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Weight Values, Scoring Rules and Abnormally Low Tenders Criteria in Multidimensional Procurement: Effects on Price*

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Abstract

Many weighing functions are currently in use for evaluating quotes in multidimensional procurement. Although these may affect the price paid by the buyer, their design has been little informed by empirical considerations. This paper studies the impacts of weight values, scoring rules, and abnormally low tenders criteria (ALTC) in an original sample of procurements of services. The panel dimension of the data and the rich set of procurement characteristics are exploited to show that convex price scoring rules (PSRs) and increases in the weight for price yield lower prices, whereas independent/interdependent PSRs and different ALTC result in the same price.

Keywords: Multidimensional procurement, weighing function, scoring rule, abnormally low tenders criterion, Spanish Armed Forces.

JEL Classification: H57

1. Introduction

Procurement often involves considerations other than price. Aspects of the good to be procured such as design, delivery time, technical characteristics, or the guarantee's term may affect the value of the procurement outcome. Procurement including non-monetary dimensions (quality for brevity) in addition to price has been called multidimensional (McAfee and McMillan, 1987), multi-attribute (Staschus *et al.,* 1991), or best value (Peckinpaugh and Goldstein, 1992). As argued by Dini, Pacini, and Valletti (2006), in multidimensional procurement it is unlikely that the same bidder offers both the cheapest price and the highest quality, as high-quality products are typically more costly to produce (although see Albano, Dini and Zampino, 2009). Hence, an essential element of multidimensional procurement is

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a method for combining a quote's price and quality dimensions into a total score. Furthermore, as price and quality are rarely observed in the same units, they must be transformed into commensurate scales before being combined into a total score (Bergman and Lundberg, 2013; Wang and Liu, 2014; Ballesteros-Pérez *et al.,* 2015).

A function that transforms a value measured on one scale (price or quality) into a measure on another scale (price score or quality score) is a scoring rule, whereas a function that combines price and quality scores into a single overall value is a weighing function (WF) (Bergman and Lundberg, 2013)[1]. This definition is a bit unconventional as functions that map a multidimensional offer into a single total value are usually called scoring rules in the theoretical literature on multidimensional bidding (see references below). However, it agrees to the purpose of this paper to distinguish between assigning a score to each individual component of a quote and combining these scores into a single overall value.

The 2004 EU's public procurement directives 2004/17/EC and 2004/18/EC, which were operating in Spain over the period covered by this analysis, included provisions to ensure that the award of public contracts complied with the principles of transparency, non-discrimination, and equal treatment. One of these provisions required contracting agencies to indicate the criterion to be used for award of the contract (either *lowest price* or *most economically advantageous tender:* MEAT) in the procurement's contract notice[2], as well as the criteria representing the MEAT and their weighing. The systematic registering of these features has contributed to advancing research on different aspects of multidimensional procurement.

The first generation of theoretical studies on multidimensional procurement (Che, 1993; Branco, 1997; David, Azoulay-Schwartz and Kraus, 2006; Asker and Cantillon, 2008 and 2010 and Nishimura, 2015) used the quasilinear WF that is linear in the price bid for selecting the winning firm. However, the transformations of the price bid exerted by certain price scoring rules (PSRs) used in many countries worldwide result in non-quasilinear WFs (Hanazono, Nakabayashi and Tsuruoka, 2015 and Ballesteros-Pérez *et al.,* 2016). Hence, recent theoretical analyses of multidimensional procurement have attempted to relax the quasilinear restriction in the WF (Wang and Liu, 2014; Hanazono, Nakabayashi and Tsuruoka, 2015; Dastidar, 2016 and Nakabayashi and Hirose, 2016). But despite the increased realism of theoretical models, the empirical work on the extent to which multidimensional bidding can be influenced by contracting authorities has been hindered by the difficulties of obtaining appropriate data (Ballesteros-Pérez *et al*., 2016).

Another strand of literature developed around the application of WFs in real procurement. Dini, Pacini and Valletti (2006) provide buyers with practical advice for choosing between some widely adopted types of PSRs. Chen (2008) and Bergman and Lundberg (2013) apply economic principles to discuss the pros and cons of several supplier selection methods. Dimitri (2013) focuses on the issue of translating a buyer's preferences into WFs. And Nishimura (2015) elucidates conditions under which the buyer's optimal WF can be additively separable in quality attributes. All these papers take more of a theoretical approach than we do here.

The empirical literature on multidimensional procurement includes comparisons of quality scores induced by different types of PSRs (Albano, Dini and Zampino, 2009), of winning bids delivered by rigid/flexible evaluation systems (Cameron, 2000) and by lowest price/MEAT awarding mechanisms (Lundberg, 2006; Lewis and Bajari, 2011), of price bids resulting from varying price weights (Iimi, 2013; Koning and van de Meerendonk, 2014), and of bidding competitiveness as influenced by price weights, PSRs, and abnormally low tenders criteria (ALTC) (Ballesteros-Pérez *et al*., 2016)[3]. But with the exception of Koning and van de Meerendonk (2014), who also provide some evidence on the effect of price weights on the winning bid, and Ballesteros-Pérez *et al*. (2016), not many authors have looked into the practical effects that characteristics of the WF such as the price weight, the PSR, and the ALTC criterion exert on the winning bid, that is, on the price paid by the buyer.

The purpose of this paper is to contribute to filling this gap by exploring the magnitudes of these effects in an original sample of multidimensional procurements of services conducted by the Spanish Armed Forces (SAF) between 2012 and 2016. Evidence on the extent to which the price paid by the buyer (and, by extension, the level of bidder competitiveness) is affected by characteristics of the WF will help contracting authorities design WFs for practical application in multidimensional procurement, making this design a less intuitive and subjective process.

A limitation of this study is that the econometric model does not include information on the specific non-monetary dimensions set out in the contract nor on the contractors' actual performance, but rather some controls for *ex ante* quality of service (Domberger, Hall, and Li 1995) and a measure of bid screening intensity by the contract awarding committee (Decarolis 2014). Although differences in non-monetary dimensions across contracts may be reflected in the price paid, finding a set of non-monetary dimensions common to all (or most) contracts in a sample of heterogeneous contracts is non-trivial. As to the risk of poor performance by contractors, to the extent that the controls included in the model are, as a matter of fact, insufficient to account for this risk, our estimates would tend to understate the true magnitude of the price paid. (Note, however, that the evidence on the effect of competition on performance is not conclusive: Domberger, Hall, and Li 1995, Cameron 2000, and Decarolis 2014.)

The empirical analyses of Lewis and Bajari (2011) and Nakabayashi and Hirose (2016) are based on the structural methodology. In contrast, this paper focuses on estimating the net effect of covariates on the winning bid, which precludes forecasting the effects of regime shifts that are not reflected in the data. Thus, for example, the results of this paper cannot be used to predict the impact on the price paid of a change in certain competition rules (e.g., an increase in the financial guarantee to be given by the contractor), or a change in market structure that goes beyond a variation in the number of bidders. Nevertheless, we provide evidence on economically advantageous determinants of the price paid, all of which are amenable to policy change.

The reminder of this paper is organized as follows. Section 2 provides background on public procurement in Spain over the period covered by this analysis and introduces the components of WFs. Section 3, which is structured into three subsections, deals with the selection of the sample, the construction of the variables representing the PSRs and ALTC identified in the sample, and the definition of the explaining and control variables. Section 4 describes the econometric framework and Section 5 presents the estimation results. Section 6 concludes.

2. Public procurement in Spain under the 2011 law

From December 16, 2011, to March 8, 2018, procurement by the Spanish government (including the armed forces) took place under Royal Legislative Decree 2011/3. During the lifetime of this law, and after receiving authorization to procure, the procurement process began with a contracting agency (CA) estimating the contracting dossier value, calculated as the estimated cost of the service (work or supply) to be procured net of VAT, plus the cost of possible options and renewals and possible bonuses paid to bidders. The estimated cost (which is often called the reserve price in the procurement literature: e.g., see Dimitri, Piga, and Spagnolo 2006) was intended to reflect the general market price of the service, and indicated to the interested providers the maximum acceptable price for the agency (excluding options, renewals, and bonuses). Often, the estimated cost was the result of a market-research study, but, when available, a previous estimated cost (possibly corrected for inflation or sectoral wage changes) could also be utilized.

Whenever a contracting dossier for a service was valued at €18,000 or more, it had to be purchased ordinarily by a procurement process open to any interested provider[4]. The winning bidder was determined by means of a sealed-bid auction in which one of two different award criteria could be employed: lowest cost or MEAT. MEAT was used whenever the procurement involved dimensions in addition to price, and it implied that each individual dimension of a quote would be evaluated and assigned a score, and the contract would be awarded to the supplier who submitted the quote with the highest total score according to a pre-specified WF.

WFs can be formulated generally as (Ballesteros-Pérez *et al.* 2015):

$$O_i = \{W_p \cdot S_i + W_q \cdot Q_i\}\ \alpha_i \tag{1}$$

where O_i is bidder *i's* total score, W_p and W_q are price and quality weights satisfying $W_p + W_q = 1$, S_i and Q_i are bidder *i's* price and quality scores calculated with the PSR and quality scoring rule given in the procurement specifications, and α_i is an abnormality indicator (defined below). The bidder with the highest total score was the bidder offering the MEAT (or best value for money). In case of a tie, the bidder with the largest proportion of disabled workers, or with the highest score in the dimension with the highest weight, was generally chosen.

In cases where it was possible to provide the service to be procured separately but it still constituted a functional unit (e.g. the maintenance of buildings located at distant military bases), the whole contracting dossier could be divided into batches. Providing that the whole dossier was valued at €18,000 or more, each batch had to be purchased ordinarily by procurement of the types described above even if the estimated cost for a batch was lower than €18,000.

Several provisions of the 2011 law were aimed at eliciting unbiased bids and limiting the risk of poor performance by contractors. One of these provisions was the possibility of requiring the winning bidder to prove the price that it offered if that price was below the economic threshold defined by the ALT criterion used in the auction. In the event that its explanations were deemed unreliable by the contract awarding committee, the winning bidder would be disqualified ($\alpha_i = 0$ in expression (1)) and typically the second best bidder would win. (If the price offered was not lower than the economic threshold or the explanations were deemed reliable, then $\alpha_i = 1$).

The bid-envelopes received were opened and examined at a public event. The awardee could be awarded a contract for at most 4 years, although contracts for services were typically of much shorter duration. With some exceptions (such as when security or public safety could be jeopardized) the identity and bid of the winning bidder were published in the contract results notice.

3. Data and variables

3.1. Data

The analysis is conducted using 477 multidimensional procurements of support services pertaining to 218 contracting dossiers undertook by the SAF in the period 2012-2016[5]. The contract notices for these dossiers were published between January 1, 2012 and December 31, 2016, and procurements were awarded before June 15, 2017. This limit on the awarding date may have excluded very few of the procurement procedures started in the last months of 2016, as 90 percent of the procurements of services are awarded within 143 days.

The 218 contracting dossiers included in this study represent 29.5 percent of the 739 service contracting dossiers awarded through competitive bidding by the SAF in the period 2012-2016. The rest of the dossiers were discarded for several reasons. In 321 cases the price paid by the government had to equal the estimated cost, either because this cost represented a commitment of funds to pay the contractor's invoices or to subsidize the price of a service, or because the dossiers involved special administrative contracts, which entail no cost for the government as the service is paid for by the end user. A further 43 dossiers were dropped on the grounds that they were canceled before awarding, 34 for having incomplete or inconsistent data, 114 for being awarded using the lowest cost criterion, and 9 for being null and void. Table 1 presents a frequency distribution of contracting dossiers by type, sample selection status, and year.

Table 1
SERVICE CONTRACTING DOSSIERS AWARDED BY THE SPANISH ARMED FORCES THROUGH COMPETITIVE BIDDING, BY TYPE OF DOSSIER, SAMPLE SELECTION STATUS, AND YEAR

	Excluded from sample									
Year	**Commit-ment of funds to pay in-voices**	**Special adminis-trati-ve contracts**	**Lump sum subsidies**	**Withdrawn before awarding**	**Inconsist-ent/ incom-plete data[a]**	**Lowest price**	**Void**	**Included in sample**	**Total**	**% in sample**
2012	33	3	1	3	13	19	6	38	116	32.8
2013	51	8	2	1	12	15	1	45	135	33.3
2014	51	4	1	2	2	26	1	46	133	34.6
2015	55	13	4	8	4	28	1	45	158	28.4
2016	55	35	5	29	3	26	0	44	197	22.3
Total	**245**	**63**	**13**	**43**	**34**	**114**	**9**	**218**	**739**	**29.5**

Notes: [a]: *Includes contracting dossiers whose legal auction information sheet was not published, whose weight for price cannot be calculated, or whose definition of the price scoring rule is missing, mistaken, or combines different price scoring rules; it does not include procurements whose savings rate was above the limit defined in the sensitivity analysis of Section 5.*

The data on the sample of procurements was compiled from the information stored in the Public Sector Contracting Platform (PLACE in Spanish), which is the national advertising website for Spanish government agencies to post contract notices, legal and technical auction information sheets, and contract results. These documents were personally read by the author and the information obtained was stored as a set of variables whose definition is outlined below. These variables contain detailed information on the procurement specifications and the winning bid. Information on all the bids for a procurement was published for about 20 percent of the procurements only.

3.2. Price scoring rules and abnormally low tenders criteria in service procurements conducted by the Spanish Armed Forces

The PSR and ALT mechanism used in each procurement were taken from the legal auction information sheet. For the sake of implementation in the econometric analysis, the variety of PSRs and ALTC identified in the sample were classified as explained in the following paragraphs.

Table 2 presents the mathematical expressions in per-unit values for the 16 PSRs identified in the sample, formulated following for the most part the notation employed by Ballesteros-Pérez *et al*. (2015). To assign a price score S_i to a bidder's price bid (b_i), $i=1,...,n$, these PSRs make use of the estimated cost of the contract (A) and/or of certain statistics calculated from the final distribution of price bids, such as order statistics ($b_{(1)} < ... < b_{(n)}$), the lowest bid ($b_{\min} \equiv b_{(1)}$), the highest bid ($b_{\max} \equiv b_{(n)}$), and the average bid (b_m)[6]. Table 2 also shows the mathematical expressions for the PSRs in terms of per-unit discounts, rebates, or drops off A (d_i), as bidder i's price bid can be expressed equivalently as:

$$b_i = (1 - d_i)\, A \tag{2}$$

or

$$d_i = 1 - \frac{b_i}{A}. \tag{3}$$

The maximum discount (d_{max}) corresponds to $b_{\min}$, and the average discount (d_m) corresponds to the average of all price bids expressed in discounts submitted. The PSRs numbered 1, 3, 4, 9, and 10 in column (1) of Table 2 can also be found (in some cases, with a slight modification) among the 124 Spanish construction auctions analyzed by Ballesteros-Pérez *et al*. (2016).

The PSRs numbered 6, 8, 9, 12, and 16 are such that S_i only depends on A and b_i, so that each bidder can calculate its price score in advance (*i.e.*, before the bidding phase takes place). This type of PSR is called independent by Albano, Dini, and Zampino (2009). The other PSRs listed in Table 2 are of the interdependent type, as the price score of any bidder

Table 2
PRICE SCORING RULES IDENTIFIED IN THE SAMPLE

ID (1)	Obs (2)	Description in terms of price bid (b_i) (3)	Description in terms of price discount (d_i) (4)	Independence (5)	Shape (6)
1	281	$S_i = b_{\min}/b_i$	$S_i = \frac{1-d_{\max}}{1-d_i}$	Interdependent	Convex
2	1	$S_i = \begin{cases} 1 \text{ if } b_i = b_{(1)} \\ 6/7 \text{ if } b_i = b_{(2)} \\ 5/7 \text{ if } b_i = b_{(3)} \\ 4/7 \text{ if } b_i = b_{(4)} \\ 3/7 \text{ if } b_i = b_{(5)} \\ 2/7 \text{ if } b_i = b_{(6)} \\ 1/7 \text{ if } b_i = b_{(7)} \\ 0 \text{ if } b_i > b_{(7)} \end{cases}$	$S_i = \begin{cases} 1 \text{ if } d_i = d_{(n)} \\ 6/7 \text{ if } d_i = d_{(n-1)} \\ 5/7 \text{ if } d_i = d_{(n-2)} \\ 4/7 \text{ if } d_i = d_{(n-3)} \\ 3/7 \text{ if } d_i = d_{(n-4)} \\ 2/7 \text{ if } d_i = d_{(n-5)} \\ 1/7 \text{ if } d_i = d_{(n-6)} \\ 0 \text{ if } d_i < d_{(n-6)} \end{cases}$	Interdependent	Linear[a]
3	68	$S_i = \frac{A-b_i}{A-b_{\min}}$	$S_i = d_i/d_{\max}$	Interdependent	Linear
4	3	$S_i = 2-\frac{b_i}{b_{\min}}$	$S_i = 2-\frac{1-d_i}{1-d_{\max}}$	Interdependent	Linear
5	27	$S_i = 1-\frac{b_i-b_{\min}}{b_{\text{m}}}$	$S_i = 1-\frac{d_{\max}-d_i}{1-d_{\text{m}}}$	Interdependent	Linear
6	1	$S_i = \begin{cases} 1 \text{ if } b_i = A-3000 \\ 59/60 \text{ if } b_i = A-2700 \\ 58/60 \text{ if } b_i = A-2400 \\ \vdots \\ 50/60 \text{ if } b_i = A \end{cases}$	$S_i = \begin{cases} 1 \text{ if } \text{d}_i = 3000/A \\ 59/60 \text{ if } \text{d}_i = 2700/A \\ 58/60 \text{ if } \text{d}_i = 2400/A \\ \vdots \\ 50/60 \text{ if } \text{d}_i = 0 \end{cases}$	Independent	Linear
7	1	$S_i = 1-\frac{b_i-b_{\min}}{2b_{\text{m}}}$	$S_i = 1-\frac{d_{\max}-d_i}{2(1-d_{\text{m}})}$	Interdependent	Linear
8	11	$S_i = \left(\frac{A-b_i}{A}\right)^{1/2}$	$S_i = d_i^{1/2}$	Independent	Concave
9	70	$S_i = \frac{A-b_i}{A-b_t}$	$S_i = d_i/d_t$	Independent	Linear
10	2	$S_i = 1-.156\frac{b_i-b_{\min}}{b_{\text{m}}-b_{\min}}$	$S_i = 1-.156\frac{d_{\max}-d_i}{d_{\max}-d_m}$	Interdependent	Linear
11	2	$S_i = 1-.517\frac{b_i-b_{\min}}{A}$	$S_i = 1-.517(d_{\max}-d_i)$	Interdependent	Linear
12	1	$S_i = \begin{cases} .833+.167\frac{A-b_i}{A} \text{ if } b_i < .95A \\ 16.667\frac{A-b_i}{A} \text{ if } b_i \geq .95A \end{cases}$	$S_i = \begin{cases} .833+.167d_i \text{ if } d_i > .05 \\ 16.667d_i \text{ if } d_i \leq .05 \end{cases}$	Independent	Concave
13	1	$S_i = (b_{\min}/b_i)^{1/2}$	$S_i = \left(\frac{1-d_{\max}}{1-d_i}\right)^{1/2}$	Interdependent	Convex

(*Continued*)

14	2	$S_i = \begin{cases} 1 \text{ if } b_i = b_{\min} \\ 11/12 \text{ if } 1 < b_i/b_{\min} \le 1.05 \\ 5/6 \text{ if } 1.05 < b_i/b_{\min} \le 1.10 \\ 3/4 \text{ if } 1.10 < b_i/b_{\min} \le 1.15 \\ 2/3 \text{ if } 1.15 < b_i/b_{\min} \le 1.20 \\ 1/3 \text{ if } 1.20 < b_i/b_{\min} \end{cases}$	$S_i = \begin{cases} 1 \text{ if } d_i = d_{\max} \\ 11/12 \text{ if } 1 < (1-d_i)/(1-d_{\max}) \le 1.05 \\ 5/6 \text{ if } 1.05 < (1-d_i)/(1-d_{\max}) \le 1.10 \\ 3/4 \text{ if } 1.10 < (1-d_i)/(1-d_{\max}) \le 1.15 \\ 2/3 \text{ if } 1.15 < (1-d_i)/(1-d_{\max}) \le 1.20 \\ 1/3 \text{ if } 1.20 < (1-d_i)/(1-d_{\max}) \end{cases}$	Interdependent	Concave
15	5	$S_i = \dfrac{A - b_i}{A - \min(b_{\min}, b_t)}$	$S_i = d_i/\max(d_{\max}, d_t)$	Interdependent	Linear
16	1	$S_i = \begin{cases} 3/5 \text{ if } b_i/A = 1 \\ 4/5 \text{ if } 1 > b_i/A \ge .97 \\ 9/10 \text{ if } .97 > b_i/A \ge .95 \\ 1 \text{ if } .95 > b_i/A \ge .93 \\ 9/10 \text{ if } .93 > b_i/A \ge .90 \\ 7/10 \text{ if } .90 > b_i/A \ge .88 \\ 3/5 \text{ if } .88 > b_i/A \ge .85 \\ 1/2 \text{ if } .85 > b_i/A \ge .83 \\ 1/10 \text{ if } .83 > b_i/A \ge 0 \end{cases}$	$S_i = \begin{cases} 3/5 \text{ if } d_i = 0 \\ 4/5 \text{ if } 0 < d_i \le .03 \\ 9/10 \text{ if } .03 < d_i \le .05 \\ 1 \text{ if } .05 < d_i \le .07 \\ 9/10 \text{ if } .07 < d_i \le .10 \\ 7/10 \text{ if } .10 < d_i \le .12 \\ 3/5 \text{ if } .12 < d_i \le .15 \\ 1/2 \text{ if } .15 < d_i \le .17 \\ 1/10 \text{ if } .17 < d_i \le 1 \end{cases}$	Independent	Concave

Notes: [a]*: Deemed as linear since the falls in* S_i *are of equal size.* b_t *and* d_t *are reference values calculated as a percentage of A and as a drop off A, respectively.*

also depends on the other bidders' price offers and hence cannot be calculated in advance. At stake is which of these two types of PSRs induces more aggressive bidding and thus lowers the price paid for the service. For Dini, Pacini, and Valletti (2006), interdependent PSRs give stronger incentives on price competition as participants, by bidding low, increase the likelihood to get the highest score reducing at the same time the score obtained by all other competitors. On the other hand, Albano, Dini, and Zampino (2009) argue that independent PSRs stimulate price competition as a supplier can calculate in advance the change in its score associated to a variation in the offered price. This issue is explored in the sample using the dummy variable *PSR_indep*, which indicates that the PSR is of the independent type. This occurs in 17.6 percent of the sample.

PSRs can also be classified attending to the shape of their gradient, *i.e.* to the sign of their second derivate with respect to b_i (Dini, Pacini, and Valletti 2006, Ballesteros-Pérez *et al.* 2015). A linear PSR is such that, as b_i decreases, S_i increases at a constant rate. But in a convex (respectively, concave) PSR, S_i increases at an increasing (decreasing) rate as b_i decreases, thus giving stronger (milder) rewards to high discounts and progressively favoring bidding on price (quality). Figure 1 depicts an example of each type. It is expected that, ceteris paribus, convex PSRs may lower the price paid for the service. This conjecture is investigated in the sample with the dummy variable *PSR_convex*, which indicates that the PSR is of the convex type. This occurs in 59.1 percent of the sample. Since only 15 procurements used a concave PSR, this type has been aggregated with the linear type[7].

ALTC set a threshold price bid (b_{abn}) or discount (d_{abn}) such that a price bid that fulfils the condition $b_i < b_{abn}$ or, equivalently, $d_i > d_{abn}$, will be disqualified unless the bidder can

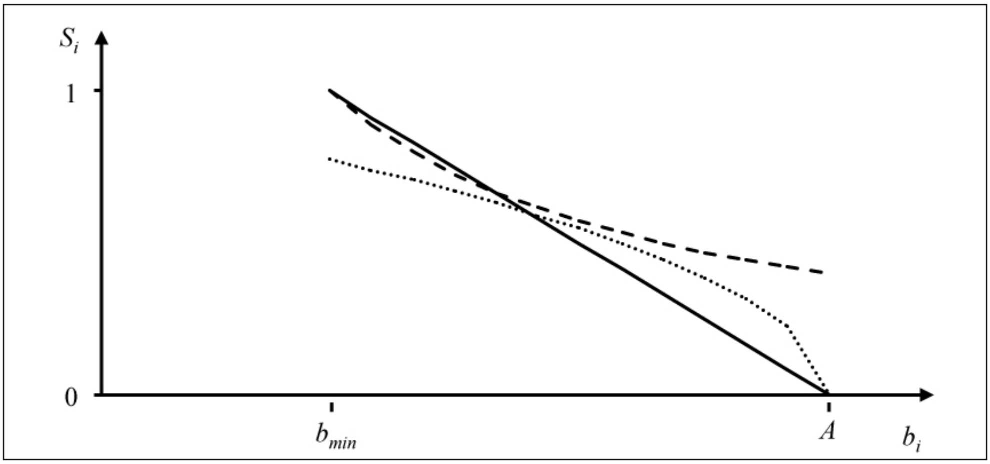

Figure 1: Linear, convex and concave price scoring

Notes: The figure depicts an example of three types of price scoring rules identified in the sample: linear ($S_i = \frac{A - b_i}{A - b_{\min}}$ *represented with a solid line), convex (* $S_i = \frac{b_{\min}}{b_i}$ *, dashed line), and concave (* $S_i = \left(\frac{A - b_i}{A}\right)^{1/2}$ *, dotted line).*

prove the price that it offered. Two types of ALTC can be identified in the sample attending to their independence from the final distribution of bids. The type of independent ALTC may be expressed generally as:

$$b_{abn} = (1 - \delta)A \tag{4}$$

or

$$d_{abn} = \delta, \tag{5}$$

where the parameter δ ranges between 0.10 and 1.00 (*i.e.* no threshold price bid) in the sample. Independent ALTC allow the economic threshold to be known in advance, so that bidders are certain about the bid below which they will be requested to prove the price that they offered.

Interdependent ALTC are such that the economic threshold cannot be known in advance. Interdependent ALTC are given in two basic forms. There are those that are a function of the average of the price bids submitted:

$$b_{abn} = (1 - \varepsilon)b_m, \tag{6}$$

or, equivalently,

$$d_{abn} = 1 - (1 - \varepsilon)(1 - d_m), \tag{7}$$

where the parameter ε ranges between 0.10 and 0.30 in the sample. There are also those whose expression depends on both the number of bidders (n) and b_m, as in[8]:

$$b_{abn} = (1 - .25)A \quad \text{if } n = 1 \tag{8}$$

$$b_{abn} = (1 - .20)b_{\max} \quad \text{if } n = 2 \tag{9}$$

$$b_{abn} = \max\left((1 - .10)b_{\mathrm{m}}, (1 - .25)A\right) \quad \text{if } n = 3 \tag{10}$$

$$b_{abn} = (1 - .10)b_{\mathrm{m}} \quad \text{if } n \geq 4 \tag{11}$$

or, equivalently,

$$d_{abn} = .25 \quad \text{if } n = 1 \tag{12}$$

$$d_{abn} = 1 - (1 - .20)(1 - d_{\min}) \quad \text{if } n = 2. \tag{13}$$

$$d_{abn} = \min(1 - (1 - .10)(1 - d_m), .25) \quad \text{if } n = 3 \tag{14}$$

$$d_{abn} = 1 - (1 - .10)(1 - d_m) \quad \text{if } n \geq 4. \tag{15}$$

While it might seem that the threshold defined in (6) is unrelated to n, the optimal price strategy in multidimensional bidding is a function of n (e.g., see Che 1993, David, Azoulay-Schwartz, and Kraus 2006, and Nakabayashi and Hirose 2016), so that when forming an expectation of b_m a rational bidder will have a belief about n. Hence, both basic forms may be grouped into a common type of ALTC. The ALTC in expressions (4) and (6) are among the six generic ALTC identified by Ballesteros-Pérez *et al*. (2015) in construction auctions worldwide.

Engel *et al*. (2006) have argued that interdependent ALTC will lead firms to bid higher so as to avoid the risk of exclusion, whereas Ballesteros-Pérez *et al*. (2015) have found that, when competing under independent ALTC, most bidders place their bids just before crossing the economic threshold, so as to avoid losing as much economic score as possible despite giving up some profits. These results seem to suggest that the price paid by the buyer might be higher under interdependent ALTC, but investigating this conjecture in the data requires controlling for differences in economic thresholds between ALTC. Unfortunately, this is not possible with the data at hand. Although δ and ε can be observed whenever the ALT criterion is of the type (4) or (6) (as it occurs in 18.9 and 30.4 percent of the sample, respectively), these parameters are not commensurate as they have different origins. Furthermore, the precise economic threshold that bidders may have taken into account in ALTC of the types (6) and (8)-(11) is not known.

For these reasons, the estimating model will allow for price differences among three groups of ALTC: those given by expressions (4), (6), and (8)-(11). The type of interdependent ALTC is split into two groups to account for possible differences in economic thresholds between them. The groups represented by (6) and (8)-(11) will be indicated with the dummy variables *ALTC_g2* and *ALTC_g3*, respectively. The estimates on these dummies measure not only the difference in price paid caused by the existence of different economic thresholds between groups of ALTC, but also the difference relative to the situation where bidders know the threshold in advance.

3.3. Other variables

The explaining variable is the natural logarithm of the price paid for the service. In the sample, prices for the contracts totaled €170.2 million, ranged between €997 and €5.3 million, and averaged €356,867. The main explanatory variables are the percentage weight given to price in the WF and the characteristics of the PSR and ALT criterion highlighted in the previous section. The regressions also include a quadratic function of the natural logarithm of the estimated cost[9], the contract term measured in months, an indicator for urgent procurement, an indicator for whether bidders must give a provisional guarantee, an indicator for whether the contract allows for extra payments, a measure of bid screening intensity by the contract awarding committee, service-type fixed effects, the procurement's vintage as represented by year dummies, and, for a subsample of 356 procurements, a quadratic function of *n*. A detailed definition of these variables plus some descriptive statistics are given in Table 3.

Table 3
DEFINITIONS OF VARIABLES AND DESCRIPTIVE STATISTICS

Variable	Description	Full sample, *N*=477 Mean (S.D.)	Subsample, *N*=356 Mean (S.D.)
Price	Winning bid inclusive of VAT, measured in €10,000s.	35.69 (79.20)	43.30 (87.23)
Weight for price	Percentage weight given to price in the award criteria, divided by 10.	6.19 (1.61)	6.34 (1.52)
PSR_indep	Dummy variable = 1 if bidders can calculate their price score before the bidding phase takes place.	0.18	0.17
PSR_convex	Dummy variable = 1 if the second derivate of the PSR with respect to the price bid is strictly positive.	0.59	0.63
ALTC_g1	Dummy variable = 1 if the ALT criterion is a discount off the estimated cost.	0.19	0.21
ALTC_g2	Dummy variable = 1 if the ALT criterion is a discount off the average price bid.	0.30	0.21
ALTC_g3	Dummy variable = 1 if the ALT criterion is a function of the number of bidders and the average price bid.	0.51	0.58

(*Continued*)

Variable	Description	Full sample, *N*=477 Mean (S.D.)	Subsample, *N*=356 Mean (S.D.)
Estimated cost	Estimated cost of the service inclusive of VAT, measured in €10,000s.	41.39 (90.83)	50.09 (100.05)
Contract term	Measured in months and exclusive of possible extensions.	10.51 (8.53)	11.06 (9.09)
Urgent	Dummy variable = 1 if procurement was deemed to be urgent.	0.15	0.08
Provisional guarantee	Dummy variable = 1 if bidders must guarantee that the quotes submitted will be maintained until the contract is awarded.	0.10	0.08
Price amend	Dummy variable = 1 if the contract price could be altered for reasons stated in the contract.	0.30	0.36
Screening	Difference (in weeks) between when the bids were opened by the awarding committee and when the identity of the winning bidder was announced.	6.02 (4.50)	6.29 (4.93)
Maintenance of vehicles/ apparatuses	Dummy variable = 1 if the contract was for repair or maintenance of vehicles or apparatuses.	0.21	0.21
Maintenance of buildings	Dummy variable = 1 if the contract was for repair or maintenance of building installations.	0.12	0.10
Hotel/Restaurant	Dummy variable =1 if the contract was for hotel, accommodation, restaurant, or food-serving services.	0.34	0.39
Other personal services	Dummy variable =1 if the contract was for transportation, telecommunications, or training services.	0.13	0.13
Refuse	Dummy variable = 1 if the contract was for sewage, refuse, or cleaning services.	0.06	0.03
Miscellaneous	Dummy variable =1 if the contract was for other services.	0.14	0.14
Year 2012	Dummy variable = 1 if the contract notice was published in 2012.	0.19	0.08
Year 2013	Dummy variable = 1 if the contract notice was published in 2013.	0.24	0.21
Year 2014	Dummy variable = 1 if the contract notice was published in 2014.	0.17	0.18
Year 2015	Dummy variable = 1 if the contract notice was published in 2015.	0.18	0.24
Year 2016	Dummy variable = 1 if the contract notice was published in 2016.	0.22	0.29
No. of bidders	Includes price bids above the estimated cost, presenting formal errors, or deemed abnormally low.		3.04 (2.83)

Notes: Author's calculations based on data compiled from the Spanish Public Sector Contracting Platform.

When procurement is deemed urgent, the period for sending bids is reduced from 15 days to just 8 days, which might induce the participation of fewer bidders and/or less intense bidding between participants, resulting in higher prices. A provisional guarantee (typically, 3 percent of the price bid) is intended to ensure that the quotes submitted will be maintained until the contract is awarded, so that quality reduction during the period between when the bids are submitted and when the contract is awarded is discouraged. As argued by Decarolis (2014), it is conceivable that an increase in the time taken to award the contract by the awarding committee comes with the benefit of selecting a more reliable contractor, but also at the expense of increasing the price paid. In the same vein, cost overruns were generally not allowed by the 2011 law except for reasons stated in the contract itself. In these cases, the prospect of extra payments might induce more aggressive bidding at the time of contracting and therefore lower prices (Goldberg 1977).

For a subsample of 356 procurements, the information stored in PLACE includes the number of bidders, *i.e.* the number of firms who submitted a quote. A firm may be a joint venture (Unión Temporal de Empresas), in whose case the companies that form the joint venture cannot submit offers individually. Firms who belong to the same group (which occurs when one company controls, directly or indirectly, the other/s) may submit offers individually, but only the group's lowest bid is used for calculating b_m in ALTC. Apparently, the information on n was missed exogenously, as (the log of) the price paid is redundant in a probit model for the observability of n, and the average number of bidders (3.0) closely resembles that observed in other samples of government service procurements (e.g. see Li and Zheng 2009). The quadratic specification allows the possibility of a varying empirical effect of n on the price paid.

4. Econometric model

The empirical strategy followed in this paper is to specify, estimate, and test a regression model for the natural logarithm of the price paid for the service (denoted by $\ln y$) on the set of explanatory variables ($\mathbf{x}$). The cancellation/completion outcome of a procurement is not analyzed due to the small number of canceled procurements in the sample (only 26 procurements were declared null and void).

Variable $\ln y$ for procurement k conducted by CA j is modeled as

$$\ln y_{jk} = \beta_0 + \mathbf{x}_{jk}\boldsymbol{\beta} + c_j + u_{jk}, \tag{16}$$

where u_{jk} is an idiosyncratic error term satisfying

$$E(u_{jk} \,|\, \mathbf{x}_{j1},\ldots,\mathbf{x}_{jK_j}, c_j) = 0, \quad k = 1,\ldots,K_j. \tag{17}$$

Vector $\mathbf{x}_{jk}$ contains the characteristics of procurements discussed in the previous section. Since most of these characteristics are decided by the CA, they might be influenced by unobserved features of the CA such as conventional practice, managerial ability, or workload. Suppose, for example, that busier agencies tended to overestimate the estimated cost of the contract so as to attract more bidders and reduce the probability that procurements be declared null and void (and consequently have to either repeat the process or negotiate the purchase with well-known, but perhaps not very efficient, potential providers: Albano *et al.* 2006). If, moreover, busier agencies tended to use PSRs of the independent type, then the estimate on PSR_indep would be biased in the positive direction. To control for the presence of unobserved factors at the CA level that were constant over the period of analysis, variable c_j is allowed to be arbitrarily correlated with the elements of $\mathbf{x}_{jk}$.

Procurements conducted by the same CA are ordered in the data file from the oldest to the most recent according to the date of publication of the contract notice. Thus, the contract notice of procurement $k = t$ was published no later than the contract notice of procurement $k = s$ for $s > t$. Batches belonging to the same contracting dossier (which have the same date of publication) are ordered in ascending order according to their official numbering in the dossier: batch #1, batch #2, etc. Note that in the case of batches belonging to the same dossier, potential feedback from y_{jt} to $\mathbf{x}_{js}$ for $s > t$ is ruled out.

The number of sample procurements managed by each CA (K_j) ranges between 1 and 97, with an average of 10.1 (*S.D.* = 17.4). With 47 CAs included in the sample and $\bar{K}_j = 10.1$, inference will be based on asymptotic properties of estimators assuming independence across CAs, but allowing for arbitrary heteroskedasticity and serial correlation in $\{u_{jk}: k = 1,...,K_j\}$. (As discussed below, the evidence suggests the $\{u_{jk}: k = 1,...,K_j\}$ are serially uncorrelated.)

The beta parameters of model (16) are estimated with the fixed effects (FE) estimator. The need for fixed effects estimation was assessed using the regression-based approach to the Hausman (1978) test of endogeneity proposed by Wooldridge (2010, p. 332), which is easier to robustify to the presence of non-spherical disturbances than the original procedure devised by Hausman (1978). Thus, to determine if c_j was correlated with $\mathbf{x}_{jk}$, the model

$$\ln y_{jk} = \gamma_0 + \boldsymbol{x}_{jk}\boldsymbol{\beta} + \bar{\mathbf{x}}_j\gamma + v_{jk}, \tag{18}$$

where $\bar{\mathbf{x}}_j = K_j^{-1}\sum_{k=1}^{K_j}\mathbf{x}_{jk}$, was estimated by the method of ordinary least squares (OLS) with clustering on CA. The null hypothesis of exogeneity was tested using a Wald test of $\gamma = \mathbf{0}$. (Previewing the results, the test's p-value was lower than 0.01 in all cases considered).

It is important to assess the degree of serial correlation in the idiosyncratic errors. Under strict exogeneity, if the u_{jk} are serially uncorrelated the FE estimator is more efficient than the first-difference (FD) estimator, whereas the opposite occurs whenever u_{jk} follows a random walk. If strict exogeneity fails but $E(\mathbf{x}'_{jk}u_{jk}) = \mathbf{0}$ holds and the u_{jk} are weakly dependent, the FE estimator can be expected to have less inconsistency for larger K_j. Furthermore, since the (unobserved) non-monetary dimensions set out in the contract tend to be constant across contracts of the same contracting dossier, they will induce serial correlation among idiosyncratic errors of contracts belonging to the same contracting dossier. Finding that the u_{jk} are serially uncorrelated would suggest that this omitted variable bias may be small.

To assess the degree of serial correlation in the u_{jk}, we used the result that if $\text{Corr}(u_{jt}, u_{js}) = 0$ for $t \neq s$, then $\rho \equiv \text{Corr}(\ddot{u}_{jt}, \ddot{u}_{js}) = -1/(K_j - 1)$ for all $t \neq s$ and $\rho_1 \equiv \text{Corr}(\Delta u_{jt}, \Delta u_{j,t-1}) = -0.5$, where $\ddot{u}_{jk}$ and Δu_{jk} denote, respectively, the time-demeaned and first-differenced errors (Wooldridge 2010, Ch. 10). To estimate ρ and ρ_1, $\hat{\ddot{u}}_{jt}$ was regressed on $\hat{\ddot{u}}_{j,t-1}$ and $\Delta\hat{u}_{jt}$ on $\Delta\hat{u}_{j,t-1}$ for $t \geq 3$ using residuals pertaining to CAs who managed at least three sample procurements (see Wooldridge 2010, pp. 311 and 320, for details). Results suggest that the u_{jk} are serially uncorrelated. In the full sample, $\hat{\rho} = 0.06$ (*S.E.* = 0.05) and $\hat{\rho}_1 = -0.51$ (*S.E.* = 0.03), whereas $\hat{\rho} = -0.07$ (*S.E.* = 0.12) and $\hat{\rho}_1 = -0.48$ (*S.E.* = 0.06) in the subsample with n included in **x**. The FE estimator was implemented in Stata using the command *xtreg* ln *y* x, *fe vce(cluster ID_CA)*, where *ID_CA* is the CA identifier.

5. Results

For comparison purposes, Table 4 presents OLS estimates of equation (16). Column (1) contains the estimates obtained in the full sample, whereas columns (2) and (3) show, respectively, the estimates obtained in the subsample first excluding and then including the quadratic function of n in order to assess the impact of including in **x** n. The three sets of estimates of $\boldsymbol{\beta}$ suggest that, with other factors being unchanged, the price paid for the service varies proportionally with the estimated cost[10], increases approximately by 0.40 percent with each additional month of contract, is about 9 to 10 percent higher when a provisional guarantee must be given, is 15 to 19 percent higher for hotel/restaurant services, and diminishes with n and with the weight given to price in the contract award criterion (an additional bidder reduces the price paid by 4.32 percent on average[11], while an increase of 10 percentage points in W_p reduces the price paid by around 3 percent). The estimate on *ALTC_g3* is statistically different from zero at 10 percent, and suggests that ALTC of the type (8)-(11) reduce the price paid for the service by around 6 percent on average. The estimate on *PSR_convex* is consistently positive but imprecisely measured when n in included in **x**. The estimates on *PSR_indep* and *ALTC_g2* tend to be small and do not attain statistical significance.

The last row of Table 4 shows the *p*-value of the regression-based method described in the previous section to test whether c_j and $\mathbf{x}_{jk}$ are correlated. The *p*-value is lower than 0.01 in the three cases shown, so that the assumption of zero correlation between c_j and $\mathbf{x}_{jk}$ is questioned. Therefore, OLS applied to equation (16) may be biased as a result of unobserved heterogeneity at the CA level.

The results of the FE estimation of equation (16) are presented in Table 5, whose layout is analogous to Table 4. The estimated coefficients on the quadratic functions of the log of the estimated cost and of n lie close to the OLS estimates. The implied elasticity of the price paid with respect to the estimated cost ranges between 1.00 and 1.02 (*S.E.* = 0.01 in all cases), whereas, with other factors being unchanged, an additional bidder reduces the price paid by 4.29 percent on average (*S.E.* = 0.56). The estimates on *Contract term* and *Hotel/ Restaurant* become much smaller and statistically insignificant once the unobserved effect is eliminated by using FE. Therefore, in the OLS estimation, the main part of the additional price paid as indicated by these two variables reflected unobserved features of CAs that were constant over the period of analysis. The effect of *Provisional guarantee* is still positive and large, but it is measured less precisely.

Table 4
ORDINARY LEAST SQUARES ESTIMATES OF LN OF PRICE PAID

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Weight for price (÷10)	-.0331***	.0083	-.0402***	.0113	-.0310***	.0114
PSR_indep	-.0053	.0392	.0070	.0505	-.0330	.0484
PSR_convex	.0511*	.0276	.1017**	.0396	.0573	.0412
ALTC_g2	.0119	.0319	-.0211	.0360	-.0300	.0317
ALTC_g3	-.0653**	.0301	-.0761**	.0327	-.0606*	.0301
Ln of estimated cost	1.0530***	.1003	1.0969***	.0800	1.1405***	.0843
(Ln of estimated cost)^2	-.0022	.0044	-.0040	.0035	-.0055	.0036
Contract term (months)	.0044**	.0021	.0040*	.0023	.0040*	.0023
Urgent	.0306	.0332	-.0200	.0569	-.0287	.0510
Provisional guarantee	.1021**	.0434	.1048**	.0508	.0904**	.0375
Price amend	-.0161	.0346	-.0281	.0419	-.0405	.0361
Screening (weeks)	-.0051	.0052	-.0035	.0042	-.0004	.0036
Maintenance of buildings	.0723	.0534	.0490	.0503	.0940	.0583
Hotel/Restaurant	.1918***	.0511	.1825***	.0527	.1499***	.0478
Other personal services	.0576	.0498	.0407	.0555	.0049	.0539
Refuse	.0992	.0719	.0610	.0667	.0518	.0600
Miscellaneous services	.0481	.0422	.0421	.0406	.0360	.0402
Year 2013	-.0425	.0324	-.0308	.0446	-.0401	.0448
Year 2014	.0427	.0290	.0622	.0476	.0263	.0518
Year 2015	.0334	.0380	.0763*	.0454	.0262	.0479
Year 2016	.0034	.0315	.0309	.0415	.0125	.0441
No. of bidders					-.0517***	.0105
(No. of bidders)^2					.0014***	.0004
Intercept	-.4234	.5811	-.6808	.4476	-.8387*	.4863

(*Continued*)

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	**S.E.**	**Coefficient**	**S.E.**	**Coefficient**	**S.E.**
R-squared	.9892		.9900		.9918	
Observations	477		356		356	
Wald test of $\gamma = \mathbf{0}$ *in eq. (18)*	[.0041]		[.0000]		[.0000]	

*Notes: Standard errors are clustered on contracting agency. Figures in brackets are p-values. Unreported categories: ALTC_g1, Maintenance of vehicles/apparatuses, Year 2012. * Significant at 10%; ** significant at 5%; *** significant at 1%.*

The estimated coefficient on W_p has decreased slightly, suggesting a -2.45 percent effect (*S.E.* = 1.41) in the subsample with n included in $\mathbf{x}$. The size of this effect is substantially smaller than the 8 to 16 percent negative effect found by Koning and van de Meerendonk (2014) in a large sample of public procurements of services for reintegrating unemployed and disabled workers in the Netherlands. In unreported regressions, we found no evidence of a non-linear effect of the weight for price on ln y.

The estimates on *PSR_indep* remain small and statistically insignificant, although the estimate obtained in the subsample with n included in $\mathbf{x}$ hints at the possibility suggested by Albano, Dini, and Zampino (2009), that independent PSRs lower the price paid by inducing more aggressive bidding. The estimates on *PSR_convex* have become negative and, in the case presented in column (3), larger in absolute value and statistically significant at 5 percent. This estimate suggests that, with other factors being unchanged, convex PSRs lower the price paid for the service by 8.81 percent on average (*S.E.* = 3.83). The direction of this effect concurs with Dini, Pacini, and Valletti (2006), who expected lower prices from "competitive" PSRs, and with Ballesteros-Pérez *et al*. (2016), who provide some evidence that concave PSRs cause a higher level of bidding conservativeness and bid concentration.

The estimates on *ALTC_g2* remain small and statistically insignificant. The estimates on *ALTC_g3* have become positive, smaller in absolute value, and statistically insignificant. Overall, therefore, we find no evidence that the type of ALT criterion influences the price paid for the service, although, admittedly, the effects are estimated somewhat imprecisely. Ballesteros-Pérez *et al*. (2016) provide some evidence that bidding aggressiveness and bid dispersion increase with ε in a sample of 124 construction auctions whose unique generic mathematical expression of ALTC is given by (6).

In this and the following paragraphs, the previous findings are subjected to further scrutiny. Consistency of the FE estimator hinges on the strict exogeneity assumption (17). Wooldridge (2010, p. 325) proposes a test of strict exogeneity using fixed effects when $K_j > 2$, which consists in testing $\boldsymbol{\mu} = \mathbf{0}$ in the expanded equation

$$\ln y_{jk} = \beta_0 + \boldsymbol{x}_{jk}\boldsymbol{\beta} + \boldsymbol{w}_{j,k+1}\boldsymbol{\mu} + c_j + u_{jk}, \quad k = 1,\ldots,K_j - 1, \tag{19}$$

where $\boldsymbol{w}_{j,k+1}$ is a subset of $\boldsymbol{x}_{j,k+1}$ and the test is carried out using FE estimation. Included in $\boldsymbol{w}_{j,k+1}$ are the weight for price plus the characteristics of the PSR and the ALT mechanism highlighted in Section 3.2, as the CA could (in principle) easily adjust these factors for future procurements based on shocks to y_{jk}. The last row of Table 5 presents the *p*-value of this test. In all cases, the claim of strict exogeneity is within confidence bounds.

Table 5
FIXED EFFECTS ESTIMATES OF LN OF PRICE PAID

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Weight for price (÷10)	-.0203***	.0075	-.0285**	.0139	-.0245*	.0141
PSR_indep	-.0021	.0551	-.0086	.0689	-.0539	.0529
PSR_convex	-.0391	.0503	-.0566	.0531	-.0881**	.0383
ALTC_g2	.0093	.0617	-.0365	.0755	-.0239	.0519
ALTC_g3	.0191	.0668	.0065	.0755	.0392	.0657
Ln of estimated cost	1.0112***	.0728	1.0857***	.0551	1.0907***	.0512
(Ln of estimated cost)^2	-.0007	.0032	-.0033	.0022	-.0031	.0020
Contract term (months)	.0010	.0024	.0002	.0035	-.0001	.0025
Urgent	.0321	.0255	.0524	.0512	.0305	.0500
Provisional guarantee	.0758	.0733	.0601	.0986	.0686	.0827
Price amend	-.0083	.0394	-.0008	.0590	-.0269	.0433
Screening (weeks)	-.0117*	.0059	-.0090	.0062	-.0047	.0038
Maintenance of buildings	.0376	.0620	.0511	.0778	.0356	.0450
Hotel/Restaurant	.0767	.0692	.0232	.0728	.0149	.0486
Other personal services	.0499	.0462	.1008	.0923	.0076	.0855
Refuse	.0370	.0590	-.0213	.1109	-.0309	.0854
Miscellaneous services	.0099	.0433	.0183	.0517	-.0141	.0384
Year 2013	-.0332	.0388	-.1141	.0681	-.1197**	.0513
Year 2014	.0394	.0261	.0169	.0544	-.0044	.0463
Year 2015	.0273	.0387	-.0093	.0585	-.0312	.0510
Year 2016	-.0216	.0295	-.0751	.0591	-.0526	.0502
No. of bidders					-.0531***	.0076
(No. of bidders)^2					.0017***	.0004
Intercept	-.0792	.4366	-.4633	.3435	-.4144	.3408
R-squared (within)	.9727		.9699		.9747	
Observations	477		356		356	

(*Continued*)

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Wald test of $\boldsymbol{\mu = 0}$ *in eq. (19)*	[.4115]		[.1140]		[.1238]	

*Notes: Standard errors are clustered on contracting agency. Figures in brackets are p-values. Unreported categories: ALTC_g1, Maintenance of vehicles/apparatuses, Year 2012. * Significant at 10%; ** significant at 5%; *** significant at 1%.*

Contracting agencies may have an incentive to overestimate the estimated cost to reduce the probability that the contract be declared null and void. But, by the "exclusion principle" (Albano *et al*. 2006), an agency pursuing the minimization of the price paid may find it in its interest to set a low estimated cost[12]. In any case, mismeasuring the general market price of the service introduces measurement error into the estimated cost, which will also induce biases on coefficients of any regressors correlated with this variable. To account for this possibility, procurements whose savings rate (calculated as the difference between the estimated cost and the winning bid, both inclusive of VAT, divided by the estimated cost) was above 57 percent (the 97.5th percentile of the savings rate distribution) have been removed from the sample, as very high savings rates are likely to be the result of mismeasuring the general market price of the service. The results of re-estimating equation (16) by the FE procedure are presented in Table 6. The main findings of this study are preserved. (The implied elasticity of the price paid with respect to the estimated cost still ranges between 1.00 and 1.02, *S.E.* = 0.01 in all cases.)

Table 6
FIXED EFFECTS ESTIMATES OF LN OF PRICE PAID. EXCLUDING SAVINGS RATES ABOVE 57%

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Weight for price (÷10)	-.0238***	.0083	-.0290**	.0131	-.0255**	.0126
PSR_indep	.0054	.0492	.0145	.0640	-.0265	.0455
PSR_convex	-.0255	.0473	-.0486	.0548	-.0816**	.0381
ALTC_g2	.0081	.0577	-.0588	.0768	-.0490	.0523
ALTC_g3	-.0007	.0607	-.0158	.0771	.0123	.0624
Ln of estimated cost	.9837***	.0681	1.0585***	.0513	1.0733***	.0405
(Ln of estimated cost)^2	.0008	.0027	-.0022	.0021	-.0025	.0016
Contract term (months)	.0010	.0019	.0001	.0031	-.0003	.0022
Urgent	.0232	.0211	.0625	.0503	.0424	.0450
Provisional guarantee	.0869	.0698	.0575	.0909	.0638	.0763

(*Continued*)

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Price amend	-.0114	.0446	-.0186	.0652	-.0376	.0512
Screening (weeks)	-.0094**	.0047	-.0071	.0050	-.0033	.0033
Maintenance of buildings	.0312	.0459	.0217	.0730	.0077	.0612
Hotel/Restaurant	.0533	.0527	-.0161	.0596	-.0247	.0358
Other personal services	.0302	.0372	.0286	.0713	-.0555	.0699
Refuse	.0350	.0452	-.0775	.0832	-.0854	.0527
Miscellaneous services	.0117	.0307	-.0102	.0442	-.0360	.0359
Year 2013	-.0190	.0268	-.0952	.0651	-.1020**	.0496
Year 2014	.0314	.0258	.0137	.0553	-.0059	.0482
Year 2015	.0309	.0398	-.0133	.0575	-.0329	.0511
Year 2016	-.0072	.0307	-.0527	.0581	-.0344	.0500
No. of bidders					-.0416***	.0087
(No. of bidders)^2					.0009	.0005
Intercept	.0682	.4407	-.2404	.3229	-.2671	.2849
R-squared (within)	.9825		.9805		.9846	
Observations	466		347		347	

*Notes: Standard errors are clustered on contracting agency. Unreported categories: ALTC_g1, Maintenance of vehicles/apparatuses, Year 2012. * Significant at 10%; ** significant at 5%; *** significant at 1%.*

PSRs numbered 1, 6, 11, 13, 14, and 16 in column (1) of Table 2 do not assign the complete price score differential because they always award a strictly positive score to the highest price bid. Suppose, for example, that $W_p = 80\%$, $S_i = b_{\min}/b_i$, $b_{(1)} = 70$, and $b_{(n)}=100$. Then, the percentage price score received by the lowest bid would be 80, and that received by the highest bid would be 56, so that the price score differential would be 24 percentage points rather than 80 percentage points as suggested by W_p. This is an example of a phenomenon called *apparent or phony economic bid weighting* by Ballesteros-Pérez *et al*. (2015), which occurs not only whenever a fraction of S_i is always awarded, but also whenever a fraction of S_i is unreachable. A practical implication of this phenomenon is that the actual weight given to price will be generally lower than stated in the procurement specifications (see Ballesteros-Pérez *et al*. 2015), so that if bidders perceived this circumstance in advance, their incentives to bid on price would be reduced.

To account for this possibility, equation (16) was re-estimated by the FE procedure with the weight given to price interacted with the dummy variable *PSR_LPSD*, which indicates that the PSR limits the price score differential. We would expect the coefficient on *Weight for price*PSR_LPSD* to be positive, so that an increase of 10 percentage points in W_p would have a smaller effect (in absolute value) in procurements whose PSR limited the price score differential. Table 7 shows the results. The estimate on *Weight for price*PSR_LPSD* is positive though small, so that the variable is economically and statistically insignificant. Note also that for these data it is difficult to disentangle the separate effects of *PSR_convex* and *Weight for price*PSR_LPSD*, as both variables are highly correlated (the Pearson's product-moment correlation coefficient equals 0.95). Therefore, when both are included in **x** the estimates become very imprecise.

Table 7
FIXED EFFECTS ESTIMATES OF LN OF PRICE PAID. INCLUDING *WEIGHT FOR PRICE* INTERACTED WITH *PSR_LPSD*

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	S.E.	Coefficient	S.E.	Coefficient	S.E.
Weight for price (÷ 10)	-.0239**	.0118	-.0337*	.0200	-.0331**	.0128
Weight for price PSR_LPSD*	.0064	.0130	.0086	.0222	.0144	.0165
PSR_indep	-.0033	.0544	-.0101	.0662	-.0565	.0505
PSR_convex	-.0757	.0516	-.1081	.1207	-.1742	.1079
ALTC_g2	.0114	.0599	-.0361	.0744	-.0230	.0505
ALTC_g3	.0194	.0662	.0052	.0762	.0372	.0681
Ln of estimated cost	1.0102***	.0725	1.0875***	.0554	1.0931***	.0491
(Ln of estimated cost)^2	-.0007	.0032	-.0033	.0022	-.0031	.0019
Contract term (months)	.0012	.0025	.0004	.0035	.0003	.0025
Urgent	.0329	.0256	.0523	.0515	.0304	.0508
Provisional guarantee	.0787	.0755	.0579	.0984	.0650	.0813
Price amend	-.0081	.0399	.0006	.0590	-.0247	.0438
Screening (weeks)	-.0116*	.0060	-.0090	.0062	-.0047	.0037
Maintenance of buildings	.0389	.0610	.0519	.0778	.0364	.0448
Hotel/Restaurant	.0744	.0681	.0201	.0721	.0096	.0485
Other personal services	.0478	.0468	.0978	.0948	.0022	.0866
Refuse	.0430	.0587	-.0129	.1134	-.0169	.0903
Miscellaneous services	.0129	.0430	.0222	.0530	-.0080	.0404
Year 2013	-.0335	.0389	-.1150	.0684	-.1209**	.0530
Year 2014	.0396	.0255	.0173	.0550	-.0035	.0467
Year 2015	.0284	.0387	-.0084	.0595	-.0295	.0519
Year 2016	-.0203	.0287	-.0738	.0594	-.0502	.0502

(*Continued*)

Explanatory variable	(1) Full sample		(2) Subsample		(3) Subsample	
	Coefficient	**S.E.**	**Coefficient**	**S.E.**	**Coefficient**	**S.E.**
No. of bidders					-.0538***	.0078
(No. of bidders)^2					.0017***	.0004
Intercept	-.0599	.4349	-.4534	.3330	-.3934	.3288
R-squared (within)	.9727		.9699		.9747	
Observations	477		356		356	

*Notes: Standard errors are clustered on contracting agency. Unreported categories: ALTC_g1, Maintenance of vehicles/apparatuses, Year 2012. * Significant at 10%; ** significant at 5%; *** significant at 1%.*

6. Conclusion

Using dimensions in addition to price in supplier selection guarantees more effectiveness when managing the trade-off between price and quality, but it adds complexity to the procedure. Mathematical criteria need to be included in the procurement specifications whenever there is need for converting a quote's price and quality dimensions into commensurate scales before being combined into a total score. Many weighing functions are currently in use for evaluating bid proposals in multidimensional procurement. These may affect the price paid by the buyer by promoting or discouraging competition between bidders, but their design has been little informed by empirical considerations.

This paper contributes to filling this gap by providing econometric evidence on the extent to which properties of weighing functions such as the weight given to price, the type of price scoring rule, and the type of abnormally low tenders criterion, affect the price paid in a sample of 477 procurements of support services conducted by the Spanish Armed Forces in the period 2012-2016. The main limitation of this study is that the econometric model does not include information on the specific non-monetary dimensions set out in the contract nor on the contractors' actual performance, although some evidence suggests that the consequences of omitting the former may be small.

We find that using a convex price scoring rule lowers the price paid by about 9 percent, and that increasing the weight given to price in 10 percentage points reduces the price paid by approximately 2.5 percent. The latter effect is substantially smaller to that found by Koning and van de Meerendonk (2014) in a different industry, whereas the former concurs with Dini, Pacini, and Valletti's (2006) insightful analysis of PSRs and provides additional support for Ballesteros-Pérez *et al.*'s (2016) finding that concave PSRs cause a higher level of bidding conservativeness and bid concentration.

On the other hand, we find little evidence of price adjustments in response to the independent/interdependent character of the price scoring rule and to the type of abnormally low tenders criterion used in the auction. The latter result may be driven by the fact that, with the data at hand, it is not possible to isolate the effect related to the type of ALT criterion by controlling for different economic thresholds across ALTC. On the other hand, all these results are not driven by contracting agencies' unobserved heterogeneity and are robust to a variety of alternative specifications.

Notes

1. A WF can also combine two or more quality scores into a single overall quality score, or two or more prices into a single overall price (Bergman and Lundberg 2013).
2. The word *tender* is used in this paper as a noun meaning "the bid" or "the price bid".
3. ALTC are commonly used mechanisms to identify unreasonably low tenders in many EU countries. Engel et al. (2006) discuss other methods used to mitigate the problem of risky bids.
4. Whenever the contracting dossier value was difficult to pinpoint, there was overriding urgency, or a previous procurement for the service had been declared void, among other circumstances, the purchase could be negotiated with at least three potential providers.
5. As pointed out by Fernández Roca (2011), the purchasing of support services by the SAF spread as a consequence of the reduction in the number of recruits that followed the abolition of compulsory military service on January 1, 2002. Coinciding with this reduction in personnel, the SAF budget for procuring commercial services grew from €48.6 million in 1997 to €122.9 million in 2002, and to €145.5 million in 2017 (all figures expressed in 2016 euros), accounting respectively for 0.61, 1.48, and 1.93 percent of the Spanish Ministry of Defense budget.
6. These statistics are called scoring parameters by Ballesteros-Pérez, González-Cruz, and Cañavate-Grimal (2012).
7. A PSR indicating the price-score reduction is smaller near b_{min} is called convex in Ballesteros-Pérez *et al.* (2016) but concave in this paper. The results of Ballesteros-Pérez *et al.* (2016) cited here have been adapted to the terminology used in this paper.
8. Article 85 of Spanish Royal Decree 2001/1098.
9. The fit of the regressions was much better with the estimated cost entering in log form.
10. The elasticity of the price paid with respect to the estimated cost is given by $\beta_{\text{ln cost}} + 2\beta_{\text{ln cost}^2}$ (*Ln of estimated cost*), where $\beta_{\text{ln cost}}$ and $\beta_{\text{ln cost}^2}$ are the coefficients associated to *Ln of estimated cost* and (*Ln of estimated cost*)^2, respectively. This elasticity ranges between 1.00 and 1.01 (*S.E.* = 0.01 in all cases) for the regressions shown in Table 4. Note that when the marginal effect of an explanatory variable depends on some regressor, it is calculated averaging across procurements.
11. The semielasticity of the price paid with respect to n is given by $\beta_{\text{bidders}} + 2\beta_{\text{bidders}^2}$ (*No. of bidders*), where β_{bidders} and β_{bidders^2} are the coefficients associated to *No. of bidders* and (*No. of bidders*)^2, respectively.
12. Thus, any supplier efficient enough to be able to submit an acceptable offer knows that competitors will belong to a pool of efficient participants, so that submitted offers will be more aggressive than in the presence of a higher estimated cost.

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Resumen

Muchas son las fórmulas en uso para evaluar proposiciones en adjudicaciones basadas en varios criterios. Aunque estas fórmulas pueden repercutir en el precio, su diseño ha estado poco influido por consideraciones empíricas. Este trabajo evalúa el impacto que las ponderaciones, las funciones de puntuación del precio (FPP), y las normas sobre ofertas con valores anormales (NOVA) ejercen sobre el precio en una muestra novedosa de contratos de servicios. La riqueza y estructura de panel de los datos permiten mostrar que FPP convexas y una mayor ponderación del precio reducen este, mientras que FPP independientes/interdependientes y las NOVA lo dejan inalterado.

Palabras clave: Adjudicaciones basadas en varios criterios, fórmulas para evaluar proposiciones, funciones de puntuación, normas sobre ofertas con valores anormales, Fuerzas Armadas españolas.

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