\to\infty} 1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} = 1$. Because $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} = 1 - \frac{\partial x_s^*}{\partial O} - \left(1 + \frac{\partial x_s^*}{\partial m_t}\right) \cdot \frac{\partial m_s^*}{\partial O}$, using the respective expressions for each of these terms derived above it is easy to show $\lim_{O\to\infty} \frac{\partial x_s^*}{\partial O} = 0$, $\lim_{O\to\infty} \left(1 + \frac{\partial x_s^*}{\partial m_t}\right) = 0$, and $\lim_{O\to\infty} \frac{\partial m_s^*}{\partial O} = 0$. Therefore, $\lim_{O\to\infty} 1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} = 1$.
- $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}$ is continuous in $R$ for all $R > R_{m_s^*=0}$.

2. Because Lemma A4 shows the numerator of Equation A3 is strictly increasing in $R$, $\overline{R}$ is unique. Furthermore, $1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} > 0$ for any $R > \overline{R}$. ■

I now formally state Proposition 2, part b.2. Case 1 characterizes $\frac{d\overline{\sigma}}{dO}$ if $R$ is small enough that $m_s^*$ hits a corner solution of 0. The sign of the effect in this parameter range differs from the corresponding part b.1 case because oil only affects equilibrium strong-period government expenditures relative to the size of the prize through its effect on institutional quality if $m_s^* = 0$; thus, the sign of the effect of oil on institutions determines the sign of $\frac{d\overline{\sigma}}{dO}$. Case 2 states that $R > \overline{R}$ is a sufficient condition for $\frac{d\overline{\sigma}}{dO} < 0$ if $\frac{d\theta}{dO} > 0$. Below, Lemma A9 proves that $\overline{R}$ is strictly less than the threshold $\overline{\overline{R}}$ that is stated in the article. Therefore, the proof for part b proves a more general claim than stated in the article. Only for Case 3, in which $\frac{d\theta}{dO} < 0$, is $R > \overline{\overline{R}}$ needed to yield $\frac{d\overline{\sigma}}{dO} < 0$.

**Proposition 2. Part b.2.** Assume $\frac{d\theta}{dO} \neq 0$.

**Case 1.** Assume $R < R_{m_s^*=0}$. If $\frac{d\theta}{dO} > 0$, then $\frac{d\overline{\sigma}}{dO} < 0$. If $\frac{d\theta}{dO} < 0$, then $\frac{d\overline{\sigma}}{dO} > 0$.

**Case 2.** Assume $R > R_{m_s^*=0}$. If $\frac{d\theta}{dO} > 0$ and $R > \overline{R}$, then $\frac{d\overline{\sigma}}{dO} < 0$.

**Case 3.** Assume $R > R_{m_s^*=0}$. If $\frac{d\theta}{dO} \in \left(\underline{d\theta}, 0\right)$, for $\underline{d\theta} < 0$ defined in the proof, and $R > \overline{\overline{R}}$, then $\frac{d\overline{\sigma}}{dO} < 0$.

***Proof of Case 1.*** The proof for Proposition 2, Part b.1, Case 1 demonstrates that $\frac{d\overline{\sigma}}{dO} = -\frac{\delta}{[\delta(1+\theta)]^2} \cdot \frac{d\theta}{dO}$ if $R < R_{m_s^*=0}$. This term is negative if $\frac{d\theta}{dO} > 0$ and positive if $\frac{d\theta}{dO} < 0$. ■

The structure of the proof for Proposition 2, Part b.2, Part 2 resembles the proof for Proposition 2, Part b.1, Part 2. The difference is that Part b.2 requires characterizing another indirect effect of oil: its effect on $\theta$. Applying the implicit function theorem to calculate $\frac{d\overline{\sigma}}{dO}$ for the general $\frac{d\theta}{dO}$ case yields:

$$\frac{d\overline{\sigma}}{dO} = \frac{1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO}}{\frac{d}{d\sigma}\left(x_s^* + m_s^*\right)} \tag{A4}$$

The sign of the additional indirect effect in Equation A4, $-\frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO}$, depends on the sign of $\frac{d\theta}{dO}$. The numerator of Equation A4 is larger than the numerator of Equation A3 if $\frac{d\theta}{dO} > 0$, and smaller if $\frac{d\theta}{dO} < 0$. Lemma A5 provides a formal result that implies this claim. The intuition for Lemma A5 is that if $G$ is more efficient at translating patronage spending into actual consumption for $C$, then $G$ needs to spend less to buy off $C$.

**Lemma A5.** $\frac{d}{d\theta}\left(x_s^* + m_s^*\right) < 0$.

***Proof.*** Because $\frac{d}{d\theta}\left(x_s^* + m_s^*\right) = \frac{\partial x_s^*}{\partial \theta} + \left(1 + \frac{\partial x_s^*}{\partial m_t}\right) \cdot \frac{\partial m_s^*}{\partial \theta}$, it suffices to prove $\frac{\partial x_s^*}{\partial \theta} < 0$, $1 + \frac{\partial x_s^*}{\partial m_t} > 0$, and $\frac{\partial m_s^*}{\partial \theta} < 0$.

- Because $\theta$ only enters directly as a positive term in the denominator of $x_s^*$, $\frac{\partial x_s^*}{\partial \theta} < 0$.
- Lemma A1 proves $1 + \frac{\partial x_s^*}{\partial m_t} > 0$.
- Using the implicit function theorem yields:
  $\frac{\partial m_s^*}{\partial \theta} = -\frac{(1+m_s^*)\left[(1-\delta)(1+m_s^*)+\delta\sigma\right](R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta(R-\sigma m_s^*)+(R+\sigma)\left[(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)\right]} < 0$. ■

Because of Lemma A5, the proof of Proposition 2, Part b.2, Case 2 follows directly from Proposition 2, Part b.1, Case 2.

***Proof of Proposition 2, Part b.2, Case 2.*** If $\frac{d\theta}{dO} > 0$, then $R > \overline{R}$ is a sufficient condition for $1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO} > 0$. Proposition 2, Part b.1, Case 2 demonstrates that $R > \overline{R}$ implies $1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} > 0$. Because Lemma A5 demonstrates $\frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) < 0$, it follows that $\frac{d\theta}{dO} > 0$ and $R > \overline{R}$ imply $1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO} > 0$. ■

Even if $\frac{d\theta}{dO} < 0$, a qualitatively similar result as Proposition 2, Part b.1, Case 2 also holds for Proposition 2, Part b.2, Case 3 if $\frac{d\theta}{dO}$ is not too large in magnitude because then the numerator of Equation A4 is a strictly increasing function of $O$. Formally, $\frac{d}{dO}\Big\{1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO}\Big\} > 0$ if $\frac{d\theta}{dO} \in (\underline{d\theta}, 0)$, for $\underline{d\theta} < 0$ defined below. The left-hand side of this inequality can be restated as $-\Big[a \cdot \Big(\frac{d\theta}{dO}\Big)^2 + 2b \cdot \frac{d\theta}{dO} + \Big(c + d \cdot \frac{d^2\theta}{dO^2}\Big)\Big]$, for:

$$a = \frac{d^2}{d\theta^2}\Big(x_s^* + m_s^*\Big)$$
$$b = \frac{d^2}{dOd\theta}\Big(x_s^* + m_s^*\Big)$$
$$c = \frac{d^2}{dO^2}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0}$$
$$d = \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big)$$

The following three lemmas characterize the signs of each component in this quadratic function of $\frac{d\theta}{dO}$.

**Lemma A6.** $a > 0$.

Because $\frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) < 0$ (see Lemma A5), Lemma A6 implies that greater efficiency at translating patronage spending into actual consumption for $C$ lowers total strong-period expenditures at a decreasing rate. Intuitively, the effect of marginally increasing $\theta$ greatly enhances $G$'s ability to buy off $C$ when $\theta$ is low—because then $G$ has almost no ability to transfer government revenues to $C$—but has a lesser effect when $G$ is better able to dispense patronage.

Unfortunately, whether using expressions that substitute in the explicit solution for $m_s^*$, or evaluating terms that are a function of parameters and $m_s^*$, is it not possible to unambiguously formally characterize the sign of $a$. I therefore used simulations to evaluate

$a$ for different parameter values. A total of 18,000 simulations were run that computed all permutations of (a) allowing $\sigma$ to vary between 0.1 and 1 in steps of 0.1, (b) allowing $\theta$ to vary between 0.1 and 1 in steps of 0.1, (c) allowing $\delta$ to vary between 0.1 and 0.9 in steps of 0.1,[1] and (d) allowing $R$ to vary between 1 and 20 in steps of 1. The sign of $a$ is indeed positive for every combination of parameters such that $R > m_s^*$, which is a necessary condition for a peaceful equilibrium to exist.

**Lemma A7.** $b < 0$.

Lemma A7 states that higher ability to deliver patronage dampens the positive effect that more oil has on total strong-period expenditures. This is intuitive because higher $\theta$ implies $G$ can more efficiently meet $C$'s higher fighting constraint that results from higher $O$.

***Proof.*** Because $\frac{d^2}{dOd\theta}\left(x_s^*+m_s^*\right) = \frac{d}{d\theta}\left[\frac{\partial x_s^*}{\partial O}+\left(1+\frac{\partial x_s^*}{\partial m_t}\right)\frac{\partial m_s^*}{\partial O}\right] = \frac{\partial^2 x_s^*}{\partial O \partial \theta}+\left(1+\frac{\partial x_s^*}{\partial m_t}\right)\cdot\frac{\partial^2 m_s^*}{\partial O\partial\theta}+\frac{\partial^2 x_s^*}{\partial m_s \partial O}\cdot\frac{\partial m_s^*}{\partial\theta}+\frac{\partial m_s^*}{\partial O}\cdot\frac{d}{d\theta}\left(1+\frac{\partial x_s^*}{\partial m_t}\right)$, it suffices to demonstrate $\frac{\partial^2 x_s^*}{\partial O\partial\theta} < 0$, $1+\frac{\partial x_s^*}{\partial m_t} > 0$, $\frac{\partial^2 m_s^*}{\partial O\partial\theta} < 0$, and $\frac{\partial^2 x_s^*}{\partial m_s\partial O}\cdot\frac{\partial m_s^*}{\partial\theta}+\frac{\partial m_s^*}{\partial O}\cdot\frac{d}{d\theta}\left(1+\frac{\partial x_s^*}{\partial m_t}\right) < 0$.

- $\frac{\partial^2 x_s^*}{\partial O\partial\theta} = -\frac{\delta[(1-\delta)(1+m_s^*)+\delta\sigma]}{[(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)]^2} < 0$.
- Lemma A1 proves $1+\frac{\partial x_s^*}{\partial m_t} > 0$.
- Using explicit solutions, $\frac{\partial^2 m_s^*}{\partial O\partial\theta} = -\frac{\delta^2[4(1-\delta)(R+\sigma)+2\delta\sigma^2(2+\theta)]}{2[\delta(4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2)]^{\frac{3}{2}}} < 0$.
- Using explicit solutions, $\frac{\partial^2 x_s^*}{\partial m_s\partial O}\cdot\frac{\partial m_s^*}{\partial\theta}+\frac{\partial m_s^*}{\partial O}\cdot\frac{d}{d\theta}\left(1+\frac{\partial x_s^*}{\partial m_t}\right) = -\frac{4[(1-\delta)(R+\sigma)\theta+\delta\sigma(2+\theta)]-\sigma\sqrt{\delta\left[4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2\right]}}{\left\{[4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2]\cdot\left(\delta\sigma\theta+\sqrt{\delta[4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2]}\right)\right\}^2}$. The denominator is clearly positive because it is squared. To demonstrate the numerator is positive, it suffices to show that $4(1-\delta)(R+\sigma)\theta+4\delta\sigma(2+\theta) > 4(1-\delta)(R+\sigma)\theta+\delta\sigma^2(2+\theta)^2$, which easily reduces to $4 > \sigma(2+\theta)$, which is true because $\sigma < 1$ and $\theta < 1$. ■

Finally, establishing the sign of $c + d\cdot\frac{d^2\theta}{dO^2}$ requires assuming $\frac{d\theta}{dO}$ either increases, or does not decrease too steeply, in $O$.

**Assumption A3.** $\frac{d^2\theta}{dO^2} > \underline{d^2\theta} \equiv -\frac{\frac{d^2}{dO^2}\left(x_s^*+m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}}{\frac{d}{d\theta}\left(x_s^*+m_s^*\right)}$.

[1] I do not consider $\delta = 1$ because this violates the assumption that the player's utility functions are continuous at infinity.

It can easily be shown that $\underline{d^2\theta} < 0$. The intuition for the signs of $c$ and $d$ follow from Lemmas A4 and A5.

**Lemma A8.** $c + d \cdot \frac{d^2\theta}{dO^2} < 0$.

***Proof.*** Because of Assumption A3, it suffices to show that $\frac{d^2}{dO^2}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} < 0$ (Lemma A4) and $\frac{d}{d\theta}\left(x_s^* + m_s^*\right) < 0$ (Lemma A5). ■

Assumption A4 imposes two additional technical conditions that are used to prove Proposition 2, Part b.2, Case 3. The first captures the natural idea that an increase in oil revenues exerts diminishing marginal effects on institutional quality. Given the first, the second is equivalent to stating that either $\frac{d^2\theta}{dO^2}$ does not converge to 0, or if it does then its rate of convergence exceeds the rate of convergence for the first derivative.

**Assumption A4.** $\lim_{O\to\infty} \frac{d\theta}{dO} = 0$ and $\lim_{O\to\infty} -\frac{\frac{d^2\theta}{dO^2}}{\frac{d\theta}{dO}} \neq 0$.

***Proof of Proposition 2, Part b.2, Case 3.*** Assume $\frac{d\theta}{dO} < 0$. Following the structure of the proof for Proposition 2, Part b.1, Case 2, to demonstrate the existence of the desired $\overline{\overline{R}}$ threshold, it suffices to demonstrate that
$1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO}$ converges monotonically to a positive number as $O$ increases. Specifically, I will show:

(1) $\frac{d}{dO}\left\{1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO}\right\} > 0$ if $\frac{d\theta}{dO} \in (\underline{d\theta}, 0)$, and

(2) $\lim_{O\to\infty} 1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO} > 0$. Note that the first result is the analog to Lemma A4 for $\frac{d\theta}{dO} < 0$.

1. To demonstrate $\frac{d}{dO}\left\{1 - \frac{d}{dO}\left(x_s^* + m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\left(x_s^* + m_s^*\right) \cdot \frac{d\theta}{dO}\right\} > 0$, need to show $a \cdot \left(\frac{d\theta}{dO}\right)^2 + b \cdot \frac{d\theta}{dO} + \left(c + d \cdot \frac{d^2\theta}{dO^2}\right) < 0$. Because this expression is quadratic in $\frac{d\theta}{dO}$ and because $a > 0$ (Lemma A6), this will be true for any $\frac{d\theta}{dO} \in \left(\underline{d\theta}, \frac{-b+\sqrt{b^2-4a(c+d\cdot\frac{d^2\theta}{dO^2})}}{2a}\right)$, with $\underline{d\theta} \equiv \frac{-b-\sqrt{b^2-4a(c+d\cdot\frac{d^2\theta}{dO^2})}}{2a}$. Because $a > 0$, $b < 0$ (Lemma A7), and $c + d \cdot \frac{d^2\theta}{dO^2} < 0$ (Lemma A8), the upper bound of this range is strictly positive. To establish $\underline{d\theta} < 0$, need to show $-b < \sqrt{b^2 - 4a(c + d \cdot \frac{d^2\theta}{dO^2})}$.

Because both sides of the inequality are positive, we can square each side to get $b^2 < b^2 - 4a(c + d \cdot \frac{d^2\theta}{dO^2})$, which simplifies to $4a(c + d \cdot \frac{d^2\theta}{dO^2}) < 0$. Because $a > 0$ and $c + d \cdot \frac{d^2\theta}{dO^2} < 0$, this inequality is true.

2. The proof of Proposition 2, Part b.1, Case 2 establishes $\lim_{O\to\infty} 1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} > 0$ if $\frac{d\theta}{dO} = 0$. The same proof applies to the general $\frac{d\theta}{dO}$ case because $\theta \in (0, 1]$ for all $R$. Therefore, to prove $\lim_{O\to\infty} 1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)\Big|_{\frac{d\theta}{dO}=0} - \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO} > 0$, it suffices to demonstrate that $\lim_{O\to\infty} \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO} = 0$. Using explicit solutions, it can be demonstrated that $\lim_{O\to\infty} \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) = -\infty$ and $\lim_{O\to\infty} \frac{d^2}{d\theta dO}\Big(x_s^* + m_s^*\Big) = 0$. Therefore, because $\lim_{O\to\infty} \frac{d\theta}{dO} = 0$ (the first part of Assumption A4), $\lim_{O\to\infty} \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO} = \lim_{O\to\infty} \frac{\frac{d}{d\theta}\left(x_s^* + m_s^*\right)}{\frac{1}{\frac{d\theta}{dO}}} = -\frac{\infty}{\infty}$. By L'Hopital's rule, this limit is equivalent to $\lim_{O\to\infty} \frac{\frac{d^2}{d\theta dO}\left(x_s^* + m_s^*\right)}{-\frac{\frac{d^2\theta}{dO^2}}{\left(\frac{d\theta}{dO}\right)^2}}$. Because the numerator goes to 0, a sufficient condition for the entire limit to go to 0 is that $\lim_{O\to\infty} -\frac{\frac{d^2\theta}{dO^2}}{\left(\frac{d\theta}{dO}\right)^2} \neq 0$. This can be rewritten as $\lim_{O\to\infty} \frac{-\frac{\frac{d^2\theta}{dO^2}}{\frac{d\theta}{dO}}}{\frac{d\theta}{dO}}$. Using both parts of Assumption A4 implies $\lim_{O\to\infty} \frac{-\frac{\frac{d^2\theta}{dO^2}}{\frac{d\theta}{dO}}}{\frac{d\theta}{dO}}$ diverges to negative or positive infinity. ■

Lemma A9 proves $\overline{\overline{R}} > \overline{R}$. This is intuitive because the numerator of Equation A3 is strictly larger than the numerator of Equation A4 if $\frac{d\theta}{dO} < 0$.

**Lemma A9.** $\overline{\overline{R}} > \overline{R}$.

***Proof.*** Define $f(R) = 1 - \frac{d}{dO}\Big(x_s^* + m_s^*\Big)$ and $g(R) = \frac{d}{d\theta}\Big(x_s^* + m_s^*\Big) \cdot \frac{d\theta}{dO}$. Proposition 2, Part b.1, Case 2 demonstrated there exists a unique $\overline{R}$ such that $f\Big(\overline{R}\Big) = 0$, and Proposition 2, Part b.2, Case 3 proves there exists a unique $\overline{\overline{R}}$ such that $f\Big(\overline{\overline{R}}\Big) - g\Big(\overline{\overline{R}}\Big) = 0$. These results produce two additional conclusions:

- $f\Big(\overline{R}\Big) = f\Big(\overline{\overline{R}}\Big) - g\Big(\overline{\overline{R}}\Big)$.

- Because $g(\cdot) > 0$ if $\frac{d\theta}{dO} < 0$ (see Lemma A5), $f\left(\overline{\overline{R}}\right) - g\left(\overline{\overline{R}}\right) < f\left(\overline{\overline{R}}\right)$.

Combining these two results yields $f\left(\overline{\overline{R}}\right) > f\left(\overline{R}\right)$. Because $f(\cdot)$ is a strictly increasing function of $R$ (see Lemma A4), this implies $\overline{\overline{R}} > \overline{R}$. ■

***Proof of Proposition 3.*** Because $\frac{d\theta}{d\omega} < 0$, it suffices to demonstrate $\frac{d\overline{\sigma}}{d\theta} < 0$. If $R < R_{m_s^*=0}$, then $\frac{d\overline{\sigma}}{d\theta} = -\frac{1}{\delta(1+\theta)^2} < 0$. If $R > R_{m_s^*=0}$, then $\frac{d\overline{\sigma}}{d\theta} = -\frac{\frac{d}{d\theta}\left(x_s^*+m_s^*\right)}{\frac{d}{d\sigma}\left(x_s^*+m_s^*\right)}$. Lemma A5 proves $\frac{d}{d\theta}\left(x_s^* + m_s^*\right) < 0$ and Lemma A1 proves $\frac{d}{d\sigma}\left(x_s^* + m_s^*\right) < 0$, rendering the overall term negative. ■

## A.3 Modified Model with Commitment to Choosing Military Spending to Minimize Lifetime Expenditures

This section modifies one aspect of the model setup. In the article, $G$ chooses $m_t$ in every period $t$. In the modified model considered here, prior to the first period of the game the initial governing actor sets a state-dependent military spending schedule $(\hat{m}_s, \hat{m}_w)$ that denotes a military spending amount $\hat{m}_s$ for the incumbent government in every future strong period and $\hat{m}_w$ in every future weak period. Therefore, after this initial choice, military expenditures cease to be a choice variable for $G$ because it is assumed that $G$ can commit to the originally determined military schedule.

The purpose of studying this modified model, as mentioned on page XX of the article, is to explain why a non-obvious source of inefficiency arises in the original model, specifically, why it is possible for low enough $R$ that $m_s^*$ has an interior solution but $\frac{d\overline{\sigma}}{dO} > 0$ even if oil does not negatively impact bureaucratic capacity. This occurs because in the original model $G$ cannot commit to choose a strong-period military spending amount that is consistent with minimizing lifetime strong-period expenditures. That is, this parameter range is not generated by any specific features of oil production and instead exists for any revenue source. By contrast, no parameter range exists in the modified model such that $m_s^* > 0$, $\frac{d\theta}{dO} \geq 0$, and $\frac{d\overline{\sigma}}{dO} > 0$. To differentiate notation in the modified model from that in the original model, strong-period military spending is denoted as $\hat{m}_s$, optimal strong-period patronage as a function of the chosen strong-period military spending amount is $x_s^*(\hat{m}_s)$, equilibrium strong-period military spending is $\hat{m}_s^*$, equilibrium strong-period patronage is $\hat{x}_s^*$, and the $\sigma$ threshold that determines whether or not fighting will occur is $\hat{\overline{\sigma}}$.

Deriving the equilibrium strong-period patronage offer for the modified model follows the same procedure as for the original model. In a strong period, $G$ makes a positive patronage offer to $C$ because $C$ prefers to fight rather than to forgo consumption in the current period. The single deviation in which $C$ fights in period $t$ in response to $G$'s offer $x_t$ generates expected lifetime utility of $\text{E}\big[U_C(\text{fight} \mid \text{strong}, \hat{m}_s)\big] = \delta\Big[\frac{1}{1+\hat{m}_s}V^G + \frac{\hat{m}_s}{1+\hat{m}_s}V^C\Big]$ because no consumption occurs in the fighting period, and $C$ receives the government's equilibrium continuation value from winning whereas losing yields the challenger's equilibrium continuation value. This term differs from the corresponding one in the article because $m_t$ is set at the previously-chosen level $\hat{m}_s$, rather than is chosen in the current period. Solving the continuation values, $C$ will consume $\theta \cdot x_s^*(\hat{m}_s)$ in the $\sigma$ percentage of future periods it is strong and nothing when weak. This implies $V^C = \frac{\sigma \cdot \theta \cdot x_s^*(\hat{m}_s)}{1-\delta}$. Furthermore, because $G$ consumes $R$ in periods $C$ is weak and $R - \hat{m}_s - \theta \cdot x_s^*(\hat{m}_s)$ in periods $C$ is strong,[2] $V^G = \frac{R-\sigma[\hat{m}_s+\theta \cdot x_s^*(\hat{m}_s)]}{1-\delta}$. Because in equilibrium $\theta \cdot x_s^*(\hat{m}_s)$ solves $\theta \cdot x_s^*(\hat{m}_s) + \delta V^C = \text{E}\big[U_C(\text{fight} \mid \text{strong}, \hat{m}_s)\big]$, substituting in the continuation values and rearranging yields the optimal strong-period patronage offer as a function of $\hat{m}_s$:

$$\hat{x}_s^*(\hat{m}_s) = \frac{\delta(R - \sigma\hat{m}_s)}{(1-\delta)(1+\hat{m}_s)\theta + \delta\sigma(1+\theta)} \tag{A5}$$

Anticipating $C$'s calculus, $G$ chooses strong-period military spending at the outset of the game to maximize its lifetime expected utility, anticipating that in the future it will make the optimal patronage offer in each period conditional on the military spending amount. Therefore, $G$ solves:

$$\hat{m}_s^* \equiv \arg\max_{\hat{m}_s} \; \frac{R - \sigma\big[\hat{m}_s + \hat{x}_s^*(\hat{m}_s)\big]}{1-\delta} \tag{A6}$$

s.t. (C1) $x_s^*(\hat{m}_s) \geq \frac{\delta(R-\sigma\hat{m}_s)}{(1-\delta)(1+\hat{m}_s)\theta+\delta\sigma(1+\theta)}$, (C2) $\hat{x}_s \geq 0$, (C3) $\hat{m}_s \geq 0$, (C4) $\hat{m}_s + x_s^*(\hat{m}_s) \leq R$.

The important difference between the optimization problems in Equation A6 and Equation 2 is that in the modified model, $G$ chooses $m_t$ for *all* strong periods at the only information set for which it chooses military spending. Therefore, $G$ internalizes the consequences of higher military spending in all strong periods and the first-order condition satisfies:

$$1 = -\frac{\partial \hat{x}_s^*}{\partial \hat{m}_s} \tag{A7}$$

Equation A7 states that in equilibrium the marginal cost of increasing military capacity by one unit, which is 1, equals the marginal benefit of military spending—reducing the required patronage offer. For this reason, in the modified model, equilibrium strong-period

[2] For the same reasons as discussed in the article, $\hat{m}_w^* = 0$.

military spending corresponds with the amount that minimizes lifetime strong-period expenditures. That is, equilibrium strong-period expenditures are $\hat{m}_s^* + \frac{\delta(R-\sigma\hat{m}_s^*)}{(1-\delta)(1+\hat{m}_s^*)\theta+\delta\sigma(1+\theta)}$ and $G$ chooses $\hat{m}_s$ for every strong period to minimize this exact expenditure function, $\hat{m}_s + \frac{\delta(R-\sigma\hat{m}_s)}{(1-\delta)(1+\hat{m}_s)\theta+\delta\sigma(1+\theta)}$. Therefore, by standard envelope theorem logic the indirect effect of $O$ on $\hat{m}_s^*$ cancels out and $\frac{d}{dO}\left(R - \hat{x}_s^* - \hat{m}_s^*\right)\Big|_{\frac{d\theta}{dO}=0} = 1 - \frac{\partial\hat{x}_s^*}{\partial O}$.[3]

By contrast, in the original model, $G$ chooses $m_t$ in the current period while *holding fixed* the future equilibrium choice $m_s^*$. That is, although equilibrium strong-period expenditures are $m_s^* + \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)}$, in every strong period $G$ chooses current-period military expenditures $m_t$ to minimize $m_t + \frac{\delta(R-\sigma m_s^*)}{(1-\delta)(1+m_s^*)\theta+\delta\sigma(1+\theta)} \cdot \frac{1+m_s^*}{1+m_t}$. This enables $G$ to profitably deviate from the lifetime strong-period expenditure-minimizing choice $\hat{m}_s^*$.[4] In other words, in the original model, the current-period $G$ faces a commitment problem with any future-period governing player (even if that player is itself): $G$'s utility in a peaceful equilibrium would be higher (and the range of $\sigma$ values low enough that fighting occurs would be smaller) if $G$ could commit to spend an amount on the military consistent with minimizing lifetime strong-period expenditures, but deviating to a higher military spending amount in any strong period is profitable because $G$ does not internalize having to choose higher military spending in *all* strong periods (because, in equilibrium, $m_t$ is constant across strong periods). The result is that $1 > -\frac{\partial x_s^*}{\partial m_t}$, meaning the marginal benefit of arming exceeds the marginal cost of arming at $m_s^*$ because of these time-inconsistent incentives.[5] Furthermore, the indirect effect of $O$ on $m_s^*$ does not cancel out and $\frac{d}{dO}\left(R - x_s^* - m_s^*\right)\Big|_{\frac{d\theta}{dO}=0} =$ $1 - \frac{\partial x_s^*}{\partial O} - \left(1 + \frac{\partial x_s^*}{\partial m_t}\right) \cdot \frac{\partial m_s^*}{\partial O}$, which is the numerator of Equation A3. The indirect effect enables the possibility that $\frac{d\overline{\sigma}}{dO} > 0$ for small $R$, even if $m_s^*$ has an interior solution and $\frac{d\theta}{dO} = 0$. However, as proven above, for large enough $R$ the direct effect always dominates the indirect effect.

All the corresponding lemmas and propositions for the modified model can be proven using a similar setup as the Section A.2 proofs. I formally state and prove the corresponding version of Proposition 2, Parts b.1 and b.2 (Part a is nearly identical to the proof above) for the modified model because the comparisons with the proofs from the original model

[3]Recall that $\frac{d}{dO}\left(R - x_s^* - m_s^*\right)\Big|_{\frac{d\theta}{dO}=0}$ is the numerator of Equation A3. Also, Proposition A2, Part b.2 below covers the $\frac{d\theta}{dO} \neq 0$ case.

[4]Formally, the proof for Proposition 1 demonstrated that $G$ has a profitable deviation from any $m_t \neq m_s^*$ in a strong period, including $\hat{m}_s^*$.

[5]It can be shown that $m_s^* = \hat{m}_s^*$ if and only if $\sigma = 0$, which is intuitive given the present explanation. If there will never be another strong period, and therefore setting $m_t$ in the current period is equivalent to setting $m_t$ for all strong periods, then the incentives in the original model are identical to those in the modified model.

are instructive.[6] In particular, they demonstrate that $G$'s ability to commit to the military spending amount that minimizes lifetime strong-period expenditures eliminates the parameter range in which $\frac{d\overline{\sigma}}{dO} > 0$ despite $m_s^*$ having an interior solution (as long as the effect of oil on institutional quality is not negative and large in magnitude).

**Proposition A2. Part b.1.** Assume $\frac{d\theta}{dO} = 0$.

**Case 1.** If $R < R_{m_s^*=0}$, then $\frac{d\hat{\overline{\sigma}}}{dO} = 0$.

**Case 2.** If $R > R_{m_s^*=0}$, then $\frac{d\hat{\overline{\sigma}}}{dO} < 0$.

***Proof of Case 1.*** Same proof as for Proposition 2, Part b.1, Case 1.

***Proof of Case 2.*** The threshold $\hat{\overline{\sigma}}$ is implicitly defined by $\hat{x}_s^*(\hat{\overline{\sigma}}) + \hat{m}_s^*(\hat{\overline{\sigma}}) = R$. Therefore, because $1 + \frac{\partial \hat{x}_s^*}{\partial \hat{m}_s} = 0$ (Equation A7), the term equivalent to $\frac{d\hat{\overline{\sigma}}}{dO}\Big|_{\frac{d\theta}{dO}=0}$ in Equation A3 is:

$$\frac{d\hat{\overline{\sigma}}}{dO}\Bigg|_{\frac{d\theta}{dO}=0} = \frac{1 - \frac{\partial \hat{x}_s^*}{\partial O}}{\frac{\partial \hat{x}_s^*}{\partial \sigma}}. \tag{A8}$$

Substituting $\hat{x}_s^*$ in for $x_s^*$, Lemma A2 implies $1 > \frac{\partial \hat{x}_s^*}{\partial O}$ and the proof for Lemma A1 implies $\frac{\partial \hat{x}_s^*}{\partial \sigma} < 0$, rendering the overall term negative. ■

**Proposition A2. Part b.2.** Assume $\frac{d\theta}{dO} \neq 0$.

**Case 1.** Assume $R < R_{m_s^*=0}$. If $\frac{d\theta}{dO} > 0$, then $\frac{d\hat{\overline{\sigma}}}{dO} < 0$. If $\frac{d\theta}{dO} < 0$, then $\frac{d\hat{\overline{\sigma}}}{dO} > 0$.

**Case 2.** Assume $R > R_{m_s^*=0}$. If $\frac{d\theta}{dO} > \underline{d\hat{\theta}}$, for $\underline{d\hat{\theta}} < 0$ defined in the proof, then $\frac{d\hat{\overline{\sigma}}}{dO} < 0$.

***Proof of Case 1.*** Same as proof for Proposition 2, Part b.2, Case 1.

***Proof of Case 2.*** The threshold $\hat{\overline{\sigma}}$ is implicitly defined by $\hat{x}_s^*(\hat{\overline{\sigma}}) + \hat{m}_s^*(\hat{\overline{\sigma}}) = R$. Therefore, because $1 + \frac{\partial \hat{x}_s^*}{\partial \hat{m}_s} = 0$ (Equation A7), the term equivalent to $\frac{d\hat{\overline{\sigma}}}{dO}$ in Equation A4 is:

$$\frac{d\hat{\overline{\sigma}}}{dO} = \frac{1 - \frac{\partial \hat{x}_s^*}{\partial O} - \frac{\partial \hat{x}_s^*}{\partial \theta} \cdot \frac{d\theta}{dO}}{\frac{\partial \hat{x}_s^*}{\partial \sigma}}. \tag{A9}$$

[6] Even though I do not also formally state a Proposition A1, I call the following statement Proposition A2 to emphasize its relationship to Proposition 2.

Substituting $\hat{x}_s^*$ in for $x_s^*$, Lemma A2 implies $1 > \frac{\partial \hat{x}_s^*}{\partial O}$, the proof for Lemma A5 implies $\frac{\partial \hat{x}_s^*}{\partial \theta} < 0$, and the proof for Lemma A1 implies $\frac{\partial \hat{x}_s^*}{\partial \sigma} < 0$. Therefore, because the denominator is negative, the result holds if the numerator is positive, i.e., if $\frac{d\theta}{dO} > \underline{d\hat{\theta}} \equiv \frac{1-\frac{\partial \hat{x}_s^*}{\partial O}}{\frac{\partial \hat{x}_s^*}{\partial \theta}}$, with $\underline{d\hat{\theta}} < 0$. ■

# B Data Details for Table 4

## B.1 Coding Center-Seeking Civil War Onsets with UCDP/PRIO Data

The coding procedure for UCDP/PRIO data used in Table 4 departs in two major ways from the conventional operationalization procedure for center-seeking civil war onsets. First, it does not code military coups with large death tolls as civil wars. Second, it uses a different procedure than the widely used two-year lapse rule to code onsets.

These choices are consequential. Demonstrating a previously unnoticed trend, 18 of the 34 (53%) center-seeking civil war onsets in country-years producing at least $100 of oil income per capita[7]—as coded by conventional procedures in the Table 3 sample—are either coup attempts or a continuation of an existing civil war. Stated slightly differently, whereas among oil-rich countries a new center-seeking civil war occurs in 2.3% of country-years in Table 3, this figure drops to 1.4% in Table 4.[8] By contrast, the corresponding percentage for oil-poor countries is nearly unchanged between Table 3 (1.7%) and Table 4 (1.6%). Because existing scholarship has not accounted for this crucial trend, it is important to scrutinize existing onset coding rules and assess whether statistical models that appear to support the conflict resource curse are in part driven by over-counting onsets in oil-rich countries.

### B.1.1 Distinguishing Military Coups from Civil Wars

Table B1 shows six UCDP/PRIO "civil wars" in oil-rich country-years in the sample are military coup attempts, but involved a high enough number of deaths to meet UCDP/PRIO's civil war criteria.[9]

[7] Below, "oil-rich country-years" refers to country-years producing at least $100 in per capita oil income.

[8] Table 4 also features considerably fewer country-years by omitting certain countries and by dropping years within ongoing civil wars, as described below.

[9] Iraq 1963 also fits this description, except this observation is dropped from the regressions because of missing income data.

**Table B1. UCDP/PRIO Coup Attempts in Oil-Rich Countries**

| Country | Year |
|---|---|
| Azerbaijan | 1993 |
| Cameroon | 1984 |
| Gabon | 1964 |
| Trinidad and Tobago | 1990 |
| Venezuela | 1962 |
| Venezuela | 1992 |

Results that exclude UCDP/PRIO military coup attempts should be favored for two reasons. First, consistent with the broader literature on oil and civil wars, the formal model in this article implicitly assumes the government has perfect control over its military and derives implications for an interaction between a government and a *non-state* challenger. Therefore, civil wars and military coup attempts are theoretically distinct phenomena. Second, the UCDP/PRIO database only includes military coup attempts with a sufficiently high death toll. This means results including UCDP/PRIO coups implicitly provide insight into how oil wealth affects military coup attempts *that create high death tolls*—but not other coup attempts. At the very least, at present we lack a theoretical defense for why this is a relevant hypothesis.

I consulted two sources to identify which country-years in UCDP/PRIO reached the 25 death threshold because of a military coup attempt. First, I examined whether the "SideB" actor in the UCDP/PRIO Armed Conflict Dataset (ACD)—which, for civil wars, is the rebel group(s)—was a military faction, as well as consulted the accompanying UCDP/PRIO conflict encyclopedia. Second, I used Powell and Thyne's (2011) coup database to match coup and civil war years. I consulted additional sources for any matched years. Almost all of the coup cases are identified as a military faction by UCDP/PRIO. Two additional coup attempts in Sudan in the 1970s were found using Powell and Thyne (2011) and verified using McGowan's (2007) codebook. In FL, two cases in the core sample were identified as coup attempts because the case name in the FL dataset states either "Mil. coup" or "Mil. faction."

### B.1.2 Coding Civil War *Onsets*

Commonly used procedures for translating UCDP/PRIO conflict incidence data into civil war *onsets* raise two types of issues.[10] First, commonly used procedures count multiple

[10] The ACD provides information on conflict incidence (that is, whether or not an incompatibility produces at least 25 battle deaths in a particular year) but does not distinguish between unique civil wars.

civil wars in some cases for what is more naturally thought of as a single war. Second, in other cases conventional procedures count multiple wars as a single one.

Elaborating on the first concern, to convert UCDP/PRIO incidence data into onset data, scholars frequently use a two-year lapse rule to code a conflict year as an "onset." That is, if the 25 death threshold is not met for at least two years after being met in the past, any future year with at least 25 battle deaths is coded as a new civil war. Fearon and Laitin (2013) discuss the conceptual problems with this approach:[11]

> "They apply a criterion of one year (or two, or ten, for different codings) with no conflict above their 25 death threshold. This has the advantage of being relatively definite, but the disadvantage of making many long-running, low level conflicts that flit above and below the 25 dead threshold look like many distinct civil wars. In our view they often are more naturally seen as a single, long-running but low level civil conflict, that happens often by chance to get above or below the threshold in some years" (25).[12]

Exemplifying the conceptual problems posed by using short lapse rules for low-intensity periodic conflicts, Kreutz's (2010) dataset codes five distinct civil wars between Iran and the rebel group MEK (1979-82, 1986-88, 1991-93, 1997, 1999-2001). Throughout this entire period, however, the conflict consisted of hit-and-run bombings by MEK and repressive retaliation by the government. In some years, MEK successfully struck big targets, and in other years they failed to do so (Global Security 2014). Coding 1986, 1991, 1997, and 1999 as onset years for new civil wars conflates conceptual considerations about civil war onset and civil war continuation. At the very least, it is of interest to know how the oil coefficient estimate changes when using a coding procedure that guards against overcounting onsets for periodic conflicts.

Second, the two-year lapse rule may undercount civil war onsets in other cases. For example, consider the UCDP/PRIO Conflict Encyclopedia's description of civil wars in the Democratic Republic of the Congo in the 1990s: "In 1996-1997 an armed rebellion led by AFDL and supported by Rwanda and Uganda managed to topple President Mobutu in May 1997. However the new regime was soon at war again [in 1998], this time against

Kreutz (2010) and Walter (2014) do explicitly code different civil war episodes using UCDP/PRIO data. However, neither of their coding procedures resolve the issues discussed here. Walter (2014) uses a higher death threshold (ACD codes whether at least 1000 battle deaths occurred in addition to the 25 death threshold). Therefore, using her data would depart considerably from the goal here of retaining the 25 death threshold that scholars in the conflict resource curse literature often use, but revising the civil war initiation/termination coding procedure. Kreutz's (2010) coding rule for for distinguishing civil war episodes raises the same conceptual issues as the standard procedure discussed below. He uses an even lower fighting lapse threshold of one year.

[11] See also Sambanis (2004, 818-9).

[12] Fearon and Laitin (2013, 21-6) provide additional discussion about their coding rules as well as how the two-year lapse rule for coding UCDP/PRIO onsets tends to overstate the number of distinct conflicts.

RCD and MLC." Although the violence involved with toppling Mobutu and subsequently to remove Kabila are usually considered two distinct civil wars, the two-year lapse rule does not code a new onset in 1998 because more than 25 battle deaths in a center-seeking conflict occurred in the previous year. Concerns about undercounting civil war onsets are even more relevant for separatist civil wars. The government may be involved with a war in one region while another ethnic group in a different part of the country seeks separation, and hence initiates a clearly distinct civil war.

Two of Fearon and Laitin's (2003) coding rules help to guard against these issues. First, "War ends are coded by observation of a victory, wholesale demobilization, truce, or peace agreement followed by at least two years of peace" (Fearon and Laitin 2003, 76, fn. 4; see this page and footnote for their full set of rules). This directly addresses the concern about overcounting onsets for periodic conflicts because war ends are marked by clear signals of intent to end the current episode of fighting. Importantly, this rule still allows for the possibility of repeated civil wars with the same rebel group. Second, "If a main party to the conflict drops out, we code a new war start if the fighting continues (e.g., Somalia gets a new civil war after Siad Barre is defeated in 1991)." This addresses the problem of undercounting onsets in cases such as the Democratic Republic of the Congo in the 1990s.

I use the following two rules to translate incidence years of at least 25 battle deaths from ACD into distinct civil war episodes. First, I match UCDP/PRIO war years with wars from FL. For wars identified in both datasets, I recode UCDP/PRIO onset year(s) after the first one—as coded using the two-year lapse rule—if FL suggests the two-year lapse rule either overcounts or undercounts civil war onsets. Specifically, if FL codes a series of conflict-years as one war but the two-year lapse rule yields multiple UCDP/PRIO onsets because the war is only periodically active, I recode all two-year lapse UCDP/PRIO onsets after the first year of the conflict to "0." In contrast, if FL codes a series of conflict-years as multiple wars but the two-year lapse rule yields only one onset because a new civil war follows a previous war with at most a one-year gap between hostilities, I use the FL onset years to code additional UCDP/PRIO onsets. In some cases, both UCDP/PRIO (using the two-year lapse rule) and FL agree there were multiple onsets, but two-year lapse UCDP/PRIO codes some non-onset years as onsets because of periodic conflict and codes some onset years as non-onset because distinct civil wars were not separated by at least two years of hostilities. In such cases, the first UCDP/PRIO conflict year is coded as an onset and all subsequent UCDP/PRIO onsets are coded using FL. Finally, if FL code a new center-seeking civil war in a country where one is already occurring, I follow UCDP/PRIO coding procedures by not coding a new onset. The UCDP/PRIO codebook explicitly states that according to their coding rule, "a state can only experience one intrastate conflict over government in a given year" (Themner 2015, 8).

Second, for conflicts included in UCDP/PRIO but not in FL, I apply Fearon and Laitin's two coding rules discussed above. For any center-seeking conflict after the first one—as would be coded using the two-year lapse rule—I first consulted ACD and its accompanying conflict encyclopedia to assess whether the same or a different rebel group participated. I referred to additional secondary sources if the ACD conflict encyclopedia was inconclusive. If it was the same group (or a faction of the original group), the conflict encyclopedia was used to code whether the group had concluded a peace treaty with the government. If so, following Fearon and Laitin, the first conflict-year involving that group at least two years after the peace treaty (if one occurred) was coded as a new civil war onset.

Cases in which casualties drop below the 25 death threshold for several years without any agreement between the government and rebels pose the greatest conceptual difficulties. Consistent with the goal of guarding against overcounting low-intensity periodic conflicts, I use a six-year lapse rule for these cases (specifically, among ones included in UCDP/PRIO but not in FL). Without using any lapse rule, such cases pose an unsolvable right-censoring problem. That is, for any country that has ever experienced a center-seeking civil war, we could never conclude definitely that a civil war is not currently ongoing. If the same rebel group caused 25 deaths sometime in the future, regardless of years elapsed since the last 25 battle-death year, a long-running periodic war would be coded for the entire period if there is no lapse rule. As an empirical example, the Naxalites' center-seeking conflict with the Indian government is coded in ACD from 1969 to 1971—after which no treaty was concluded—and not again until conflict resumed in 1990 (and has lasted through at least 2015). Without applying a lapse rule, India would be coded as engaged in a long-running periodic center-seeking war since 1969 despite fewer than 25 battle deaths in every year for 18 consecutive years. Given that any lapse rule is necessarily arbitrary in the exact choice of years, I specifically chose six years because this incorporates all but two empirical cases in which the same rebel group participated in non-consecutive fighting spells that did not include an intervening peace treaty. India (18 year lapse) and Malaysia (13 year lapse) are the only cases listed below in Table B2 that meet all the criteria for being coded as a single periodic war but are coded as multiple wars because of the six-year lapse rule for cases lacking a peace treaty.

As a final note, this revised procedure for coding civil war episodes does not *exclude* UCDP/PRIO conflicts that are not also coded by FL (nor do I add FL conflicts to UCDP/PRIO for the handful of cases FL codes as a civil war but UCDP/PRIO does not), but instead applies an onset coding rule to ACD that better fits the concept of civil war "onset." Table B2 presents country-by-country scores for all countries coded either by UCDP/PRIO using the two-year lapse rule or by FL as having multiple onsets, and were active post-1960 (because Ross' 2012 sample begins in 1960).

**Table B2. Countries with Multiple Center-Seeking Civil Wars as Coded By Either UCDP/PRIO Two-Year Lapse Rule or FL**

| **Country** | **UCDP Center Episodes (2-yr lapse rule)** | **FL Center Episodes** | **Notes** | **Revised UCDP Center Episodes** |
|---|---|---|---|---|
| Afghanistan | 1978- | 1979-1992; 1992-2001; 2003- | FL codes three distinct conflicts, I use FL onset years for latter two. | 1978-1992; 1992-2001; 2003- |
| Algeria | 1991- | 1962-3; 1992- | FL includes a post-decolonization war, both include 1990s war. | unchanged |
| Angola | 1975*-95; 1998*-2002 | 1975-2002 | FL and UCDP/PRIO include the same conflict (which featured a temporary lapse in hostilities). UCDP/PRIO recoded to only one onset. | 1975-2002 |
| Burundi | 1991- | 1972; 1988; 1993-2006 | Both agree on 1990s conflict, FL includes two earlier conflicts not coded by UCDP/PRIO. | unchanged |
| Cambodia | 1967-75; 1978-98 | 1967-75; 1978-98 | FL and UCDP/PRIO include the same conflicts and agree on onset years. | unchanged |
| Chad | 1966-72; 1976-94; 1997-2002; 2005*- | 1965-; 1992-8 | FL and UCDP/PRIO include the same conflicts, FL onset year not used for a distinct post-Deby civil war because a center-seeking war was already occurring. Recoded to one onset. | 1965- |
| Central African Republic | 2001-2; 2006 | n.a. | Although 2001 was recoded as not a center-seeking war because it was a coup attempt, it was followed in 2002 by a challenge from a non-government group. Distinct rebel group in 2006. | 2002; 2006 |
| Colombia | 1964- | 1948-62; 1963- | FL and UCDP/PRIO agree on conflict by leftist groups in 1960s. FL also codes an earlier conflict. | unchanged |

| Country | UCDP Center Episodes (2-yr lapse rule) | FL Center Episodes | Notes | Revised UCDP Center Episodes |
|---|---|---|---|---|
| Congo, DR | 1964-7; 1977-8; 1996-2001; 2006- | 1977-8; 1996-7; 1998- | UCDP/PRIO conflict in 1960s is not included by FL, all subsequent conflicts are included for both. FL onset years are used for 1990s conflicts. | 1964-7; 1977-8; 1996-8; 1998- |
| Congo, Rep. | 1993*; 1997*-9; 2002* | 1997-9 | UCDP/PRIO and FL both include 1997-9, the period of major fighting that led to government overthrow. UCDP/PRIO encyclopedia refers to 1993 and 2002 as the first and final phases, respectively, of intrastate conflict following 1992-3 elections. | 1993-2002 |
| Djibouti | 1991-4; 1999 | n.a. | Rebel group in early 1990s conflict signed a peace agreement and more than two years of peace followed, implying 1999 is a second war. | unchanged |
| Eritrea | 1997-99; 2003 | n.a. | Periodic conflict with same rebel group. | 1997-2003 |
| Georgia | 1991-3 | 1992-4 | The coup in 1991 is recoded as not a UCDP/PRIO onset. However, a two-year center-seeking challenge began in 1992. This is distinct from the conflict FL codes over Abkhazia (which FL codes as both center and separatist; UCDP/PRIO codes Abkhazia as only separatist). | 1992-3 |
| India | 1969-71; 1990- | 1998- | FL and UCDP/PRIO both code conflicts against communist rebel groups. Although the Naxalites that emerged in the late 1960s also engaged in violence in the 1990s (which is when FL starts coding a civil war because violence increased), a new war is coded in 1990 because there were more than 6 years of inactivity and no intervening peace treaty. | unchanged |

| Country | UCDP Center Episodes (2-yr lapse rule) | FL Center Episodes | Notes | Revised UCDP Center Episodes |
|---|---|---|---|---|
| Iran | 1979*-82; 1986*-88; 1991*-93; 1997*-01; 2005*- | 1978-9 | No agreement between UCDP/ PRIO and FL. UCDP/PRIO: Periodic conflict with MEK 1979-2001. New rebel PJAK in 2005. FL: Iranian Revolution. | 1979-2001; 2005- |
| Iraq | 1982*-4; 1987*; 1991*-6; 2004- | 1991; 2004- | No agreement between UCDP/ PRIO and FL before U.S. occupation. UCDP/PRIO: Periodic conflict with SCIRI. FL: Shi'a uprising after 1991 Persian Gulf war. | 1982-96; 2004- |
| Laos | 1959-73; 1989-90 | 1960-73 | First UCDP/PRIO spell corresponds with FL war. Different challenger in 1989 after Pathet Lao won first civil war and then lost power. | unchanged |
| Liberia | 1989-90; 2000-3 | 1989-96; 2000-3 | FL and UCDP/PRIO include same conflicts and agree on onset years. | unchanged |
| Malaysia | 1958-60; 1974*-5; 1981* | n.a. | All the conflict-years are a continuation of the decolonization struggle by the Communist Party of Malaya, which featured periodic fighting after independence. It is coded as two distinct wars because of at least 6 years of inactivity without an intervening peace treaty. | 1958-60; 1974-81 |
| Nepal | 1960-2; 1996-2006 | 1997-2006 | Second UCDP/PRIO spell corresponds with FL war. | unchanged |
| Nicaragua | 1977-9; 1982-90 | 1978-9; 1981-8 | FL and UCDP/PRIO code the same distinct conflicts. | unchanged |

| Country | UCDP Center Episodes (2-yr lapse rule) | FL Center Episodes | Notes | Revised UCDP Center Episodes |
|---|---|---|---|---|
| Niger | 1991-2; 1997 | n.a. | The original rebel group signed a peace agreement in 1993, but a faction initiated a separatist civil war that was ended by peace agreements in 1994 and 1995. Further rebel fractionalization engendered another year of center-seeking fighting in 1997. Although more than two years lapsed between the peace agreement to terminate center-seeking fighting in 1993 and the next year of at least 25 center-seeking battle deaths (1997), 1997 is not coded as a new onset because a faction from the original rebel group most recently signed a peace agreement in 1995. | 1991-7 |
| Pakistan | 1990; 95-6 | n.a. | Same rebel group involved in periodic conflict. | 1990-6 |
| Peru | 1965; 1982*-99 | 1981-95 | Second UCDP/PRIO spell corresponds with FL war. Distinct left-wing challengers in UCDP/PRIO spells. | unchanged |
| Rwanda | 1990-2002 | 1962-5; 1990-2002 | Both agree on 1990s conflict, FL includes an earlier conflict not coded by UCDP/PRIO. | unchanged |
| Somalia | 1982-96; 2001-2; 2006 | 1981-91; 1991- | FL and UCDP/PRIO include the same conflicts. FL years are used to distinguish pre- and post-1991 conflicts. | 1982-91; 1991- |
| Sri Lanka | 1971; 1989-90 | 1971; 1987-9 | FL and UCDP/PRIO code the same distinct conflicts. | unchanged |

| Country | UCDP Center Episodes (2-yr lapse rule) | FL Center Episodes | Notes | Revised UCDP Center Episodes |
|---|---|---|---|---|
| Sudan | 1983- | 1983-2005 | UCDP/PRIO and FL agree on 1983 war. | unchanged |
| Turkey | 1991-2; 2005 | 1977-80 | No agreement between UCDP/PRIO and FL. The two UCDP/PRIO conflict spells involve distinct leftist rebel groups. | unchanged |
| Uganda | 1972; 1979- | 1981-88; 1988- | FL does not include the short conflict in 1972. Both UCDP/PRIO and FL code the same subsequent wars. FL's distinction between pre- and post-1988 conflict spells is used. | 1972; 1979-88; 1988- |
| Uzbekistan | 1999*-2000; 2004* | n.a. | Second conflict initiated by a splinter group of the first Islamic challenger (see citation for National Counterrorism Center 2014). | 1999-2004 |
| Yemen | 1962-70; 1979-82 | 1962-9; 1994; 2004- | FL and UCDP/PRIO agree on war between republicans and royalists in the 1960s but not the subsequent conflicts. Challenger in second UCDP/PRIO war was a leftist group. FL codes war with former South Yemen in 1994 (FL codes this war as both center and separatist; UCDP/PRIO codes only separatist) and war with al-Houthi in 2004. | unchanged |

*Notes*: Table B2 excludes cases that were determined to be coup attempts. UCDP/PRIO center-seeking conflict episodes are coded using the UCDP/PRIO database. Semi-colons are used whenever there is a multiple-year gap in fighting to denote that a distinct conflict would be coded using the two-year lapse rule. There are several minor discrepancies with Ross' (2012) coding of center-seeking civil wars using the two-year lapse rule, presumably from using a revised version of UCDP/PRIO. FL center-seeking conflict episodes are coded using FL's database, which explicitly codes distinct civil wars that are separated by a semicolon.

* Country produced at least $100 in oil income per capita in a UCDP/PRIO onset year using the two-year lapse rule.

## B.2 Sample

For reasons discussed on pages XX and XX of the article, the samples for the Table 4 regressions exclude two types of country-years that are included in the Table 3 regressions. First, Table 4 excludes countries that joined the OECD in the 1960s and 1970s, which includes Western Europe and offshoots (Australia, Canada, New Zealand, United States) and Japan. However, more recent OECD members—such as South Korea—*are* included for the duration of the sample because of the possibility that their successful late development was endogenous to not having oil. This corresponds directly with dropping all country-years with OECD=1 in Ross' (2012) dataset. For example, his variable codes New Zealand as OECD=1 in all years despite New Zealand not joining the OECD until 1973, and codes South Korea as OECD=0 in all years despite South Korea joining the OECD in 1996.

Second, Table 4 excludes non-sovereign countries. Specifically, never-colonized countries and former Western European colonies achieving independence before 1960 are included in all years. Countries that gained independence after 1960 are included starting from their year of independence. Former Soviet republics are included from 1991 onward. Non-Soviet former Eastern bloc countries and Mongolia are included starting in 1990. Former Yugoslavian countries are included from their year of independence onward. Albania, USSR/Russia, and Yugoslavia/Serbia would be included for the entire period, however, they are effectively dropped for most of the Cold War period because of missing income data. Occupied territories as coded by Geddes, Wright, and Frantz (2014) are excluded: Afghanistan 2001-, Bosnia and Herzegovina 1996-, Dominican Republic 1965-6, Iraq 2003-, and Lebanon 1976-2005.

## B.3 Additional Modifications

### B.3.1 Dropping Years with Ongoing Civil Wars and Generating Temporal Dependence Controls

When estimating models with a binary event as the dependent variable, researchers "typically set ongoing years of the event to zero or instead set ongoing years to missing," albeit usually without including justification for their coding decision (McGrath 2015, 534). McGrath presents econometric theory and simulation results demonstrating that setting ongoing years to missing yields less biased and more efficient results than setting ongoing years to 0, which is why Table 4 uses the former procedure. The basic idea is that the probability of a new civil war is systematically different in years with ongoing civil wars than in peace years, but setting ongoing years to 0 does not detect this distinction.

Additionally, the temporal dependence controls (peace years and cubic splines) are generated from a center-seeking civil war termination variable. A related problem with coding ongoing civil wars as 0 on the onset variable is that researchers typically generate temporal dependence controls from the civil war onset variable, which is problematic because the counters for peaceful years and splines reset even though a war has just begun. Calculating peace years and cubic splines from the last rather than first year of a civil war avoids this problem (and, because years with ongoing center-seeking civil wars are excluded, the peace year and spline variables do not count upward during the war).

#### B.3.2 Imputing Missing Income Data

Ross' (2012) per capita income variable is missing for 500 of the 5555 observations in the core sample. As described in his book and in his STATA replication data, he uses World Development Indicators data when available and data from Penn World Table 6.2 when not. Fortunately, Penn World Table 6.2 does contain *rgdpch* data for 270 of the 500 missing observations. Using data for 2005 and 2006 from Penn World Table 7.1 yields an additional 4 data points.

An additional 131 data points are imputed. Imputing missing income data will undoubtedly introduce some measurement error, although this appears better than the alternative of dropping a large number of country-years. Many countries lack income data in early years in the dataset. To address this, I imputed income data from the first measured year backwards up to 10 years. In sum, these additions to Ross' (2012) dataset recover 405 of the 500 country-years that were originally missing income data.

#### B.3.3 Excluding/Modifying Other Post-Treatment Covariates

Rather than control for annual population data, to avoid post-treatment issues that arise because oil booms tend to raise population (e.g., Cotet and Tsui 2010), I instead control for log population measured in 1950. Note that the issue of *controlling* for pre-oil population—in which the goal is to distinguish the effect of oil from the effect of what population would have been had the country not discovered oil—is distinct from the issue of *dividing* oil income by annual population to generate an oil income per capita variable. For the purpose of assessing the effects of the oil income per capita "treatment," the issue of how much per capita oil income a country would have had if oil had not raised population is not relevant because one typical consequence of gaining the oil treatment is, indeed, raising population.

Additionally, Columns 3 and 4 in Table 4 exclude three of Fearon and Laitin's covariates because they are post-treatment: Polity, Polity squared, and political instability.

### B.3.4 Coding Region Dummies

Ross (2012) includes region dummies for the Middle East and North Africa, Sub-Saharan Africa, Latin America, and East Asia. Even after dropping OECD countries, this still yields a somewhat heterogeneous omitted basis region of Eastern European and former Soviet states, South Asia, and Mediterranean and Pacific islands. To minimize heterogeneity in the basis region I include additional dummies for (1) South Asia and (2) Mediterranean and Pacific islands, which leaves Eastern Europe and former Soviet states as the omitted basis region.[13]

# C Additional Empirical Results

## C.1 Controlling for Counterfactual Non-Oil Income

The key comparative statics prediction from Proposition 2 about the effects of increases in oil revenues is premised on holding non-oil revenues fixed. Implementing this control using statistical modeling, however, is extremely difficult. Controlling for factual income per capita data induces post-treatment bias, and constructing counterfactual income measures poses severe concerns. One author who has attempted this task acknowledges "any possible calculation of counterfactual GDP requires major, perhaps heroic, assumptions" (Herb 2005, 302). There are two important impediments to accurately estimating counterfactual non-oil income per capita data for oil-rich countries. The first is that we have only scant income per capita data prior to initial oil discoveries for most major oil producers. This poses difficulties for estimating a country's pre-oil discovery income per capita. The second is that even with such data, it would still be extremely difficult to estimate how these countries' economies would have evolved over time had they not become major oil producers. This poses difficulties for estimating an oil-rich country's post-oil discovery counterfactual non-oil income per capita. Given the issues raised below, there is no reason to believe regressions that control for a counterfactual non-oil income variable will be less biased than regressions that omit an income control.

Even acknowledging these caveats, however, the results do not support a center-seeking conflict resource curse when controlling for any of four different counterfactual non-oil income estimates. The first procedure uses one data point for each country that estimates pre-oil discovery income, the second procedures uses one data point for each country that estimates post-oil discovery income in 1970, and the last two procedures estimate annual counterfactual non-oil income for the entire time frame. All these results use the coding

[13] I also change the coding of Djibouti to belong to Sub-Saharan Africa, and Israel and Turkey to belong to the Middle East and North Africa.

modifications described in Appendix B. The five columns in Tables C1 through C4 correspond to the odd-numbered specifications from Table 4 except the different counterfactual income per capita variables described below replace the per capita income variable in Table 4.

As one possibility for estimating pre-oil income, Angus Maddison provides the most comprehensive estimates available for historical income per capita. But although some data points go back as far as 0 CE, the global sample only becomes broad starting in 1950. By this date, many major oil-producers had already began oil production and had achieved considerably higher incomes per capita than would have been imaginable had they not become major oil producers (see Table 1 from Alexeev and Conrad 2009, 587). Unfortunately, the next most recent year that has any degree of coverage for the non-European world, 1913, still has considerable missing data (only 30% of the countries from the Table 4 sample have data). However, because this variable is truly pre-oil for almost every country, it may be useful to consider how controlling for this initial income estimate impacts the results after imputing missing data. One possible procedure is, for each country/colony with missing 1913 data, to set its income equal to the lowest income value among countries/colonies in its geographic region. The idea here is that countries with missing data are, on average, more likely to be poor. Table C1 shows the negative correlation remains when controlling for imputed 1913 income per capita.

**Table C1. Counterfactual Non-Oil Income: 1913 Income**

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Panel A. DV: UCDP/PRIO Center CW Onset | | | | |
| Log oil income per capita | −0.057 | −0.066 | −0.054 | −0.146** | −0.087* |
| | (0.046) | (0.048) | (0.052) | (0.073) | (0.048) |
| Country-years | 4790 | 4208 | 4646 | 4083 | 4738 |
| | Panel B. DV: FL Center CW Onset | | | | |
| Log oil income per capita | −0.101 | −0.130* | −0.089 | −0.209** | −0.129* |
| | (0.063) | (0.067) | (0.064) | (0.106) | (0.074) |
| Country-years | 4819 | 4238 | 4675 | 4112 | 4731 |
| | Peace years and cubic splines? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | 1913 income per capita covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Additional covariates/Sample modifications | | | | |
| | None | FL | Region FE | Drop MENA | Drop IRN/IRQ |

*Notes:* Table C1 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 126 and 145 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

Alexeev and Conrad (2009) suggest one possibility for estimating post-oil income per capita. They regress income per capita in 1970 on "strongly exogenous variables" (specifically, region dummies for Latin America and East Asia, and absolute latitude) and use the predicted values for income per capita to estimate pre-oil income. Given constraints on the set of possible strongly exogenous variables, it appears reasonable to focus on regional location because this is known to be a strong predictor of income per capita. Alexeev and Conrad chose the year 1970 because most economic growth regressions use 1970 as the initial year (presumably because Penn World Table data only gains widespread global coverage starting in 1970), but this year is also attractive specifically for estimating counterfactual income for oil-rich countries. Although many oil-rich countries had achieved high incomes per capita by 1970, only after 1973 did oil transform *neighboring* countries' economies as well, particularly in the Middle East and North Africa. Therefore, there are no strong spillover effects from pre-1973 data that impede using neighboring countries' income per capita as counterfactuals for oil-rich countries (which is largely what using regional dummies to generate fitted values achieves). In the regressions in Table C2, each country's counterfactual income is generated using fitted values from the following model (which was itself estimated by regressing log 1970 income on these covariates using OLS):

$$Y_i = 6.065 + 0.024 \cdot \text{ABSLAT}_i + 1.096 \cdot \text{LATIN}_i - 0.108 \cdot \text{EASIA}_i$$

Table C2 shows that although the strength of the negative correlation weakens in all the specifications (except the region fixed effects ones) when controlling for the fitted 1970 income values, the results continue to be more consistent with authoritarianism than conflict resource curse arguments.

**Table C2. Counterfactual Non-Oil Income: Estimated 1970 Income**

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Panel A. DV: UCDP/PRIO Center CW Onset | | | | |
| Log oil income per capita | −0.041 | −0.041 | −0.085 | −0.119* | −0.070 |
| | (0.044) | (0.047) | (0.057) | (0.071) | (0.046) |
| Country-years | 4790 | 4208 | 4646 | 4083 | 4738 |
| | Panel B. DV: FL Center CW Onset | | | | |
| Log oil income per capita | −0.063 | −0.070 | −0.131* | −0.154 | −0.089 |
| | (0.059) | (0.060) | (0.071) | (0.097) | (0.068) |
| Country-years | 4819 | 4238 | 4675 | 4112 | 4731 |
| | Peace years and cubic splines? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Estimated 1970 income per capita covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Additional covariates/Sample modifications | | | | |
| | None | FL | Region FE | Drop MENA | Drop IRN/IRQ |

*Notes:* Table C2 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 126 and 145 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

Additional problems arise when attempting to estimate annual counterfactual non-oil income per capita. One possibility is to subtract oil income per capita from total income per capita to generate non-oil income data. That is, an oil-rich country's actual non-oil income is used to estimate what its income per capita would have been had it not become a major oil producer. Herb (2005) critiques this approach: "Rent or oil wealth can not be subtracted from existing per capita GDP figures, thus 'unmixing' the two types of wealth. The effect of oil (or other rents) on the economies of the rich rentier states is transformative, not additive. The non-oil economy that Kuwait might have had without oil is no longer there. Oil destroyed it" (302). However, this may still provide an attractive strategy—among a set of imperfect alternatives—because the bias generated by this procedure is ambiguous. On the one hand, to the extent that oil creates "Dutch disease" by substituting economic activity away from non-oil industries, the true counterfactual non-oil income may be higher than estimated when subtracting oil income from total income. On the other hand, especially in countries with minimal economic activity prior to oil (e.g., the Arabian peninsula), there was not much of a non-oil economy to substitute away from—implying oil either creates no, or even positive, spillovers. In this scenario, the true counterfactual non-oil income is lower than estimated when subtracting oil income from total income. The implausibly high non-oil income estimates generated by this procedure for many Arabian peninsula countries suggests the latter effect may dominate, i.e., this procedure biases against finding a negative relationship between oil income per capita and center-seeking civil war onset. Still, as Table C3 shows, after controlling for this variable the negative correlations are very similar in magnitude to the coefficient estimates in Table 4.

**Table C3. Counterfactual Non-Oil Income: Actual Non-Oil Income**

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Panel A. DV: UCDP/PRIO Center CW Onset | | | | |
| Log oil income | $-0.058$ | $-0.061$ | $-0.058$ | $-0.139^{**}$ | $-0.086^{*}$ |
| per capita | (0.045) | (0.047) | (0.055) | (0.070) | (0.047) |
| Country-years | 4790 | 4208 | 4646 | 4083 | 4738 |
| | Panel B. DV: FL Center CW Onset | | | | |
| Log oil income | $-0.106^{*}$ | $-0.118^{*}$ | $-0.106$ | $-0.201^{**}$ | $-0.128^{*}$ |
| per capita | (0.063) | (0.064) | (0.073) | (0.098) | (0.071) |
| Country-years | 4819 | 4238 | 4675 | 4112 | 4731 |
| | Peace years and cubic splines? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Non-oil income per capita covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Additional covariates/Sample modifications | | | | |
| | None | FL | Region FE | Drop MENA | Drop IRN/IRQ |

*Notes:* Table C3 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 126 and 145 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01,^{**} p < 0.05,^{*} p < 0.1$.

Herb (2005) uses a different procedure for estimating annual counterfactual non-oil income. Rather than estimate the counterfactual non-oil economy using actual non-oil income data, he instead compares oil-rich countries to selectively chosen baskets of neighbors.[14] Alexeev

[14]See Herb (2005, 302) and his footnote 25 for his coding discussion, which was used to code the counterfactual income variable for Table 4. I changed the estimates for three countries. Botswana was the only country Herb lists as a rentier country that is not oil-rich (see his Table 1 on page 299) and therefore I used Botswana's factual income levels, and I generated counterfactual comparison baskets of countries for Brunei and Equatorial Guinea, which his Table 1 lists as rentier states.

and Conrad (2009, 593) note that Herb uses a "somewhat similar approach" to theirs, although critique it for being "somewhat ad hoc" relative to their "fitted values of GDP [which] are obtained according to a significantly more rigorous procedure." Another problem with Herb's procedure relates to an issue noted above: after 1973, oil-rich countries' neighbors were also transformed by their proximity to oil-rich countries. Although Herb acknowledges this point (302), it is not clear why he dismisses the non-oil income estimation procedure used in Table C3 when the same critique applies to his own procedure. The non-oil economy that the Middle East and North Africa, and other oil-rich regions, might have had without oil is no longer there because oil destroyed it. Furthermore, in contrast to the discussion preceding Table C3, Herb's measure almost certainly biases against finding a negative relationship between oil and center-seeking civil war onset by overestimating oil-rich countries' counterfactual incomes because of remittances and aid to neighboring oil-poor countries (e.g., Jordan). However, contrary to the conflict resource curse hypothesis, Table C4 shows the oil coefficient estimates continue to be negative, although considerably smaller in magnitude than in Table 4.

**Table C4. Counterfactual Non-Oil Income: Estimates from Neighbors**

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | Panel A. DV: UCDP/PRIO Center CW Onset | | | | |
| Log oil income per capita | −0.027 | −0.043 | −0.049 | −0.093 | −0.055 |
| | (0.045) | (0.049) | (0.053) | (0.074) | (0.047) |
| Country-years | 4790 | 4208 | 4646 | 4083 | 4738 |
| | Panel B. DV: FL Center CW Onset | | | | |
| Log oil income per capita | −0.038 | −0.072 | −0.073 | −0.121 | −0.066 |
| | (0.062) | (0.068) | (0.065) | (0.104) | (0.071) |
| Country-years | 4819 | 4238 | 4675 | 4112 | 4731 |
| | Peace years and cubic splines? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Counterfactual non-oil income per capita covariate? | | | | |
| | Yes | Yes | Yes | Yes | Yes |
| | Additional covariates/Sample modifications | | | | |
| | None | FL | Region FE | Drop MENA | Drop IRN/IRQ |

*Notes:* Table C4 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 126 and 145 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

## C.2 Additional Endogeneity Concerns and Country Fixed Effects Models

This article argues the common practice in the conflict resource curse literature of regressing civil war onset on oil income per capita while controlling for total income per capita will produce incorrect counterfactual comparisons—specifically, upwardly biased estimates for the oil coefficient. Other recent contributions focus on a different reason that conventional model specifications may produce incorrect counterfactual comparisons: oil income is not exogenous "manna from heaven" but is instead affected by systematic factors such as fiscal needs (Haber and Menaldo 2011; Menaldo 2014), the form of oil ownership (Jones Luong and Weinthal 2010), and industrialization (Brooks and Kurtz 2015). This section discusses why the results from the article are relevant despite these endogeneity concerns, and presents additional statistical results.

The empirical results from the present article are informative because it was previously believed that regression models estimated with pooled time-series cross-sectional data at the country-level supported a conflict resource curse—in fact, these are the foundational empirical results for the conflict resource curse hypothesis. Regression results considered here demonstrate existing regressions consistently find evidence for a conflict resource curse in large part because they include a theoretically problematic control variable, income per capita (see Table 3 for center-seeking wars, Table C8 for separatist wars, and Table C9 for both types of civil wars). Regarding Table 4, it is striking that statistical models thought to strongly support a conflict resource curse instead more consistently support the *opposite* hypothesis for center-seeking wars when addressing a major theoretical concern and several other seemingly minor specification choices—a distinct concern than country-level oil production being correlated with other determinants of civil war.

Furthermore, although research that critiques models based on pooled time-series cross-sectional variation delivers valuable insights, there are still three crucial issues that require attention in future research.

First, contributions such as Haber and Menaldo (2011), Menaldo (2014), and Brooks and Kurtz (2015) control for country fixed effects to eliminate bias from unobserved time invariant factors. Although reasonable, this strategy also raises econometric concerns. It is well known that including country fixed effects in a logit model in with rare events data will drop many countries from the regressions (Beck and Katz 2001; Wiens, Poast, and Clark 2014). Therefore, country fixed effects models reduce efficiency (Clark and Linzer 2014) and produce biased estimates when treatment effects are heterogeneous (Imai and Kim 2014). Thus, although country fixed effects regressions address one specific concern with models that base identification off pooled time-series cross-sectional variation, pending further analysis it is not clear that results from country fixed effects models should necessarily be preferred over pooled models for the purposes of evaluating the conflict resource curse.

Having acknowledged these limitations, Table C5 presents country fixed effects results for center-seeking civil wars. Because nearly every regressor used in Table 4 is time invariant, C5 presents results from specifications that regress center-seeking civil war onset (either UCDP/PRIO or FL) on oil income per capita, temporal dependence controls, and country fixed effects, as well as incorporate other modifications described in Appendix B. The sample is identical to that in Column 2 of Table 4, except countries that never experience a center-seeking civil war are dropped because their country fixed effect perfectly predicts the outcome. The results continue to support oil-authoritarianism implications more than oil-civil war arguments, and the UCDP/PRIO specification achieves statistical significance despite losing considerable statistical power by dropping 63% of the country-year observations (the Fearon regression drops 78%).

**Table C5. Country Fixed Effects Regressions**

| | (1) | (2) |
|---|---|---|
| | DV: UCDP/PRIO | DV: FL |
| Log oil income per capita | $-0.613^{***}$ | $-0.209$ |
| | (0.182) | (0.253) |
| Country-years | 1794 | 1084 |
| | Country fixed effects? | |
| | Yes | Yes |
| | Peace years and cubic splines? | |
| | Yes | Yes |

*Notes:* Table C5 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The two specifications contain at least one country-year observation from between 37 and 58 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

Second, some recent work analyzes the effects of oil field *discoveries* rather than annual oil income.[15] Although the instruments and matching strategies used in these works produce perhaps the strongest identification strategies in the literature (see each article/manuscript for a lengthier discussion), oil discoveries do not provide a relevant "treatment" variable for empirically assessing the theory presented in this article. Scrutinizing the effects of

[15]These analyses have reached mixed findings. Lei and Michaels 2014 find a strong positive effect of oil discoveries on civil war onset, Cotet and Tsui 2013 report null results, and Blair 2014 shows onshore oil discoveries in densely populated areas raise the probability of separatist civil wars but not the probability of center-seeking wars.

oil discoveries is only appropriate for learning about the short-term effects of oil production. For example, Lei and Michaels' (2014) core results analyze civil war onset within 4, 6, and 8 years of a major oil field discovery. The theory presented here assumes the government has consolidated control over the country's oil production—which corresponds with much existing research—but this assumption may not be valid within a small window after major oil discoveries. Therefore, aggregate oil production is the more theoretically relevant variable for evaluating the debate between oil-authoritarianism and oil-civil war mechanisms.

Third, an even smaller body of work endogenizes annual oil production to systematic factors (Menaldo 2014; Brooks and Kurtz 2015). These two contributions reach contradictory conclusions about non-geological causes of oil production. Whereas Brooks and Kurtz (2015) argue and provide evidence that more industrialized countries tend to produce more oil, Menaldo (2014) shows revenue-starved rulers in weak states have often provided strong incentives for international oil companies to extract oil at above-optimal rates. Additional research is needed to reconcile these competing arguments. As a preliminary consideration, among the weakly institutionalized and mainly authoritarian states analyzed here, many countries that provide strong support for the article's theory by having high levels of oil production and no major center-seeking civil war onsets in the time period analyzed—including Bahrain, Brunei, Equatorial Guinea, Gabon, Kuwait, Libya, Qatar, and Saudi Arabia—had minimal industrial capacity prior to becoming oil-rich (see the discussion accompanying assumption #2). Instead, international oil companies facilitated oil discoveries and production. Therefore, it does not appear likely that the conditions needed for the oil coefficient estimates from Table 4 to be negatively biased—oil-rich countries in the sample had higher industrialization levels prior to major oil discoveries—are true.

### C.3 Sample Alterations

Re-running the regressions on two subsamples of the main dataset provides a "hard" test for the present argument. Following Ross' (2012) arguments for where the conflict resource curse should be strongest, Table C6 only includes post-1990 data points and Table C7 only includes country-years with less than $5,000 in income per capita. The two main conclusions from Table 4 hold under each alteration: the oil coefficient estimate is negative in regressions that omit income per capita, and dropping the income control greatly impacts the estimates. However, the magnitude of the negative effect estimate varies across the two sample alterations. These specifications incorporate the modifications described in Appendix B.

Ross' (2012, 154) disaggregated data analysis shows oil wealth appears to exert a stronger positive effect on civil war onset from 1990 onward. Table C6, however, shows the sign of the oil coefficient estimate remains negative for center-seeking civil war onsets when excluding the income control. Comparing the magnitude of the coefficient estimates in the even-numbered specifications to the corresponding Table 4 regression, the UCDP/PRIO estimates are each larger in magnitude in the post-1990 sample whereas most of the FL estimates are smaller. Notably, dropping three decades of data diminishes the sample size considerably.

**Table C6. Only Post-1990 Years**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO Center-Seeking Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.016 | −0.131 | 0.023 | −0.110 | −0.023 | −0.083 | 0.003 | $-0.204^{*}$ | −0.013 | $-0.164^{*}$ |
| | (0.095) | (0.084) | (0.111) | (0.088) | (0.107) | (0.087) | (0.118) | (0.121) | (0.099) | (0.089) |
| Country-years | 2077 | 2077 | 1732 | 1732 | 1798 | 1798 | 1814 | 1814 | 2066 | 2066 |
| | Panel B. Dependent Variable: FL Center-Seeking Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.225 | −0.054 | 0.168 | −0.054 | 0.138 | −0.017 | 0.016 | −0.276 | 0.202 | −0.098 |
| | (0.149) | (0.111) | (0.162) | (0.124) | (0.157) | (0.104) | (0.228) | (0.233) | (0.170) | (0.128) |
| Country-years | 2124 | 2124 | 1667 | 1667 | 1419 | 1419 | 1853 | 1853 | 2093 | 2093 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C6 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 96 and 140 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

Ross (2012, 154) also provides evidence the conflict resource curse is stronger in poor countries, i.e., those with income per capita below $5,000. Intriguingly, the model anticipates this prediction in two different ways. First, low income is partly endogenous low amounts of oil. The smaller the amount of oil, the higher the probability of center-seeking civil war should be. Second, to the extent that oil wealth does not translate effectively into income per capita or government revenues, again, the higher the probability of center-seeking civil war should be. Therefore, we should not expect to find a strong negative relationship in this subsample even if oil does exert an overall effect of reducing center-seeking civil war propensity. Strikingly, the coefficient estimates remain negative although, as expected, the magnitude of the negative coefficient estimates is smaller in even-numbered specifications in Table C7 compared to those in Table 4.

## Table C7. Only Poor Countries

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO Center-Seeking Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.001 | −0.028 | −0.027 | −0.039 | −0.012 | −0.021 | −0.106 | −0.152* | −0.037 | −0.070 |
| | (0.062) | (0.059) | (0.063) | (0.059) | (0.067) | (0.065) | (0.092) | (0.088) | (0.068) | (0.063) |
| Country-years | 4048 | 4048 | 3651 | 3651 | 3948 | 3948 | 3642 | 3642 | 4000 | 4000 |
| | Panel B. Dependent Variable: FL Center CW Onset | | | | | | | | | |
| Log oil income per capita | 0.009 | −0.057 | −0.041 | −0.081 | −0.017 | −0.41 | −0.127 | −0.207* | −0.018 | −0.087 |
| | (0.078) | (0.070) | (0.080) | (0.072) | (0.075) | (0.070) | (0.126) | (0.118) | (0.091) | (0.084) |
| Country-years | 4123 | 4123 | 3726 | 3726 | 4023 | 4023 | 3671 | 3671 | 4039 | 4039 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C7 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 123 and 135 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

### C.4 Results for Separatist and All Civil Wars

The argument that controlling for per capita income induces post-treatment bias should also apply to regressions that use either separatist wars only or all civil wars as the dependent variable. The four tables presented below support this implication. Tables C8 and C9 present the same specifications as Table 3 except they change the dependent variable (and calculate the temporal dependence controls from that variable). That is, they do not incorporate the modifications discussed in Appendix B. Tables C8 and C9 demonstrate a strong positive relationship between oil and civil war onset in regressions that control for per capita (odd-numbered columns) but either a weak positive correlation or a negative correlation in regressions that omit the per capita income control (even-numbered columns).

### Table C8. Oil Wealth and Separatist Civil War Onset, Existing Models

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO Separatist Civil War Onset | | | | | | | | | |
| Log oil income per capita | $0.158^{***}$ | 0.022 | $0.141^{**}$ | 0.011 | $0.178^{***}$ | 0.059 | $0.170^{**}$ | 0.019 | $0.137^{**}$ | −0.004 |
| | (0.060) | (0.056) | (0.059) | (0.053) | (0.060) | (0.057) | (0.068) | (0.068) | (0.067) | (0.064) |
| Country-years | 6426 | 6426 | 5538 | 5538 | 5211 | 5211 | 5771 | 5771 | 6351 | 6351 |
| | Panel B. Dependent Variable: FL Separatist Civil War Onset | | | | | | | | | |
| Log oil income per capita | $0.133^{*}$ | −0.011 | 0.049 | −0.029 | 0.109 | −0.013 | 0.135 | −0.031 | 0.087 | −0.062 |
| | (0.073) | (0.068) | (0.071) | (0.065) | (0.085) | (0.074) | (0.089) | (0.085) | (0.082) | (0.074) |
| Country-years | 6426 | 6426 | 5538 | 5538 | 6426 | 6426 | 5771 | 5771 | 6351 | 6351 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C8 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 150 and 169 countries, among a broad global sample of oil and non-oil producers. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

### Table C9. Oil Wealth and All Civil War Onset, Existing Models

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO All Civil War Onset | | | | | | | | | |
| Log oil income per capita | $0.133^{***}$ | 0.010 | $0.109^{***}$ | 0.026 | $0.119^{***}$ | 0.027 | $0.124^{***}$ | −0.009 | $0.100^{***}$ | −0.025 |
| | (0.038) | (0.036) | (0.038) | (0.034) | (0.040) | (0.037) | (0.043) | (0.043) | (0.038) | (0.035) |
| Country-years | 6426 | 6426 | 5538 | 5538 | 6426 | 6426 | 5771 | 5771 | 6351 | 6351 |
| | Panel B. Dependent Variable: FL All Civil War Onset | | | | | | | | | |
| Log oil income per capita | $0.115^{**}$ | −0.038 | 0.074 | −0.032 | 0.074 | −0.044 | 0.049 | $-0.112^{*}$ | 0.060 | $-0.094^{*}$ |
| | (0.056) | (0.050) | (0.059) | (0.051) | (0.054) | (0.046) | (0.061) | (0.059) | (0.055) | (0.049) |
| Country-years | 6426 | 6426 | 5538 | 5538 | 6426 | 6426 | 5771 | 5771 | 6351 | 6351 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C9 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 150 and 169 countries, among a broad global sample of oil and non-oil producers. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

The discussion accompanying assumption #6 also implies location should condition the relationship between oil wealth and separatist civil wars. Therefore, because standard cross-national oil measures do not take location into account, oil income should not correlate systematically with either separatist war onset or all civil war onset in cross-national regressions. Tables C10 and C11 demonstrates the inconsistent relationship from Tables C8 and C9 remains in models that impose the modifications described in Appendix B although, notably, the sign of the oil coefficient is negative in every Table C11 regression that omits the income control.[16]

**Table C10. Oil Wealth and Separatist Civil War Onset, Revised Models**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO Separatist Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.109 (0.075) | 0.050 (0.068) | 0.059 (0.085) | 0.032 (0.075) | 0.043 (0.077) | 0.033 (0.074) | 0.093 (0.087) | 0.033 (0.085) | 0.043 (0.080) | −0.011 (0.076) |
| Country-years | 4982 | 4982 | 4398 | 4398 | 3800 | 3800 | 4337 | 4337 | 4943 | 4943 |
| | Panel B. Dependent Variable: FL Separatist Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.076 (0.081) | 0.020 (0.072) | 0.026 (0.086) | −0.009 (0.080) | 0.017 (0.079) | 0.003 (0.077) | 0.030 (0.088) | −0.029 (0.083) | −0.010 (0.081) | −0.063 (0.073) |
| Country-years | 4923 | 4923 | 4341 | 4341 | 4923 | 4923 | 4271 | 4271 | 4878 | 4878 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C10 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 124 and 142 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

[16] Implementing the Appendix B modifications required recoding repeated UCDP/PRIO civil war onsets for separatist wars using the guidelines discussed in Section B.1.2. The case-by-case onset codings are available upon request.

**Table C11. Oil Wealth and All Civil War Onset, Revised Models**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| | Panel A. Dependent Variable: UCDP/PRIO All Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.060<br>(0.045) | −0.016<br>(0.040) | 0.015<br>(0.050) | −0.040<br>(0.041) | 0.022<br>(0.046) | −0.016<br>(0.042) | 0.020<br>(0.057) | −0.054<br>(0.055) | 0.027<br>(0.044) | −0.047<br>(0.039) |
| Country-years | 4454 | 4454 | 3883 | 3883 | 4454 | 4454 | 3848 | 3848 | 4421 | 4421 |
| | Panel B. Dependent Variable: FL All Civil War Onset | | | | | | | | | |
| Log oil income per capita | 0.039<br>(0.057) | −0.042<br>(0.050) | −0.002<br>(0.059) | −0.057<br>(0.053) | −0.008<br>(0.055) | −0.038<br>(0.051) | −0.002<br>(0.068) | −0.092<br>(0.064) | −0.000<br>(0.059) | −0.081<br>(0.052) |
| Country-years | 4543 | 4543 | 3972 | 3972 | 4543 | 4543 | 3916 | 3916 | 4499 | 4499 |
| | Peace years and cubic splines? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Log population covariate? | | | | | | | | | |
| | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| | Income per capita covariate? | | | | | | | | | |
| | Yes | No | Yes | No | Yes | No | Yes | No | Yes | No |
| | Additional covariates/Sample modifications | | | | | | | | | |
| | None | None | FL | FL | Region FE | Region FE | Drop MENA | Drop MENA | Drop IRN/IRQ | Drop IRN/IRQ |

*Notes:* Table C11 summarizes a series of logistic regressions by presenting the coefficient estimate for log oil income per capita (lagged one year) and the standard error in parentheses. Other coefficient estimates are suppressed for expositional clarity. The various specifications contain at least one country-year observation from between 121 and 141 countries, among a broad global sample of oil and non-oil producers that excludes OECD countries and occupied countries. The table incorporates the coding decisions discussed in Appendix B. $^{***}p < 0.01, ^{**}p < 0.05, ^{*}p < 0.1$.

# D Case Evidence

In addition to regression evidence that rejects a conflict resource curse for center-seeking civil wars, it is also striking how few major center-seeking wars have occurred in oil-rich countries and that most of these conflicts do not strongly suggest a conflict-inducing role for oil. Table D1 presents a $2 \times 2$ tabulation for whether or not oil income per capita is above \$100 in a given country-year, and whether or not a FL center-seeking conflict begins. Among cases with oil income per capita above the \$100 threshold in the revised sample (see Appendix B), a center-seeking civil war began in only eleven country-years in FL's dataset. Oil income per capita in the onset year is listed in parentheses.

**Table D1.** $2 \times 2$ **Table of Oil and FL Center-Seeking Civil War Onset**

| | **Onset = 0** | **Onset = 1** | **Percentage** |
|---|---|---|---|
| **Oil = 0** | 3804 observations | 45 observations | 1.17% |
| **Oil = 1** | 1129 observations | Algeria 1962 ($147)<br>Algeria 1992 ($431)<br>Angola 1975 ($321)<br>Argentina 1973 ($108)<br>Congo, Rep. 1997 ($556)<br>Iran 1978 ($1,763)<br>Iraq 1991 ($142)<br>Peru 1981 ($295)<br>Syria 1979 ($553)<br>Yemen 1994 ($150)<br>Yemen 2004 ($246) | 0.96% |

The possible causal role for oil on civil war onset is unclear in four cases. Civil wars broke out in Argentina in 1973 and Peru 1981 in two countries that only recently crossed the $100/capita threshold. Qualitative sources do not suggest oil motivated the left-wing insurgencies in either case (Skidmore and Smith 2005, 94-8 and 214). Additionally, civil wars that "began" in the first year of independence in Algeria and Angola were continuations of anti-colonial struggles. Oil likely played an important role in the *continuation* of conflict in Angola (Le Billon 2007), but this is distinct from oil contributing to civil war *initiation*. Oil perhaps could have contributed to France's and Portugal's attempts to maintain colonial control of Algeria and Angola, respectively. This is an intriguing possibility for future research to analyze. Still, even if oil did contribute to the post-independence civil wars in these two cases, it did so through a channel that lies outside the broad array of existing theories that connect oil to civil war initiation.

Perhaps surprising, qualitative evidence from four other oil-rich/civil war cases highlights the coercive possibilities afforded by oil. Uprisings occurred after Iraq lost the Persian Gulf War. The Iraqi government had defeated the uprisings by the end of 1991, backed by its large and modern army (Fearon and Laitin 2006, 11). The Syrian army grew rapidly in size during the 1970s (Correlates of War dataset), and took the additional step in 1979 of arming party members to root out urban guerrillas (Seale 1988, 327). The government was successful at recruiting citizens to fight against the rebels (327), and the military eradicated the Muslim Brotherhood in 1982 by shelling the city of Hama for a month. The Algerian war lasted considerably longer than the previous two, at nearly a decade, but Fearon and Laitin (2006a, 27-9) stress the army was strong in this case as well. Even in Iran, where the government actually fell, Skocpol (1982, 270) asserts that "all of the [many] vulnerabilities of the prerevolutionary Iranian regime could well have had little significance. The Shah, after all, had both munificent wealth and ominous repressive power at his disposal."

The final two countries, which account for three civil war onsets, exhibit different patterns. Yemen was weak, but Yemen also had a long history of state weakness and civil wars prior to discovering oil in the 1980s—plus, Yemen does not have much oil. Furthermore, it is not clear why FL code the 1994 onset as including center-seeking aims because the goal of the rebel leaders was to recreate the state of South Yemen that had existed until four years prior. Congo-Brazzaville provides the most notable exception. Rebels' desire to capture oil wealth clearly motivated a fight that eventually led to government overthrow. However, even this case lies outside the scope of the formal model because "booty futures" finance from an international oil company (contrary to Assumption #1 from the article about government control of oil revenues) proved crucial to the rebels' success (Ross 2012, 174-6).

# E Fearon and Laitin's *Relative* State Weakness Hypothesis

To support the argument that existing theoretical discussions often conflate overall and relative resource curse hypotheses, pages XX and XX of the article discuss how scholars often misinterpret Fearon and Laitin's (2003) influential hypothesis as concerning the overall effects of oil rather than the effects of oil relative to other types of resources. Fearon and Laitin argue: "Oil producers tend to have weaker state apparatuses than one would expect *given their level of income* because the rulers have less need for a socially intrusive and elaborate bureaucratic system to raise revenues—a political 'Dutch disease' " [emphasis added] (Fearon and Laitin 2003, 81). Their case study on Iraq reiterates: "To be exact, the effect [of oil] is (at least in our theory) conditional on level of income. One can think of oil as having two effects. On the one hand, it means greater per capita income, which may favor peace by giving the state more resources to deal with rebellion. But on the other hand, *given the level of income*, for purposes of avoiding civil war it would be better to come from income taxes than oil because this signifies greater state capability, penetration, and information on the population" [emphasis added] (Fearon and Laitin 2006, 2). Scholars citing their work, however, usually drop the caveat about conditioning on the level of income:

- "The 'weak state' mechanism draws on the harmful effects of resource abundance on the quality of state institutions (corruption, clientelism), which in turn makes internal violent conflict more likely (see Fearon, 2005; Fearon & Laitin, 2003)" (Basedau and Lay 2009, 759).
- "Fearon and Laitin's theoretical approach pits frustrations and inequalities along

ethnic lines against their own favored explanations, which highlights insurgency as an unemotional technology that relies on the state's weakness in peripheral areas and on the corrupting influence of oil production" (Cederman, Gleditsch, and Buhaug (2013, 16).

- "A related mechanism is the weak state mechanism, through which oil extraction weakens the state because governments that rely on natural resources have less incentive to create strong bureaucratic institutions (Fearon and Laitin 2003)" (Cotet and Tsui 2013, 52).
- "One explanation consistent with this correlation is that countries with oil wealth have less incentive to build state capacity—"a political 'Dutch Disease'," and the resulting lower state capacity increases the likelihood of civil war (Fearon and Laitin, 2003; Fearon, 2005)" (Glynn 2009, 1).
- "Moreover, resource abundance or dependence may result in the weakening of state institutions since resource wealth typically relieves governments from establishing a socially intrusive apparatus for levying taxes. This in turn often implies a level of bureaucratic capacity that is too low to manage social peace (Auty, 2001,a,b; Torvik, 2002; Fearon & Laitin 2003; Snyder and Bhavnani, 2005)" (Koubi et al. 2014, 232).
- "In addition, abundant resources may have a detrimental effect on a state's fighting capacity, as argued, for example, by Fearon and Laitin (2003)" (Lujala 2009, 52).
- "Some scholars have suggested that oil-producing states may systematically have lower state capacity (e.g., Karl 1997; Fearon and Laitin 2003) ..." (Morrison 2012, 18).
- "Fearon and Laitin, among others, have suggested a fourth mechanism: that resource wealth—in particular, oil—causes 'state weakness,' which in turn increases the probability of civil war" (Ross 2004, 42).
- "Fearon (2005) provided a partial test of the effect of state capacity. Rather than directly using a proxy of state capacity, he argued that the reliance on lootable resources signaled the weakness of a state" (Sobek 2010, 269).
- "Fearon and Laitin provide an alternative perspective ... Oil revenues are causally significant not because they can finance insurgency, but because they weaken state institutions" (Waldner and Smith 2015).

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