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Political Donations, Public Procurement and Government Efficiency

Abstract

Firms' political donations can induce distortions in the allocation of public procurement contracts. In this article, we employ an advanced non-parametric efficiency model to study the public sector (cost) efficiency implications of such distortions. Using a unique dataset covering the Czech regions over the 2007-2014 period, we find that the efficiency of public good provision is lower when a larger share of public procurement contracts is awarded to firms donating to the party in power ('party donors'). We link this efficiency difference to two underlying mechanisms: i.e. shifts in procurement contract allocations from firms with previous procurement experience to party donors, and the use of less restrictive allocation procedures that benefit party donors.

JEL-Codes: H570, D720, C230.

Keywords: political connections, non-parametric efficiency analysis, benefit-of-the-doubt.

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1. Introduction

Public procurement contracts awarded by public-sector institutions account for 15 to 20% of GDP in most developed countries (OECD, 2013). Given their large share in the economy, recent literature increasingly investigates sources of potential distortions in the allocation of such contracts away from the normative principles that ideally drive them (Goldman et al., 2013; Coviello and Mariniello, 2014; Palguta and Pertold, 2017; Baltrunaite, 2019; Titl and Geys, 2019). The subsequent question about such distortions' impact on the nature and quality of public good provision has received little attention. From both an academic and public interest point of view, however, it is essential to understand whether and why distortions in the allocation of procurement contracts have broader economic implications.

Using novel methodological tools, we address this research gap by studying the public sector (productive) efficiency implications of favouritism in procurement processes towards politically connected firms (Goldman et al., 2013; Baltrunaite, 2019; Titl and Geys, 2019).[1] Theoretically, favouritism towards politically connected firms causes that the 'best' firm is not necessarily the winner or recipient of procurement contracts. That is, those favoured by political connections might not provide more or better services compared to those failing to get contracts due to a lack of political ties (Cingano and Pinotti, 2013; Fisman and Wang, 2015; Geys, 2017).[2] The ensuing misallocation of resources can induce provision at excessive cost compared to the situation where contracts are allocated optimally (i.e. to the most competitive bidders). This is a matter of productive efficiency (i.e. maximum production at minimum cost), and directly leads to the proposition that distortions in procurement contract allocations due to firms' donations to political parties may have a negative impact on public sector efficiency.

Furthermore, we propose two mechanisms that can explain this effect on observed efficiency levels. The first builds on the idea that distortions in public procurement contract allocations are affected by the design of the relevant decision-making rules – and, more specifically, by the discretionary power of public officials in this process (Coviello and Mariniello, 2014; Coviello et al., 2018; Palguta and Pertold, 2017; Titl and Geys, 2019). The underlying argument is that such discretionary power increases "the risk that dishonest officials will collude with some suppliers and ultimately misallocate public resources" (Palguta and Pertold, 2017, p. 294). The use of procedures with extensive discretionary power thus may increase preferential treatment of connected firms and lead a higher share of procurement contracts to be awarded to such firms. This, as argued above, is expected to be associated with a decline in the efficiency of public good provision. Such mechanism is consistent with, for instance, empirical evidence documenting a decrease in corruption in Argentina after external controls and reduced discretion of officials was introduced in the procurement of medical goods (Di Tella and Schargrodsky, 2003).

Our second mechanism reflects the fact that public procurement contract allocations are a zero-sum game: i.e. if one supplier is offered the contract, all others lose out on this contract. Awarding contracts to donating firms then pushes out alternative suppliers. If these alternative suppliers have previous procurement experience (which we will refer to as 'frequent suppliers'), this might depress public sector efficiency. The reasons is that,

[1] We understand political connections broadly as any link – whether through personal ties, board memberships or financial transactions (including ownership stakes or donations) – between politicians and private-sector firms. In the empirical analysis, we focus more specifically on ties established through direct donations by firms to political parties. We therefore use the terms donating firms and connected firms interchangeably throughout the remainder of the paper.

[2] Such a negative link between political connections and service provision can easily be formalised in a selection model with positive substitutability between political connections and firm quality. We refrain from this formalisation to preserve space.

from a theoretical perspective, such 'frequent suppliers' can be expected to provide services faster and more efficiently due to their experience in dealing with, and providing services to, the public sector (Coviello et al., 2018).

In order to asses these theoretical propositions, we exploit a unique new dataset covering poliitcal donations, public procurement contracts and public good provision in the Czech regions over the period 2007-2014. We start by calculating the productive efficiency of the Czech regional governments. In our application, efficiency corresponds to a relative measure relating outputs (measured by indicators of public good provision discussed in detail below) to employed inputs (measured as public procurement spending) for each regional government. We estimate this using an advanced non-parametric robust conditional efficiency model (Daraio and Simar, 2005, 2007a; De Witte et al., 2013). To reduce the number of output indicators in the model and avoid the curse of dimensionality, we proceed in two steps to obtain efficiency scores. First, we combine a wide range of output measures into a composite output indicator for three key policy domains covering the vast majority of public goods produced by the Czech regions. Second, we use these composite indicators as outputs in the efficiency model. This two-step approach avoids the (strong) assumption of an overall production technology applicable to distinct public goods, and allows for different production functions in each policy area. Our robust and conditional efficiency model furthermore possesses several convenient properties for the analysis. Its fully non-parametric nature mitigates specification bias as no assumptions are made about the underlying production function (Yatchew, 1998), while the 'robust' order-m specification of the model corrects for measurement errors and outlying observations (Cazals et al., 2002). Finally, the 'conditional' specification allows adjusting the efficiency scores for heterogeneity across jurisdictions by comparing each region to other observations sharing a similar environment (which builds on insights of Banker and Morey, 1986). Although often ignored in earlier work on public sector efficiency (inducing inference problems due to unobserved heterogeneity across units, see Narbón-Perpiñá and De Witte, 2018a,b), this approach provides the critical opportunity to examine the link between productive efficiency and regions' dependency on politically connected firms in public procurements.

Our main findings show that a larger share of public procurement contracts awarded by the Czech regions to politically connected suppliers – operationalised via firms' donations to political parties – is associated with lower efficiency of regional public good provision. We also provide evidence indicating that our two proposed mechanisms help explain this efficiency difference. On the one hand, favouring political donors in procurement allocations reduces the efficiency of the public sector as it implies the exclusion of other more experienced suppliers. On the other hand, a larger share of contracts assigned under less restrictive allocation procedures is associated with an increase in politically connected suppliers, which are characterised by lower efficiency.[3] Taken together, these findings have important implications for the design and oversight of procurement allocation processes, and also highlight the relevance of our methodological approach for managers within oversight bodies in the public sector. Furthermore, our findings may have significant macro-level implications since government (in)efficiency has previously been shown to impact the aggregate efficiency of the economy (Méon and Weill, 2005) and overall productivity growth (Olson et al., 2000).

Our analysis contributes to two main strands of literature. The first literature studies the level and determinants of public sector efficiency (for a recent review, see Narbón-Perpiñá and De Witte, 2018a,b).[4] A wide range

[3]This finding is consistent with a considerable literature highlighting a negative association between political accountability and corruption (Adsera et al., 2003; Lederman et al., 2005; Yan and Oum, 2014).

[4]A substantial part of this vast literature concentrates explicitly on local governments' (in)efficiency. This includes studies from,

of potential drivers of observed variation in efficiency across jurisdictions has been brought forward including, for instance, political competition (Ashworth et al., 2014; Sørensen, 2014), voter turnout and civic engagement (Borge et al., 2008; Geys et al., 2010), and the presence of direct democratic citizen initiatives (Asatryan and De Witte, 2015). To the best of our knowledge, no study has investigated whether firms' political donations have implications for public sector efficiency due to their distortive effect on the allocation of public procurement contracts. The second relevant literature studies the link between political discretion and the outcome of public procurement processes. Coviello and Mariniello (2014) find that more stringent publicity requirements increase the number of bidders, which in turn reduces the costs of public procurements by rationalizing public spending. Using Czech data, Palguta and Pertold (2017) find that higher political discretion in public procurement processes is associated with increased contract prices especially when contracts are allocated to anonymously owned companies. Importantly, these analyses do not examine the efficiency implications induced by such distortions in public procurement allocations.

The remainder of the paper is organized as follows. The next section discusses the institutional setting in the Czech Republic and the data employed in our analysis. Section 3 presents the methodology for computing conditional efficiency scores and assessing the influence of politically connected suppliers on the efficiency of public good provision. Section 4 summarizes our main findings. Finally, in Section 5, we provide a concluding discussion and some avenues for further research.

2. Institutional setting and data

2.1. Institutional setting

Our analysis focuses on the Czech regional governments, which were devised in 1997 (Act no. 347/1997 Coll.) and have been functioning since January 2000. There are 13 regions (plus the capital of Prague, which constitutes its own region) that have considerable competences in economic policies including transport, education, health care and regional development or tourism (Hooghe et al., 2016). The political power within each region is concentrated in the Regional Council ("Zastupitelstvo kraje"; henceforth 'Council') and the Board of Councillors ("Rada kraje"; henceforth 'Board'). The Council is the legislative body of a region, and is elected every four years using a system of proportional representation. The Board is the executive body of a region, and its members are appointed by – and selected among the councillers of – the parties holding a majority in the Council. Both public bodies are chaired by the *Hejtman*, which is a position equivalent to a state Governor in the US setting. The Boards and *Hejtmans* are directly accountable to the Councils (Act no. 129/2000 Coll.).

The Czech Republic provides two critical advantages for our analysis. First, although the general framework for public procurement is established in national legislation (Act No. 137/2006 Coll. on Government Procurement), local policy-makers have significant influence on the procurement allocation process – and the civil servants administering this process – via a number of mechanisms. This is crucial if we want to quantify the efficiency implications of politically driven distortions in the procurement allocation process. In practice,

for instance, Belgium (Geys and Moesen, 2009; Ashworth et al., 2014), Germany (Geys et al., 2010; Kalb, 2010; Geys et al., 2013), Norway (Borge et al., 2008; Sørensen, 2014; Helland and Sørensen, 2015), Portugal (Cordero et al., 2017a), Spain (Arcelus et al., 2015; Pérez-López et al., 2015) and the United Kingdom (Revelli, 2010; Andrews and Entwistle, 2015). There is also a smaller but growing literature using data from Central and Eastern Europe, which includes studies about Croatia (Alibegovic et al., 2013), Czech Republic (Šťastná and Gregor, 2015), Russia (Hauner, 2008), and Slovenia (Pevcin, 2014).

this influence first of all works through contracting authorities' ability to (mis)use the detailed prerequisites set out in the legislative framework. Examples of such activities include the imposition of unnecessarily rigid technical requirements, the need to have specific certificates, or the requirement that contractors should have an annual turnover multiple times the value of a contract (Bezkorupce, 2015). Such constraints drastically reduce the number of firms that qualify for a specific procurement contract, and thereby guide the process in the direction of a preferred firm. Furthermore, contracting authorities often are free to set the evaluation criteria – as well as the weights given to the various criteria – employed during the allocation process. Bezkorupce (2015) shows that contracting authorities can use this flexibility to award procurement contracts to firms performing exceptionally well on one specific criterion (e.g., a fine due by the firm in case of delay), even though their overall bid may not otherwise have been the most beneficial.

Second, direct corporate donations to political parties are allowed in the Czech Republic, but parties have to disclose full lists of donors and amounts. Since parties that do not comply can be fined and lose part of their operational allowance from the government, they are incentivized to adhere to this regulation and existing evidence suggests that all do so.[5] As such, we are able to observe *all* party donations, which account for up to 33% of the budget of big parliamentary parties in the studied period (Titl et al., 2015). This provides a unique opportunity to operationalize politically connected suppliers based on firms' donations to the party in power (Titl and Geys, 2019). We thereby consider the main party in power to be the one that holds the Hejtman position, since this makes it the most powerful party in the regional Council and Board. Table OA.1 in the appendix presents the distribution of Hejtmans across parties and time in the three legislative periods of relevance to our analysis.

2.2. Data

The paper exploits a unique balanced region-level panel dataset covering the period from 2007 to 2014. There are 13 regions in the Czech Republic, which means that our dataset consists of 104 observations. Three sets of data are necessary to empirically assess the propositions set out in the introduction.

First, to measure public sector efficiency, we need information on the inputs and outputs in regional governments' production process. As we are particularly interested in the efficiency implications of distortions in the procurement process, we use regional expenditures through procurement contracts per capita as our central input (see Panel I of Table 2). The mean yearly expenditure on these contracts was approximately 2.443 billion CZK (equivalent of $122.15 million). This includes expenditures by the regions themselves as well as expenditures by region-owned enterprises (which administer substantial levels of procurement spending especially in healthcare).[6] Since the Czech regions allocate approximately 76% of their expenditures on education (i.e. schools and other educational facilities), health and social care, and infrastructure (i.e. roads and rail transport), we focus on outputs in these three key service domains.[7] Table 1 provides summary statistics on the output variables available for these policy areas. All variables except for transport outputs are presented

[5] We naturally cannot observe other possibly illegal payments by firms to parties. This should be kept in mind, as it implies that any observed efficiency implications in our analysis should best be viewed as independent of such corruptive practices.

[6] Note that public procurement has been estimated to be about 76% of total purchases by regional governments (for details, see `https://www.zindex.cz/category/detail/KRAJ/8/`). As such, it constitutes a critical share of regional public expenditures.

[7] According to the Czech statistical office, about 70% of secondary schools are region-run. Similarly, the Czech regions (in)directly own medical facilities that account for roughly two thirds of beds in all healthcare facilities. Finally, Czech regions own and take care of all roads of class 2 and 3 (which are regional and local roads; Act No. 13/1997 Coll. on Roads), and order rail transport from railway companies for the provision of local and regional transport services.

per 1000 inhabitants. The mean of the variable “Nr. of kindergartens” thus reflects that there were, on average, 0.55 kindergartens per 1000 inhabitants across all regions and years. Similarly, the mean of the variable “Nr. of doctors” indicates that there were, on average, 4.5 doctors per 1000 inhabitants across all regions and years. The variables reflecting outputs related to transport infrastructure are expressed in kilometres per 1000 inhabitants. The variable “Roads class 2” indicates an average of 6.5 kilometres of regional motorways per 1000 inhabitants. We should note that these measures do not reflect the actual quality of provided services, which is unfortuantely unavailable. Hence, our analysis can only capture the extent of service provision, but not its quality.

Overall, we have 17 indicators for outputs across our three policy areas. A high number of indicators triggers the curse of dimensionality: i.e., the discriminatory power of the model is reduced such that a substantial proportion of the decision-making units would be considered efficient. Hence, we reduc the number of outputs measures by calculating one composite indicator (CI) per policy area (see Panel II of Table 2). The technical details of the calculation of these composite indicators is described in Section A of the Appendix. In short, we employ a Benefit-of-the-Doubt model (Melyn and Moesen, 1991) where the weights for the different measures within each policy area are assigned endogenously (this avoids discrimination of regions based on their observed preferences). These weights are restricted to lie within specific parameter bounds, which avoids assigning unreasonably large weights to policy areas where the region at hand does not invest much money. The implemented model specification is robust and conditional, such that the resulting composite indicators are robust to measurement errors and outliers, as well as taking into account the regions' socio-economic environment (such as revenues per capita and the shares of young and elderly in the population).

Table 1: Indicators of outputs in three service fields

All variables per 1000 inhabitants	N	Mean	St. Dev.	Min	Max
Panel I: Indicators of education output					
Nr. of kindergartens	104	0.552	0.274	0.209	1.289
Nr. of teachers in kindergartens	104	2.828	1.529	1.124	7.130
Nr. of children in kindergartens	104	36.042	19.164	14.652	88.505
Nr. of primary schools	104	0.468	0.216	0.186	0.929
Nr. of teachers in primary schools	104	6.541	3.602	2.751	16.783
Nr. of students in primary schools	104	90.483	50.217	39.009	240.397
Nr. of secondary schools	104	0.145	0.075	0.061	0.380
Nr. of teachers in secondary schools	104	4.682	2.615	1.749	11.936
Nr. of students in secondary schools	104	54.543	31.253	19.596	145.725
Panel II: Indicators of healthcare output					
Nr. of doctors	104	4.526	2.588	1.862	12.387
Nr. of hospitals	104	0.020	0.014	0.007	0.070
Nr. of beds hospitals	104	6.350	3.942	2.392	17.316
Nr. of spec. medical institutes	104	0.017	0.011	0.004	0.043
Nr. of beds in medical institutes	104	2.303	1.250	0.320	5.032
Panel III: Indicators of transport infrastructure output					
Roads class 2 in kms	104	6.530	2.855	2.688	13.898
Roads class 3 in kms	104	4.053	1.851	1.516	9.200
Length railways in kms	104	1.160	0.636	0.514	3.407

Notes: All figures are per thousand inhabitants. Dataset covers 2007 to 2014. *Source*: Authors' elaboration based on Czech Statistical Office.

Second, to assess the role of firm donations in procurement allocations, we require detailed information about both firms' donations and public procurement contracts. As mentioned in the previous section, the Czech institutional setting allows us to observe *all* party donations. The value of the average donation in a given year to the party in power among donating firms is 109,570 CZK (circa \$4,384). Interestingly, while only 1.1% of all Czech firms donate to political parties, 12.9% of donating firms supply procurement contracts (see also Titl and Geys, 2019). We also have access to information on *all* public procurement contracts above a relatively limited threshold.[8] The available data include, among other aspects, the details of the winner of the contract, the allocation procedure and criteria, the decision date, and the value of the contract. Crucially, the data on firm donations and public procurement allocations can be linked via unique firm identifiers. As such, we can calculate for each region i in year t the share of the value of public procurement contracts awarded by a region to firms that donated to the party in power (henceforth "politically connected firms"), which will be our central independent variable explaining variation in regional public-sector efficiency (henceforth referred to as *Dependency*):

$$Dependency_{it} = \frac{value\ of\ contracts\ supplied\ to\ region\ i\ by\ connected\ firms\ in\ year\ t}{value\ of\ all\ contracts\ supplied\ to\ region\ i\ in\ year\ t} \tag{1}$$

Summary statistics on the *Dependency* measure are provided in Panel III of Table 2. On average, 1.3% of the total value of procurement contracts in region i in year t is allocated to politically connected firms, with standard deviation of 3.8% and a maximum value just over 25%. This panel also includes summary statistics for our contol variables. The $Share\ of\ left\ parties$ is defined as the number of left-wing regional councillors in the total number of councillors, and used as a control variable in our efficiency model. The variables $Revenues\ per\ capita$, $Share\ of\ young\ residents$ (under the age of 16) and $Share\ of\ elderly$ (65 years or older) are used while calculating our composite output indicators to control for different conditions in different regions (in line with Asatryan and De Witte, 2015, see also above).

Finally, we require information to explore the mechanisms underlying the efficiency implications of political donations. Panel IV of Table 2 includes information about the share of contracts assigned to suppliers with different levels of procurement experience, and the shares of contracts assigned under more or less stringent allocation procedures. We return to these variables and their theoretical motivation in more detail in Section 4.2 below.

3. Methodology

In order to estimate Czech regional public sector efficiency scores, we conduct a non-parametric estimation of the production frontier. This relates per capita expenditures on public procurement as our main input to a number of outputs (i.e. proxies for public goods and services listed in Table 1).[9] The resulting efficiency scores are relative measures indicating how much output each regional government produces given its spending

[8] The threshold is set by Act No. 137/2006 Coll. on Government Procurement, and differs depending on the type of contract. It is 2,000,000 CZK (excluding VAT; circa \$80,000) for public service contracts, 6,000,000 CZK (excluding VAT; circa \$240,000) for public works.

[9] A possible concern with using expenditures as input indicator is that not all units may face the same prices. Such resource price differentials are unlikely in our setting for two reasons. First, all firms bidding for procurement contracts can bid in all regions (and they often do so). This ensures that there is competition on the procurement market across regions, and limits variation in prices. Second, our sample does not include Prague due to its institutional differences with other regions. As Prague is the only region where price levels are significantly different, all other regions can be credible viewed as price takers in their input markets.

Table 2: Summary statistics

Statistic	N	Mean	St. Dev.	Min	Max
Panel I: Input					
Spending in public procurement contracts	104	2,443.446	2,773.788	130.123	20,929.050
Panel II: Output					
Composite Indicator on education	104	0.978	0.041	0.825	1.000
Composite Indicator on health	104	0.988	0.028	0.873	1.000
Composite Indicator on transport	104	0.943	0.072	0.768	1.000
Panel III: Environmental variables					
Dependency	104	0.013	0.038	0.000	0.252
Share of left parties	104	0.501	0.091	0.267	0.636
Revenues per capita	104	41.507	22.604	17.812	105.132
Share of young residents	104	0.147	0.007	0.122	0.168
Share of elderly	104	0.161	0.014	0.130	0.190
Panel IV: Mechanisms					
Share of frequent suppliers	104	0.918	0.109	0.451	1.000
Share of contracts below the threshold	104	0.498	0.230	0.011	0.988
Share of contracts under econ. adv. criteria	104	0.564	0.267	0.069	0.986

Notes: Spending is in millions of CZK (20 CZK is equivalent to approximately 1$). Outputs are calculated as composite indicators using a robust and conditional Benefit-of-the-Doubt procedure (see section 3). "Dependency" is defined as the combined value of contracts supplied by politically connected firms over the combined value of all contracts supplied to region i in year t. The share of left-wing parties equals the number of left-wing regional councillors in the total number of councillors. Regional government revenues per capita are expressed in thousand CZK. The shares of young (under 15) and elderly (over 65) residents are relative to the total population. Finally, the mechanism variables are measured as their share in the combined value of particular contracts (more details below).

level. We more specifically implement an advanced version of the canonical free disposal hull model (Deprins et al., 1984), which is robust to outlying observations and takes into account heterogeneity between observations (Daraio and Simar, 2005, 2007a).

In our model, no functional form of the production function is a priori assumed. We only impose that a set of p inputs $x \in \mathbb{R}_+^p$ is transformed into a set of q outputs $y \in \mathbb{R}_+^q$ within a production set that is defined as follows (we follow notation by Daraio and Simar, 2007a):

$$\Psi = \{(x, y) \in \mathbb{R}_+^{p+q} \mid x \text{ can produce } y\} \tag{2}$$

The efficiency of each decision-making unit – in our case, the Czech regions – is then calculated as the Farrell-Debreu output-oriented efficiency score (Debreu, 1951; Farrell, 1957). The output orientation means that we evaluate what the output shortfall is compared to the best performing observations while keeping the inputs fixed.[10]. Our efficiency scores effectively measure the distance of each region from the best practice frontier.

$$\lambda(x, y) = \sup\{\lambda \in \mathbb{R} \mid (x, \lambda y) \in \Psi_{FDH}\}\tag{3}$$

This measure maps the characteristics of each region to an output efficiency scalar λ. Efficient regions would thereby be assigned a $\lambda = 1$ (these are on the production frontier), inefficient regions a $\lambda > 1$. However, for the sake of a more intuitive interpretation, we will invert these scores so that the scores are between 0 and 1. For instance, $\lambda = 0.8$ can then be interpreted as a region performing at 80% of its potential (inefficient). The best practice frontier is in our case is defined as Ψ_{FDH} (for details, see Fried et al., 2008):

$$\Psi_{FDH} = \{(x, y) \in \mathbb{R}_+^{p+q} \mid y \leq Y_i, x \geq X_i, i = 1, \ldots, n\}\tag{4}$$

3.1. The robust and conditional model

As the conditional FDH model is deterministic, it may be very sensitive to the presence of outliers or measurement errors. To accommodate this, we implement the so-called order-m robust specification. For a given input (x_0), we draw B times a sample of size m ($< n$) with replacement among those Y_i such that $X_i \leq x_0$.[11] Then, we compute efficiency scores on each of these B samples and define the final efficiency score as a mean across all samples – i.e. $\tilde{\lambda}_{m,n}(x, y) \sim \frac{1}{B}\sum_{b=1}^{B} \tilde{\lambda}_m^b(x, y)$.[12]

Still, this procedure can be sensitive to the fact that regions with certain characteristics may be better reference points for each other (see e.g. Banker and Morey, 1986). The influence of such exogenous factors can be captured by implementing a conditional efficiency estimator (Daraio and Simar, 2005, 2007a). The key difference with respect to the approach outlined thus far is that, for a given x_0, we draw a sample of size m ($< n$) with a given probability determined by a Kernel function around relevant exogenous characteristics (z).

[10]For input-oriented scores, one would instead evaluate how excessive the input is compared to the best performing observations while keeping outputs fixed.

[11]In our case, the sample size m is set to 40. This sample size is determined such that the number of super-efficient observations is stable (Daraio and Simar, 2005).

[12]In practice, to save computational time, we will use the integral formulation given by: $\tilde{\lambda}_{m,n}(x, y) = \int_0^\infty (1 - \hat{F}_{X,n}(ux \mid y))^m du$. For details, see (Simar, 2003).

Hence, similar observations in term of the exogenous characteristics z are drawn with a higher probability than other observations. Then, as before, we compute B efficiency scores for each region and define the final scores as the mean value: i.e. $\tilde{\lambda}_{m,n}(x, y \mid z) \sim \frac{1}{B}\sum_{b=1}^{B} \tilde{\lambda}_m^{b,z}(x, y)$.

3.2. Accounting for the influence of exogenous factors on efficiency

Since we are mainly interested in how firms' donations to the party in power are associated with the efficiency of public good provision, we must evaluate the direction of the correlation of these factors with the efficiency scores calculated above. We thereto apply a non-parametric bootstrap procedure previously used in De Witte and Geys (2011), De Witte et al. (2013) and Asatryan and De Witte (2015). The impact of exogenous factors is retrieved by regressing the ratio of conditional and unconditional efficiency scores against the vector of exogenous factors in a local linear regression.

$$\frac{\tilde{\lambda}_{m,n}(x, y \mid z)_i}{\tilde{\lambda}_{m,n}(x, y)_i} = f(z_i) + \varepsilon_i \tag{5}$$

Non-parametric naive bootstrap is used to test significance and obtain p-values. Although the magnitude of the coefficients deriving from these regressions cannot be easily interpreted, their sign reveals whether larger values of characteristics z are favourable or unfavourable to efficiency (De Witte et al., 2013). For causal interpretation of these results, characteristics z should be exogenous to the outcomes of interest. We are aware that this is unlikely to hold for our measure of political connections (i.e. donations to the party in power). For instance, underlying levels of corruption might affect both firms' donations and public sector efficiency. Hence, throughout the analysis below, we will refer to associations or correlations rather than causal effects.

4. Results

4.1. The efficiency implications of political donations

In Table 3, we present our main results on how the efficiency of Czech regional governments is related to the allocation of public procurement contracts to politically connected firms. As explained in Section 2.2, the central variable in our analysis – *Dependency* – is defined as the share of public procurement contracts supplied by politically connected firms for each region and each year. The results of an unconditional FDH estimation are given in column (1), whereas columns (2) to (5) present a set of conditional models including *Dependency*. The latter allows us to examine the direction of the correlation between efficiency and the dependency on politically connected firms in public procurements. We further show results without (column (2)) and with (columns (3) to (5)) region fixed effects (for similar approach see Cordero et al., 2017b). In column (5), we also include a linear time trend to accommodate any general trend in public sector efficiency across all regions. In columns (3) to (5), we also include a measure for the left-right orientation of the regional government, which has previously been shown to matter for public sector efficiency. The underlying proposition is that the preference of left-wing parties for higher public spending might be associated with lower efficiency (see e.g. Revelli and Tovmo, 2007; Borge et al., 2008; Kalb, 2010; Ashworth et al., 2014; Helland and Sørensen, 2015).[13]

[13]One might consider adding other control variables to mitigate the potential for missing variable bias. We refrain from doing so here due to the limited number of degrees of freedom available, and the very demanding nature of our fully non-parametric estimation model for small sample sizes. Moreover, several socio-economic characteristics of the regions – including their per capita revenues and age composition – are indirectly controlled for via their inclusion in the BoD model constructing our composite output indicators. Note also that the region fixed effects capture all time-invariant region-specific effects.

The results in column (1) show a mean efficiency score of 0.994, which indicates that the average region can only increase its provision of outputs by 0.6% when working equally efficiently as the best practice observations. This very high level of observed efficiency in part reflects that our estimates account for regional heterogeneity (measured by revenues per capita and the shares of young and elderly, for details see Appendix A) in the conditional BoD procedure generating the composite output indicators. However, there remains important variation in these efficiency scores. The least efficient region could improve its efficiency by 8.4%. In the remaining columns of Table 3, the inclusion of environmental variables gradually reduces the variation in the efficiency scores – reflecting the role of heterogeneity sources not captured by the unconditional model. As seen in the table, the mean efficiency also progressively edges towards 1 as the model gets more saturated, which reflects that more of the underlying (unobserved) heterogeneity between the observations is captured.

Importantly, the conditional efficiency models strongly suggest that *Dependency* is negatively correlated with efficiency once we include region fixed effects. In other words, an increase in the share of public procurement contracts supplied by politically connected firms is associated with a decrease in regional government efficiency. Given the fully non-parametric nature of the model and the demanding specification (i.e. 14 to 16 explanatory variables – including 13 region fixed effects – in a model with 104 observations), the consistency and statistical significance in most models with region fixed effects provides strong support for our central theoretical proposition. Before turning to the underlying mechanisms of this effect, we should note that *Dependency* shows considerable variation in the cross-sectional dimension. Hence, the regional structure of the Czech political environment appears to matter for public sector efficiency, and one key element of this regional political environment indicated by our data is the level of public procurement contracts allocated to politically connected firms. In partial contrast to earlier results by, e.g., De Witte and Geys (2011) and Asatryan and De Witte (2015), we do not observe a clear direction for the correlation of regional governments' left-wing ideology with public sector efficiency.

Table 3: Unconditional and conditional FDH efficiency scores

Variable	(1)	(2)	(3)	(4)	(5)
Mean efficiency score	0.994	0.994	0.997	0.998	0.998
St. dev. efficiency score	0.017	0.017	0.011	0.009	0.008
Min score	0.916	0.920	0.934	0.934	0.946
Max score	1.003	1.001	1.000	1.000	1.000
Observations	104	104	104	104	104
Dependency	-	Favourable	Unfavourable ***	Unfavourable	Unfavourable **
Share of left wing parties	-			Unfavourable	Favourable
Region FE	NO	NO	YES	YES	YES
Year time trend	NO	NO	NO	NO	YES

*** p<0.01, ** p<0.05, * p<0.1

Notes: The table shows the results from a set of unconditional and conditional FDH models as formally specified in Section 3. *Dependency* is defined as the share of public procurement contracts supplied by politically connected firms in region i and year t. The share of left-wing parties is measured by the share of left-wing regional councillors in the total number of councillors. Year time trend is a continuous variable across the years in our sample.

4.2. Mechanisms driving the efficiency implications of political donations

In this section, we investigate two potential mechanisms underlying the negative efficiency implications of *Dependency* observed in the previous section.

The first of these is linked to suppliers' experience with procurement contracts. From a theoretical perspective, firms with previous procurement experience can be expected to have built up relations to politicians via their procurement experience. This may not only establish a relationship of trust between firms and politicians (as well as administrators), it also gives these firms valuable experience in providing services to the public sector, which can work to smooth the cooperation and increase the efficiency of public good provision (Witko, 2011; Goldman et al., 2013). In line with this line of argument, Coviello et al. (2018, p. 5) find strong evidence that "contractors who have won in the past systematically deliver current works faster". This is important since public procurement contract allocations are a zero-sum game in the sense that only one firm can win a given contract. Hence, contracts allocated to politically connected firms can no longer be allocated to experienced firms,[14] which would work to depress public sector efficiency. This line of argument leads to the hypothesis that the share of experienced firms and the share of party donors in procurement allocation contracts are negatively correlated. To examine this hypothesis, we distinguish between frequent and infrequent suppliers. The former are defined as firms that for every year in the 2007 – 2014 period were awarded at least one procurement contract from any region in the two years preceding the year of observation. All other firms are considered infrequent suppliers.

A second mechanism relates to the possibility that the exact nature of the procurement process affects the share of politically connected firms among procurement contract allocations (which, as argued above, might squeeze out more experienced providers). One important element thereby is that procedural restrictions are more stringent for contracts with a total value exceeding 4,997,000 CZK (circa $249,850) – or 20 million CZK in case of construction works (circa $1,000,000). Below this threshold, contracts are not regulated by EU law. In the Czech setting, this means that contracting authorities may use the simplified so-called "below-the-threshold" procedure and the negotiated procedure without publication. Public authorities may thereby directly ask a minimum of five firms to provide bids, and are required to publish only the final outcome (e.g., a winner of the tender). Furthermore, contracts concluded under the below-the-threshold procedure are not published in the Official Journal of the European Union, and contracting authorities can choose shorter time limits for the delivery of bids. All these elements provide a setting more tenable to favouring some firms over others. Hence, we hypothesize that a higher share of below-the-threshold procurement contracts allocated by public authorities would be associated with an increase in the share of connected firms among procurement contract winners (*Dependency*), and thereby a decline in efficiency.

Another key procedural element concerns the degree to which politicians have more discretionary power (Titl and Geys, 2019). The legislative framework in the Czech Republic provides a considerable range of possible evaluation criteria and allocation procedures available to public authorities, which vary substantially in terms of the restrictiveness and public visibility they impose (see also Palguta and Pertold, 2017). Specifically, procurement processes using the 'lowest price' framework impose a clear (and self-evident) decision criterion. In sharp contrast, contracts allocated using the so-called 'economically advantageous' criterion provide substan-

[14] Note that experienced firms are significantly less likely to be a donor of the party in power compared to inexperienced firms (0.6% versus 1.5%; $p<0.05$). Hence, party donors tend to be different from the set of experienced firms.

tially more leeway since Czech public procurement legislation does not prescribe in detail how ‘economically advantageous’ should be understood (Act No. 137/2006 Coll. on Government Procurement). This framework thus would appear particularly convenient for politicians and civil servants intent on favouring politically connected firms (Titl and Geys, 2019). We hypothesise, therefore, that a higher share of contracts allocated using this criterion is associated with a higher share of politically connected firms among procurement contract winners (*Dependency*), and thereby lower efficiency scores. Summary statistics for these three variables are presented in Table 2.

To examine these hypotheses, we run a series of non-parametric fixed-effects regressions (for similar approach, see Czekaj and Henningsen, 2013; Czekaj, 2013) where *Dependency* is the dependent variable. The share of contracts supplied by frequent suppliers, share of below-the-threshold contracts and the share of contracts using the ‘economic advantageousness’ criterion are the independent variables. We furthermore include region fixed effects (Column 1), year fixed effects (Column 2), or both (Column 3).

In the first two columns, the results indicate a substantively strong and statistically significant coefficient estimate for the shares of frequent suppliers and below-the-threshold contracts. While their signs remain robust in the last column, they loose significance due to the (overly) demanding nature of this specification.[15] Overall, our findings provide evidence to suggest the following plausible mechanisms behind the efficiency implications of political connections. The share of politically connected firms among procurement contract allocations increases with less regulated and less transparent allocation procedures (allowing distortions away from the most competitive bidders), and is associated with a reduction in the share of frequent, experienced suppliers (whose experience benefits effective public good provision).

Table 4: Non-parametric fixed effects estimation results examining the mechanisms driving the efficiency implication of political donations

VARIABLES	(1) FE Dependency	(2) FE Dependency	(3) FE Dependency
Share of frequent suppliers	-0.149***	-0.096***	-0.041
	(0.000)	(0.008)	(0.914)
Share of below the threshold	0.038**	0.044***	0.023
	(0.027)	(0.007)	(0.130)
Share of economic advantageousness	-0.014	-0.001	-0.009
	(0.366)	(0.932)	(0.638)
Observations	104	104	104
R-squared	0.338	0.287	0.980
Number of regions	13	13	13
Region FE	YES	NO	YES
Year FE	NO	YES	YES

P-values in parentheses; *** p<0.01, ** p<0.05, * p<0.1

Notes: The table shows the results from a set of non-parametric fixed-effects regressions. *Dependency* is defined as the share of public procurement contracts supplied by politically connected firms in region i and year t. The shares of frequent suppliers, below the threshold contracts, and contracts using the economic advantageousness criterion are measured as the shares of all public procurement contracts within these categories.

[15] It includes three main variables, 12 regional dummies and 7 year dummies – i.e. 22 variables for a model with 104 observations. This induces overfitting issues as evidenced by the very sharp increase in both R-squared and the standard errors of our coefficient estimates.

5. Conclusion

This paper studies the public sector efficiency implications of distortions in the allocation of public procurement contracts. There is a large body of literature studying public sector efficiency and its various determinants (Narbón-Perpiñá and De Witte, 2018a,b). However, none of these studies investigates the efficiency implications of political donations and different public procurement allocation procedures. While politically connected firms might foster local efficiency, it is also plausible that such donations have negative efficiency implications due to their distortive effect on the allocation of public procurement contracts and their association with less restrictive allocation procedures in public procurement contracts (Titl and Geys, 2019).

To assess the relation between firm's political donations, public procurement outcomes and public sector efficiency, we develop an advanced efficiency model that accounts for outlying observations and regional heterogeneity. We apply this model to a unique dataset with detailed information of public good provision, political donations and public procurement contracts in Czech regions in the period from 2007 to 2014. Our analysis shows substantial evidence that a larger share of public procurement contracts awarded to politically connected suppliers is associated with a decline in the cost efficiency of public good provision. In testing the underlying mechanisms, we observe that politically connected firms obtain more contracts in regions and time periods with higher use of below-the-threshold procedures. In other words, the share of politically connected firms among procurement contract allocations seems to increase with less regulated and less transparent allocation procedures. Moreover, it is associated with a reduction in the share of frequent, experienced suppliers.

The policy implications of our results are relevant beyond our Czech setting. As the quality of governance and government efficiency are crucial determinants of productivity growth and the aggregate efficiency of the economy (see e.g. Olson et al., 2000; Méon and Weill, 2005), our findings have important implications for the design and oversight of procurement allocation processes. In particular, the negative influence of firm donations on public sector efficiency through distorted procurement allocations indicates that the ties between firms and politicians should be reduced. This is consistent with earlier evidence showing that firms' political connections result in favoritism in public procurement contracts, which can be resolved by a stricter legal framework (see e.g. Baltrunaite, 2019; Titl and Geys, 2019).

Although this paper, to the best of our knowledge, provides the first empirical evidence on the efficiency implications of politically connected firms, further research is necessary. First, we need to assess the external validity of our results by replication studies on other sectors and countries. To facilitate such replications and extensions, the R-code underlying the present analysis is available from the authors upon request. This can also increase the value of our methodological approach for managers within oversight bodies in the public sector. Second, it is worth investigating how our findings translate to settings with more rigid procurement procedures. Therefore, one could study the impact of politically connected firms using more stringently legislated EU-contracts. Third, the proposed model specification is a fully non-parametric specification, which is relatively demanding on the data. While this avoids specification bias, it limits the number of control variables and fixed effects that can be included in the analysis. Thus, further research can test the robustness of the results by relying on more parametric assumptions of the production frontier. Finally, an alternative empirical approach would be to structurally estimate an auction model with asymmetric bidders (i.e. political donors or not), which allows assessing whether less efficient donor firms (with a higher cost structure) systematically win or are granted the provision of public goods. This, however, requires detailed firm-level data on (the participants and outcomes of) the procument process.

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A. Dimensionality reduction using Benefit of the Doubt model

In the estimation of a production frontier and efficiency scores, a high number of inputs and outputs in the analysis can result in a curse of dimensionality. Given the large number of output indicators available in our dataset, we reduce the dimensionality of our output space via a set of composite output measures (i.e. one for each service domain). This automatically also implies that we do textitnot assume that there exists an overall technology combining all outputs, but rather three separate technologies across the public policy areas. These three composite indicators are computed via a fully non-parametric Benefit of the Doubt model (BoD) (Melyn and Moesen, 1991), which keeps the number of final output variables low without losing information. The basic idea behind BoD weighting is that the weights should be high (low) for aspects where the decision-making unit provides more services compared to other decision-making units in the dataset. The underlying argument is that different decision-making units might focus on different services within one policy area, which should be accommodated in the analysis. For instance, some decision-making units might provide more services in specialized healthcare centers rather than hospitals. By choosing arbitrary exogenous weights one would unavoidably "discriminate" some decision-making units for the preferences they exhibit.

The traditional BoD model can be represented by the following linear programming problem (Melyn and Moesen, 1991). For a certain region i (from all regions $1, ..., n$)

$$CI_i = \max \quad \sum_{r=1}^{s} y_{i,r} w_{i,r} \tag{A.1}$$

$$s.t. \sum_{r=1}^{s} y_{j,r} w_{i,r} \leq 1, \quad \text{for j=1,...,n,} \tag{A.2}$$

$$w_{i,r} \geq 0, \qquad \text{for r=1,...,s} \tag{A.3}$$

where CI_i is the composite indicator value for region i, $y_{i,r}$ is the measured service provision of region i on indicator r and $w_{i,r}$ is the set of most favorable weights for the region i. We tailor this BoD model in three ways. First, we implement the order-m methodology proposed by Cazals et al. (2002) to make our composite indicators robust to outlying observations. This is an application of Monte Carlo simulations on the usual computation of the composite indicator. m observations are drawn with replacement B times from the original sample and then B composite indicators for every observation are computed. The value of the composite indicator is then set to the mean of the B indicators.

Second, when environmental variables affect the location and shape of the production frontier, not including such variables in the analysis leads to biased estimation results (Daraio and Simar, 2005, 2007b; Simar and Wilson, 2011). We therefore adjust the unconditional BoD model outlined above using three environmental variables captured in the vector Z: i.e. per capita regional government revenues, the share of residents younger than 15, and the share of residents older than 65.[1] Specifically, the probability a region is drawn into the reference group of another region depends on their respective characteristics captured in vector Z (using the estimation of a kernel function around Z). This directly implies that in the construction of the composite indicator for region i in year t, this region is more likely to be compared with other regions that are similar in terms of their revenue per capita and age composition.

Third, in the traditional BoD model the weights $w_{i,r}$ are endogenously determined in such a way that the composite indicator is maximized for each region. This might, however, unreasonably assign very high weights to policy areas where the region does not invest much money. To accommodate this, we implement weight restrictions based on observed spending shares across the relevant policy areas – which can be interpreted as preference expressions (Cusack, 1997, 1999; Potrafke, 2017). Thereby, we avoid the situation where some regions would be assessed only on a small subset of their outputs (while all other outputs are effectively ignored bybeing assigned zero weight as explained in Dyson and Thanassoulis, 1988) even though they spend money also on other outputs (for an in-depth discussion about interpretation of the optimal weights, see Podinovski, 2016). Specifically, we set a lower and an upper bound for the weights in each policy area – i.e. so-called Assurance Regions type I (or ARI) restrictions (for a review of different types of weights restrictions, see Khalili et al., 2010) – by adding the following constraint to the linear problem defined in equations A.1, A.2 and A.3:

$$\phi_r \leq w_{i,r} \leq \psi_r \quad \text{for r=1,...,s,} \tag{A.4}$$

We set these weight restrictions based on available information about public expenditures in our three main policy areas in 2011 (i.e. the middle of our period of analysis 2007–2014). Specifically, we impose bounds at

[1] Note that these variables are subsequently no longer used in the conditional efficiency model estimated in the main text. The reason is that their effects are already controlled for here.

$\pm 50\%$ around the average expenditure share of each output indicator at the country level (i.e. our restrictions do not differ across regions due to lack of more detailed information). For education, we use official statistics from OECD (2014) to calculate the share of total education spending allocated to pre-primary (17.0%), primary (41.4%) and secondary (41.6%) education.[2] As we have for each education level three output variables, we assume weight restrictions that correspond to one third of the spending per education level. Further, we allow for 50% bounds to reach our final weight restrictions. For example, for pre-primary education, the weight for each indicator is bounded between 2.83% and 8.5% (i.e. a 50% range around 17.0/3, or 5.67%). For healthcare, information provided by ÚZIS ČR (2012) indicates a spending division across hospitals (59.2%), medical institutes (6.6%), doctor salaries (19.4%) and salaries of other medical workers (18.84%). As before, we split these spending shares equally across the available output indicators in each category and allow for 50% bounds to reach our final weight restrictions. Finally, for transport we use OECD data on road (64.21%) and rail (35.79%) expenditure (OECD, 2018), and split the former across road types proportionally to the lengths of these roads.[3] The final weight restrictions are summarized in Table OA.2.

Summary statistics for the resulting robust and conditional BoD composite indicators with weight restrictions – which are our three main outputs in the efficiency analysis – are presented in Panel II of Table 2 in the main text. The mean of the composite indicator for education (0.978) suggests that on average the regions operate at 97.8% of their best-performing comparison group. Similar high performance – with small levels of variation across regions – is documented for healthcare and transport. This reflects the fact that our composite indicators directly take into account the regions' socio-economic environment (as we use the conditional specification with three environmental variables as highlighted above).

[2]We derive total spending in each education category by multiplying the expenditure per student by the number of students.

[3]Although not ideal since roads of class 2 are presumably more expensive than roads of class 3, this is our best approximation given the data available. The total length of roads of class 2 is 63.6% longer than that of roads of class 3.

B. Summary statistics on output measures and the allocation of political power in the Czech regions

Table OA.1: The parties in power in the Czech regions (2004-2016)

Region	Party in power 2004-08	Party in power 2008-12	Party in power 2012-16
Středočeský	ODS	ČSSD	ČSSD
Jihočeský	ODS	ČSSD	ČSSD
Plzeňský	ODS	ČSSD	ČSSD
Karlovarský	ODS	ČSSD	KSČM/ČSSD
Ústecký	ODS	ČSSD	KSČM
Liberecký	ODS	ČSSD	STAN
Královehradecký	ODS	ČSSD	ČSSD
Pardubický	ODS	ČSSD	ČSSD
Olomoucký	ODS	ČSSD	ČSSD
Moravskoslezský	ODS	ČSSD	ČSSD
Jihomoravský	KDU-ČSL	ČSSD	ČSSD
Zlínský	ODS	ČSSD	ČSSD
Kraj Vysočina	ODS	ČSSD	ČSSD

Note: The party in power is determined as the party of a Hejtman (State Governor in the Czech setting) in the particular region. This party also corresponds in the vast majority of cases to the strongest party in the Regional Council and Board. *Source*: Authors based on various official sources.

Table OA.2: Outputs' weight restrictions for the Benefit-of-the-Doubt model in the three service fields

	Mean	Min	Max
Panel I: Education output			
Nr. of kindergartens	5.67%	2.83%	8.5%
Nr. of teachers in kindergartens	5.67%	2.83%	8.5%
Nr. of children in kindergartens	5.67%	2.83%	8.5%
Nr. of primary schools	13.79%	6.895%	20.685%
Nr. of teachers in primary schools	13.79%	6.895%	20.685%
Nr. of students in primary schools	13.79%	6.895%	20.685%
Nr. of secondary schools	13.88%	6.94%	20.82%
Nr. of teachers in secondary schools	13.88%	6.94%	20.82%
Nr. of students in secondary schools	13.88%	6.94%	20.82%
Panel II: Healthcare output			
Nr. of doctors	19.37%	9.685%	29.055%
Nr. of medical workers	18.84%	9.42%	28.26%
Nr. of hospitals	29.62%	14.81%	44.43%
Nr. of beds hospitals	29.62%	14.81%	44.43%
Nr. of spec. medical institutes	3.275%	1.6375%	4.9125%
Nr. of beds in medical institutes	3.275%	1.6375%	4.9125%
Panel III: Transport infrastructure output			
Roads class 2	40.86%	20.43%	61.29%
Roads class 3	23.35%	11.675%	35.025%
Length railways	35.79%	17.895%	53.685%

Note: The weight restrictions for the Benefit-of-the-Doubt model are set based on available information about public expenditures in our three main policy areas in 2011 (i.e. the middle of our period of analysis 2007–2014). We impose bounds at ±50% around the average expenditure share in each output indicator at the country level (see columns Min and Max for the bounds). *Source*: Authors