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Contract Duration and the Costs of Market Transactions†

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The optimal duration of a supply contract balances the costs of re-selecting a supplier against the costs of being matched to an inefficient supplier when the contract lasts too long. I develop a structural model of contract duration that captures this trade-off and provide an empirical strategy for quantifying (unobserved) transaction costs. I estimate the model using federal supply contracts for a standardized product, where suppliers are selected by procurement auctions. The estimated transaction costs are substantially greater than consumer switching costs and a significant portion of total buyer costs. Counterfactuals illustrate the importance of accounting for the duration margin. (*JEL* D22, D23, D44, D86, H57, L14)

When buyers select sellers, they select not only *who* but also *how long*. For the supply of services, the duration of a buyer–seller relationship is often formalized in a contract. At the expiration of the contract, the buyer returns to the market to re-select a supplier. The buyer bears some costs for doing so: each time the buyer goes to market, he must identify potential sellers, negotiate, determine the new seller, and draw up a contract (Coase 1937, 1960). These *market transaction costs* all occur before an agreement is reached.[1] As noted by Coase (1937), a key motivation for a longer contract is to avoid these costs. Choosing a two-year contract instead of sequential one-year contracts can cut these ex ante costs in half.

Though going to the market frequently can be costly, there are typically gains from doing so. Consider a procurement auction that selects the efficient (lowest-cost) supplier. Over a longer contract, the lowest-cost supplier at the beginning of the contract may not have the lowest cost by the end. By running more frequent auctions

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[1] Market transaction costs correspond to the first two categories of the taxonomy of transaction costs suggested by Dahlman (1979): (i) search and information costs and (ii) bargaining and decision costs. The third category, (iii) policing and enforcement costs, relates to the principal-agent problem and incomplete contracts, which have been the focus of the literature on transaction costs following Williamson (1979).

with shorter contracts, the buyer can switch among the lowest-cost suppliers and pay a lower price. In the absence of transaction costs, the buyer would prefer a spot market that allocates the lowest-cost supplier in every instant.

In the exchange of intermediate goods, market transactions can be especially costly. Unlike the retail sector, markets for intermediate goods are typically not well established. When running a procurement auction, a buyer brings together potential sellers and administers a mechanism to determine the winner and the price. In a sense, the buyer bears the cost of creating his market. Each transaction typically requires market research, advertising the opportunity, and additional steps to ensure that the process complies with the buyer's internal policies. Even for standardized goods and services (e.g., raw materials, electricity, paper products, and accounting services), market transaction costs can be large, as exact specifications vary among buyers. Indeed, as inputs, these products are typically sold on fixed-price, fixed-duration contracts, rather than in spot markets.[2] Thus, we may infer that market transaction costs are meaningful for a wide variety of goods and services.

Despite their importance, quantifying transaction costs has proven challenging, as they are typically unobserved. To address this, I develop an empirical model where a buyer selects a seller from an imperfectly competitive market. The buyer chooses the duration of the contract to balance ex ante transaction costs against the benefit from selecting more efficient suppliers. This trade-off is intuitive and corresponds to a common real-world contracting problem. The model also provides a novel strategy for the direct estimation of transaction costs. By revealed preference, transaction costs may be identified from the duration of the contract and the price schedule faced by the buyer. I apply the model in the context of federal procurement for a standardized product, and I find that market transaction costs are large and economically meaningful, comprising 10.9 percent of total buyer costs. Thus, I provide the first estimates of transaction costs in intermediate goods markets.[3]

The estimates suggest that transaction costs can be quite large. Though a detailed analysis of what comprises these costs is limited by the available data, a rough breakdown indicates that several components may be meaningful sources of ex ante transaction costs, including formal due diligence procedures and the direct costs of the bidding mechanism. Further work is needed to better understand these costs and their impacts in various settings.

The model of optimal contract duration and its main implications are presented in Section I. One novel prediction of the model is that, in equilibrium, a longer contract would result in a higher price. This arises from straightforward economic logic: if a longer duration would reduce the price, the buyer would prefer it, because it would

[2] In their seminal NBER survey, Stigler and Kindahl (1970) found that about half of the commodities in their sample were purchased with fixed-term contracts. A more recent comprehensive survey has not been conducted and would be welcome. Fixed-price contracts constitute over 90 percent of US federal government contracts (source: Federal Procurement Data System), and, anecdotally, remain predominant in the private sector.

[3] Economists studying the effects of transaction costs have primarily pursued the testable implications of these costs, rather than their direct estimation (see, e.g., Monteverde and Teece 1982, Walker and Weber 1984). One recent exception is Atalay et al. (2019), who construct a measurement of external transaction costs by examining input flows between integrated and nonintegrated firms across sectors. Likewise, the literature on contract duration has also focused on testable implications rather than structural modeling.

also reduce the burden of transaction costs.[4] A price schedule that is increasing with duration is the natural result of time-varying supply costs among suppliers, and it may arise from, e.g., capacity constraints or idiosyncratic outside options. Therefore, the model provides an economic justification for shorter contracts.

I focus on a setting in which goods and services are standardized, there is little uncertainty, and relationship-specific investments are negligible. Even in these straightforward economic environments, the duration decision is nontrivial, depending on (i) the magnitude of the transaction costs; (ii) the degree of competition; and (iii) the stochastic properties of the underlying supply costs. I present a simplified version of the model to provide some intuition about these features. Perhaps surprisingly, the optimal duration is nonmonotonic in the degree of competition. With few suppliers, the benefit of re-selecting a supplier is small, and long-term contracts are optimal. This benefit increases as the number of suppliers increases, leading to short-term contracts at moderate levels of competition. With many suppliers, the buyer can find a seller that provides a low-enough price over many periods, so long-term contracts are once again optimal. Likewise, higher variance in supply costs could lead to longer or shorter contracts. These ambiguous predictions help motivate a structural approach to estimation.

For an empirical application, I select a specialized setting that allows me to isolate the trade-off described above and recover estimates of market transaction costs. I construct a unique dataset of 1,046 contracts for building cleaning services for the U.S. federal government. Consistent with the model, duration is determined ex ante by the local government agency, and, for the contracts I analyze, the government is required to go to market at the expiration of the previous contract. Importantly, building cleaning services are standardized, supply-side conditions are stable, and relationship-specific investments are small.[5] This suggests that abstracting away from other contracting concerns may be reasonable, and it allows me to focus on identifying the direct (Coasian) costs of going to the market.

For context, each year the federal government manages over one million contracts that have an annual value of less than \$1 million. These constitute 97 percent of all federal contracts and are disproportionately made with fixed-price, fixed-duration contracts through competitive procedures.[6] In prices, the estimation sample is roughly comparable to these contracts and closely resembles the full set of building cleaning contracts. The data are presented in Section II, along with descriptive regressions that are used to motivate the structural model. I verify a core prediction of the model: in the data, longer contracts are more expensive. Therefore, time-varying supply costs appear to outweigh potential supplier-side benefits from a longer contract (e.g., learning), which are likely small in this setting.

[4] By contrast, the prevailing wisdom in the transaction costs literature is that longer contracts tend to reduce supply costs by solving ex post incentive problems.

[5] Ex post incentive problems, which are a large focus of the contract literature, are not a first-order concern here. Contracts have detailed specifications, performance is observable, and contracts are rarely canceled. I discuss this further in Section IIB. Hyytinen, Lundberg, and Toivanen (2018), who study cleaning contracts in Sweden, make similar observations and also assume complete contracts. As is common in the auction literature, Hyytinen, Lundberg, and Toivanen (2018) do not analyze the duration margin.

[6] Source: Federal Procurement Data System.

Section III presents the specific modeling assumptions and parameterizations used to take the model to data. Consistent with the empirical setting, the model takes three stages: (i) the buyer's duration decision; (ii) a participation decision by suppliers; and (iii) a first-price auction. Thus, compared to a standard auction model with endogenous entry, the model also incorporates a strategic decision by the buyer (duration). As in Krasnokutskaya (2011), I allow for unobserved heterogeneity across auctions. I show that the joint distribution of private costs and unobserved heterogeneity are identified when only the winning bid and the number of bidders are observed, thus extending identification to data that are more broadly available.[7] Intuitively, variation in the number of bids shifts the distribution of the private component in a known way, while the distribution of auction-specific heterogeneity is unaffected.

Section IV presents the model estimates. The median estimated transaction cost is \$10,400, representing a meaningful portion of total buyer costs. Though providing a detailed breakdown of these costs lies beyond the scope of the paper—in part due to the fact that they are not directly observed—I provide some discussion of what comprises these costs using supplementary data. In magnitude, the estimates are roughly comparable to back-of-the-envelope calculations of the labor costs of procurement specialists. Further, I find that these costs are correlated with other observables in ways that align with our intuition. For example, the costs are positively correlated with the complexity of the facility: contracts for medical buildings have much higher transaction costs than those for offices.

In Section V, I provide two counterfactual exercises to illustrate the impact of the duration margin, which is generally not accounted for in empirical studies of procurement or other business-to-business settings. First, I consider the value to the buyer of the strategic ability to change the duration of the contract. Instead of optimizing for each contract, I consider an alternative policy where all contracts are issued with a standard duration. Mandating more frequent transactions could be costly. For instance, issuing only one-year contracts would increase total costs by 37 percent. Of standard contracts with full-year durations, the four-year standard term has the lowest impact, increasing total costs by 1.4 percent. Therefore, a poorly chosen standard could substantially increase costs, but an informed standard may have modest effects.

As a second counterfactual, I demonstrate the impact of endogenous contracts on the estimation of welfare effects. To illustrate the importance of this margin, I consider the effects on cost pass-through. When buyers can adjust duration, the pass-through of supply costs to prices is reduced by 10 percent, compared to a world in which duration is held fixed. Thus, appropriately modeling contract duration can change the interpretation of observed price changes and the estimation of welfare effects. Further, transaction costs may be a sizable portion of total costs and should be accounted for in addition to any price effects.[8]

[7] Previous approaches relied on observing either multiple bids per auction (Krasnokutskaya 2011; Hu, McAdams, and Shum 2013) or a reservation price (Roberts 2013). Aradillas-López, Gandhi, and Quint (2013) exploit variation in the number bids for second-price auctions, though the identification results of their paper are limited to constructing bounds on surplus. Concurrent work by Quint (2015) exploits variation in the number of bidders in a model with additively separable unobserved heterogeneity. That identification strategy does not translate to the more common multiplicative structure examined here.

[8] Carlton and Keating (2015) emphasize the role of transaction costs in welfare analysis when the affected variable is not simply the price level, through the effect on a firm's ability to implement nonlinear pricing.

A novel contribution of this paper is an empirical model of optimal contract duration. To the best of my knowledge, this is the first model to illustrate a general cost of longer contracts, which arises from suboptimal buyer–seller matching over time. The previous literature on contract duration has focused on ex post coordination problems, primarily through costly renegotiation (Masten and Crocker 1985) and relationship-specific investments (Joskow 1987). Recent empirical work on these features (e.g., Decarolis 2014; Bajari, Houghton, and Tadelis 2014) focuses on one-time projects and therefore does not model repeated demand. As discussed above, I abstract away from such ex post incentive problems and focus on "recurrent spot contracting," in the terminology of Williamson (1979). For commodities and simple products, finding the lowest-cost supplier is often more important than whether buyer and seller incentives are properly aligned. I am also able to test for and abstract away from incumbency advantage, which is often a concern in settings with repeated contracts (see, e.g., Greenstein 1993). My work is complementary to models with these features.[9]

A related empirical literature measures switching costs in consumer markets (e.g., Dubé, Hitsch, and Rossi 2010; Handel 2013; Honka 2014; Luco 2019). These studies also use a revealed-preference approach to recover switching costs, using a different identification strategy that is made possible by the economic environment. Conceptually, switching costs in consumer markets can be inferred from posted prices,[10] whereas contract prices for intermediate goods are idiosyncratic to the buyer–seller match. Additionally, the switching costs literature tends to take supply costs as exogenous, whereas variation in supply costs is a key factor in the decision to switch suppliers in my setting. As one might expect, I find that transaction costs in intermediate goods markets are substantially higher than consumer switching costs.[11]

The theoretical and empirical analysis of the costs of market transactions has typically been cast in light of the decision to vertically integrate (for a summary, see Lafontaine and Slade 2007). Through integration, buyers and sellers can avoid the costs of going to the market, in addition to realizing other benefits. Supply contracts provide an intermediate option, lying between arms-length transactions and vertical integration. As noted by Coase (1960), conditions that favor longer contracts are also likely to favor vertical integration. Thus, the mechanisms studied in this paper may also be relevant for the analysis of vertical integration.

[9] The trade-off in this paper between transaction costs and price is closely related to the models of contract duration of Dye (1985) and Gray (1978), who take the stochastic price process as given. An innovation of this paper is to use tools of industrial organization to model primitives of the price process and explore its implications. The contract duration decision is also theoretically linked to a simultaneous bundling problem, where the contract bundles demand over time. In the bundling literature, Zhou (2017) and Palfrey (1983) provide the most closely related analogues. Compared to Zhou (2017) and Palfrey (1983), I allow for intermediate degrees of bundling and introduce transaction costs. I demonstrate that the smaller variance induced by bundling reduces total surplus when there are no transaction costs. Relatedly, Salinger (1995); Bakos and Brynjolfsson (1999); and Cantillon and Pesendorfer (2006) note that bundling affects prices by reducing the variance of average valuations.

[10] See, for example, Dubé, Hitsch, and Rossi (2010) for orange juice and margarine, or Elzinga and Mills (1998) for wholesale cigarettes. The wholesale market in the analysis of Elzinga and Mills (1998) mirrors a consumer market in that pricing, though nonlinear, is uniformly applied.

[11] A key feature of consumer markets is an inability to contract on future prices, leading to models that weigh an "investing" effect versus a "harvesting" effect (Klemperer 1995; Rhodes 2014; Cabral 2016). When buyers and sellers agree on future prices, as in this paper, these effects are competed away.

I. Model

Consider a buyer that demands a good or service for many periods. In the model I introduce, the buyer chooses the duration of the contract, balancing the per period payment to suppliers with the market transaction costs realized at the beginning of each contract. I provide a numerical example to illustrate this key trade-off. A central empirical implication of the model is that it may be applied to recover unobserved transaction costs. The model does not capture every real-world consideration, but its representation of a key contracting trade-off provides the basis for an empirical investigation into the magnitudes of transaction costs.

A. *The Buyer's Problem*

A risk-neutral buyer has inelastic demand for a good or service over many (infinite) future periods. The buyer selects a single seller and can commit to buy from that seller for multiple periods with a contract. The buyer announces the duration of the contract (T) in advance of implementing a market mechanism, which is used to select the seller and determine price. Each time the buyer uses the mechanism, it costs the buyer δ.

The game proceeds in three stages. First, the buyer determines duration T after observing characteristics of the service x, market conditions m, and the mechanism cost. Second, N suppliers decide to participate in the market mechanism after observing (T,x,m). Contract characteristics (T,x) affect the per period supply costs, while market conditions affect entry costs (through outside options). Third, a supplier is selected with a per period stochastic price $P(N,T,x,m)$, where the price distribution may depend on the duration of the contract and the number of sellers.

Let $\bar{P}$ denote the ex ante expected price conditional on (T,x,m), so that $\bar{P}(T,x,m) = \sum_{n=1}^{\mathbb{N}}\left(E\left[P(n,T,x,m)\right] \cdot \Pr(N = n \mid T,x,m)\right)$. The buyer expects market conditions and mechanism costs to remain the same in future periods. With this assumption, we use $\bar{P}(T)$ in the exposition below as shorthand, suppressing (x,m).

The value function for the buyer in period τ who has not yet chosen a seller can be expressed as

$$V(\tau) = \min_T \delta + \sum_{k=1}^{T} \beta^{k-1}\bar{P}(T) + \beta^T V(\tau + T). \tag{1}$$

After incurring the cost to determine the seller and the price (δ), the buyer pays $\bar{P}(T)$ for T periods and returns to the decision problem in period $\tau + T$. The buyer discounts future periods at rate β.

For an optimal T, it must be that, for any other duration S:

$$\delta + \sum_{l=1}^{T} \beta^{T-1}\bar{P}(T) + \beta^T V(\tau + T) \le \delta + \sum_{l=1}^{S} \beta^{S-1}\bar{P}(S) + \beta^S V(\tau + S). \tag{2}$$

We can expand each side of the equation by iterating forward to period $\tau + T \cdot S$. As the buyer expects market conditions to persist, the problem is stationary. If T is optimal in period τ, the buyer expects T to be optimal at the expiration of a contract in a future period, e.g., in period $\tau + T$. Plugging in a sequence of contracts of duration T and S, we obtain

$$\sum_{l=1}^{S}\beta^{T(l-1)}\left(\delta + \sum_{k=1}^{T}\beta^{k-1}\bar{P}(T)\right) + \beta^{T\cdot S}V(\tau + T\cdot S) \leq \sum_{l=1}^{T}\beta^{S(l-1)}\left(\delta + \sum_{k=1}^{S}\beta^{k-1}\bar{P}(S)\right) + \beta^{T\cdot S}V(\tau + T\cdot S). \tag{3}$$

That is, the buyer may pay a per period price of $\bar{P}(T)$ while running the market mechanism S times in $S \cdot T$ periods, or a per period price of $\bar{P}(S)$ while running the mechanism T times over the same horizon.

Rearranging,[12] we obtain the optimality condition

$$\bar{P}(T) - \bar{P}(S) \leq \frac{\delta}{\sum_{k=1}^{S}\beta^{k-1}} - \frac{\delta}{\sum_{k=1}^{T}\beta^{k-1}}. \tag{4}$$

This formulation has straightforward interpretation. The left-hand side is the per period savings by choosing contract S instead of T. The right-hand side is the increase in amortized transaction costs from choosing S instead of T. Thus, at the optimal contract, potential savings in the per period price by selecting a different (shorter) duration are less than increased transaction costs from using the market mechanism more frequently.

Given realizations for contract and market characteristics x and m, the optimal duration, T^*, is therefore given by

$$T^* = \arg\min_{T\in\mathcal{T}} \bar{P}(T,x,m) + \frac{\delta}{\sum_{k=1}^{T}\beta^{k-1}}, \tag{5}$$

where $\mathcal{T}$ is the set of allowable durations. Intuitively, this expression shows that the buyer's objective is to minimize the sum of the per period supply price and amortized transaction costs.

The optimality condition generates two fundamental results, which we express as our first propositions.

PROPOSITION 1: *If the optimal contract is not the maximum allowable duration* (*i.e.*, *an interior solution exits*), *then the expected per period price is increasing with the duration of the contract.*

PROPOSITION 2: *If an interior solution exits*, *then the optimal duration is increasing with transaction costs.*

PROOF:

See Appendix A. ∎

[12] The substitutions $\delta = \sum_{k=1}^{T}\beta^{k-1}\left(\delta/\sum_{k=1}^{T}\beta^{k-1}\right)$ on the left-hand side and $\delta = \sum_{k=1}^{S}\beta^{k-1}\left(\delta/\sum_{k=1}^{S}\beta^{k-1}\right)$ on the right-hand side allow us to factor out the common aggregate discount factor $\sum_{k=1}^{TS}\beta^{k-1} = \sum_{l=1}^{S}\beta^{T(l-1)}\sum_{k=1}^{T}\beta^{k-1} = \sum_{l=1}^{T}\beta^{S(l-1)}\sum_{k=1}^{S}\beta^{k-1}$ and simplify.

The second proposition is intuitive, and it helps motivate the empirical approach of using variation in contract duration to recover transaction costs. The first proposition is a direct result of having ex ante costs for the market mechanism. The buyer can always reduce these (amortized) costs by choosing a longer contract. Therefore, if the buyer chooses something other than the maximum duration, it must be that the buyer expects the marginal increase in the per period price to offset the decline in transaction costs. As illustrated below, the per period price will be increasing when suppliers have idiosyncratic variation in supply costs. This variation causes the low-cost supplier changes over time and provides a benefit of shorter contracts.

B. *Illustrative Example*

To illustrate the key trade-off of this model and its implications, consider a stylized example. Suppose there are N symmetric suppliers in the market, and the set of suppliers stays the same in every period. The buyer can only issue single-period or two-period contracts. Under these conditions, the buyer only has to consider the effects of his decision over the next two periods. Thus, the analysis of the infinite-horizon problem in this example is equivalent to that of a two-period problem, and I describe it as such for clarity.

Suppliers are risk-neutral and participate in an auction to win the contract. Every supplier participates in the auction (entry is exogenous). Thus, the mechanism is efficient. The per period cost to each supplier is the random variable c. The distribution of c is stable across periods, but the realizations for each supplier may vary over time. When the buyer issues single-period contracts, the per period cost of the winning supplier is $c_{1:N}$, which is the minimum of N draws of c. When the buyer issues a two-period contract, the average per period costs for each supplier is the average of two draws, $\tilde{c} = (1/2)\left(c^{(1)} + c^{(2)}\right)$, and the cost to the winning supplier is $\tilde{c}_{1:N}$. This brings us to a key feature about costs in the model.

REMARK 1: *By changing the duration of the contract, the buyer changes the effective per period cost structure faced by suppliers.*

As long as the per period costs c are not perfectly correlated across periods, $\tilde{c} \neq c$ and $\text{var}(\tilde{c}) < \text{var}(c)$. As suppliers' bids will reflect the average cost over the duration of the contract, the distribution of per period costs changes with contract duration. When the distribution of supply costs is stable over time, this serves to reduce the variance of cost draws. The cost of a longer contract is that the low-cost supplier may not be selected in each period. In the absence of transaction costs, short-term contracts would be optimal.

Risk-neutrality and symmetry generate the standard auction result that the expected winning bid is equal to the second-order statistic from the cost draws. Thus, the buyer-optimal contract solves

$$\min\Big\{\underbrace{2E[c_{2:N}] + 2\delta}_{\text{short-term}}, \underbrace{2E[\tilde{c}_{2:N}] + \delta}_{\text{long-term}}\Big\}. \tag{6}$$

The buyer will pick the long-term contract if the increase in expected supply costs is less than the reduction in (amortized) transaction costs, i.e., if

$$E[\tilde{c}_{2:N}] - E[c_{2:N}] < \frac{\delta}{2}. \tag{7}$$

This condition mirrors the optimality condition in equation (4).

This simple example illustrates a second key feature of the model.

REMARK 2: *The intensity of supply-side competition, in terms of the number of participating suppliers, affects the optimal contract by changing the per period cost structure.*

Variation in N affects the left-hand side of (7), changing the marginal effect of a longer contract on the per period price. This marginal effect is nonmonotonic in N, so an increase in the number of suppliers has an ambiguous effect on equilibrium contracts. Therefore, the optimal duration may be decreasing, increasing, or U-shaped with N. I describe the intuition for this result below along with the numerical example.

To illustrate the above features, I present a numerical example in which the per period costs are drawn independently over time from a beta distribution with shape parameters $(0.5, 0.5)$. Recall that the beta distribution has support $[0,1]$. With shape parameters $(1,1)$ it is equivalent to a uniform distribution, and as the shape parameters approach zero it approaches a Bernoulli distribution.

Figure 1 illustrates how expected buyer costs vary with contract duration and the degree of competition. Panel A plots the expected supply price for one-period contracts and a two-period contract. For $N = 3$, the expected prices are the same, and for $N > 3$ the single-period contracts always have a lower expected price. The blue line in panel B plots the difference between these two lines. This difference is equivalent to the left-hand side of equation (7) and is nonmonotonic in the number of suppliers. The dashed line indicates a transaction cost of 0.20, which is amortized by two periods. When the blue line falls above this dashed line, the increase in the expected supply price exceeds the savings in transaction costs, and one-period contracts are optimal. Panel C plots the U-shaped buyer-optimal duration as a function of N. Short-term contracts are optimal for moderate level of competition; in this case, when $N \in \{6, \ldots, 21\}$.

This stylized example conveys a general insight from the model. For low levels of competition, the benefit of switching suppliers is low, and long-term contracts are preferred. At moderate levels of competition, there is an increased benefit of switching among suppliers more frequently. When competition is intense, the expected costs of both long-term and short-term contracts approach the lower bound of costs, and therefore long-term contracts, which minimize transaction costs, are optimal.

Thus, even in a stylized example, the directional predictions of the model are empirical, depending on particular contract or market conditions. This finding further motivates the use of a structural approach to assess the impacts of transaction costs. The model also generates a set of predictions related to the stochastic properties of per period costs to suppliers. Higher autocorrelation in supply costs will lead

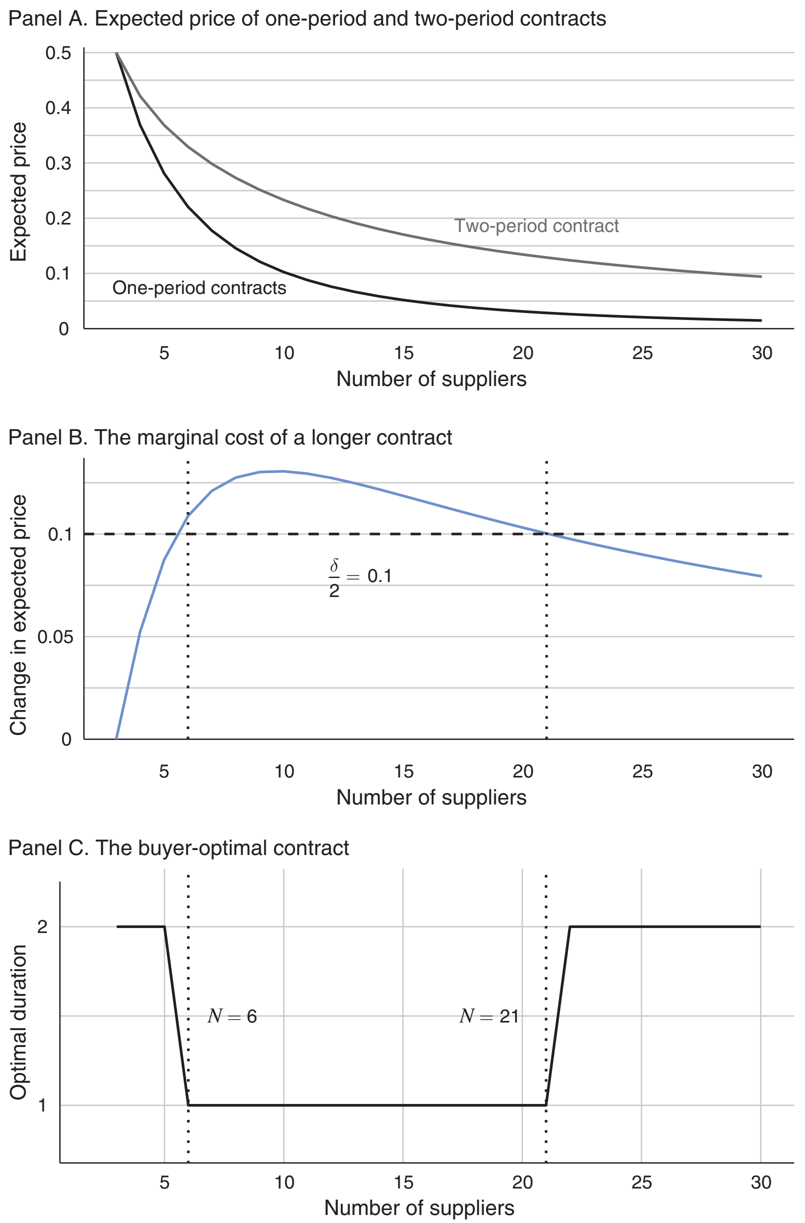

Figure 1. Competition, Costs, and Contract Duration: A Numerical Example

Notes: Panel A plots the expected per-period costs for separate one-period contracts and a bundled two-period contract, as a function of the number of bids. The blue line in panel B is the difference between the two, which is the expected price increase to the buyer. The dashed line in panel B reflects a transaction cost of 0.2 amortized over two periods, which is the amount saved by issuing a two-period bundled contract. For values of N where the blue line is above the dashed line ($N \in \{6, \ldots, 21\}$), short-term contracts are optimal, as the increase in supply costs from the long-term contract is greater than the savings in transaction costs. Panel C plots the buyer-optimal contract duration.

to longer contracts, while higher variance in per period supply costs can have effects in either direction. As the remainder of the paper focuses on the structural approach, I discuss these in more detail in Appendix A.

C. *Identification of Transaction Costs via Revealed Preference*

I now discuss the empirical strategy for recovering unobserved transaction costs. The optimality condition for the buyer generates a contract-specific implied value for δ that rationalizes the observed duration. Thus, by revealed preference, we can recover δ for each contract.

If a contract T is optimal, then, relative to contract S, we have the following optimality condition, which is a re-expression of equation (4):

$$\left(\sum_{k=1}^{T}\beta^{k-1}\right)\left(\sum_{k=1}^{S}\beta^{k-1}\right)\left(\bar{P}(T) - \bar{P}(S)\right) \leq \sum_{k=1}^{T}\beta^{k-1}\delta - \sum_{k=1}^{S}\beta^{k-1}\delta. \tag{8}$$

By comparing the optimal contract to one that is one period shorter ($S = T-1$), we obtain the inequality

$$\delta \geq \frac{1}{\beta^{T-1}}\left(\sum_{k=1}^{T}\beta^{k-1}\right)\left(\sum_{k=1}^{T-1}\beta^{k-1}\right)\left(\bar{P}(T) - \bar{P}(T-1)\right). \tag{9}$$

Likewise, a comparison of of T to $S = T+1$ obtains

$$\delta \leq \frac{1}{\beta^{T}}\left(\sum_{k=1}^{T+1}\beta^{k-1}\right)\left(\sum_{k=1}^{T}\beta^{k-1}\right)\left(\bar{P}(T+1) - \bar{P}(T)\right). \tag{10}$$

In principle, one could generate an inequality for every element in the duration choice set $\mathcal{T}$, excluding the chosen contract. The minimum upper bound and the maximum lower bound are sufficient bounds on δ.

The right-hand sides of the inequalities depend only on the discount rate and the expected per period price function. The next result follows immediately:

PROPOSITION 3: *When β and $\bar{P}(T)$ are known, bounds for contract-specific realizations of δ are identified.*

The tightness of the bounds depends on the slope of the function $\bar{P}(T)$ and the discount rate β. Effectively, these parameters are governed by the length of each period. Thus far, we have used a discrete formulation where one unit separates each period. If the length of this unit became infinitesimally short, then T is chosen from a continuous set and we obtain point identification of δ. We present this as a corollary.

COROLLARY 1: *When T is continuous (i.e., the duration of each period approaches zero), then δ is point identified for each contract.*

PROOF:

See Appendix A. ∎

Even in the discrete case, the full distribution of δ can be identified from additional assumptions on the relationship between δ and x or m. This distribution can be used as a prior over the bounds. Recall that $\bar{P}(T)$ is shorthand for $\bar{P}(T,x,m)$.

PROPOSITION 4: *Assume that there exists a special covariate w that is an element of x or m. This covariate meets the following two conditions:* (i) *δ and w are independent and* (ii) *$\bar{P}(T,x,m)$ varies continuously with w. Local variation in w provides local identification of the distribution of δ. When sufficient variation in $\bar{P}(T,x,m)$ is generated by variation in w, the distribution of δ is identified.*

PROOF:

Under the above conditions, the bounds (9) and (10) vary continuously with w. Therefore, the cumulative distribution function of δ is identified. ■

The model provides a straightforward way to recover unobserved contract-specific transaction costs. A key input to this procedure is the duration-dependent function $\bar{P}(T,x,m)$. Intuitively, observed variation in contract duration will be necessary to estimate this function when not known a priori. I discuss the empirical strategy to estimate this object in Section III.

D. *Discussion*

Setting.—The model captures a setting in which the buyer chooses fixed-price, fixed-duration contracts. This type of contract is quite common. They constitute the vast majority of government service contracts, and, anecdotally, are used quite frequently in the private sector. Why might a buyer choose this type of contract instead of allowing for options to renew or a less formal structure? Some possible reasons are (i) to maximize the number of suppliers that bid on the next contract and (ii) to protect against favoritism by the buyer's agent (at the expense of the buyer) through increased transparency. I take the contract structure as given; an analysis of why fixed-duration contracts are prevalent lies outside the scope of this paper.

The model abstracts away from ex post transaction costs that arise from the principal–agent problem and incomplete contracts. These considerations have been studied in detail by the transaction costs literature. The focus of this paper is on illustrating a new fundamental mechanism, which may be important even when ex post transaction costs are negligible or when complete contracts are possible. In some settings, a richer model that accounts for both ex ante and ex post transaction costs may be appropriate. Ex post transaction costs tend to suggest longer contracts, as longer contracts can offer a greater return and better align incentives. By contrast, the model of this paper illuminates that time-varying supply costs make longer contracts more costly, providing a reason why shorter contracts may be preferred. Even with ex post transaction costs, the per period price should be increasing in duration at the equilibrium. If not, the buyer will opt for a longer contract that reduces the supply price and (both ex ante and ex post) transaction costs.

One simplifying assumption of the model is that the market transaction costs are fixed. In some settings, we may expect that these costs vary with the duration of the contract. Longer contracts could require more market research by the buyer or additional costs of drawing up the contract. This is not a first-order consideration for simple goods and services. The language of the contracts in the empirical application is not duration-dependent, likely reflecting the standardized nature of the service and stable market conditions.

Another simplification is that suppliers have perfect foresight about future costs. A more general setup with imperfect information shares the same qualitative features of this model. For example, consider the case in which a supplier has no information about future costs. Then the bid for a longer-duration contract will reflect an average between the current-period realization and the expected mean of future costs, which is the same across suppliers. Therefore, longer contracts shrink the variance of prices across suppliers. In this scenario, there is an additional ex post inefficiency arising from imperfect information, though fixed-price contracts eliminate this risk to the buyer.

Efficiency.—The duration of a fixed-price, fixed-duration contract is typically determined by the buyer. Thus, the analysis of this paper focuses on buyer-optimal contracts, though the intuition translates to efficient contracts as well. The buyer is concerned with the supply price, which shares similar stochastic properties to supply costs in most settings. For example, in the illustrative example above, expected supply cost and expected price are the first and second order statistics from the cost distribution, and they generate similar directional predictions. In the efficient case, supply-side frictions such as entry costs should also be incorporated for when analyzing welfare.

Though the qualitative features of the buyer-optimal and the efficient contract are similar, they are not identical. In online Appendix C, I provide an analysis of these two outcomes, as well as the seller-optimal contract. Contracts that are determined by market participants (buyers and sellers) may be too long or too short, resulting in wasteful social costs. Counterintuitively, these extra costs may increase as a market becomes more competitive. Therefore, when ex ante costs are taken into account, highly competitive markets may be of more concern for regulators than those that are more concentrated.[13]

II. Empirical Application: Data and Reduced-Form Analysis

A. *Data*

To estimate the costs of market transactions, I construct a dataset of 1,046 competitive contracts for building cleaning services for the United States federal government. This market provides a relatively clean case study to analyze the duration decision and estimate costs. For many commodity products and services, the buyer (a government agency) is compelled to run a sealed-bid auction for a contract of a predetermined duration at the expiration of the previous contract. Thus, for many

[13] This result occurs because market participants care about price rather than cost, and the price responds more quickly to a change in contract duration when the number of bidders is large. If we think of expected price as the expected second-order statistic, and the cost as the first-order statistic, then we have some intuition for why this could be true. The second-order statistic responds more strongly to a change in variance (or mean) than the first-order statistic when the number of draws is large and the cost distribution is bounded from below. The buyer (or seller) internalizes the duration's effect on the second-order statistic rather than its effect on the first-order statistic.

federal procurement contracts, the empirical setting is aligned with the model of the previous section. The sample period is October 2003 to May 2017.

A general empirical challenge is that procured goods and services may have heterogeneity that is multidimensional and difficult to quantify. Thus, I focus on commodity-like goods and services with standard cost structures. Indeed, products of this sort are numerous in procurement and make up a significant portion of all transactions.[14] From the set of commodity-like products, building cleaning services were chosen because they are numerous, cost factors are easily quantified, and there is a lot of variation in contract duration. Finally, demand is inelastic, as there were no significant substitutes. The market for such services is sizable; the federal government spent $1.2 billion annually on such services. In addition, the standardized nature of the work, along with the fact that the contracts are rarely terminated, mitigates concerns about the impact of relationship-specific investments in this setting.

To the best of the author's knowledge, this is the first dataset on contracts to combine measures price, duration, and competition, which are the key outcomes of the model. To construct this dataset, I combined detailed location, price, and vendor information maintained in the Federal Procurement Data System (FPDS) with contract-specific documents downloaded from the Federal Business Opportunities (FedBizOpps) website. By law, the FPDS keeps public records of all contracts for the U.S. federal government, and its data have been used in recent empirical work (e.g., Kang and Miller 2017; Bhattacharya 2021; Decarolis et al. 2020). I was able to identify 4,119 contracts that appeared in both sources (USAspending.gov 2000–2017; FedBizOpps 2003–2017). The final sample was restricted to competitive contracts in the United States that received more than one bid, had an annual price of less than $1 million, and included square footage in the text of the contract documents, which is a key cost factor. These contracts span different types of facilities and government agencies. For additional details on the construction of the sample, see online Appendix D. Replication files are available from the AEA Data and Code Repository (MacKay 2020).

The FPDS data is subject to measurement error. To correct for this, I collected a subsample of 75 finalized contracts and cross-validated price, duration, and total contract value to what was reported in FPDS. The cross-validation approach allowed me to generate accurate measures for the value of the contract. For details, see online Appendix E.

I matched the contract-specific dataset with auxiliary datasets of (i) government contracting expenditures at the same location in related products and (ii) local labor market conditions. Local labor market conditions include county-level unemployment from the Local Area Unemployment Statistics (Bureau of Labor Statistics 2000–2017) and the number of NAICS code–level establishments in the same three-digit ZIP code from the County Business Patterns data (Census Bureau 2004, 2012).

[14] For context, 97 percent of federal government contracts during the period had an annual price of under $1 million. A counterexample of the ideal setting for this sort of analysis might be a customized, large-scale computer software system for an agency.

B. *Institutional Details*

Competitive contracts are posted publicly and allow open competition from registered vendors.[15] Many of these contracts are posted on the centralized web portal FedBizOpps.gov, from which I collected the data in this analysis. On the website, a prospective supplier can view the contract details, including contract duration and the square footage of the building, requirements for the job, and a list of interested suppliers. From the portal, a supplier submits a bid to the contracting office that includes the total price over the duration of the contract. The contracting office determines the winning supplier primarily based on the lowest price. By law, the contracting office must provide a justification when selecting a supplier that does not have the lowest-price offer.

Importantly, contract duration is determined by the local contracting office and varies from contract to contract, even within an agency. As several industry personnel described to the author, contract duration is a balance between minimizing the administrative costs of re-contracting and realizing lower supply costs from re-competing more frequently. Administrative costs include market research, drafting contract specifications, ensuring compliance with policies and regulations, advertising the opportunity, administering a supplier selection mechanism, and concluding the contract. Ex ante transaction costs and the competitive benefits of shorter contracts are key factors for the duration decision, motivating this market as a case study.

Contracts include specifications for the tasks to be done and their frequencies. For building cleaning, tasks include mopping, vacuuming carpets, picking up debris, dusting, and emptying trash cans. For an example list of specifications, see online Appendix D. Contract documents are extensive, and multiple documents are often posted for each solicitation. The median primary document runs 49 pages.

As mentioned previously, one motive for using this market as a case study is that ex post incentive problems are not a significant concern, based on the nature of the service (Hyytinen, Lundberg, and Toivanen 2018), conversations with contracting officers, and the data. The extensive list of contract specifications combined with observable performance means these contracts are more or less complete; ex post incentives might be of more concern in a different equilibrium with less rich contracts. Service contracts under $1 million annually are rarely canceled, and the use of fixed-price contracts limits the ability of firms to drive up costs, compared to cost-plus contracts.

In particular, building cleaning services have lower rates of terminations, change orders, and additional work than the average federal government service contract.[16] For these contracts, the best estimate of what the buyer pays is the initial stated value

[15] These contracts fall under three categories: Full and Open Competition, Full and Open Competition after the Exclusion of Sources, and Competed Under Simplified Acquisition. Note that 86 percent of the contracts deemed Full and Open Competition after the Exclusion of Sources are listed as a small business set-aside. As 96 percent of the contracts are won by small businesses (as determined by the contracting officer), I ignore this distinction for the purposes of analysis. See Federal Acquisition Regulation (FAR) Part 5.

[16] On a contract-by-contract basis, these outcomes occur at higher rates for building cleaning services. However, after correcting for the fact that building cleaning contracts last three to four times longer than the typical service contract, the comparison is reversed. In other words, other services require, on average, three to four contracts to cover the same period as a typical building cleaning contract.

Table 1—Summary Statistics

	Mean	Min.	p25	Median	p75	Max.
Price (annual, $)	43,870	1,112	7,259	13,180	26,731	976,538
Contract value ($)	190,200	2,914	28,500	50,550	102,000	4,882,692
Duration (years)	4.2	0.4	3.0	5.0	5.0	6.5
Square footage	25,701	145	3,700	7,000	14,500	2,031,842
Price per square foot	2.91	0.16	1.32	2.01	3.14	33.02
Number of bids	6.5	2.0	4.0	5.0	8.0	40.0
Weekly frequency	3.5	0.1	2.0	3.0	5.0	7.0
Number of employees (winner)	61.5	1.0	3.0	14.0	75.0	650.0
Observations	1,046					

Notes: The table displays summary statistics for key variables in the contract data. Included are outcomes (price, duration, and number of bids), as well as cost characteristics such as the number of square feet and the frequency of cleaning. The last variable is the size of the winning firm, in terms of number of employees.

of the contract. I provide empirical support for this in online Appendix E, where I show that the median overrun, defined as the difference between the total payments made and the initial value of a contract, is zero.[17]

C. *Summary Statistics*

Summary statistics for the contracts are displayed in Table 1. Contracts vary in price, duration, and the number of bids. As shown later in this section, much of the variation in price can be captured by the square footage of the building and the weekly cleaning frequency. For the sample, which removes contracts greater than $1 million per year, the mean annual contract price is $43,870 and the median is $13,180. The sample contains 76 contracts with an annual price greater than $100,000.

Overall, the estimation sample compares favorably to the broader set of cleaning contracts in the FPDS. For fiscal years 2004 to 2016, an average of 6,366 cleaning contracts are in effect each year. Of these, 95 percent have an annual value of less than $1 million. Within this subset, the mean annual value is $70,322 and the median is $11,200. Cleaning contracts are larger in value than the average contract. For all contracts with an annual value of less than $1 million, the mean annual value is $35,469 and the median is $4,731. These contracts comprise 97 percent of all contracts in the relevant period, or an average of 1,004,991 each year.

One important source of variation in the analysis is in the number of bids received. The median is 5 bids, and the maximum is 40. Thus, there is a good deal of competition for these contracts. The variation in the number of bids will help to disentangle the effect of private costs from unobserved heterogeneity in the structural analysis.

In the last row, the table provides the number of employees for the winning firms. The winning firms in this dataset are typically small, with a median of 14 employees. Over 25 percent of the winning suppliers have 3 or fewer employees.

[17] The initial entry for total contract value in FPDS is systematically underreported, as shown in online Appendix E. Thus, comparing total payments to the initial entry in FPDS would systematically overstate the degree of cost overruns.

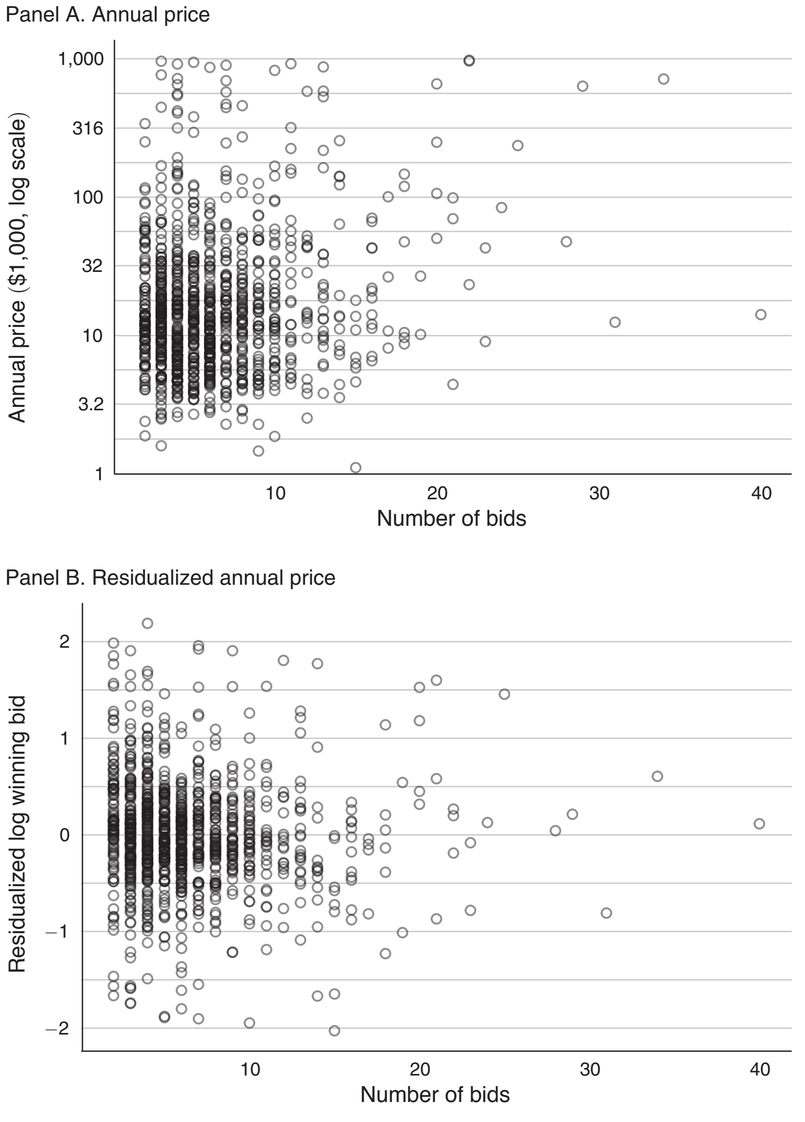

Figure 2. Price versus Number of Bids

Notes: The figure plots the log annual price against the number of bids received for each contract. There is a great deal of variation in the annual price, much of which cannot be explained by observable variables. This is illustrated by the residualized bids in the lower panel. The R^2 of the regression used to construct the residuals, which includes duration, square footage, frequency, baseline unemployment, and fixed effects for facility type, is 0.74. It is notable that some of the highest and lowest prices are realized with few bidders.

Figure 2 plots the logged values of the winning bids on the y-axis against the number of bidders on the x-axis. The second panel displays residualized values for the (log) winning bids. The residuals were constructed from a regression of price on duration, square footage, cleaning frequency, baseline unemployment, and fixed effects for facility type. Even after controlling for observable characteristics, there is large variation in prices for auctions with many bidders. The pattern observed in the figure—large variation in prices with clustering at the median price, rather than the

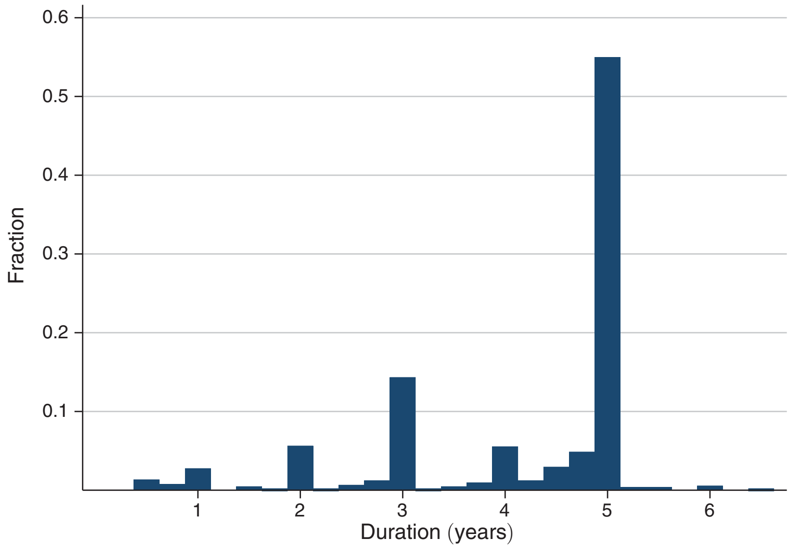

Figure 3. Contract Duration

Notes: The figure displays a histogram of contract duration in three-month bins. Over half of the contracts have a five-year duration, which is the maximum duration (by regulation) without specifically requesting an extension. Contracts are clustered in yearly intervals, though the support in between full years is relatively well covered.

minimum—motivates the assumption of unobserved auction-specific heterogeneity used in the model. Though much of the variation in prices can be explained by observables, there is still residual variation that is inconsistent with an independent private values model; the model with multiplicative common costs fits far better.

The contracts in the dataset have a good deal of variation in duration, ranging from 5 months to 6.5 years, though contracts tend to cluster at yearly increments. Figure 3 provides a histogram of duration in three-month intervals. Empirical variation in contract duration occurs within and across government agencies. Consistent with the model, larger facilities and more frequent cleaning are correlated with shorter contracts, after conditioning on department and facility type.[18] As a first-pass check that transaction costs matter, contracts issued under the government's simplified acquisition protocol, which reduces the ex ante transaction costs, are 15 percent shorter than other contracts, after conditioning on square footage, cleaning frequency, and facility type.

Notably, 53 percent of contracts are for 5 years, which is the typical maximum contract duration imposed by federal budgeting regulations. Longer durations require the contracting officer to request and justify an extension. The observed variation in duration, combined with the five-year cap on contract duration, help motivate the counterfactual analysis of Section V, where I consider the value of the duration decision compared to standard-duration contracts.

The buildings to be cleaned for each contract are categorized into offices (694), research facilities (111), medical facilities (61), service centers (59), visitor centers

[18] For summary statistics by agency, see online Appendix D.

(41), airports (30), technical facilities (19), accommodations (18), and industrial facilities (13). Offices are split into standard offices (424) and field offices (270), which also have an auxiliary building, such as an exercise room, a bunkhouse, or a small warehouse. Online Appendix D provides a breakdown by the issuing agency and by subcategory.

D. *Descriptive Regressions*

To demonstrate the fit between the model and the data and to motivate specific assumptions made in estimation, I present descriptive regressions. Table 2 provides regressions of the log annual price on the number of bids, duration, and controls. The first three columns display the results from ordinary least squares (OLS) regressions. Square footage alone, as reported in the first specification, captures 62 percent of the variation in prices.

To account for endogenous entry, I instrument for the number of bidders using time-series and cross-sectional variation in local labor market conditions, as well as variation in the type of bidders permitted to compete for the contract. The first instrument is the (log) ratio of county-level unemployment relative to a 2004 baseline. This generates a time-varying county-specific unemployment shock. The second instrument is the number of establishments for NAICS code 561720 (corresponding to building cleaning services) in the same three-digit ZIP code.[19] It is plausible that an increase in unemployment or the presence of more firms in the broader geographic area are not driven by unobservable characteristics of these contracts, yet they are likely to generate increased entry.

A third instrument is developed from the federal government practice of "setting aside" certain contracts for firms with particular types of owners. Specialized set-asides include women-owned and veteran-owned small businesses. As we have removed economically disadvantaged set-asides (e.g., for Economically Disadvantaged Women-Owned Small Business) from the sample, it is plausible that the ownership type is uncorrelated with the underlying cost structure of the participating firms. If the cost structure is independent of ownership for these firms, then the type of set-aside is a valid instrument for price (by affecting entry). This instrument is implemented as a binary variable with the value of 1 if the contract is open to all small businesses, i.e., the set-aside does not restrict entry based on characteristics of the owner.

The last two columns report the estimated coefficients from instrument variables regressions. Consistent with endogenous entry, I find a larger negative effect of the number of bidders on price compared to the corresponding OLS specifications. In the structural model of Section IV, I explicitly model entry to account for this endogeneity. The main motivating specification is IV-1, which uses square footage, weekly cleaning frequency, and baseline (2004) unemployment as controls. To capture variation in the types of buildings and cleaning required, IV-1 includes an indicator for "high intensity" cleaning of airports and medical buildings. IV-2 includes indicators for all building types. The inclusion of fixed effects for all types

[19] I add 1 to the raw value to use the logged value in estimation, as a few contracts have zero in the raw value.

Table 2—Descriptive Regressions: ln(Annual Price)

	OLS-1	OLS-2	OLS-3	IV-1	IV-2
ln(square footage)	0.730	0.658	0.658	0.689	0.687
	(0.018)	(0.017)	(0.017)	(0.024)	(0.024)
Number of bids		−0.014	−0.009	−0.053	−0.047
		(0.005)	(0.005)	(0.022)	(0.022)
Duration (years)		0.041	0.032	0.043	0.033
		(0.015)	(0.015)	(0.016)	(0.015)
ln(weekly frequency)		0.459	0.394	0.467	0.407
		(0.039)	(0.038)	(0.041)	(0.040)
ln(2004 unemployment)		0.054	0.037	0.080	0.060
		(0.012)	(0.012)	(0.019)	(0.018)
High-intensity cleaning		0.586		0.559	
		(0.071)		(0.075)	
Building type fixed effects			X		X
Observations	1,046	1,046	1,046	1,046	1,046
R^2	0.62	0.71	0.74	0.69	0.73

Notes: The table displays estimated coefficients from regressions of log annual price on auction characteristics. The variables from specification IV-1 are included in the structural model. These regressions show that square footage, cleaning frequency, and market characteristics explain much of the variation in prices. Once square footage, cleaning frequency, and market characteristics are accounted for, fixed effects for location type add little explanatory power. Specifications IV-1 and IV-2 are two-stage least squares regressions, where the instruments for the number of bids are monthly (log) county-level unemployment relative to 2004, the (log) number of NAICS code 561720 establishments in the same 3-digit ZIP code in 2004, and an indicator for whether the set-aside was for generic small businesses. Standard errors in parentheses.

have low in specifications OLS-3 and IV-2 have a low per-variable impact on R^2 and do not have a substantial effect on the estimated coefficients. Therefore, I omit them from the structural estimation and proceed with the variables used in IV-1.[20]

Though the linear model does not account for the offsetting effects of duration on price and (via profits) on entry, the regressions capture a positive relationship between price and duration. This reduced-form correlation is consistent with Proposition 1, which predicts that a longer duration generates a higher per period price. Thus, the data match a key distinguishing feature of the model of optimal duration. In the structural estimation, I also find a positive and significant direct relationship between duration and price.

In Table 3, I display regressions of the number of bids on auction characteristics and local measures of unemployment. Specification (3) is equivalent to the first-stage regression in IV-1, with an F-statistic of 25.9. All three instruments—the unemployment shock, the presence of existing firms, and a generic set-aside—have the expected positive signs. Though current unemployment is associated with more bids, higher baseline levels, which are used as a control, are associated with fewer bids. I interpret the negative correlation between higher 2004 unemployment and fewer bids as a reflection of local labor market frictions, leading to reduced competition and higher wages.

[20] Each of the three instruments, when used by itself, pushes the coefficient on number of bids more negative and has little impact on the other coefficients. If separate indicators are estimated for medical buildings and airports, the coefficients on the indicators are very similar and the coefficients on the other variables are unchanged.

Table 3—Descriptive Regressions: Number of Bids

	(1)	(2)	(3)	(4)
Duration (years)	0.104	−0.017	−0.002	−0.002
	(0.104)	(0.099)	(0.099)	(0.100)
ln(square footage)	0.760	0.779	0.834	0.825
	(0.111)	(0.106)	(0.106)	(0.112)
ln(weekly frequency)	0.487	−0.081	0.009	0.137
	(0.254)	(0.247)	(0.253)	(0.257)
ln(2004 unemployment)		−0.832	−0.794	−0.793
		(0.239)	(0.238)	(0.238)
ln(unemployment)		1.415	1.420	1.356
		(0.232)	(0.231)	(0.231)
ln(number of firms in ZIP3)		0.241	0.257	0.276
		(0.148)	(0.148)	(0.147)
Generic set-aside			1.134	0.987
			(0.350)	(0.361)
High-intensity cleaning			−0.294	
			(0.475)	
Building type fixed effects				X
Observations	1,046	1,046	1,046	1,046
R^2	0.06	0.16	0.17	0.19
F-statistic	22.2	32.0	25.9	14.7

Notes: The table displays estimated coefficients from regressions of the number of bids on auction characteristics and local labor-market variables. Specification (3) is equivalent to the first-stage regression of IV-1 in Table 2. Specification (4) includes fixed effects for each building type. Standard errors in parentheses.

III. Empirical Implementation

A. *Supplier Participation, Bidding, and Equilibrium*

In the general model of Section I, the buyer decides on the contract T with knowledge of $\bar{P}(T,x,m)$, the expected price conditional on contract and market characteristics. The buyer's expectation is taken over the number of bidders N and the cost realizations for these bidders.

For the empirical analysis, we separately model the participation decision that determines $\Pr(N = n \mid T,x,m)$ and the market mechanism that determines $E[P(n,T,x,m) \mid N = n]$. The model thus proceeds in three stages: the first stage reflects the buyer's problem, the second stage is the participation decision of suppliers, and the third stage is the market mechanism that determines the chosen supplier and the price.

1st Stage: Duration Decision.—The buyer observes (x,m,δ) and sets T to minimize the expected per period price plus the amortized transaction cost. The buyer's objective function is:

$$\min_{T \in \mathcal{T}} \sum_{n=1}^{\mathbb{N}} \Big(E\big[P(n,T,x,m)\big] \cdot \Pr(N = n \mid T,x,m) \Big) + \frac{\delta}{\sum_{k=1}^{T} \beta^{k-1}}. \tag{11}$$

2nd Stage: Participation.—Potential entrants observe (T, x, m), entry costs $k(m) \cdot \varepsilon$, and contract cost-shifters $h(x)$. These costs are common across bidders. $k(m)$ and $h(x)$ are observed by all parties (including the econometrician), but the entry shock ε is only observed by suppliers.

Bidders enter if expected profits exceed entry costs. Profits conditional on participation depend on total contract costs $C_i \cdot U \cdot h(x)$. Thus, we make the usual assumption that cost components are multiplicative. Bidders do not observe the private cost C_i or the common cost U until after they decide to participate. In this context, U captures unobserved auction-specific heterogeneity.

Let π_n denote proportional profits for the nth marginal entrant, which depends on C_i, the distribution of C, and the number of participating suppliers. Total profits are $\pi_n \cdot U \cdot h(x)$. The entry condition is given by

$$E\left[\pi_n \cdot U \cdot h(x) \,|\, n, T\right] - k(m) \cdot \varepsilon > 0 \quad \Leftrightarrow \quad N \geq n. \tag{12}$$

3rd Stage: Bidding.—Participating suppliers realize their private (proportional) cost C_i and the common cost U. They then engage in a supplier selection mechanism. For the empirical application, we model the mechanism as a first-price auction, though the model can be generalized to other structures.

We assume bidders are risk neutral. Therefore, in equilibrium, each bidder submits a bid of $b_i \cdot U \cdot h(x)$, where b_i represents the proportional bid for bidder i. The lowest proportional bid, B, is the winning bid.

Equilibrium is characterized by the buyer choosing duration to minimize expected buyer costs, potential suppliers entering if expected profits exceed entry costs, and participating suppliers bidding optimally in the market mechanism. The model is summarized in Figure 4. Relative to a standard auction model with entry, the model also allows for a strategic decision by the buyer (duration).

B. *Identification of Participation and Bidding*

As shown in Section I, contract-specific transaction costs are identified conditional on $\bar{P}(T, x, m)$, using the optimality condition of the buyer. The function $\bar{P}(T, x, m)$, which captures the participation and bidding game, is identified separately. It is identified even if T is not set optimally, or if T is chosen from a restricted set (e.g., capped a maximum duration).

Identification of Supply Price.—The econometrician observes the transaction price $P = B \cdot U \cdot h(x)$ as well as (N, T, x, m). The cost shocks U, ε, and C are unobserved by the buyer and the econometrician, but their distributions are common knowledge. To achieve nonparametric identification, we make restrictions on the distributions of unobservables. Assume

(i) *Independence of Unobservables:* C_i, U, and ε are independent conditional on (N, T, x, m).

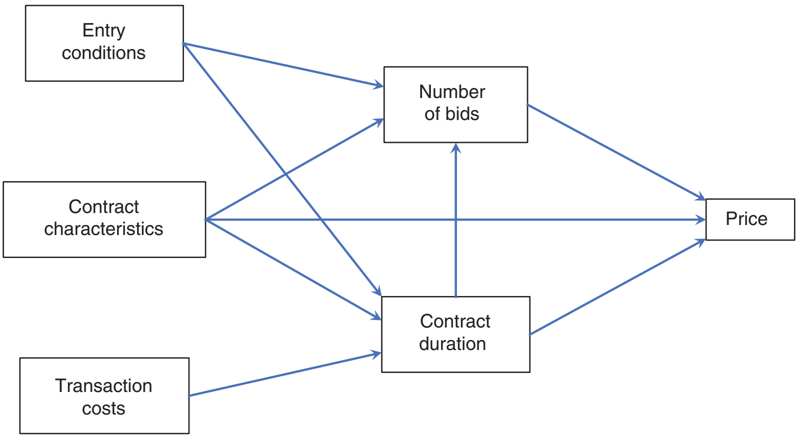

FIGURE 4. SUMMARY OF MODEL

Notes: The figure summarizes the causal assumptions embedded in the empirical model. The three sets of variables on the left: entry conditions, contract characteristics, and transaction costs, are taken as given. Price, number of bids, and contract duration are jointly determined in the model. Arrows indicate the direction of causality.

(ii) *Mean Independence of Common Shocks:* $E[\varepsilon|T,x,m] = E[\varepsilon]$ and $E[U|T,x,m] = E[U]$. $E[U]$ is normalized to 1.

(iii) $h(\cdot)$ and $k(\cdot)$ are continuous, and the range of $h(\cdot)$ or $k(\cdot)$ has broad support. $h(x_0)$ and $k(m_0)$ are normalized to 1 for specific values of x and m.

Importantly, we assume that U is independent of T. In the timing of the model, U is realized after T is chosen; T depends endogenously on (x,m,δ) but not U. Under these assumptions, it is straightforward to obtain identification of the expected proportional bid (B), the cost-shifter functions, and relative profits.

PROPOSITION 5: *When (P,N,T,x,m) is observed, the following components of the model are identified:*

1. $E[B|N,T,x,m]$;
2. $h(x)$ *and* $k(m)$;
3. *relative profits for* N *and* N' *participants:* $\frac{E[\pi_N|N,T]}{E[\pi_{N'}|N',T]}$;
4. *relative profits for* T *and* T' *with* N *participants:* $\frac{E[\pi_N|N,T]}{E[\pi_N|N,T']}$.

PROOF:
See Appendix B. ∎

The first two components are sufficient to identify $\bar{P}(T,x,m)$ and transaction costs, without using the auction structure for the third stage of the game. Thus, even when the underlying selection mechanism is unknown, the model can be useful for counterfactual analysis of the impact of duration and transaction costs.

Identification of Profits and the Joint Distribution of Costs.—With additional structure on the final mechanism, seller surplus can be identified. This can be useful in analyzing efficiency and for identification of the full joint distribution of costs. To this end, in addition to (i)–(iii), we further assume:

(iv) Bidders are symmetric: $C_i \sim F_i$, with $F_i = F$ for all i.

(v) F is continuous with positive support. $U \sim G$, where G has positive support.

(vi) Auctions with sequential values of $N \in \{\underline{N}, \ldots, \bar{N}\}$ are observed, with $\underline{N} < \bar{N}$.

PROPOSITION 6: *When the supplier selection mechanism is an auction with symmetric bidders, seller surplus is identified.*

PROOF:

See Appendix B. ∎

Variation in N, combined with identification of relative profits, allows for identification of seller surplus in the auction model.

Once seller surplus (or expected profit) is identified, the distribution of ε is identified from equation (12), using variation in h or k. Further, we can pin down properties of the private cost distribution.

PROPOSITION 7: *The first $(\bar{N} - \underline{N} + 2)$ expected order statistics of $\bar{N}$ draws from F are identified, providing $(\bar{N} - \underline{N} + 2)$ restrictions on the private cost distribution.*

PROOF:

See Appendix B. ∎

Intuitively, exogenous variation in N shifts the private cost component of the winning bid, but not the common costs. Restrictions on the order statistics of F have additional power in that they may reject many classes of flexible distributions with $(\bar{N} - \underline{N} + 2)$ parameters. Because the full set of expected order statistics approximates the quantile function when $\bar{N}$ is large, exact nonparametric identification of F is obtained when $\underline{N} = 2$ and $\bar{N} \to \infty$.

COROLLARY 2: *The distribution of unobserved heterogeneity is obtained after F is identified.*

PROOF:

By independence, we can use the characteristic function transform to write $\varphi_{\ln W_N}(z) = \varphi_{\ln B_N}(z) \cdot \varphi_{\ln U}$, where $W_N = B_N \cdot U$ is the observed winning bid scaled by the observables and B_N is the distribution of the proportional winning bid conditional on N. We can construct this function for two different values of N. Once the characteristic function of F is obtained, either by exact identification $\left(\bar{N} \rightarrow \infty\right)$ or by flexible estimation methods, G is pinned down. ∎

Discussion of Identification Assumptions.—In my empirical setting, it is important to account for unobserved auction-specific heterogeneity. The baseline set of assumptions map to the setting of Krasnokutskaya (2011), while allowing for endogenous entry. Symmetry is a typical assumption in auction models of unobserved heterogeneity (see also, e.g., Aradillas-López, Gandhi, and Quint 2013). To relax symmetry, one could consider alternative restrictions to pin down costs and the joint distribution of outcomes. One alternative would be to employ supplementary data on profits for one (N, T) pair. This would identify the expected profit function, which could then be used to identify the joint distribution of costs.[21]

Another common assumption is that bidders realize some auction-specific costs after deciding to participate. This assumption may be a reasonable approximation in this setting because observables explain roughly 70 percent of the variation in prices. One could relax this assumption by allowing bidders to select into the auction based on observing U beforehand. In this case, it is straightforward to extend the identification results. Additional steps would be required if potential bidders observed their private cost draw before deciding to participate.

These assumptions are sufficient for the nonparametric identification of the auction model. In effect, the entry cost shifters m serve as instruments for the number of bidders. Exogenous variation in N is then used to separately identify private costs from unobserved heterogeneity. With no instruments, these distributions can still be separately identified. Therefore, with unobserved heterogeneity and conditionally independent private values (i.e., the setting of Krasnokutskaya 2011), the model is identified with data on only the winning bid and variation in the number of bidders. I provide this alternative identification approach in Appendix B.

Under maintained assumptions, the buyer makes the duration decision based on observable characteristics of the contract (e.g., square footage and cleaning frequency) and the market (e.g., unemployment and local number of firms), as well as transaction costs. Thus, after conditioning on contract and market characteristics, variation in contract duration is driven by unobserved transaction costs. We can then estimate the price schedule by observing contracts with similar characteristics that vary in duration. For example, if we observe two contracts for similar office buildings at different agencies, we attribute a difference in duration to different transaction costs between the agencies, rather than different costs of cleaning the facilities.

C. *Parameterizations*

For the empirical application, I estimate the model of Section IIIB, where bidders are symmetric and participate in a first-price auction. I employ a parametric

[21] I test for the presence of asymmetry in Section IVD. The results are consistent with symmetric bidders.

approach for parsimony. The nonparametric identification results provided earlier, along with robustness checks, suggest that first-order features of the estimated distributions are not entirely driven by functional form. In this application, there is an added complication of estimating a duration-dependent distribution of private costs, which would increase the number of parameters needed for any nonparametric approach.

The parameterizations are given in Table 4. A central consideration of this paper is that the distribution of the average per period private cost, C_i, changes with the duration of the contract. One approach to estimation would be to estimate a microfounded model where the per period cost shocks are governed by an autocorrelation parameter. Instead, I estimate the average per period cost distribution as a primitive, allowing the mean of the average per period cost and the variance to vary with T. As I am not taking a stand on the underlying cost process, I estimate a "reduced-form" primitive for the cost distribution. By picking an appropriately flexible distribution, this approach may better approximate a wider range of per period distributional families.[22]

For private costs, the Weibull distribution is chosen for tractability and flexibility, as it allows the estimated probability density functions to be either convex or concave. It is governed by the mean parameter $\mu(T) = \mu_0 + \mu_1 T$ and the shape parameter $\alpha(T) = \alpha_0 + \alpha_1 T$. I allow the parameters of the private cost distribution to vary linearly with duration to capture the first-order effects of interest in this model. A finding of $\alpha_1 > 0$ is consistent with autocorrelation in cost shocks, which results in reduced variance in average per period costs over longer contracts. For the distribution of unobserved heterogeneity, the log-normal distribution was chosen because it best fit the model out of several choices.[23]

In the first step, I estimate the parameters for participation and bidding using maximum likelihood. In this step, I do not require that the choice of duration is optimal. The fact that observed duration may be capped or otherwise chosen from a restricted set does not affect the parameter estimates. Entry cost shocks are parameterized as increasing linearly with the duration of the contract.[24] Note that square footage and weekly frequency can affect the entry decision by both increasing supply costs (price) and affecting entry costs.

One of the challenges in the estimation of auction models arises from the computational burden of inverting the bid function. I employ a simple innovation—a change of variables—which circumvents this step and greatly speeds up estimation in the presence of unobserved heterogeneity. This innovation and details of the likelihood function are in online Appendix G.

In a second step, I use the estimated parameters in the buyer's objective function (equation (11)) to generate nonparametric bounds on δ for each contract, following the steps in Section IC. In my data, contracts are either set to the nearest monthly or nearest yearly increment, providing a set of tight and loose bounds, respectively.

[22] For an example microfounded model, see online Appendix F.

[23] Other estimated distributions of unobserved heterogeneity were the gamma distribution and the Weibull distribution. Both have the desirable properties of support on $(0, \infty)$ and can be normalized to have a mean of 1.

[24] This has the interpretation that entry costs are borne annually and could reflect the opportunity costs of other contracts. Allowing a free parameter on the entry costs in estimation generates a coefficient close to one.

Table 4—Empirical Parameterizations

Cost component	Notation	Parameterization
Private costs	C_i	$\sim$ *Weibull*$(\mu_0 + \mu_1 T, \alpha_0 + \alpha_1 T)$
Unobserved heterogeneity	U	$\sim \ln\mathcal{N}(-\sigma_U^2/2, \sigma_U^2)$
Entry shock	ε	$\sim \ln\mathcal{N}(\mu_\varepsilon, \sigma_\varepsilon^2)$
Observed heterogeneity	$h(x)$	$= square\_footage^{\gamma_1} \times weekly\_frequency^{\gamma_2} \times 2004\_unemployment^{\gamma_3} \times \gamma_4^{\mathbf{1}[high\text{-}intensity\_cleaning]}$
Entry costs	$k(m)$	$= T \times square\_footage^{\kappa_1} \times weekly\_frequency^{\kappa_2} \times unemployment\_shock^{\kappa_3} \times establishments^{\kappa_4} \times \kappa_5^{\mathbf{1}[generic\_set\text{-}aside]}$

That is, if I observe a 16-month contract, I assume it was preferred to 15-month and 17-month contracts, whereas I assume that a 24-month contract was preferred to other yearly increments (12-month and 36-month contracts). I estimate these transaction costs with an annual discount rate of $\beta = 0.97$. A smaller value of β would imply higher transaction costs. This is because the β-dependent components of the bounds given by equations (9) and (10) are decreasing in β, for $\beta < 1$. Intuitively, a smaller β corresponds to a smaller marginal benefit for a longer contract. This, in turn, implies that a greater transaction cost is needed to rationalize the chosen duration.

Finally, to construct expected market transaction costs and conduct counterfactuals, I apply a uniform prior for the density between the distribution-free bounds.[25] Using the prior, I construct point estimates by taking the expectation. In practice, many of the contracts in my data face a cap on maximum duration of five years, due to federal regulation. For contracts affected by the cap, only a lower bound for δ can be obtained without additional assumptions. I make the assumption that the chosen duration at five years is optimal. This generates a relatively conservative upper bound on transaction costs. Many of the optimal contracts under a higher cap would likely be longer, implying a larger upper bound and larger point estimates for the transaction costs.

IV. Results

Estimation of the structural model proceeds in three steps. First, I use a parametric maximum likelihood to perform joint estimation of entry and bidding. Second, using the duration decision of the buyer and estimated parameters from the first step, I construct distribution-free bounds for transaction costs. Third, I construct estimates of transaction costs by applying a prior over the bounds. These estimates are inputs to the policy counterfactuals presented in Section V. The estimation steps are described in more detail in Section III.

[25] The uniform prior is appealing for its transparency and also because the observed duration is optimal at the mean transaction cost when buyers can issue contracts in monthly increments. If a left triangular prior were used instead, the optimal monthly-increment contract would be shorter than the observed value for contracts observed in yearly increments.

A. *Estimated Supply Costs and Entry Costs*

Table 5 displays the parameter estimates from the first step. Square footage, weekly frequency, and 2004 unemployment are scaled by the mean, so that the estimate of μ_0 is interpreted as the mean annual private cost draw for a zero-duration contract at a typical location. The mean annual private cost is \$18,521 and increases by 2.9 percent per contract year (μ_1/μ_0).[26] Thus, I find that the data are consistent with the equilibrium prediction from Proposition 1. Prices increase with duration due to both the increase in mean costs and the reduction in variance, as we would expect if cost shocks are not perfectly correlated over time. The reduction in variance is captured by the positive coefficient α_1.

As expected, higher values for square footage and weekly frequency increase costs. Consistent with the findings from the descriptive regressions, baseline unemployment and high-intensity buildings have higher costs. For entry, higher current unemployment and the presence of more local establishments lower entry costs. Generic set-asides also have lower entry costs, relative to demographic-specific set-asides. Square footage has a net positive effect on entry, as $\gamma_1 > \kappa_1$. Supply costs, which are positively correlated with profits, increase by more than entry costs for square footage. Weekly frequency, on the other hand, has a net negative effect on entry, as $\gamma_2 < \kappa_2$. This is consistent with capacity constraints, as some firms may be limited in the days they are available to clean. The mean per-bidder entry cost estimate is \$573.

The model fits the data well. In Figure 5, I display actual values for annual prices compared to the predicted values. The R^2 for the structural model is 0.71, which compares favorably to the linear model IV-1 in Table 2. Unobserved heterogeneity is important to match the distribution of prices. Unobserved common costs are economically meaningful, in that they capture approximately 30 percent of the variance of log prices.

B. *Estimated Transaction Costs*

Market transaction costs are significant in this setting, comprising 10.9 percent of annual costs. These costs capture the marginal costs to the federal government of running a procurement auction, and they are summarized in Table 6. To obtain the aggregate share of costs attributable to transaction costs, I divide the mean annualized transaction costs by the mean total annual cost (the sum of annualized transaction costs and the price). Also displayed in the table are the median values for transaction costs, the annualized values, and the corresponding medians for contract value and price. The median transaction cost is estimated to be \$10,400, and the median share of costs attributable to transaction costs is 15 percent across contracts. The estimated magnitudes seem reasonable based on some back-of-the-envelope calculations, which I discuss in the following section.

[26] For a visual representation of how costs depend on duration, I plot the density of private cost draws for a one-year and a five-year contract in online Appendix H.

TABLE 5—PARAMETER ESTIMATES

Group	Parameter	Variable	Estimate	95 percent C.I.
Private costs	μ_0	–	18.521	[16.658, 20.861]
	μ_1	Duration	0.546	[0.085, 0.979]
	α_0	–	4.807	[3.527, 6.969]
	α_1	Duration	0.386	[0.053, 0.674]
Heterogeneity	σ_U	–	0.608	[0.572, 0.646]
	γ_1	Square footage	0.664	[0.627, 0.701]
	γ_2	Weekly frequency	0.488	[0.411, 0.556]
	γ_3	2004 unemployment	0.087	[0.070, 0.106]
	γ_4	High-intensity cleaning	0.302	[0.187, 0.437]
Entry	μ_ε	–	−0.459	[−0.749, −0.173]
	σ_ε	–	0.649	[0.608, 0.687]
	κ_1	Square footage	0.543	[0.488, 0.599]
	κ_2	Weekly frequency	0.542	[0.420, 0.650]
	κ_3	Unemployment shock	−0.309	[−0.449, −0.208]
	κ_4	Establishments	−0.066	[−0.108, −0.025]
	κ_5	Generic set-aside	−0.262	[−0.390, −0.148]

Notes: The table displays maximum likelihood parameter estimates from the structural model. The first group of coefficients indicate how the mean and shape of the private cost distribution change with the duration of the contract. The second set of coefficients indicate the distribution of unobserved auction-specific heterogeneity and how auction-specific common costs vary with observable cost characteristics. The third set of coefficients pertain to entry costs in the model. Ninety-five percent confidence intervals are displayed in the last column. As minor data cleaning steps (de-meaning) are data-dependent, confidence intervals are constructed via 500 bootstrap samples.

The sequential, revealed-preference approach has the benefit of providing testable implications of the model via the unconstrained estimates presented here. A finding of negative transaction costs, which would arise with private costs that fall with duration, would suggest that the trade-off in this paper is not first-order to contract duration. Instead, the 95 percent confidence intervals of μ_1 and α_1 have positive support, implying positive transaction costs only, which is consistent with the model. As previously, the premium on duration can arise simply from averaging cost draws across multiple periods. The premium captures opportunity costs as well, reflecting the seller's beliefs about the arrival rate of more profitable options.

In some cases, the market transaction costs are quite large as a percent of total costs. The ninety-fifth percentile of share of costs attributable to market transaction costs is 32.3 percent. For these estimates, this is driven by moderate transaction costs realized by low-price projects, rather than very high absolute costs. For example, contracts with a portion of transaction costs in the ninety-fifth percentile or above (greater than 32.3 percent) have a mean price of \$9,000, which is much smaller than the full-sample mean of \$43,900.

C. *Verifying the Estimated Transaction Costs*

To check the magnitudes of the estimated transaction costs, we can perform a back-of-the-envelope calculation by looking at the labor costs of employees who

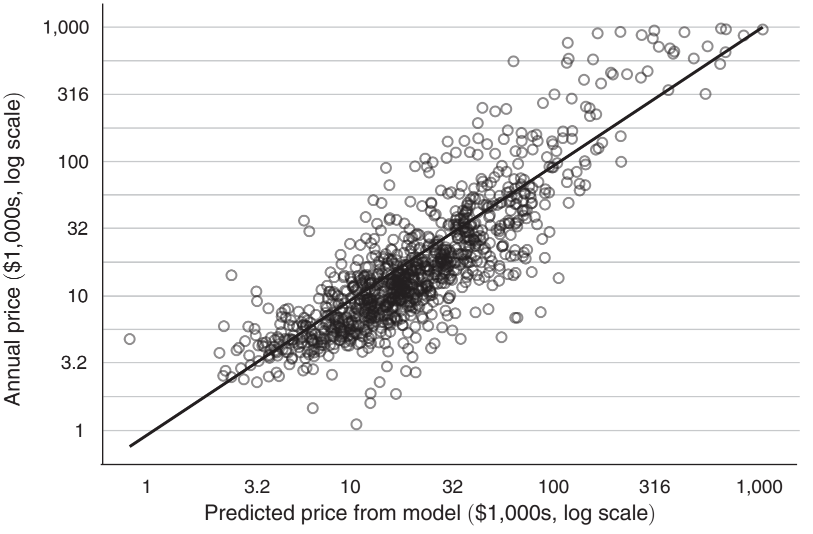

Figure 5. Model Fit: Actual versus Predicted Annual Price

Notes: The figure plots observed prices against predicted prices from the model. The R^2 of the predicted values is 0.71, which compares favorably to the R^2 of 0.69 from the linear instrumental variables model.

specialize in contracting, purchasing, and procurement.[27] Considering only these specialists will understate the full labor costs of managing contracts, as employees in several other categories are involved in issuing contracts. These other categories include supply program managers, logistics managers, and support service administrators, who all perform other tasks in addition to procurement, as well as specialists with industrial, engineering, or scientific knowledge who assist in determining the requirements for the contract.[28]

The labor costs for contract specialists provide a reasonable benchmark for relatively simple products and services. Large and complex contracts require more input from subject-matter experts and from consumers of the product or service, which will not be captured by the measure. Another consideration is that some contracts require ongoing maintenance by the contracting agency, so a portion of labor costs do not reflect the one-time costs of new contracts. For simple contracts, the labor costs of contracting officers are likely a close approximation of these otherwise hidden costs, as most of the effort is made up front.

For a new contract, the contracting officer must draft the requirements,[29] survey the market and decide terms, ensure compliance with existing regulations, post the solicitation, communicate with interested bidders, determine the winning bidder, and conclude the contract. A senior contracting officer for the federal government

[27] These correspond to GS-1102, GS-1105, and GS-1106 in the federal government's classification system.

[28] The classifications corresponding to the administrative roles are GS-2003, GS-0346, GS-0342. Those pertaining to industrial, engineering, or scientific knowledge are GS-1150, GS-0800, GS-1300, and GS-0400. The standards may be obtained here: https://www.opm.gov/policy-data-oversight/classification-qualifications/classifying-general-schedule-positions/.

[29] The main document for the median contract runs 49 pages. See online Appendix D for example pages.

Table 6—Estimated Market Transaction Costs ($)

Contract-specific measure	Median	95 percent C.I.	p25	p75	Mean
Transaction costs	10,400	[3,000, 17,400]	5,100	21,400	24,500
Annualized	2,500	[700, 4,100]	1,300	4,800	5,400
Contract value	50,500	[46,900, 54,100]	28,500	102,000	190,200
Price (annual)	13,200	[12,300, 13,900]	7,300	26,700	43,900
Percent share of costs	15.2	[5.0, 22.0]	9.9	20.8	16.3
Aggregate measure	Estimate	95 percent C.I.			
Percent share of costs	10.9	[3.8, 20.3]			

Notes: Estimated transaction costs are the expectation taken with a uniform prior over the distribution-free bounds identified from the duration decision of the buyer. For $T = 5$, conservative upper bounds are projected by assuming that the duration is optimally chosen. Transaction costs are also expressed as a share of total (buyer) costs. The aggregate share of total costs attributable to transaction costs is 10.9 percent, which is calculated by comparing the mean annualized transaction costs to the mean price. Confidence intervals are constructed via the bootstrap. For display, values are rounded to the nearest 100.

estimated that the simplest cleaning contract would take about three weeks of full-time work for a contracting employee. For fiscal years 2004 to 2016, the average salary for contracting specialists was $78,253. Three weeks of full-time work, assuming a 50-week year, provides a cost estimate of $4,695, which is roughly in line with the twenty-fifth percentile estimate.

Looking at the contracts in aggregate provides another rough benchmark for transaction costs. To compare to the estimation sample, I examine all contracts for fiscal years 2004 to 2016 that had an annual price of less than $1 million. Approximately 750,000 of these contracts are issued each year, or 99 percent of all new contracts. These contracts are modestly less expensive than the estimation sample, with a mean annual price of $35,469 and a median of $4,731. Overall, 19.8 new contracts per contracting specialist are issued each year. Allocating the salaries of contracting specialists across new contracts obtains an average labor cost of $3,953 per contract, or 7.7 percent of total buyer costs.[30] These two calculations provide benchmarks that are the same order of magnitude of the transaction costs recovered in estimation.

As an additional exercise to check the plausibility of the estimated market transaction costs, I project the estimates on other variables not used in the structural estimation. First, I calculate the median transaction costs by facility type and by department in Table 7. As expected, the highest transaction costs are among facilities with relatively complicated or technical requirements, such as medical centers, airports, and technical facilities (e.g., power plants). Simpler settings such as office cleaning have the lowest estimated transaction costs. In the second panel, I calculate the median by government department. The Department of Homeland Security has the highest

[30] If we instead assumed that new contracts over $1 million required five times the contracting specialists compared to these smaller contracts, then the labor cost per new contract is $3,805, or 7.4 percent of total buyer costs. Seventy-five percent of contracts under $1 million and 29 percent of larger contracts are new each year, based on the latter half of the data (FY 2011–16). Larger contracts likely involve significant resources from other employee classifications. The 3 percent of contracts over $1 million annually (31,929 each year on average) comprise over 90 percent of all dollars obligated. When including this long tail, the aggregate labor cost per contract from contracting specialists is less than one percent of total costs. Larger contracts likely involve significant resources from other employee classifications.

TABLE 7—ESTIMATED MARKET TRANSACTION COSTS BY CATEGORY

Type	Transaction costs	Contract value	Square footage	Count
Panel A. Median values by location type				
Medical	38,900	206,000	10,000	61
Airport	32,850	254,300	7,850	30
Technical	30,600	84,000	15,600	19
Industrial	28,500	60,000	27,900	13
Accommodations	28,350	121,200	32,000	18
Services	18,400	87,700	8,300	59
Research	13,400	58,500	6,000	111
Visitors	13,000	189,700	6,500	41
Field office	9,250	48,500	8,450	270
Office	7,300	35,550	4,750	424
Panel B. Median values by department				
Department				
Homeland security	40,000	269,300	14,600	45
GSA	29,350	223,100	12,750	40
Veterans affairs	24,450	143,550	8,800	80
Other	20,900	69,200	11,650	24
Commerce	14,150	59,750	5,500	78
Interior	11,900	71,000	8,900	43
Agriculture	9,500	46,700	9,300	347
Defense	7,300	35,900	4,200	389

Notes: The table displays the median estimated market transaction costs. Also displayed are the median contract value, the median square footage of the facility, and the count of observations. In panel A, observations are grouped by location type. In panel B, observations are grouped by contracting department. For display, underlying values are rounded to the nearest 100.

median transaction costs, at $40,000 per contract. This might be expected given the high levels of security required at their facilities and the relative lack of institutional knowledge at the recently formed department.[31] Conversely, Agriculture and Defense have low median transaction costs, at $9,500 and $7,300, respectively. After controlling for square footage, cleaning frequency, and facility type in a regression, Homeland Security has the highest fixed effect for (log) transaction costs, 91 percent larger than Defense. Agriculture has the lowest fixed effect, 17 percent smaller than Defense. The regression table is provided in online Appendix H.

In Table 8, I regress the estimated market transaction costs on variables excluded from the structural model. Included covariates are the number of pages in the contract, related expenditures and contract modifications[32] in the same nine-digit zip code, and an indicator for whether the contract falls under the simplified acquisition protocol. One would expect that lengthier contracts and busier agencies are reflective of higher opportunity costs, and that the simplified acquisition label would reflect lower market transaction costs. After controlling for square footage and cleaning frequency, high-expenditure locations are associated with higher transaction costs. Economic theory could rationalize a sign in either direction, as economies of scale lead to a positive association and capacity constraints produce a negative one. The

[31] Contracting employees in Homeland Security had lower length of service and were less likely to have a bachelor's degree compared to the average contracting employee.

[32] Related expenditure categories are other housekeeping services, maintenance, and office furniture.

TABLE 8—PROJECTING MARKET TRANSACTION COSTS ON VARIABLES OUTSIDE OF THE MODEL

	(1)	(2)	(3)	(4)	(5)	(6)
High-intensity cleaning	1.448	1.189	1.140	1.022	1.158	0.612
	(0.144)	(0.133)	(0.134)	(0.138)	(0.138)	(0.108)
ln(word count)	0.085				0.124	0.043
	(0.024)				(0.023)	(0.018)
ln(related expenditures)		0.096			0.073	0.054
		(0.011)			(0.023)	(0.017)
ln(related modifications)			0.283		0.047	−0.141
			(0.035)		(0.071)	(0.055)
Simplified acquisition indicator				−0.692	−0.715	−0.429
				(0.090)	(0.089)	(0.069)
ln(square footage)						0.577
						(0.026)
ln(weekly frequency)						0.510
						(0.057)
Observations	1,046	1,046	1,046	1,046	1,046	1,046
R^2	0.09	0.14	0.13	0.13	0.20	0.54

Notes: The table displays estimated coefficients from regressing estimated log transaction costs on variables outside of the model. These variables are (i) the (log) number of pages in the contract; (ii) log government procurement expenditures at the same nine-digit ZIP code for maintenance, office furniture, and other housekeeping services; (iii) the count of contract actions for these expenditures; and (iv) an indicator for whether the contract falls under the federal government's simplified acquisition protocol. Standard errors in parentheses.

negative coefficient on contract modifications in the fifth specification may reflect economics of scale or simply that lower transaction costs lead to more contract modifications.

Though understanding the precise composition of these costs requires further study, the analysis above and some supplemental data allow us to (very roughly) speculate about the breakdown of these costs. Table 8 shows that, controlling for other features of the contract, simplified acquisition procedures are associated with 35 percent lower transaction costs (−0.429 log points). Contracts made under simplified acquisition do not require (i) market research, (ii) conducting discussions with potential vendors, and (iii) establishing a formal source selection team. The bidding mechanism itself (iv) may constitute around 1 percent of the total value of the contract, or 10 percent of the transaction costs. For example, FedBid, one of the leading auction platform providers for the federal government, charges "no more than 3 percent of the winning bid." Anecdotally, a fee of around 1 percent of the winning bid is not unusual for external auction platforms. Therefore, these first four components may make up about 45 percent of transaction costs. The bulk of the residual is likely allocated to (v) deciding on contract terms and (vi) checking that the contract specifications are up to date with current regulations. Contracting officers may copy and paste text from a previous contract, but there is a significant burden to ensure that the contracts are in compliance. The Federal Acquisition Regulation (FAR) that governs these contracts is over 2,000 pages and frequently references other legal code, such as federal employment regulations. Finally, other costs include (vii) marketing the opportunity to potential vendors and (viii) documenting each step according to FAR.

Again, I emphasize that these breakdowns are very rough. They are intended to illustrate the different components that constitute ex ante transaction costs, rather than providing exact empirical quantities. The literature on this front is sparse, and deeper analysis of these costs is a potential avenue for future research.

D. *Robustness*

To estimate the model, I have followed typical practice in the empirical literature, such as assuming multiplicative separability in cost components. For the discount rate and for five-year contracts, I have made assumptions that generate conservative estimates of transaction costs. For a discussion of these choices and sensitivity to specific parameterizations, see Section IIIC.

For the empirical model, we have proceeded under the assumption that bidders are symmetric with respect to private supply costs. In a dynamic setting, the procurement process might result in asymmetry between bidders that would invalidate this assumption. One common source of asymmetry in procurement is the presence of an incumbent bidder who may have an advantage via a relationship-specific investment (e.g., through learning-by-doing or lowered transaction costs of retaining the same supplier). Additionally, competing bidders may retain some information about competitors if costs are correlated over time.

I check for the presence of asymmetries by comparing the expected win rate under symmetry (based on the number of bidders) to the win rate for incumbent suppliers in follow-on contracts. I identify follow-on contracts in the analysis sample by finding contracts that have a single active supplier on another contract in the same 9-digit zip code within the prior year (and starting at least 30 days before). Note, nine-digit zip codes are geographically narrow, typically corresponding to a city block or an individual company. The prior contract may be any of the approximately 11,000 cleaning contracts in the FPDS data. I also construct a broader set of follow-on contracts from the extended FPDS sample.

Table 9 compares the expected win rate for symmetric bidders to the actual win rate for incumbent bidders in identified follow-on contracts. There is no significant difference between the two, suggesting that the incumbency advantage is not first-order in this setting. I obtain similar results for the 175 contracts in the estimation sample and the 845 contracts from the broader FPDS sample.[33] There are a priori reasons to believe that the incumbency advantage is not large for competitive federal procurement, as, per regulation, the agencies are mandated to seriously consider all qualified bidders and, in most cases, select the lowest price. The degree of relationship-specific investments in facility cleaning is likely to be low, as the menu of services tend to be standardized.

Transaction costs do not appear to depend on whether the winning bidder is an incumbent. For follow-on contracts, there is no difference in mean (log) transaction costs between contracts that are won by an incumbent $(2.277, N = 38)$ and those that are not $(2.274, N = 137)$. As an additional test, I include a dummy for whether

[33] As I only observe winning bidders, I am unable to adjust for when a supplier does not bid on a follow-on to the supplier's current contract.

TABLE 9—TEST FOR ASYMMETRY: DO INCUMBENTS HAVE AN ADVANTAGE?

Follow-on contracts	Symmetric win rate	Incumbent win rate	Observations	t-statistic
Estimation sample	0.224	0.217	175	(0.20)
Extended FPDS sample	0.278	0.263	845	(1.00)

Notes: The table displays the results of a test for asymmetry in performance by incumbent bidders. The expected win rate for symmetric bidders, based on the number of bids, is compared to the observed win rate by incumbent bidders. The t-statistics indicate no significant difference in either sample. The first sample is follow-on contracts in the estimation sample, and the second sample uses the same criteria for all FPDS building cleaning contracts. Follow-on contracts are identified as contracts that have a single leading contract for the same agency in the same nine-digit ZIP code. A leading contract is one that is active in the year prior to the start of the follow-on contract and begins at least 30 days prior to the start of the follow-on contract.

the contract is an identified follow-on contract in the descriptive regressions from Section II to determine if variation in prices and entry are explained by the presence of an incumbent bidder. None of the coefficients on the dummy are significant, and its inclusion does not meaningfully change any of the coefficients of interest. For these regressions, see online Appendix H. The results of these tests are consistent with the maintained assumption of no endogenous asymmetries.

An additional general concern might be that there is heterogeneity in supplier types. The above tests for endogenous asymmetries are also valid tests for exogenous asymmetries in supplier types. Lower-cost types would be more likely to win the first contract in the identified set of follow-on contracts, thereby generating a correlation in win rates over time. Thus, the above findings are consistent with symmetry across suppliers more generally. In contrast to many other industries, there is no great distinguishing factor that separates types of building-cleaning firms, and it is reasonable to expect that production is roughly constant returns-to-scale. This makes the empirical setting a nice fit for the model.

Finally, an implicit assumption of the empirical model is perfect alignment between the government and the contracting officer; i.e., there is no principal–agent problem in determining duration. If we instead suppose that the contracting officer does not fully internalize the price of the contract to the government (but does bear the market transaction costs), then we would obtain the same estimates. In this case, the estimated transaction costs have a shadow cost interpretation: they represent the relevant ex ante costs to the government, but they are greater than the direct transaction costs due to misaligned incentives.[34] Empirically, one could estimate

[34] In the case of ex ante misalignment, the contracting officer solves

$$\min_{T \in \mathcal{T}} \omega \cdot \bar{P}\left(T, x, m\right) + \frac{\delta}{\sum_{k=1}^{T} \beta^{k-1}}, \tag{13}$$

where $\omega < 1$ captures misalignment between the contracting officer and the government. The problem may be rewritten as

$$\min_{T \in \mathcal{T}} \bar{P}\left(T, x, m\right) + \frac{\delta/\omega}{\sum_{k=1}^{T} \beta^{k-1}}. \tag{14}$$

Applying this model, we would obtain the same estimates for transaction costs, though we would instead capture the shadow cost to the government, δ/ω, or the transaction costs scaled by the degree of misalignment between the contracting officer and the government. I thank two anonymous referees for this suggestion.

misalignment by interpreting the agency-specific fixed effects from the previous section as a measure of misaligned incentives. If we assume that the agency with the smallest fixed effect for transaction costs has perfect alignment, then the average agency treats the contract price at 82 cents on the dollar, and the direct transaction costs are 9.1 percent of total buyer costs.

E. *Generalizability*

The estimated transaction costs are obtained in a specialized setting that allows them to be isolated from other factors. It may be reasonable to translate these costs to federal contracts for other standardized products, where the buyer's problem is similar. As discussed in Section II, cleaning contracts are comparable in magnitudes to the 97 percent of federal contracts under $1 million per year. An appealing feature about the application is that the large variation in observables allows us to account for how these costs vary with project scale and complexity.

In the private sector, many of the components that make up ex ante transaction costs are still relevant: market research, writing up specifications, marketing the opportunity, and running a mechanism to determine the winner. Buyers may be able to avoid regulation and compliance costs that are specific to federal government procurement; since these appear to be sizable for federal contracts, these might result in meaningfully lower transaction costs in the private sector. On the other hand, the size of government procurement may provide some economies of scale and reduce these costs. The federal government can use an existing platform (FedBizOpps) to post each solicitation, so the estimated transaction costs exclude the fixed costs of setting up such a platform.

Another consideration is that the federal regulations are designed to minimize poor decision-marking by the contracting officers. Because the private sector has fewer restrictions, the potential for misalignment between the buying agent and the firm may be greater. This can be reflected in the terms, the type of competitive procedure, or the contractual arrangement. Based on conversations with procurement officers from several organizations, transaction costs of around 10 percent of total costs is not an unreasonable estimate in the private sector, though there is a great deal of heterogeneity in procurement efficiency and the data are sparse.[35]

The nature of supply costs is an important consideration in more general contexts. With building cleaning services, supply is stable over time, and the service represents a small fraction of overall expenditures. In other settings, such as technology-dependent firms, the per period supply costs may be falling over time. Whether or not this feature leads to shorter contracts depends on whether cost-reducing innovations are predictable and whether they are driven by high-cost or low-cost firms. Furthermore, buyers may not be risk-neutral with respect to a

[35] In one example, a private university describes a nine-month process to obtain a one-year discount contract on microscope light bulbs: "Selection of the Preferred Vendor was the culmination of 9 month process managed by Purchasing in cooperation with the Office for Research. It involved surveying the research community to assess needs, submitting an RFQ to vendors, interviewing vendors, and negotiating price and service … Purchasing was able to negotiate a substantial reduction in price for one year." https://www.northwestern.edu/procurement/about/dollars-sense-newsletter/DSFall2015.pdf.

primary input that is specialized to their business, which is another factor to consider when taking the model to other settings. Even so, the trade-off I identify here remains relevant.

V. Counterfactuals: The Impact of the Duration Margin

I now consider the implications of the endogenous duration decision on welfare. In the first counterfactual, I analyze the cost to the buyer of removing the ability to adjust duration through standard contracting terms. In the second counterfactual, I show how duration provides a margin of adjustment that mediates the pass-through of cost shocks to prices.

A. *Strategic Value of the Duration Decision*, *Compared to Standardization*

When the buyer can adjust the nonprice terms of a contract, the buyer can minimize expected costs for each transaction. This flexibility provides a cost-minimizing advantage to the firm. In many settings, contract terms are standardized. For example, a three-year contract is the industry standard for office supplies. The structural model allows us to estimate how costly it is to remove the buyer's strategic option to adjust duration. One may also interpret these impacts as the costs of going to the market more or less frequently.

Table 10 reports the impact on aggregate buyer costs by moving to standardized terms of yearly increments. Total costs would increase substantially, by 37 percent, if all contracts were issued in one-year terms. This is not surprising, as the median duration in the data is five years and standardization would result in much more frequent contracting. On the other hand, standard durations of four years or five years would have a small impact, increasing buyer costs by less than two percent.

These results suggest that flexible terms may be quite valuable, compared to a poorly chosen standard (e.g., one year or two years in this setting). Thus, knowledge of the relevant cost structure and transaction costs is important for setting nonprice terms. This analysis further highlights the importance of the duration margin.[36]

To provide additional context on these magnitudes, I calculate the reduction in transaction costs that would be required to offset the increase in supply costs from standardization. This may be relevant to a firm that is considering a new contracting process that requires standardized terms but also reduces the transaction costs for each contract. In the final column of Table 10, I report the compensating change in transaction costs that would make the standardized term policy equivalent to the flexible term policy. For a four-year standard term, the necessary reduction in transaction costs is modest. If the government could implement a process that required

[36] A related question to standardized terms is that of a cap on maximum duration, similar to the five-year cap imposed by government-wide budgeting regulations in my data. This is analogous to the imposition of standard terms on only a subset of contracts. In online Appendix H, I provide a detailed breakdown of the effects by whether duration is increased or decreased by the standard, which provides insight into the cost of the cap. Table H4 in the online Appendix reports averages by contract, rather than in aggregate, which is why the numbers differ slightly from those in Table 10.

Table 10—Effects of Standardized Terms (Percent)

$\bar{T}$	Total cost	95 percent C.I.	Price	Transaction cost	Affected	Compensating δ
1	36.7	[12.1, 55.3]	−11.9	354.0	1,018	−60.9
2	9.9	[3.2, 15.0]	−8.0	127.0	992	−33.0
3	3.2	[1.0, 4.8]	−4.2	51.3	907	−16.0
4	1.4	[0.5, 2.4]	−0.4	13.5	992	−9.4
5	1.6	[0.5, 2.8]	3.2	−9.2	496	−13.0
6	2.7	[0.8, 4.4]	6.8	−24.3	1,041	−26.6

Notes: The table displays the resulting percent changes in total costs, prices, and annualized transaction costs when all contracts are issued in standardized durations corresponding to $\bar{T}$. For a uniform duration policy of four years or less, the average price paid decreases and the amount spent on transaction costs increases. Affected contracts are the count of those that are displaced from the optimal duration. The final column displays the reduction in transaction costs that would render a uniform policy equivalent to the existing policy in terms of buyer costs. Confidence intervals are reported for total costs and are constructed via the bootstrap.

standard durations of four years and reduced transaction costs by 10 percent, it would be beneficial to do so.

B. *Contract Duration and Welfare Analysis*

Transaction costs are important to welfare analysis as they can constitute a substantial portion of total costs and affect how equilibrium prices respond to a change in the economic environment. When transaction costs are unaffected by a policy change, a welfare analysis that omits transaction costs will misstate the impact for two reasons. First, the measured impact on prices should be weighted by the share of total costs attributable to prices. That is, the impact should be discounted toward zero by the share attributable to (unaffected) transaction costs. Second, market participants adjust equilibrium behavior in response to the change. The choice of duration provides an additional margin of adjustment, mitigating the effect on prices but improving welfare compared to an analysis that takes duration as fixed.

To demonstrate these effects, I calculate equilibrium cost pass-through in the model when duration is accounted for and when duration is fixed. To measure pass-through, I simulate a 10 percent reduction in supply costs for all the contracts in my data. For comparison, I provide three specifications: one in which entry and duration are taken as given, a second in which entry is endogenous, and a third in which both entry and duration respond to the cost shock.

The results of the counterfactual are reported in Table 11. In the first specification, the aggregate price falls by exactly 10 percent, as the equilibrium bids in the model are proportional to supply costs. However, total costs to the buyer fall by only 8.7 percent, as transaction costs represent a substantial portion of total costs. Thus, these costs that are otherwise hidden may be important to account for when measuring welfare impacts.

In the second specification, I allow the participation of the bidders to respond to the change in costs. Based on the estimated parameters, lower supply costs results in less entry. Due to reduced competition, the impact on prices is 14 percent lower (8.6 percent compared to 10 percent). The impact of a reduction in supply costs is

Table 11—Effects of a 10 Percent Reduction in Supply Costs (Percent)

Specification	Price	Duration	Total cost
No response in entry or duration	−10.00	0.00	−8.67
	[−10.00, −10.00]	[0.00, 0.00]	[−9.53, −8.01]
Endogenous entry only	−8.59	0.00	−7.45
	[−8.92, −8.14]	[0.00, 0.00]	[−8.23, −6.85]
Endogenous entry and duration	−7.70	6.64	−7.50
	[−8.33, −7.04]	[5.86, 9.63]	[−8.26, −6.92]

Notes: The table reports the equilibrium changes to prices, duration, and total costs when supply costs fall by 10 percent. In the first row, entry and contract duration are treated as exogenous. In the second row, entry by potential bidders changes in response to supply costs. In the last row, both entry and contract duration respond endogenously. The first column shows the effects on price. The estimate divided by −10 percent captures the pass-through of supply costs to prices in each scenario. Confidence intervals are constructed via the bootstrap.

not perfectly passed through to prices because firms adjust their participation decision on the margin.

Likewise, the duration margin also mitigates the pass-through of supply costs to prices. In the third specification, buyers can adjust duration. In response to lower supply costs, the average contract duration increases by 6.6 percent. As supply costs increase with duration, this adjustment partially offsets the reduced supply costs. While allowing for both endogenous entry and duration to respond, supply prices fall by only 7.7 percent. Thus, the marginal effect of endogenous duration is to reduce pass-through by 10 percent. In contrast to the previous counterfactual, here the duration margin provides only a small reduction in total costs.

This counterfactual exercise illustrates that the duration margin can mitigate the pass-through of costs to prices. It can be important in a welfare analysis, as observed changes to prices may reflect a number of endogenous levers. Further, prices may only capture a portion of the total costs, so accounting for transaction costs can be important. In addition, transaction costs may be affected by a policy change. The changes should be accounted for when evaluating welfare effects, as well as the equilibrium impacts on duration and price.

VI. Conclusion

In this paper, I develop a model of optimal contract duration arising from underlying supply costs and market transaction costs. I show how latent transaction costs may be recovered from the duration decision of the buyer. Using a dataset of federal supply contracts, I find that the costs of going to the market can be a significant portion of total costs for intermediate goods.

The methods developed in this paper may prove useful for welfare analysis, especially in industries where supply contracts are prevalent. Using counterfactual analyses, I show that the ability to endogenously adjust contract duration can have meaningful impacts on welfare. In many settings, the trade-off presented in this paper may complement other concerns arising from ex post incentive problems and incomplete contracts. An appropriate model should be tailored to the industry in question.

The analysis presented here offers, albeit indirectly, one novel prediction regarding the theory of the firm. Supply contracts lie in between arms-length transactions and vertical integration. As is known, conditions favorable for long-term contracts are also favorable for vertical integration, as the end is similar and integration may result in additional benefits. I demonstrate here that long-term contracts arise when competition is sufficiently low, and also when competition is very intense. Likewise, vertical integration may be most likely for low levels and high levels of competition. When the industry is moderately competitive, a downstream firm can realize a large benefit by switching among suppliers and may have the smallest incentive to integrate upstream.

Appendix A: Model Proofs and Additional Predictions

A. *Proof of Propositions 1 and 2*

By assumption, an interior solution exists. Let the set of possible contracts be order in increasing duration: $\mathcal{T} = \{1, \ldots, R, S, T, \ldots\}$. At an interior solution S, it must be that

$$\bar{P}(R) \geq \bar{P}(S) - \left(\frac{\delta}{\sum_{k=1}^{R}\beta^{k-1}} - \frac{\delta}{\sum_{k=1}^{S}\beta^{k-1}}\right), \tag{A1}$$

$$\bar{P}(T) \geq \bar{P}(S) + \left(\frac{\delta}{\sum_{k=1}^{S}\beta^{k-1}} - \frac{\delta}{\sum_{k=1}^{T}\beta^{k-1}}\right). \tag{A2}$$

These are obtained by simple rearrangements of equation (4). Because $\beta \in (0,1]$ and $R \leq S \leq T$, both terms in parentheses are positive. Therefore, the function $\bar{P}$ is locally increasing the duration of the contract, i.e., $\bar{P}(S) \leq \bar{P}(T)$.

Further, an increase in δ decreases the right-hand side of equation (A1), so that the inequality remains satisfied. An increase in δ increases the right-hand side of equation (A2). No increase in δ could make the shorter contract R preferred to S, but a large enough increase in δ will flip the inequality in (A2), and the longer contract T will be preferred to S. ∎

B. *Additional Predictions from Model*

Recall the decision rule for the buyer from the illustrative example. The buyer will choose a long-term contract if and only if

$$E[\tilde{c}_{2:N}] - E[c_{2:N}] < \frac{\delta}{2}. \tag{A3}$$

This simple decision rule generates a number of comparative statics. I discuss two of these in the main text, and I present additional ones here.

REMARK 3: *Higher marginal costs lead to shorter contracts.*

Higher marginal costs increase the left-hand side of equation (A3). This increases the cost of long-term contracts relative to the savings in transaction costs, which shifts buyers to long-term contracts at the margin. Similarly, this increase in transaction costs affects the right-hand side only and leads to longer contracts.

REMARK 4: *The optimal duration is increasing with autocorrelation in supply costs.*

This prediction is intuitive. As the autocorrelation in marginal costs increases, there is less of a benefit from switching suppliers, and longer-term contracts are preferred. Suppose that d is a cost process with lower autocorrelation than c, but the same per period marginal distribution, i.e., $E[d_{2:N}] = E[c_{2:N}]$. Let $\tilde{d}$ denote the average cost across two periods. Then it follows that, for $N > 3$,

$$E[\tilde{d}_{2:N}] > E[\tilde{c}_{2:N}] \tag{A4}$$

$$\Rightarrow \quad E[\tilde{d}_{2:N}] - E[d_{2:N}] > E[\tilde{c}_{2:N}] - E[c_{2:N}]. \tag{A5}$$

The marginal cost of long-term contracts is decreasing with the autocorrelation of the cost process. With greater autocorrelation, long-term contracts are preferred.

REMARK 5: *The optimal duration is decreasing in the variance of costs across suppliers, provided there is sufficient competition* $(N > 3)$.

For a simple case, consider location-scale transformations of c, such that $d = a + bc$ and $E[d] = E[c]$. Under the marginal cost structure d, a longer contract is chosen if

$$b \cdot \left(E[\tilde{c}_{2:N}] - E[c_{2:N}]\right) < \frac{\delta}{2}. \tag{A6}$$

As b increases, shorter contracts become more desirable.

REMARK 6: *When costs are bounded from below, the optimal duration is U-shaped in the variance in costs, provided there is sufficient competition* $(N > 3)$.

From a starting point of zero variance across suppliers, increasing the variance of marginal costs leads to shorter contracts, as there is more to gain from selecting the low-cost supplier in each period. This holds for the buyer-optimal contract as long as there are more than three suppliers, in which case the expected second-order statistic falls below the median. When costs are bounded from below, eventually both $E[\tilde{c}_{2:N}]$ and $E[c_{2:N}]$ approach zero, and the cost of a longer duration falls with respect to transaction costs. After a certain threshold, contract duration increases.

This occurs because the expected average price and expected per period price approach the lower bound. Let c denote per period marginal costs with a lower bound at 0, and let σ represent its standard deviation. Then, when $N > 3$,

$$\lim_{\sigma \to \infty} E[\tilde{c}_{2:N}] = \lim_{\sigma \to \infty} E[c_{2:N}] = 0. \tag{A7}$$

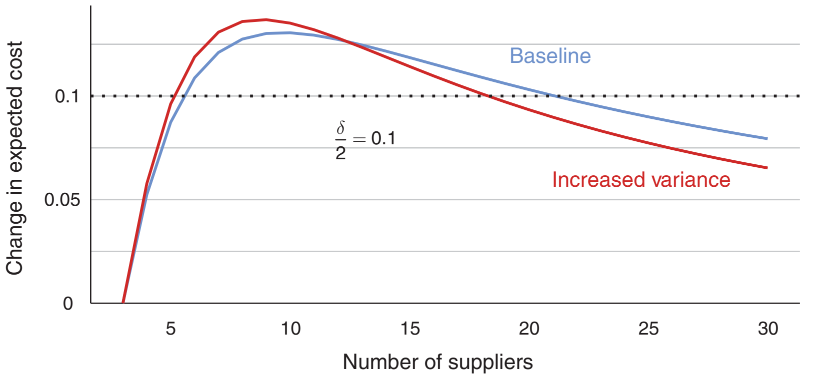

FIGURE A1. INCREASED VARIANCE IN COST

Notes: The blue line shows the marginal cost to the buyer of a two-period contract relative to one-period contracts and is equivalent to the blue line in Figure 1. The red line shows the marginal cost of a longer contract when the costs are drawn from the same distributional family (the beta distribution) with 11 percent greater variance.

As $E[\tilde{c}_{2:N}] - E[c_{2:N}] \rightarrow 0$, long-term contracts are optimal in the limit. This effect tends to dominate as N gets large, as more draws brings the minimum price closer to the lower bound.

This trade-off is illustrated in Figure A1. The blue line displays the baseline marginal cost of longer contracts, corresponding to the blue line in panel B of Figure 1 in the main text. Costs drawn from a beta distribution with shape parameters $(0.5, 0.5)$. The red line displays the change in marginal costs when variance of the cost distribution increases by 11 percent, corresponding to a beta distribution with shape parameters $(0.4, 0.4)$.

For $N < 12$, greater variance increases the cost of long-term contracts, reflecting the intuition of Remark 5. When $N \geq 13$, the winning supplier's price is close enough to the lower bound to reduce the cost, reflecting the prediction of Remark 6. Thus, the figure illustrates how an increase in variance leads to shorter contracts when competition is lower and longer contracts when competition is intense.

APPENDIX B: IDENTIFICATION PROOFS

A. *Some Lemmas*

To demonstrate the following proofs, it will be useful to first introduce several lemmas.

LEMMA 1: *For symmetric auctions with independent private values,* $E[b_{1:N}] = E[c_{2:N}]$.

This is a standard result and can be obtained directly by taking the expectation given the equilibrium bid function. I omit the proof here.

LEMMA 2: $\min b_{1:N} = E\left[c_{1:(N-1)}\right]$ *for the IPV model when the support of c is bounded from below by* $\underline{c} > -\infty$.

PROOF:

The equilibrium bid function is given by

$$\beta(c;N) = c + \frac{\int_c^\infty \left[1 - F(\xi)\right]^{N-1} d\xi}{\left[1 - F(c)\right]^{N-1}}. \tag{B1}$$

Then the minimum bid is

$$\begin{aligned}
\beta(\underline{c};N) &= \underline{c} + \frac{\int_{\underline{c}}^\infty \left[1 - F(\xi)\right]^{N-1} d\xi}{\left[1 - F(\underline{c})\right]^{N-1}} \\
&= \underline{c} + \int_{\underline{c}}^\infty \left[1 - F(\xi)\right]^{N-1} d\xi \\
&= \underline{c} + \xi \left[1 - F(\xi)\right]^{N-1}\Big|_{\underline{c}}^\infty + \int_{\underline{c}}^\infty \xi(N-1)f(\xi)\left[1 - F(\xi)\right]^{N-2} d\xi \\
&= \underline{c} + (0 - \underline{c}) + \int_{\underline{c}}^\infty \xi(N-1)f(\xi)\left[1 - F(\xi)\right]^{N-2} d\xi \\
&= E\left[c_{1:(N-1)}\right].
\end{aligned}$$

The third line comes from integration by parts. Here we require the assumption that $\lim_{\xi\to\infty} f(\xi)\left[1 - F(\xi)\right]^N = 0$, so that

$$\begin{aligned}
\xi\left[1 - F(\xi)\right]^{N-1}\Big|_{\underline{c}}^\infty &= \lim_{\gamma\to 0} \frac{\left[1 - F\left(\frac{1}{\gamma}\right)\right]^{N-1}}{\gamma} - \underline{c}\left[1 - F(\underline{c})\right]^{N-1} \\
&= \lim_{\gamma\to 0} \frac{-(N-1)f\left(\frac{1}{\gamma}\right)\left[1 - F\left(\frac{1}{\gamma}\right)\right]^{N-2}}{1} - \underline{c} \\
&= 0 - \underline{c}. \blacksquare
\end{aligned}$$

LEMMA 3: *The expected kth order statistic of N draws can be written in terms of the expected kth and* $(k+1)$*th order statistics from* $N+1$ *draws:* $E\left[c_{k:N}\right] = \left(k/(N+1)\right)E\left[c_{(k+1):(N+1)}\right] + \left((N+1-k)/(N+1)\right)E\left[c_{k:(N+1)}\right]$.

PROOF:

First, examining the difference between the kth order statistics of N and $N+1$ draws. Expressing $E\left[c_{k:N}\right] - E\left[c_{k:(N+1)}\right]$ and rearranging terms gives:

$$
\begin{aligned}
&E\left[c_{k:N}\right] - E\left[c_{k:(N+1)}\right] \\
&\quad = \int \frac{N!}{(k-1)!(N-k)!} cf(c)F(c)^{k-1}[1-F(c)]^{N-k} dc \\
&\qquad - \int \frac{(N+1)!}{(k-1)!(N+1-k)!} cf(c)F(c)^{k-1}[1-F(c)]^{N+1-k} dc \\
&\quad = \int \left(\frac{N!(N+1-k)}{(k-1)!(N+1-k)!} - \frac{(N+1)!}{(k-1)!(N+1-k)!}[1-F(c)] \right) \\
&\qquad\qquad \times cf(c)F(c)^{k-1}[1-F(c)]^{N-k} dc \\
&\quad = \int \frac{(N+1)!}{(k-1)!(N+1-k)!} cf(c)F(c)^{k}[1-F(c)]^{N-k} dc \\
&\qquad - \int \frac{kN!}{(k-1)!(N+1-k)!} cf(c)F(c)^{k-1}[1-F(c)]^{N-k} dc \\
&\quad = \frac{k}{(N+1-k)} \left(E[c_{(k+1):(N+1)}] - E[c_{k:N}] \right).
\end{aligned}
$$

Rearranging, we obtain

$$E\left[c_{k:N}\right] = \tfrac{k}{N+1} E\left[c_{(k+1):(N+1)}\right] + \tfrac{N+1-k}{N+1} E\left[c_{k:(N+1)}\right]. \blacksquare$$

B. *Proof of Proposition* 5

Consider the entry equation

$$E\left[\pi_n \cdot U \cdot h(x) | n, T\right] - k(m) \cdot \varepsilon > 0 \quad \Leftrightarrow \quad N \geq n \tag{B2}$$

$$\Rightarrow \quad E\left[\pi_n | n, T\right] \cdot \frac{h(x)}{k(m)} > \varepsilon \quad \Leftrightarrow \quad N \geq n. \tag{B3}$$

For any realization (T, x, m), there exists (T, x', m') such that $\Pr(N \geq n | T, x, m) = \Pr(N \geq n | T, x', m')$ for all N.[37] Using these values x' and m' that provide the same conditional distribution of N, calculate:

$$\frac{E\left[B \cdot U \cdot h(x) | N, T, x, m\right]}{E\left[B \cdot U \cdot h(x') | N, t, x', m'\right]} = \frac{E\left[B | N, t\right] \cdot E\left[U | N, T, x, m\right] \cdot h(x)}{E\left[B | N, t\right] \cdot E\left[U | N, t, x', m'\right] \cdot h(x')} = \frac{h(x)}{h(x')}. \tag{B4}$$

As $h(x)$ is normalized to 1 at $x = x_0$, $h(\cdot)$ is identified.

[37] Here, and once more in the proof, I rely on either $h(\cdot)$ or $k(\cdot)$ having broad support.

Once $h(\cdot)$ is identified, $E[B|N,T,x,m]$ is identified by calculating the mean of the scaled (conditional) winning bid $E[(1/h(x))B \cdot U \cdot h(x)|N,T,x,m]$. Recall that $B \cdot U \cdot h(x)$ is the observed winning bid.

Now, $k(\cdot)$ is identified by the relation

$$\text{(B5)} \qquad \frac{h(x)}{k(m)} = \frac{h(x')}{k(m')}$$

based on the identification of $h(\cdot)$ and the normalization $k(m_0) = 1$ for an arbitrary m_0.

Relative profits are identified by considering different values of x and m that generate N and $N' \neq N$ with the same probability. For every (N,T,x,m), there exists (N',T,x',m') for which $\Pr(N \geq n|T,x,m) = \Pr(N' \geq n|T,x',m')$. Here, I reuse notation; x' and m' are different from the first part of this section. Thus, the entry condition

$$\text{(B6)} \qquad E[\pi_N|N,T] \cdot \frac{h(x)}{k(m)} = E[\pi_{N'}|N',t] \cdot \frac{h(x')}{k(m')}$$

can be used to solve for $E[\pi_N|N,T]/E[\pi_{N'}|N',T]$, as $h(\cdot)$ and $k(\cdot)$ are identified. Analogously, relative profits $E[\pi_N|N,T]/E[\pi_N|N,T']$ are identified.

It is straightforward to extend these identification results to setting in which sellers observe U prior to entry. This is possible because the distribution of U conditional on (T,x,m) will be equal to the distribution of U conditional on (T,x',m'). ∎

C. *Proof of Proposition* 6

The ratio of expected profits for n and n' conditional on T is given by

$$\text{(B7)} \qquad R = \frac{E[\pi_n|n,T]}{E[\pi_{n'}|n',T]} = \frac{\frac{1}{n}\left(E[B|n,T] - E[C|n,T]\right)}{\frac{1}{n'}\left(E[B|n',T] - E[C|n',T]\right)}.$$

As shown by Proposition 5, R is identified. Let $n' = n+1$.

When the selection mechanism is a symmetric auction, $E[B|n,T] = E[C_{2:n}|T]$ and $E[C|n,T] = E[C_{1:n}|T]$. From here on I suppress notation indicating that costs are conditional on T. From Lemma (3), we have $E[C_{1:n}] = (1/(n+1))E[C_{2:(n+1)}] + (n/(n+1))E[C_{1:(n+1)}]$. Plugging this into the equation for R obtains

$$R\left(E[C_{2:(n+1)}] - E[C_{1:(n+1)}]\right) = E[C_{2:n}] - \frac{1}{n+1}E[C_{2:(n+1)}] - \frac{n}{n+1}E[C_{1:(n+1)}],$$

$$\left(R + \frac{n}{n+1}\right)E[C_{1:(n+1)}] = E[C_{2:n}] - \left(R + \frac{1}{n+1}\right)E[C_{2:(n+1)}].$$

Here, $E[C_{2:(n+1)}]$ and $E[C_{2:(n+1)}]$ are equivalent to $E[B|N = n]$ and $E[B|N = n+1]$, both of which are identified by Proposition 5. Therefore, $E[C_{1:(n+1)}]$ is identified. Note, $E[C_{1:n}]$ is obtained from equation (B7). These are sufficient to identify seller surplus.

Once seller surplus is identified, the distribution of ε is identified from equation (B3) by using variation in $h(\cdot)$ or $k(\cdot)$. ∎

D. *Proof of Proposition* 7

For each observed sequential value of $N \in \{\underline{N}, \ldots, \bar{N}\}$, the first-order and second-order statistics of N draws from the cost distribution are identified (see Propositions 5 and 6). Using the recursive relationship of order statistics shown in Lemma 3, these are equivalent to identifying the first $\bar{N} - \underline{N} + 2$ expected order statistics from $\bar{N}$ draws of C. ∎

E. *Alternative Identification and a Note on Independent Private Values*

The model in this paper allows for endogenous entry. To separate the private cost distribution from unobservable heterogeneity, I make use of entry cost shifter m to generate exogenous variation in N. Without endogenous entry, there is no entry cost shifter and the entry equation cannot be used to identify the model. When N is purely exogenous, variation in N can still be used to separately identify the private and common cost distributions.

PROPOSITION 8: *First-price, symmetric auctions with unobserved heterogeneity and conditionally independent private values are identified when only the winning bid and the number of bidders is observed. In particular, seller surplus and the first* $(\bar{N} - \underline{N} + 2)$ *expected order statistics of* $\bar{N}$ *draws from F are identified. Identification is obtained without modeling entry as long as there is no selection on unobservables.*

The result can be generalized to auction settings that are independent of the duration-setting problem. Thus, this identification result may prove practical. With only the winning bid and variation in the number of bidders, researchers can estimate a model with unobserved heterogeneity, which is far less restrictive than the assumption of independent private values (IPV) that is common in such settings with limited data.

This implies that in any setting where estimation is motivated by IPV, one could also estimate a conditional independent private-values model with unobserved heterogeneity. The econometrician may expect that unobserved heterogeneity is present, and this provides a theoretical background to test for its importance. In online Appendix G, I detail a computational approach that greatly speeds up the maximum likelihood estimation of these models.

PROOF:

The ratio of second-order statistics is identified by comparing winning bids $B \cdot U \cdot h(x)$ for different values of n and n':

$$\text{(B8)} \qquad \frac{E[B|n,T,x,m] \cdot E[U|n,T,x,m] \cdot h(x)}{E[B|n',T,x,m] \cdot E[U|n',T,x,m] \cdot h(x)} = \frac{E[C_{2:n}|T,x,m]}{E[C_{2:n'}|T,x,m]},$$

where $E[U|n,T,x,m] = E[U|n',T,x,m] = E[U|T,x,m]$ by independence and no selection on unobservables.

From here on, C_i and U may be conditional on (T,x,m). I suppress this in my notation for clarity. Normalizing $E[U] = 1$ pins down the scale of $E[C_{2:n}]$.[38]

Suppose that another $(\hat{F},\hat{G})$ rationalizes the data. Then

$$B_n \cdot U \overset{d}{=} \hat{B}_n \cdot \hat{U}, \tag{B9}$$

$$B_{n'} \cdot U \overset{d}{=} \hat{B}_{n'} \cdot \hat{U}. \tag{B10}$$

Construct $\tilde{B}_{n'}$, $\tilde{\hat{B}}_{n'}$, $\tilde{U}$, and $\tilde{\hat{U}}$ as random variables that are independent of and have the same conditional distributions as their tilde-free counterparts. Then it follows that

$$(B_n \cdot U) \cdot \left(\tilde{\hat{B}}_{n'} \cdot \tilde{\hat{U}}\right) \overset{d}{=} \left(\hat{B}_n \cdot \hat{U}\right) \cdot \left(\tilde{B}_{n'} \cdot \tilde{U}\right) \tag{B11}$$

$$\Rightarrow \quad B_n \cdot \tilde{\hat{B}}_{n'} \overset{d}{=} \hat{B}_n \cdot \tilde{B}_{n'}. \tag{B12}$$

From this relation, we may take the minimum on both sides. By independence and Lemma 2, I obtain

$$E\left[C_{1:(n-1)}\right] \cdot E\left[\hat{C}_{1:(n'-1)}\right] = E\left[\hat{C}_{1:(n-1)}\right] \cdot E\left[C_{1:(n'-1)}\right], \tag{B13}$$

$$\frac{E\left[C_{1:(n-1)}\right]}{E\left[C_{1:(n'-1)}\right]} = \frac{E\left[\hat{C}_{1:(n-1)}\right]}{E\left[\hat{C}_{1:(n'-1)}\right]}. \tag{B14}$$

That is, any $(\hat{F},\hat{G})$ that rationalizes the data has a private cost distribution with the same ratio of first-order statistics.

Finally, using the fact that $E\left[C_{1:(n-1)}\right] = (1/n)E[C_{2:n}] + ((n-1)/n)E[C_{1:n}]$, we can link together these ratios when $n' = n+1$:

$$\frac{\frac{1}{n}E[C_{2:n}] + \frac{n-1}{n}E[C_{1:n}]}{E[C_{1:n}]} = \frac{\frac{1}{n}E\left[\hat{C}_{2:n}\right] + \frac{n-1}{n}E\left[\hat{C}_{1:n}\right]}{E\left[\hat{C}_{1:n}\right]} \tag{B15}$$

$$\Rightarrow \quad \frac{E[C_{2:n}]}{E[C_{1:n}]} = \frac{E\left[\hat{C}_{2:n}\right]}{E\left[\hat{C}_{1:n}\right]}. \tag{B16}$$

As we have identified $E[C_{2:n}]$, $E[C_{1:n}]$ and $E\left[C_{1:(n-1)}\right]$ are also identified. Therefore, $\hat{F}$ and F have the same ratio of second-order statistics. With sequential values of $N \in \{\underline{N},\ldots,\bar{N}\}$, we can iterate forward from the from the identified first-order and second-order statistics using the recursive relationship between order statistics from Lemma 3. Therefore, $\hat{F}$ and F and identical up to the first $\bar{N} - \underline{N} + 2$ expected order statistics from $\bar{N}$ draws of C. ∎

[38] Note that, in practice, we may normalize $E[U|T,x,m] = 1$ for all (T,x,m) realizations. How the mean of $C_{2:n} \cdot U$ changes is captured in changes to the mean of C.

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