# Competition and subsidies in the deregulated US local telephone industry

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*The 1996 Telecommunications Act opened the monopolistic US local telephone industry to new entrants. However, substantial entry costs have prevented some markets from becoming competitive. We study various subsidy policies designed to encourage entry. We estimate a dynamic entry game using data on potential and actual entrants, allowing for heterogeneous option values of waiting. We find that subsidies to smaller markets are more cost effective in reducing monopoly markets, but subsidies to only lower-cost firms are less cost effective than a nondiscriminatory policy. Subsidies in only early periods reduce the option value of waiting and accelerate the arrival of competition.*

## 1. Introduction

■ Many telecommunication services have been deregulated in the last few decades, including the US local telephone industry. Before deregulation, services were provided by regulated monopolists, competitive entry was forbidden, and prices were set by federal and state authorities according to cost-plus, rate-of-return regulation guidelines (Hausman and Taylor, 2012). After deregulation, the markets were opened to competition and many of the pricing regulations were phased out. Consequently, on the one hand, smaller, competitive telecommunications companies were allowed to enter, even using the incumbents' unbundled network and facilities; on the other hand, incumbents enjoyed newfound freedom and greater market power if new entrants did not arrive. Given the substantial cost of entry in the telecommunications industry, ensuring a competitive market structure after deregulation is an ongoing concern for policy makers.

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In general, entry costs can hinder competition in a deregulated market. When the costs of entry are high enough, deregulation itself may not be sufficient to attract entry. The monopoly market structure from before the deregulation might remain, only now the incumbent is unregulated and may exploit its market power. To address this issue, one policy remedy could be to subsidize new firms' entry costs. Such subsidy policies have been adopted in many industries.[1] This opens up the question of how to design such subsidies as a function of the economic environment. For instance, different potential entrants may face different levels of entry costs. Also, markets differ in size, which affects post entry profit. How important is it to consider such firm heterogeneity and market heterogeneity in the design of the subsidy policy? In addition, although a subsidy lowers a firm's entry cost today, it also changes this firm's belief about the future competition level in the market it considers entering. How important is this competition effect for the design of a subsidy policy?

In this article, we address the above questions by estimating a dynamic oligopoly game of entry into the US local telephone industry. Prior to 1996, local markets were served by regulated monopolists, the so-called Incumbent Local Exchange Carriers (ILECs), who were mostly Baby Bell companies. After the 1996 Telecommunications Act (henceforth, the Act), the markets were opened to new entrants, referred to as Competitive Local Exchange Carriers (CLECs). In this study, we focus on facilities-based CLECs, which build their own fiber-optic networks and digital switches[2] and are deemed by industry experts to represent true competition to ILECs (Crandell, 2001, 2005; Economides, 1999).[3]

We use a comprehensive panel data set, which records all facilities-based CLECs' entry decisions into local telephone markets between 1998 and 2002. With this data set, we observe the identity of CLECs providing local telephone services to each local market each year. We also observe the set of CLECs with certification to enter in each state each year. In this industry, a CLEC needs to obtain certification from a state to operate in a market within the state. After receiving state certification, a CLEC may wait years to actually enter. Based on this industry feature, we define potential entrants into a local market as CLECs with certification from the respective state. With information on the identity of potential and actual entrants, we are able to observe how long a potential entrant waits to enter a market and several crucial firm-level attributes associated with the cost of entry.

We set up a dynamic oligopoly game and incorporate the timing of entry and firm heterogeneity in the game. In our model, a potential entrant is a long-run player that decides whether to enter or wait in each period. When making this decision, the potential entrant compares the value of entry, minus entry costs, to the value of waiting. This is in contrast to most other entry studies, in which a firm either enters or perishes and the value of waiting is set to zero.[4] Moreover, we allow potential entrants to be heterogeneous in entry costs. For example, a more experienced potential entrant may face lower entry costs. To estimate our model, we follow the recent development in two-step estimation strategies for dynamic oligopoly entry games. That is, we first obtain the conditional choice probabilities at each state from the data. We then match the empirical conditional choice probability with its counterpart predicted by the model.

The estimation of the model gives results that are consistent with basic economic intuition. For instance, we find that a CLEC's post entry profit is decreasing in competition and increasing

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[1] This practice is especially common in service industries, where the service is considered "essential for the basic well-being of consumers." For example, from 1978 to the present day, federal government programs have subsidized entry of dentists, physicians, and mental health specialists into geographic areas designated as Health Professional Shortage Areas (Dunne, Klimek, Roberts, and Xu, 2013).

[2] Facilities-based CLECs also lease some networks from ILECs to locations not served by the CLECs' own networks; and, more importantly, they need to interconnect with ILECs' networks to exchange voice and data traffic.

[3] CLECs that resell ILECs' service or CLECs that rent ILECs' networks and provide value-added services only yield thin profit margins. They are considered as unsustainable (Crandall, 2001).

[4] One exception is Fan (2006), who makes the distinction between a short-run and long-run player in a dynamic entry game.

in market size, as measured by the overall number of business establishments in a market. This finding is in line with the conventional wisdom that a larger market is necessary to support more competitors (e.g., Bresnahan and Reiss, 1991). In addition, we find that entry costs play an important role in determining whether a potential entrant enters a local market. Overall, the estimated model fits the data rather well—the predicted numbers of monopoly, duopoly, triopoly, and more competitive markets are similar to those observed in the data.

With the estimated model parameters, we then study various subsidy policies designed to encourage entry into monopolistic markets. We compare subsidy policies that would cost the same in terms of the total subsidy spent and examine which policy leads to fewer monopoly markets. Through counterfactual analyses, we find that a subsidy amounting to 5% of the average entry cost reduces the fraction of monopoly markets to 32% by the end of 1998 (compared to 52% in the data), and to 7% by the end of 2001 (compared to 23% in the data). Doubling such a subsidy would reduce this fraction to 14% by the end of 1998 and to 1% by the end of 2001. However, we also show that such subsidies can be more effective at reducing monopolies if offered only in smaller markets. Although applied to small markets only, such a subsidy policy in general also leads to a reduction in the number of customers stuck with monopoly markets, as measured by the sum of market size over all monopoly markets. This suggests that subsidy policies should exploit market heterogeneity. A subsidy policy that exploits firm heterogeneity in entry costs, however, is not as effective at reducing monopoly markets as a nondiscriminatory policy. This is because of the following trade-off: a subsidy to low-cost firms is more conducive to entry than the same subsidy per firm to high-cost firms; however, by applying to fewer potential entrants, it may also lead to less overall entry. According to our estimation, the latter effect dominates the former.

More importantly, we quantify the influence of the option value of waiting on how quickly a market becomes competitive. We find that subsidies intended to reduce the option value of waiting, as expected, change the timing of firms' entry behavior. Specifically, a 10% subsidy that is offered only in 1998 reduces the number of monopoly markets to 9% by the end of 1998, as opposed to 14% when such a subsidy is applied in all years. This is because of a direct effect of changing the timing of the subsidy and an indirect competition effect that potential entrants anticipate less entry in the future due to the lack of the subsidy in the future. The direct effect reduces the option value of waiting, whereas the indirect competition effect increases the expected value of entry. Further investigation through decomposition exercises indicates that both effects contribute to the overall results, but the indirect competition effect is slightly larger.

Our counterfactual exercises focus on the reduction of monopolistic local markets. We do not conduct a full welfare analysis in this article. Measuring welfare would require detailed data describing demand. To the best of our knowledge, the data that would allow us to estimate the demand for local telephone services is not available at the national level. Although we are unable to gauge the total welfare gain of the counterfactual subsidies, previous work in the literature has demonstrated a substantial gain associated with increased competition. Increased competitiveness of a market typically leads to lower prices (e.g., Bresnahan and Reiss, 1991; Nevo, 2000; Basker, 2005) and even better quality or wider variety (e.g., Mazzeo, 2003; Economides, Seim, and Viard, 2008; Matsa, 2011; Fan, 2013).

This article contributes to several strands of the literature. First, it is related to the literature on dynamic entry game estimation. Several studies have made significant progress in this area since Hotz and Miller (1993) proposed a two-step estimation strategy that does not require solving for equilibrium in a complex dynamic model (Aguirregabiria and Mira, 2007; Bajari, Benkard, and Levin, 2007; Pakes, Ostrovsky, and Berry, 2007; Pesendorfer and Schmidt-Dengler, 2008). Nonetheless, due to lack of data, there are some limitations to applications utilizing this approach (Ryan, 2012; Collard-Wexler, 2013; Dunne, Klimek, Roberts, and Xu, 2013). For example, researchers are usually unable to observe the identities of potential entrants and therefore have to assume that potential entrants are *ex ante* homogeneous, short-run players. The players in these

dynamic games face the short-run decision of either entering or perishing.[5] In most industries, however, the decision that a potential entrant faces is to enter or to wait. By identifying potential entrants for a market, we are able to incorporate more information from entry timing of these firms to recover the distribution of entry costs. Specifically, the players in our game face a long-run decision of entering or waiting. We allow them to take into account the option value of delaying their entry.[6] When we compare our model to a model where the identity of potential entrants is ignored, we find that our model fits the data better and, more importantly, generates different effects of counterfactual subsidies.

This article is also related to the literature on competition in the local telephone markets. Within this body of literature, Greenstein and Mazzeo (2006) study CLEC entry decisions into differentiated categories using a static entry model. In another study, Economides, Seim, and Viard (2008) measure the consumer welfare effects of the increase in local telephone competition after the Act, using household-level data from New York state. Finally, Goldfarb and Xiao (2011) emphasize the importance of heterogeneity in managerial ability, which they back out from entry behavior.[7] Our article complements these studies by emphasizing the importance of market heterogeneity and the competition effect of entry in the design of subsidy policies.

This article proceeds as follows. Section 2 provides relevant background information on the US telephone market. Section 3 introduces our data set. Sections 4 and 5 describe in detail our model and estimation strategy, respectively. Section 6 reports our estimation results, and Section 7 presents the results from our counterfactual experiments. Section 8 concludes.

## 2. Industry background

■ Access to telephone service is widely recognized as a fundamental part of public infrastructure. Increased access to telecommunication services creates positive network externalities for individual consumers and enhances democratic participation and public safety. Equal access to such infrastructure has been considered by regulators as essential in narrowing socioeconomic gaps across different regions.

The Act marked the end of a long, monopolistic era in the US local telephone industry. Before the Act, ILECs enjoyed regulated monopoly power for decades on the grounds of substantial economies of scale. Since the 1990s, however, dramatic reduction in the cost of fiber-optic technology has made competitive entry possible. The Act's primary goal was to promote competitive entry. Specifically, Section 253(a) of the Act eliminates a state's authority to erect legal entry barriers in local-exchange markets. More importantly, Section 2.1 mandates that ILECs must offer interconnections and lease part or all of their network facilities to any new entrant at "rates, terms, and conditions that are just, reasonable, and nondiscriminatory."[8]

□ **Local telephone industry after the act.** After the Act, ILECs remained major players in the local telephone industry, but CLECs started to erode the ILECs' market power in some local markets. These CLECs come from various backgrounds. Some CLECs are ILECs in other markets (e.g., CTC Exchange Services, an ILEC with a history of over 100 years, started a CLEC division in 2000), some are long-distance carriers trying to enter the local exchange market (e.g., AT&T obtained certification from every US state right after the Act), and others are *de novo*

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[5] For example, Doraszelski and Satterthwaite (2010) make this assumption explicit: "They (potential entrants) are short lived and base their entry decisions on the net present value of entering today; potential entrants do not take the option value of delaying entry into account."

[6] On this front, our article is connected to the literature on investment and uncertainty. A key insight of this literature is that there is a value of delaying the investment in the presence of investment irreversibility and uncertainty about the future (see Pindyck, 1991, for an overview, and Kellogg, 2014, for a recent empirical study).

[7] Other studies of the US local telephone industry include Ackerberg et al. (2009), Alexander and Feinberg (2004), Mini (2001), and Miravete (2002).

[8] Economides (1999) provides an overview of the Act and its impact on the U.S. telecommunications industry.

entrants catering to a targeted clientele (e.g., PaeTec Communications, founded in 1998, targeted medium- and large-size businesses, government entities, and universities). These CLECs differ substantially in ownership structure, financial resources, and experiences in the local telephone markets.

The pace of new entry after the Act was slower than what policy makers had anticipated back in 1996 (Economides, 1999; Young, Dreazen, and Blumenstein, 2002). Although around 40% of medium-size markets experienced entry by the end of 1998, about 30% of these markets did not have any CLEC operating, even by the end of 2002. One factor presumably contributing to low entry levels is the substantial cost of entry.

□ **Costs of entry.** Facilities-based CLECs must make substantial investments in building facilities such as switching and distribution centers, as well as laying out fiber-optic networks physically connecting these switching and distribution centers to the end users of telephone services. In our data, we observe annual capital expenditures for the majority of the CLECs. Dividing a CLEC's capital expenditure for a given year by the number of cities it entered next year, we get a rough measure of its entry costs per market, which amounts to $6.5 million per market on average. Furthermore, much of the investment has to be made at specific locations, so these assets are not movable (Economides, 1999).

In addition, there are "soft" entry costs (Pindyck, 2005). For example, the costs consumers face in switching from an incumbent to a new entrant, which are especially important in telecommunication industries, may create disadvantages for new entrants. To overcome these disadvantages, new entrants may need to incur substantial advertising costs. Motivated by these facts, we focus on the role of entry costs in shaping CLECs' entry decisions. A measure of total entry costs does not exist in accounting books. However, firms' strategic entry decisions reflect the size and distribution of such costs. We can back them out by combining a model of strategic entry with data on actual entry behavior. With our estimates of entry costs, we can evaluate the effects of different subsidy policies that directly reduce the costs of entry.

□ **State certification.** To identify the set of potential entrants in a local market, we make use of the requirement that CLECs must first obtain certification from state regulators before they can operate in any city within the state. To obtain state certification, a CLEC applicant needs to submit paperwork outlining the services to be offered, detailed construction plans, and an environmental impact statement. Furthermore, the applicant needs to show a certain degree of financial ability to serve. Some states require an applicant to show possession of a certain amount of cash or cash equivalent at the time of the application, whereas others use more complex formulas.[9] Overall, the consensus in the industry is that obtaining state certification is a time-consuming process, and only those with certification are likely to enter in a given year. Any CLEC without a real intent to enter any market in a state will likely not apply for certification. This consensus is also consistent with the data. As we will show in the next section, the average number of state certifications that a CLEC holds is around 10 rather than all 50 states. The rather low number of states thus suggests that the certification process is sufficiently time consuming that only firms with real entry intentions will pursue state certification. On the other hand, the data also show that firms on average wait more than two years from the time of certification to enter a local market, whereas some CLECs never enter any city in a state for which they are approved to enter during the years covered in this study (1998–2002). These data patterns indicate that firms do not wait until they are certain about entry to get state certification. Thus, we identify potential entrants in a local market as the set of CLECs with certification to operate in that state.[10]

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[9] Texas, for example, requires an applicant to show that (i) it has either $100,000 in cash or sufficient cash for startup expenses for the first two years of operation or (ii) it is an established business entity and has shown a profit for two years preceding the application date (Kennedy, 2001).

[10] Although many states give CLEC applicants authority to serve the entire state, a few states require applicants to specify each local area to be served. We deal with this potential caveat by dropping small cities – those with fewer than

In summary, for each local market, a CLEC can take on one of four (mutually-exclusive) roles: a CLEC without state certification is a "potential" potential entrant, a CLEC with state certification becomes a potential entrant, a CLEC in its first year of providing services is a new entrant, and a CLEC providing services from its second year and on is an incumbent.

## 3. Data

■ To obtain our data set, we combine data on CLECs and data on markets to create a panel data set of firms' entry decisions, firm-level characteristics, and market attributes.

□ **The NPRG annual reports on CLECs.** For our CLEC data, we use CLEC annual reports obtained from the New Paradigm Resources Group, Inc. (NPRG). This database contains information on the universe of facilities-based CLECs in the United States between 1998 and 2002.[11] For each CLEC, we observe the state certifications it held in each year. We also observe the cities that each CLEC provided with local telephone services in each year and the exact year when the service started, which we treat as the year the CLEC entered the market. NPRG also reports firm attributes, such as the year the company was founded, the zip code of the headquarters, whether the company is publicly traded or privately held, whether the company is venture capital funded, and whether the company is a subsidiary of a larger telecommunications company.

□ **Market definition, market characteristics, and sample selection.** We combine data on CLECs with data on market characteristics. The locations in the NPRG reports, that is, the cities a CLEC provides services to, are best interpreted as census "places." Therefore, we choose a census place as our market definition and refer to each as a "city" henceforth.

As most of these CLECs catered to business clientele in the early years of the industry (see, e.g., NPRG CLEC Reports, 1999–2003; Alexander and Feinberg, 2004; Greenstein and Mazzeo, 2006), the best proxy of market size is the number of business establishments in a city. To collect data on the number of business establishments for each city, we divide each city into a set of Zip Code Tabulation Areas (ZCTAs) and obtain the number of business establishments within each area from the Census' Zip Code Business Patterns.

Last, we select medium-size cities based on the number of business establishments. We drop 26 US cities, those who had more than 15,000 business establishments in 1997, from our sample because CLECs in these markets may not serve the whole market and thus may not directly compete with other CLECs in the same market.[12] Furthermore, we drop small cities (those with less than 2000 business establishments) from our data. The entry rate into these small cities is extremely low from 1998 to 2002, which suggests that these small cities may not represent realistic entry candidates. That is, a CLEC holding a state certification may not actually be a potential entrant in each small city, which makes it difficult to identify the set of potential entrants for these kinds of cities.[13] After dropping all of the markets that do not fit our criteria, we are left with 398 medium-size cities for our analysis. These cities are listed in web Appendix A.

□ **Summary statistics.** Tables 1 and 2 report the descriptive statistics from our data. Table 1 summarizes the data on firm attributes, which we argue are determinants of a CLEC's

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2000 business establishments – from our analysis because these cities are less likely to be the target areas in the early years of the competitive US local telephone industry.

[11] The NPRG reports are published a year late relative to the year of data collection. The NPRG CLEC annual reports cover 1996 to the present. However, 1998 is the year when NPRG started to report for the universe, instead of a selected sample, of facilities-based CLECs. In 2001, NPRG split facilities-based rural CLECs into another report series, which were published only for the years of 2001 and 2002. Therefore, we are only able to assemble information on the universe of facilities-based CLECs from 1998 to 2002.

[12] Altanta is the smallest city we drop based on this threshold.

[13] If we include these small cities into our analysis, we may have a biased estimate of entry costs, because a CLEC will be considered to be waiting even in markets that it never intends to enter in the first place.

**TABLE 1     Summary Statistics on Firm Attributes**

|  | 1998 | 1999 | 2000 | 2001 | 2002 |
|---|---|---|---|---|---|
| Variables | Mean (Standard Deviation) | | | | |
| Firm-level | | | | | |
|   Privately held | 0.639 | 0.595 | 0.576 | 0.598 | 0.640 |
| | (0.483) | (0.493) | (0.496) | (0.493) | (0.483) |
|   Subsidiary | 0.320 | 0.284 | 0.288 | 0.268 | 0.302 |
| | (0.469) | (0.453) | (0.455) | (0.445) | (0.462) |
|   Financed by venture capital | 0.175 | 0.190 | 0.220 | 0.206 | 0.221 |
| | (0.382) | (0.394) | (0.416) | (0.407) | (0.417) |
|   Firm age[a] | 8.897 | 8.526 | 11.017 | 13.639 | 15.616 |
| | (20.206) | (16.548) | (20.680) | (23.449) | (24.717) |
|   # states with certification | 9.237 | 10.500 | 13.217 | 13.135 | 12.917 |
| | (11.438) | (11.978) | (13.019) | (12.484) | (12.646) |
|   # cities with certification | 100.381 | 111.138 | 128.102 | 122.196 | 123.209 |
| | (108.026) | (112.298) | (116.251) | (114.220) | (117.659) |
| # observations (firm)[b] | 97 | 116 | 118 | 97 | 86 |
| Firm-market-level | | | | | |
|   Home state dummy: if market is | 0.124 | 0.110 | 0.107 | 0.096 | 0.102 |
|     in the same state as firm headquarters | (0.327) | (0.313) | (0.309) | (0.295) | (0.303) |
|   Distance from firm headquarters to market | 0.901 | 0.916 | 0.901 | 0.876 | 0.896 |
|     in 1000 km | (0.680) | (0.693) | (0.688) | (0.675) | (0.682) |
|   # observations (firm-market) | 9737 | 12892 | 15116 | 11853 | 10596 |

[a]The average firm age increases by more than one over the years, reflecting entry into Competitive Local Exchange Carriers (CLEC) business by already established firms.
[b]All CLECs in a given year.

**TABLE 2     Summary Statistics on Market Attributes**

|  | 1998 | 1999 | 2000 | 2001 | 2002 |
|---|---|---|---|---|---|
| Variables | Mean (Standard Deviation) | | | | |
| # business establishments | 3999.565 | 4022.812 | 4021.528 | 4011.382 | 4091.894 |
| | (2437.058) | (2439.516) | (2439.216) | (2427.654) | (2469.448) |
| # incumbents[a] | 1 | 1.819 | 2.661 | 3.015 | 3.231 |
| | (0) | (1.429) | (2.269) | (2.696) | (2.787) |
| # potential entrants | 24.465 | 31.573 | 36.319 | 27.766 | 24.392 |
| | (7.481) | (9.616) | (10.575) | (10.031) | (8.578) |
| # actual new entrants | 0.960 | 1.523 | 1.638 | 0.746 | 0.327 |
| | (1.618) | (1.798) | (2.306) | (1.066) | (0.597) |
| Entry rate (among potential entrants) | 0.044 | 0.056 | 0.051 | 0.030 | 0.018 |
| | (0.079) | (0.072) | (0.076) | (0.050) | (0.037) |
| $\mathbb{1}$(any competition)[b] | 0.417 | 0.688 | 0.729 | 0.711 | 0.726 |
| | (0.494) | (0.464) | (0.445) | (0.454) | (0.447) |
| # observations (market) | | | 398 | | |

[a]At the beginning of each year.
[b]At the end of each year.

entry costs. These attributes include the organizational, financial, and ownership structure of the firm, as well as the age of the firm. We also include two measures of the relationship between a firm and a market it can potentially enter. One is a dummy variable indicating whether the market is in the same state as the firm's headquarters. This variable captures a home state advantage, such as lower costs in passing zoning requirements, dealing with local administration, advertising, and public relations. The other is a measure of the distance (in 1000 kilometers) between a firm's headquarters zip code and the population centroid of a state.

TABLE 3    Summary Statistics on Entry Patterns, Including Waiting Time

|  | Observations | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|---|
| 𝟙(enter any market by 2002) | 187 firms | 0.775 | 0.418 | 0 | 1 |
| # markets entered by 2002 | 187 firms[a] | 11.053 | 18.731 | 0 | 114 |
| 𝟙(enter any market in a given year) | 514 firms/years[b] | 0.578 | 0.494 | 0 | 1 |
| # markets newly entered in a given year | 514 firms/years | 4.021 | 8.708 | 0 | 104 |
| % markets newly entered in a given year[c] | 514 firms/years | 0.050 | 0.107 | 0 | |
| 𝟙(enter the local market by 2002) | 22,192 firms/markets[d] | 0.849 | 0.358 | 0 | 1 |
| Years in waiting if entered the market by 2002[e] | 18,806 firms/markets | 2.056 | 1.081 | 0 | 4 |

[a] All CLECs in our sample.
[b] All CLEC-year combinations in our sample.
[c] The percentage of markets that firms are certified to enter.
[d] The number of unique potential entrant-market combinations that ever appear in the five years of our sample period.
[e] The time between when a firm gets state certification and when it actually enters a local market.

We can see from Table 1 that the CLECs in our sample are generally privately held (on average, 58%–64% across years), with high age variance (the Standard Deviation is about twice the mean). In addition, a small proportion of these firms are subsidiaries of large corporations (on average 27%–32% across years) and partially funded by venture capital (on average, 18%–22% across years). The average number of cities in which a CLEC has state certifications increases gradually from 1998 and peaks in 2000, right after which the telecommunication market suffered a stock market crash. The variation in the number of firms over time also reflects the rapid boom-and-bust pattern in the early years of the telecommunications industry. Overall, the statistics in Table 1, especially the summary statistics on firm attributes, show that the CLECs in our sample are heterogeneous. In the model below, we therefore allow firms to be heterogeneous in entry costs.

Table 2 describes the 398 medium-size cities that we use for our analysis. We can see that the number of business establishments is gradually increasing until 2001, reflecting the ups and downs of the macroeconomy. Note that there is only one incumbent for every market at the beginning of 1998 because only a single ILEC existed in each market at the time of the Act.[14] However, after 1998, the number of incumbents fluctuates up and down because entry and exit are frequent events.[15] A typical city in our sample has a large set of potential entrants but only a few incumbents (including the one ILEC in each market) or new entrants. Furthermore, the summary statistics show that the number of new entrants first increases during our sample period and then drops sharply, again echoing the 2001 crash in the stock market. The entry rate, defined as the number of new entrants divided by the number of potential entrants in a local market, varies from 0.018 to 0.056 across the years in our sample. As the most effective competition usually arrives with the first competitor, we also show summary statistics for the existence of any competition at the end of each year. Specifically, we see that although about 40% of the markets have at least one CLEC competing with the ILEC as early as 1998, about 30% of the markets are still monopolistic even at the end of 2002. Overall, the post-Act landscape is uneven in terms of entry and competition across the 398 markets.

Table 3 describes CLECs' entry patterns, including waiting time. A few patterns here are notable. First, firms do wait. Around 22% of the firms do not enter any market by the end of the sample period, even though they have certification from at least one state. In a given year, only

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[14] Due to the data limitation explained in footnote 10, we treat 1997 (right after the Act) and 1998 (the first year of our data) as one period.

[15] We treat bankruptcy, being acquired by another firm, or simply going out of business as an exit. In the few cases of mergers (fewer than 10 out of approximately 200 CLECs in our time period), we treat the smaller CLEC as the firm exiting from business.

58% of CLECs entered any market. The average number of local markets a CLEC enters in that year is four, accounting for about 5% of the markets that the CLEC is certified to enter. Overall, the average waiting time for a firm to enter a market after obtaining certification is about two years. This average is taken across potential entrant-market combinations conditional on the potential entrant entering the local market by 2002 (so that we observe the waiting time). The unconditional average waiting time is therefore larger than two years. Second, we find considerable variation in the waiting time across firms and the entry rates across local markets, suggesting the existence of firm-level and market-level heterogeneity. The Standard Deviation of the waiting time, reported in the last row of Table 3, is 1.081 years. The Standard Deviation of entry rates in local markets, reported in Table 2, is almost always twice the level of the entry rates across years.

## 4. Model

■ The summary statistics on firm attributes (in Table 1) and waiting time (in Table 3) indicate that potential entrants are heterogeneous and that some of them wait for several years before they actually enter a local market. To capture these aspects of the data, we use a model based on Pakes, Ostrovsky, and Berry (2007) (henceforth, POB) and add two new features. First, we assume that potential entrants are long-run players. Under this assumption, in each period, a potential entrant may choose to enter or to wait, with a potentially positive value of waiting. Second, we allow different types of potential entrants to face different entry cost distributions.

At the beginning of each year, a firm decides whether to obtain certification from a state and thereby become a potential entrant in that state's local markets, if it has not already done so. Then, entry costs for this potential entrant in each local market are realized.[16] Afterward, the potential entrant decides whether to enter a local market. Therefore, in deciding whether to obtain certification, a firm considers (i) the incumbents and potential entrants in each local market in the state, (ii) other characteristics of each local market, (iii) the pool of firms who have not obtained certification from the state but might be interested in doing so, and (iv) its own expected entry costs. Information on (i) to (iii) affects the expectation of the firm regarding the aggregate value of being eligible to enter the local markets of a state, which is state year specific. In contrast, expected entry costs in (iv) are firm specific. Once we control for the first three using state-year fixed effects, a firm's decision to obtain certification reveals its type in terms of entry costs. As the focus of this article is a potential entrant's entry decision rather than a firm's decision to become a potential entrant, we use the following simple Logit model to explain the decision to become a potential entrant and infer firms' entry cost types.

□ **A firm's decision to become a potential entrant.** The Logit model of a firm's (a "potential" potential entrant's) decision to become a potential entrant is specified as follows:

$$\Pr\left(certification_{fst}|\mathbf{z}_f, \mathbf{d}_{fs}\right) = \frac{\exp\left(\xi_{st} + \varphi_1\mathbf{z}_f + \varphi_2\mathbf{d}_{fs}\right)}{1 + \exp\left(\xi_{st} + \varphi_1\mathbf{z}_f + \varphi_2\mathbf{d}_{fs}\right)}, \quad (1)$$

where the state-year fixed effect, $\xi_{st}$, captures the information in (i) to (iii), and $\mathbf{z}_f$ and $\mathbf{d}_{fs}$ represent firm and firm-state characteristics, respectively, that affect the entry cost (i.e., the covariates affecting (iv) above). Specifically, $\mathbf{z}_f$ includes whether a firm is privately held, whether it is a subsidiary, whether it is financed by venture capital, and its age in 1998; $\mathbf{d}_{fs}$ includes whether the market is in the same state as the firm's headquarters (home state dummy), the distance between the firm's headquarters and the population centroid of the state, and that distance squared. We use these three firm-state characteristics to capture the idea that firms may face different entry costs in different geographies.

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[16] In other words, a firm's decision to obtain a state certification is assumed to be exogenous to the entry decision in a local market in the sense that it is independent of the shock to the cost of entering the local market.

As the firm characteristics $\mathbf{z}_f$ and $\mathbf{d}_{fs}$ affect the entry cost, with equation (1) estimated, we use the estimated $\varphi$ to represent the multiple dimensions of firm-level heterogeneity that affect entry costs with a single index. In other words, $\hat{\varphi}_1 \mathbf{z}_f + \hat{\varphi}_2 \mathbf{d}_{fs}$ is a scalar that denotes firm $f$'s type in state $s$. To restrict the dimensionality of the state space, we also discretize firms' types. In particular, we let $\hat{\varphi}_1 \mathbf{z}_f + \hat{\varphi}_2 \mathbf{d}_{fs}$ determine whether a firm is of type-1 or of type-2. We explain the discretization in detail in Section 5.

□ **A potential entrant's decision to enter a local market.** After obtaining state certification, a firm becomes a potential entrant and decides whether to enter a market within the state in each period. As local telephone services are rather homogeneous, market size and competition are the main driving factors of post entry profits. Therefore, we assume that post entry profits are identical across firms within a market and otherwise depend only on the size of the market and the number of incumbents.[17] Let $m_{ct}$ be the market size and $n_{ct}$ be the number of incumbents in city $c$ and year $t$. We assume that the one-period profit function has the following parametric form:

$$\pi (m_{ct}, n_{ct}) = e^{\alpha m_{ct} + \gamma n_{ct}}, \tag{2}$$

where $\alpha$ (the market-size effect) and $\gamma$ (the competition effect) are parameters to be estimated. Note that the exponential function form ensures that profits are always positive.[18]

At the beginning of each period, a potential entrant observes its entry cost. The realized entry cost, which is independently distributed across firms, markets, and time, is a potential entrant's private information. This distribution of the entry cost, which is public information, depends on a potential entrant's type. Given that firm attributes are observed by all firms, the number of potential entrants of each type in a city is common knowledge. To summarize, at the beginning of each period, a potential entrant to a market observes the number of potential entrants of each type $(T_{1ct}, T_{2ct})$ as well as the market conditions $(m_{ct}, n_{ct})$. These are the relevant state variables for firms' decisions. The market size, $m_{ct}$, evolves exogenously according to a first-order Markov process. The number of incumbents, $n_{ct}$, is endogenous: its transition follows $n_{ct+1} = n_{ct} +$ (# new entrants)$_{1ct} +$ (# new entrants)$_{2ct} -$ (# exited incumbents)$_{ct}$. The transition of the number of potential entrants is determined by $T_{\tau ct+1} = T_{\tau ct} +$ (# new potential entrants)$_{\tau ct} -$ (# exited potential entrants)$_{\tau ct} -$ (# new entrants)$_{\tau ct}$ for $\tau = 1, 2$. New potential entrants in market $c$ in year $t$ are CLECs who in year $t$ got certification in the state in which market $c$ lies. As explained, we assume that (# new potential entrants)$_{\tau ct}$ is exogenous and i.i.d. across cities and years. Note that sometimes a CLEC exits the industry as a whole and ceases to be a potential entrant in any market. That is why we need to consider (# exited potential entrants)$_{\tau ct}$ in the transition of the number of potential entrants. For notational simplicity, we suppress subscripts $c$ and $t$ for the remainder of this section. In addition, from now on, whenever it is not obvious what we mean by "state," we use the phrase "geographic state" for a US state such as California, and the word "state" for a state in the model.

If a potential entrant decides to enter a market, we assume it will start to earn profits in the next period after paying an up-front cost of entry in the current period. The value of entry is therefore the expected value of being an incumbent in the next period. Let $V^I(m, n, T_1, T_2)$ be the value of an incumbent at state $(m, n, T_1, T_2)$. Then,

$$V^I (m, n, T_1, T_2) = \pi (m, n) + \delta E_{(m', n', T_1', T_2')|(m, n, T_1, T_2)} V^I \left( m', n', T_1', T_2' \right), \tag{3}$$

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[17] The assumption of homogeneity in post entry profits across firms in a market is necessary for identification. With only entry data, we cannot determine whether different entry timing across firms reflects the heterogeneity in post entry profits or the heterogeneity in entry costs. Given that local telephone services are rather homogeneous, we have decided to allow for the latter heterogeneity while assuming that post entry profits are identical.

[18] This exponential functional form also allows for nonlinear effects of the market size and the number of incumbents in the profit function. For example, when $\gamma$, which captures the competition effect, is negative, this functional form allows the marginal effect of an additional competitor on profit to decrease with the number of competitors.

where $\delta$ is the discount factor and $E_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2)}$ is the expectation of the state in the next period $(m', n', T'_1, T'_2)$, conditional on the current state $(m, n, T_1, T_2)$.

Note that an incumbent in such a dynamic game typically also decides whether to continue operating at the end of each period. We choose not to endogenize this decision for two reasons. First, in our data, an incumbent always stays in the local market until the CLEC exits as a whole, which is consistent with the observation that the variable costs of maintaining operations are low. If exit were a firm-level endogenous decision, we could not treat the firm's entry decisions into local markets as independent across markets. This would dramatically increase the state space of our dynamic problem (and hence our data requirements). Second, during our sample period, firm exits appear to be largely due to exogenous macroeconomic shocks.[19] We thus assume that a firm exits as a whole exogenously and that all firms have the same expected probability of exit, denoted by $p^x$. Note that $p^x$ is common knowledge among all firms. Hence, $\delta$ in equation (3) is in fact the discount factor adjusted for the expected probability of exit: $\delta = \beta(1 - p^x)$, where $\beta$ is the standard discount factor.

A potential entrant decides whether to enter by comparing the value of waiting with the value of entry net of entry costs. As explained, the value of entry is the expected value of being an incumbent in the next period, that is, $\delta E^e_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} V^I(m', n', T'_1, T'_2)$, where $E^e_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)}$ is a type-$\tau$ potential entrant's expectation regarding future states conditional on itself entering.[20] The value of waiting is the expected value of being a potential entrant in the next period. Let $V^E(m, n, T_1, T_2, \tau, \zeta)$ be the value of a potential entrant of type $\tau$ with entry costs $\zeta$. Then, the value of waiting is $\delta E^w_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} E_{\zeta'|\tau} V^E(m', n', T'_1, T'_2, \tau, \zeta')$, where $E^w_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)}$ is a type-$\tau$ potential entrant's expectation regarding future states conditional on the firm itself waiting at state $(m, n, T_1, T_2)$, and $E_{\zeta'|\tau}$ is the expectation of its entry cost in the next period. A potential entrant compares the value of entry net of entry costs with the value of waiting and decides whether to enter. Thus, the value of a potential entrant satisfies the following equation:

$$V^E(m, n, T_1, T_2, \tau, \zeta) = \max \left\{ \delta E^e_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} V^I(m', n', T'_1, T'_2) - \zeta, \right.$$
$$\left. \delta E^w_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} E_{\zeta'|\tau} V^E(m', n', T'_1, T'_2, \tau, \zeta') \right\}. \quad (4)$$

Because a firm may also exit as a whole with probability $p^x$, the same discount factor $\delta$ for the incumbent is used.

A potential entrant decides to enter if the value of entry net of entry costs is larger than the value of waiting. In other words, the probability of entry for a type-$\tau$ potential entrant is

$$p^e(m, n, T_1, T_2, \tau)$$
$$= \Pr_{\tau} \left( \zeta < \delta E^e_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} V^I(m', n', T'_1, T'_2) \right.$$
$$\left. - \delta E^w_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)} E_{\zeta'|\tau} V^E(m', n', T'_1, T'_2, \tau, \zeta') \right). \quad (5)$$

We assume that a potential entrant's entry cost, $\zeta$, follows a gamma distribution with mean $\mu_1$ for type-1 firms and mean $\mu_2$ for type-2 firms. As usual in discrete choice models, we can

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[19] When we regress a firm exit dummy on firm attributes and year dummies using a linear probability model, we find that the estimated coefficients of firm attributes are small and statistically insignificant, whereas year dummies play an important role in explaining variation in exit. This finding suggests that firm exit is indeed driven by macroeconomic shocks rather than inherent firm-level heterogeneity.

[20] The expectation is type specific for two reasons: first, conditional on its own action, a type-1 potential entrant's perception regarding the number of incumbents in the next period depends on its belief about how many out of $(T1 - 1, T2)$ potential entrants will enter, whereas a type-2's perception hinges on how many out of $(T1, T2 - 1)$ potential entrants will enter; second, the same argument about type dependence also holds for a potential entrant's perception of the number of potential entrants in the next period.

only identify model parameters up to a scale. We therefore normalize the variance of the entry cost to be 1.

Following the literature on dynamic games of oligopoly competition, we assume that the data come from a Markov perfect equilibrium of our model. An equilibrium is a triple of policy and value functions $(p^e, V^I, V^E)$ such that for any potential entrant, (i) given that other potential entrants follow the policy function, $p^e$, the value functions $V^I$ and $V^E$ are the fixed point of the Bellman equations (3) and (4), and (ii) given that other potential entrants follow the policy function, $p^e$, and the value functions $V^I$ and $V^E$, $p^e$ satisfies equation (5). The expectations in these equations are formed based on a potential entrant's beliefs, which coincide with the equilibrium policies at the equilibrium.

## 5. Estimation

■ The estimation process is carried out in two main steps. In the first step, we classify each potential entrant by its type. To this end, we estimate $(\varphi_1, \varphi_2)$ in the Logit regression from equation (1). We then compute $\hat{\varphi}_1 z_f + \hat{\varphi}_2 \mathbf{d}_{fs}$ for each firm-state and divide the potential entrants into two groups: a firm $f$ is of type-1 in geographic state $s$ if and only if $\hat{\varphi}_1 z_f + \hat{\varphi}_2 \mathbf{d}_{fs}$ is below the median of all firm-states; otherwise, this firm is of type-2 in geographic state $s$. We expect type-1 potential entrants to have higher entry costs on average than type-2 potential entrants. We do not impose this restriction (i.e., the mean of the entry cost for type-1 potential entrants ($\mu_1$) be larger than the mean for type-2 ($\mu_2$)) in our estimation. However, as we will show, the estimation results confirm the expected ranking.

In the second step of our estimation, we estimate the parameters in the profit function, $(\alpha, \gamma)$, and the parameters in the entry costs distributions, $(\mu_1, \mu_2)$. As mentioned, the discount factor in the model is adjusted by the probability of exiting: $\delta = \beta(1 - p^x)$, where $\beta$ is the standard discount factor and $p^x$ is the expected exit probability. We estimate the model using different values for $\beta$ and study the robustness of our results with respect to these values. We set the mean exit probability $p^x$ to be the empirical average exit probability at the firm level, which is 23.9% from 1999 to 2002. The estimation of $(\alpha, \gamma, \mu_1, \mu_2)$ follows the procedure in POB with one modification: we need to consistently estimate the value of waiting as well as the value of entry.

To estimate the above parameters, it is convenient to rewrite equation (3) in vector form, comparable to the procedure in POB. The state in this model is a quadruple $(m, n, T_1, T_2)$. We denote the $i$th state by $(m_i, n_i, T_{1i}, T_{2i})$. With a slight abuse of notation, we let $V^I(\alpha, \gamma)$ be the vector with $V^I(m_i, n_i, T_{1i}, T_{2i})$ as its $i$th element. Similarly, the $i$th element of the vector $\pi(\alpha, \gamma)$ is $\pi(m_i, n_i)$. Using this notation, we can rewrite equation (3) in vector form:

$$V^I(\alpha, \gamma) = \pi(\alpha, \gamma) + \delta M V^I(\alpha, \gamma), \qquad (6)$$

where $M$ is the transition probability matrix, that is, its $ij$-element is the transition probability from state $(m_i, n_i, T_{1i}, T_{2i})$ to $(m_j, n_j, T_{1j}, T_{2j})$. This matrix $M$ is estimated directly from data.

To rewrite equation (4) in vector form, we define vectors $V_{e1}(\alpha, \gamma)$ and $V_{e2}(\alpha, \gamma)$ as the values of entry for a type-1 and type-2 potential entrant, respectively. Their $i$th elements are the expected value of being an incumbent in the next period, that is, $E^e_{(m', n', T_1', T_2')|(m_i, n_i, T_{1i}, T_{2i}, \tau)} V^I(m', n', T_1', T_2')$ for $\tau = 1$ and 2, respectively. In other words,

$$V_{e\tau}(\alpha, \gamma) = M^e_\tau V^I(\alpha, \gamma), \qquad (7)$$

where $M^e_\tau$ is a matrix whose $ij$-element is the transition probability from $(m_i, n_i, T_{1i}, T_{2i})$ to $(m_j, n_j, T_{1j}, T_{2j})$ conditional on a type-$\tau$ potential entrant entering.

Similarly, we define the vector $V_{w\tau}(\alpha, \gamma, \mu_\tau)$ as the value of waiting for a type-$\tau$ potential entrant, whose $i$th element is the expected value of being a potential entrant in the next period, that is, $E^w_{(m', n', T_1', T_2')|(m_i, n_i, T_{1i}, T_{2i}, \tau)} E_{\zeta'|\tau} V^E(m', n', T_1', T_2', \tau, \zeta')$. We also define the vector $p^e_\tau$ as the probability of entry for a type-$\tau$ potential entrant analogously. Then, applying the expectation

operator $E^w_{(m',n',T'_1,T'_2)|(m,n,T_1,T_2,\tau)}E_{\zeta|\tau}$ on both sides of equation (4), we have the value of waiting satisfying the following equation in vector form[21]:

$$V_{w\tau}(\alpha,\gamma,\mu_\tau) = M^w_\tau \left\{ p^e_\tau \left( \delta V_{e\tau}(\alpha,\gamma) - E \left[ \zeta | \zeta < \delta V_{e\tau}(\alpha,\gamma) - \delta V_{w\tau}(\alpha,\gamma,\mu_\tau);\mu_\tau \right] \right) \right. \quad (8)$$
$$\left. + (1 - p^e_\tau) \delta V_{w\tau}(\alpha,\gamma,\mu_\tau) \right\},$$

where $M^w_\tau$ is a matrix whose $ij$-element is the transition probability from $(m_i, n_i, T_{1i}, T_{2i})$ to $(m_j, n_j, T_{1j}, T_{2j})$ conditional on a type-$\tau$ potential entrant waiting.

To estimate $V_{e\tau}(\alpha,\gamma)$ and $V_{w\tau}(\alpha,\gamma,\mu_\tau)$, we need consistent estimates of the transition probability matrices $M$, $M^e_\tau$, and $M^w_\tau$. To obtain these estimates, we follow the procedure in POB. That is, we use empirical counterparts of these matrices. See the Appendix for the details on how we obtain $\hat{M}$, $\hat{M}^e_1$, $\hat{M}^e_2$, $\hat{M}^w_1$, and $\hat{M}^w_2$.

With $\hat{M}$, $\hat{M}^e_1$, $\hat{M}^e_2$, $\hat{M}^w_1$, and $\hat{M}^w_2$ estimated, the estimate of the value of entry is given by

$$\hat{V}_{e\tau}(\alpha,\gamma) = \hat{M}^e_\tau \hat{V}^I_\tau(\alpha,\gamma), \qquad (9)$$

where

$$\hat{V}^I_\tau(\alpha,\gamma) = (I - \delta\hat{M})^{-1}\pi(\alpha,\gamma), \qquad (10)$$

and $I$ is the identity matrix. Meanwhile, $\hat{V}_{w\tau}(\alpha,\gamma,\mu_\tau)$ is the fixed point of (8) when $V_{e\tau}(\alpha,\gamma)$ and $M^w_\tau$ are replaced by their empirical counterparts. Note that the RHS of equation (8) is a contraction mapping of $V_{w\tau}(\alpha,\gamma,\mu_\tau)$ because $\zeta$ is assumed to be a log-concave random variable (with a gamma distribution), and it follows that $0 \leq \frac{\partial E(\zeta|\zeta<d)}{\partial d} \leq 1$ (see Heckman and Honoré, 1990).

Having obtained consistent estimates of the values of entry and waiting, we can now get consistent estimates of the probabilities of entry for given parameters. As shown in equation (5), the probability of entry at state $(m_i, n_i, T_{1i}, T_{2i})$ is the probability that the entry costs for a firm are smaller than the difference between the discounted value of entry and the discounted value of waiting at the given state.

We estimate the distribution parameters $(\mu_1, \mu_2)$ and the profit parameters $(\alpha, \gamma)$ using the Generalized Method of Moments. We observe the state of each year-market combination. The model prediction of the probability of entry in this year-market is therefore determined by the element in the entry probability vector $p^e_\tau(\alpha,\gamma,\mu_\tau)$ that corresponds to this state. Its empirical counterpart is the fraction of type-$\tau$ potential entrants in this year-market that enter. The difference between the model prediction and the empirical probability of entry is the prediction error, which we compute for each firm type and year-market. We use the Euclidian norm of the prediction errors as well as the covariances between the prediction errors and the following variables as moment conditions: market size, the total number of potential entrants, the percentage of type-1 potential entrants, and a year 2001 dummy.

Identification of structural parameters $(\alpha, \gamma, \mu_1, \mu_2)$ is similar to that in POB. For example, the market size coefficient, $\alpha$, is identified by how much entry probabilities vary with market size. The competition coefficient, $\gamma$, is identified similarly; that is, by how entry probabilities change with different numbers of incumbents in a local market. The variation in the number of incumbents is affected by the variation in the number of potential entrants, which itself is largely driven by the number of new potential entrants. The year 2001 dummy captures the macroeconomic crash in that year, which presumably shrank the number of potential entrants. Last, the difference in the entry probabilities of type-1 and type-2 potential entrants identifies the difference in entry costs of these two types. Together with the levels of entry probabilities, these differences help us identify the parameters $(\mu_1, \mu_2)$.

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[21] Note that $E_\zeta \max(b - \zeta, a) = \Pr(\zeta < b - a)[b - E(\zeta|\zeta < b - a)] + [1 - \Pr(\zeta < b - a)]a$.

**TABLE 4    Firms' Decisions to Obtain State Certification[a]**

| | (1) OLS | (2) Logit | (3) OLS | (4) Logit |
|---|---|---|---|---|
| $\mathbb{1}$(the CLEC is privately held) | −0.094 | −1.154 | −0.093 | −1.229 |
| | (0.004)*** | (0.057)*** | (0.004)*** | (0.059)*** |
| $\mathbb{1}$(the CLEC is a subsidiary of a firm) | −0.054 | −0.759 | −0.056 | −0.830 |
| | (0.004)*** | (0.062)*** | (0.004)*** | (0.064)*** |
| $\mathbb{1}$(the CLEC is funded by venture capital) | 0.030 | 0.491 | 0.032 | 0.561 |
| | (0.005)*** | (0.070)*** | (0.005)*** | (0.073)*** |
| Log of firm age in 1998 | 0.004 | 0.041 | 0.002 | 0.038 |
| | (0.002)* | (0.023)* | (0.002) | (0.025) |
| Home state dummy | 0.651 | 3.252 | 0.585 | 2.879 |
| | (0.021)*** | (0.218)*** | (0.020)*** | (0.233)*** |
| Distance between headquarters and state | −0.226 | −2.407 | −0.217 | −2.500 |
| | (0.012)*** | (0.141)*** | (0.013)*** | (0.162)*** |
| Distance squared | 0.067 | 0.708 | 0.058 | 0.650 |
| | (0.005)*** | (0.057)*** | (0.005)*** | (0.065)*** |
| Constant | 0.297 | −0.220 | 0.304 | |
| | (0.008)*** | (0.089)** | (0.009)*** | |
| State-year fixed effects included | No | No | Yes | Yes |
| $R^2$ | 0.111 | | 0.160 | |
| Log Likelihood | | −6137.563 | | −5106.745 |
| Observations (firm-state-year) | 21430 | 21430 | 21430 | 20681[b] |

OLS = ordinary least squares. Standard errors in parentheses.
[a]"Potential" potential entrants' decisions to become a potential entrant.
[b]For 10 state-year combinations (corresponding to 749 observations), no firm obtained state certification and therefore these groups are dropped from the Logit regressions with state-year fixed effects.
*significant at 10%. **significant at 5%. ***significant at 1%.

# 6.    Results

☐    **State certification regression results.** Table 4 presents the results from the regressions of firm decisions to obtain state certification for the first time, as described in the first subsection of Section 4. The first two columns present the ordinary least squares (OLS) and Logit regression results, whereas the last two present their counterparts with state-year fixed effects. Comparing the results with and without state-year fixed effects, we can see that including such fixed effects significantly improves the model's fit to the data, particularly for the Logit model. This improvement suggests the importance of using state-year fixed effects to capture a general expectation of aggregate value of being eligible to enter in a given geographic state $s$ and year $t$. The results in the last two columns of Table 4 indicate that the observed firm attributes are key determinants of firm decisions to obtain state certification. CLECs that are privately held or subsidiaries of other firms are significantly less likely to obtain state certification. This finding may reflect the fact that such CLECs typically do not have a deep pocket and their opportunity cost of using capital is high. In contrast, those funded by venture capital, and thus with a higher ability to finance, are more likely to obtain certification. We also find that CLECs are significantly less likely to obtain state certification in states further from their headquarters: the home state dummy has a significant positive impact on a CLEC's obtaining state certification; the distance between a CLEC's headquarters zip code and the population centroid of the state it obtains certification from has a significant negative impact, though such a negative impact is diminishing with the distance. Overall, it seems that CLECs may have a home state cost advantage and have higher entry costs into a more distant geography.

As described in the first subsection of Section 4, we use the results from the certification regression (Column 4, Table 4) to categorize firms into two types: type-1 and type-2. Firms with $\hat{\varphi}_1 \mathbf{z}_f + \hat{\varphi}_2 \mathbf{d}_{fs}$ smaller than the median are labeled as type-1 in geographic state $s$ and those with

**TABLE 5    Summary Statistics on Potential Entrant Types**

| Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| # potential entrants | 28.903 | 10.365 | 5 | 56 |
| # type-1 potential entrants | 16.243 | 8.442 | 1 | 38 |
| # type-2 potential entrants | 12.660 | 6.794 | 1 | 34 |
| Entry rate | 0.040 | 0.067 | 0 | 0.5 |
| Type-1 potential entrants' entry rate | 0.029 | 0.066 | 0 | 0.6 |
| Type-2 potential entrants' entry rate | 0.055 | 0.100 | 0 | 1 |
| # observations (market-year) | | 1990 | | |

**TABLE 6    Estimation Results for Structural Parameters**

| | $\beta=$ | | | |
|---|---|---|---|---|
| Parameter | 0.95 | 0.90 | 0.85 | 0.80 |
| $\alpha$ (market size effect) | 0.150*** | 0.162*** | 0.173*** | 0.188*** |
| | (0.023) | (0.014) | (0.015) | (0.015) |
| $\gamma$ (competition effect) | −0.026 | −0.030*** | −0.033*** | −0.037*** |
| | (0.021) | (0.013) | (0.014) | (0.014) |
| $\mu_1$ (entry cost mean for type-1) | 10.082*** | 9.315*** | 8.693*** | 8.176*** |
| | (0.918) | (0.489) | (0.476) | (0.466) |
| $\mu_2$ (entry cost mean for type-2) | 9.671*** | 8.917*** | 8.308*** | 7.802*** |
| | (0.917) | (0.485) | (0.475) | (0.464) |

***Significant at 1%.

this measure larger than the median are labeled as type-2. Any market-year combination can now be characterized by four state variables: market size (the log of the number of business establishments), number of incumbents (including a single ILEC and CLECs), number of type-1 potential entrants, and number of type-2 potential entrants. Table 5 reports the summary statistics on the types of potential entrants. From this table, we can see that we have, on average, more type-1 than type-2 potential entrants in a local market and that there is substantial variation in the distributions of types.[22] In the data, the entry rate for type-1 potential entrants is, on average, 0.029, whereas that for type-2 potential entrants is 0.055.[23] This difference in entry probabilities helps us identify the difference in entry costs for the two types of potential entrants.

☐    **Estimates of structural parameters.**  Table 6 reports the estimation results for the four structural parameters in the model: the two parameters in the profit function (the market size effect, $\alpha$, and the competition effect, $\gamma$) and the two parameters describing the distribution of entry costs for each type of potential entrants (mean $\mu_1$ and $\mu_2$).[24] Table 6 reports, in different columns, the estimation results when the unadjusted discount factor $\beta$ is chosen to be 0.95, 0.9, 0.85, and 0.8, respectively. It is not surprising that the estimation results vary with the discount factor. For example, as the discount factor decreases, the estimated entry cost means decrease. This is intuitive: when firms discount the future payoffs more, the entry cost (that potential entrants need to pay now) must be smaller to explain the same entry behavior. As we show in web

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[22] Note that the median cutoff we use in determining a firm's type in a geographic state is the median across firm-states, whereas the local market is a city within a geographic state. Therefore, it is possible that type-1 and type-2 firms are unevenly distributed within a local market. Moreover, a firm is no longer a potential entrant after entry. Given that type-1 firms, on average, have a lower entry rate (see Table 5), it is not surprising that there are more type-1 potential entrants, on average, across year-markets.

[23] The entry rate for type-$\tau$ potential entrants is defined as the number of type-$\tau$ entrants over the number of type-$\tau$ potential entrants in the local market.

[24] To take into account the estimation error in the first-stage estimation of a firm's decision to become a potential entrant, we bootstrap to estimate the standard errors.

Appendix B, however, the fit of the model and the results from the counterfactual simulations as explained below are robust. In what follows, we focus on the estimation results in the first column of Table 6 where ß = 0.95.

The estimation results are rather intuitive. For example, market size, measured by the logarithm of the number of business establishments, has a positive effect on the incumbent's operating profit. This is in line with Bresahan and Reiss (1991), who find that a larger market size is necessary to support more competitors. It also implies that smaller markets may get stuck with a monopolistic structure, as these markets do not have sufficient demand to attract entry. Furthermore, we see that the number of incumbents negatively affects the operating profit of an incumbent. This result confirms the conventional wisdom that a higher number of incumbents in a market erodes the average profitability per firm.[25]

The estimate for the mean of a type-1 potential entrant's entry cost is 10.082, higher than that for a type-2 potential entrant. Recall that we group firms into two types based on their propensity to obtain state certification – type-1 firms have lower propensity than type-2 firms. In the estimation, we do not impose any restriction on the ranking of the entry cost mean for these two types, $\mu_1$ and $\mu_2$. We find that type-1 potential entrants indeed have higher entry costs on average than type-2 potential entrants. Put together, these results show that firms who are more likely to obtain state certification have lower entry costs. This finding is consistent with intuition.

The difference between type-1's and type-2's entry cost means is statistically significant at the 1% level. As we will show in the third subsection of Section 7, this difference has significant economic implications for firms' entry behavior.

Overall, our estimates imply that the net value of entry (the value of entry minus the average entry cost) for type-1 potential entrants varies between −0.036 and 1.430 depending on the state, and that the net value of entry for type-2 potential entrants varies between 0.376 and 1.846. Given that our rough measure of the average entry cost per market is $6.5 million, the net value of entry for type-1 potential entrants then amounts to between −$24,000 to $941,000 per market.[26] Similarly, the net value of entry for type-2 potential entrants varies between $247,000 and $1,215,000 per market. In comparison, the value of waiting varies between 0.006 and 0.090 ($4,000 and $60,000) for type-1 potential entrants and between 0.021 and 0.199 ($14,000 and $131,000) for type-2 potential entrants, around 10% of the net value of entry. Note that the value of waiting for type-1 high-cost potential entrants is smaller than that for type-2 low-cost potential entrants. This is mainly because the value of waiting is in part influenced by a potential entrant's perception of how likely it is to enter in the future. For example, at the extreme, if a potential entrant thinks that it will never enter, its option value of waiting is zero. As type-1 potential entrants have a lower probability of entry than type-2 potential entrants, their value of waiting is also smaller.

☐ **Fit of the model.** To ensure that our model captures the dynamics of entry behavior in the industry, we compare the distribution of the market structure from the observed data with the predictions from our model. Figure 1 shows the percentage of markets with $n$ CLECs from 1999 to 2002 for $n = 0, 1, 2,$ and above. The data shows that local markets become increasingly competitive over time. However, monopoly markets (markets with no CLECs) continue to represent a significant proportion of all markets. The prediction from the estimated model displays the same

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[25] This competition effect, however, is statistically insignificant. This may be due to unobserved market heterogeneity. Local markets differ in demand (e.g., the affluence level of local markets), in cost of laying out the network (e.g., various terrain conditions), and even in how hard ILECs compete with new entrants. Due to data limitations, we are unable to capture such heterogeneity. If more profitable markets (in unobservable dimensions) attract more entrants, we may underestimate the competition effect.

[26] The number –24,000 is obtained when we scale –0.036 by $6,500,000/(\frac{\hat{\mu}_1+\hat{\mu}_2}{2})$, where $\frac{\hat{\mu}_1+\hat{\mu}_2}{2}$ is the entry cost mean averaged across the two types of firms. Similarly, the number 941,000 is obtained by scaling 1.430 by the same factor.

FIGURE 1

DISTRIBUTION OF MARKET STRUCTURE: DATA VS. MODEL PREDICTION
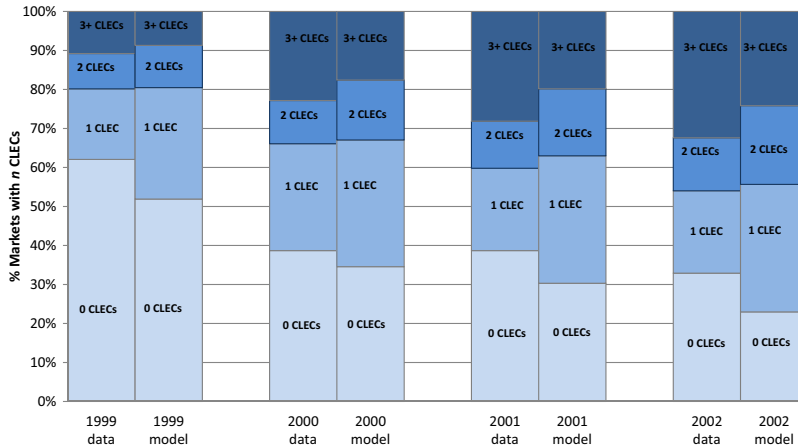


Note: CLECs = Competitive Local Exchange Carriers.

pattern. From the comparison, we can see that our estimated model fits the overall evolution of local market structures rather well. If anything, our model tends to slightly overestimate entry.

☐     **Comparison to POB.**  In this article, we take advantage of a unique feature of the US local telephone industry and identify potential entrants to a local market as CLECs with certification to operate in that state. Knowing who the potential entrants are allows us to observe how long a firm waits to enter a market and firm-level attributes associated with the cost of entry. Data indicate that potential entrants are heterogeneous and that some of them wait for several years before they actually enter a local market. To capture these features of the data, our model differs from POB to allow for a value of waiting and for firm heterogeneity. In this section, we compare the estimation results and the fit with the data using our model to the results and the fit using the original POB model, in which potential entrants are not observed and hence there is no waiting.

To this end, we estimate the POB model assuming that the number of potential entrants in each market is 20, 30, or 40.[27] We also estimate a hybrid model where we use the actual number of potential entrants in computing the empirical conditional probabilities of entry in the first-stage estimation of the POB model, but not as a state variable in the second-stage estimation. In the POB model and the hybrid model, a potential entrant either enters a market or perishes (i.e., there is no value of waiting). Table 7 presents the estimation results from our model, the POB model with different assumptions on the number of potential entrants, and the hybrid model. As shown in Table 7, these alternative models produce much larger estimates of the market size effect, the competition effect, and the entry cost means. We believe these changes are due to the information lost when we do not allow the number of different-type potential entrants to play a role. In the POB model and the hybrid model, the number of potential entrants, which affects the competitiveness of a market, is not included as a state variable. The two remaining state variables, the number of incumbents and the market size, have to explain the same variation in the probability of entry, leading to larger estimates of their coefficients. With these estimates, the profit is also larger, which in turn leads to a larger estimate of entry cost means. In the hybrid model, which uses the number of potential entrants in a limited way (no firm-level heterogeneity and not included in the state space), the overestimation of estimated coefficients is smaller, again pointing to the value of gaining information from the waiting structure.

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[27] The mean and the median of the number of potential entrants in our data is 29.

**TABLE 7  Estimation Results: Our Model vs. POB**

| Parameter | Our Model | Hybrid | POB # Potential Entrants = 20 | 30 | 40 |
|---|---|---|---|---|---|
| $\alpha$ (market size effect) | 0.150*** | 0.247*** | 0.283*** | 0.272*** | 0.264*** |
| | (0.023) | (0.048) | (0.037) | (0.036) | (0.036) |
| $\gamma$ (competition effect) | −0.026 | −0.112*** | −0.137*** | −0.131*** | −0.128*** |
| | (0.021) | (0.033) | (0.023) | (0.023) | (0.023) |
| $\mu$ (entry cost mean) | | 14.904*** | 17.961*** | 16.965*** | 16.415*** |
| | | (3.690) | (3.662) | (3.337) | (3.155) |
| $\mu_1$ (entry cost mean for type-1) | 10.082*** | | | | |
| | (0.918) | | | | |
| $\mu_2$ (entry cost mean for type-2) | 9.671*** | | | | |
| | (0.917) | | | | |

***Significant at 1%.

FIGURE 2

MODEL FIT COMPARISON: OUR MODEL VS. POB
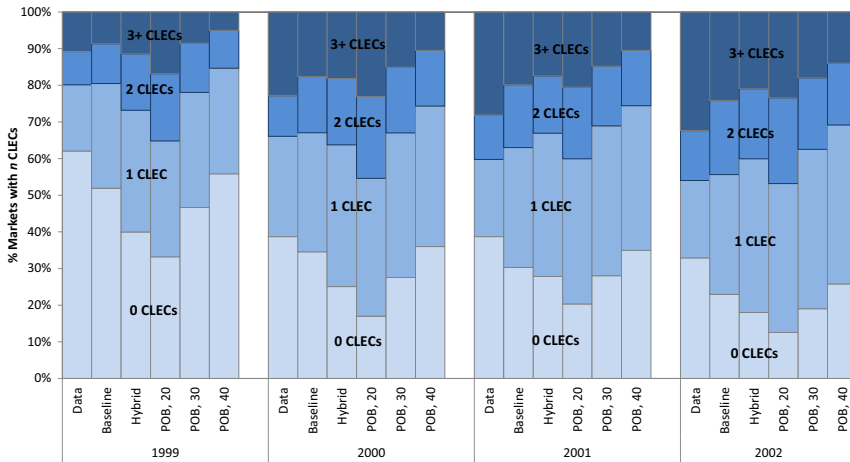


Note: CLECs = Competitive Local Exchange Carriers.

These estimates have direct consequences for the model's fit with the data. Figure 2 shows that our model fits the data better than the POB model and the hybrid model. As in the previous subsection, we show the percentage of markets with $n$ CLECs from 1999 to 2002 for $n = 0, 1, 2$, and above. We show this distribution of the market structure in the data, as predicted by our model, the hybrid model, and the POB model. Figure 2 indicates that our model fits the data the best. Let $x_{nt}^{data}$ be the share of markets with $n$ CLECs in year $t$ observed in the data and $x_{nt}^{est}$ be its estimated counterpart. We compute a measure for the fit of the market structure distribution as $\sum_{n=1,2,3+} \sum_{t=1999}^{2002} (x_{nt}^{data} - x_{nt}^{est})^2$. The value of this measure is 0.133 for our model, 0.355 for the hybrid model, and 0.472, 0.351, and 0.305 for the POB model with 20, 30, and 40 potential entrants, respectively. That is, our model predicts a market structure closest to what we observe in the data. Because in the counterfactual policy simulations below, we focus on how subsidies change market structure, we think it is assuring that our model fits the market structure in the data well. In addition, we will show in the next section that ignoring the identity of potential entrants leads to biased estimates of the effects of entry subsidies.

**TABLE 8    The Effect of Subsidies on Percentage of Monopoly Markets**

| | Percentage of Monopoly Markets (%) | | | |
|---|---|---|---|---|
| | 1999 | 2000 | 2001 | 2002 |
| **Subsidy to all markets and any entrant** | | | | |
| (1) No subsidy | 52.00 | 34.97 | 29.96 | 23.07 |
| (2) 5% subsidy, all markets, both types | 32.45 | 15.53 | 12.29 | 7.49 |
| (3) 10% subsidy, all markets, both types | 13.77 | 3.11 | 2.49 | 1.39 |
| **Subsidy to small markets only** | | | | |
| (4) 7.3% equivalent subsidy, small markets only[a] | 25.30 | 8.93 | 7.35 | 3.90 |
| (5) 12.6% equivalent subsidy, small markets only[b] | 10.17 | 1.59 | 1.54 | 0.71 |
| **Subsidy to high-cost or low-cost CLECs only** | | | | |
| (6) 10% subsidy, high-cost type only | 30.85 | 14.44 | 13.47 | 7.45 |
| (7) 10% subsidy, low-cost type only | 22.30 | 9.60 | 7.45 | 4.85 |
| (8) 13.5% equivalent subsidy, low-cost type only[c] | 14.00 | 5.12 | 4.28 | 2.77 |
| **Subsidy in 1998 only** | | | | |
| (9) 10% subsidy, 1998 only | 8.72 | 8.35 | 12.51 | 11.21 |
| (10) 16.7% equivalent subsidy, 1998 only[d] | 1.01 | 1.25 | 3.56 | 4.13 |

[a] A 5% subsidy to all markets is equivalent to a 7.3% subsidy to small markets only. Under these two subsidy schemes, the total amount of subsidy paid in these four years is the same.
[b] A 10% subsidy to all markets is equivalent to a 12.6% subsidy to small markets only.
[c] A 10% subsidy to both types of potential entrants is equivalent to a 13.5% subsidy to low-cost potential entrants only.
[d] A 10% subsidy in all years is equivalent to a 16.7% subsidy in 1998 only.

## 7.    Counterfactuals

■    Having ascertained that our model is a good fit for the data, we now study various subsidy policies for encouraging further entry after the Act. Note that the Act includes policies that can be interpreted as implicit nondiscriminatory subsidies to every entrant, most notably in the form of forcing ILECs to interconnect with CLECs and to lease their networks and facilities to CLECs at rates based on long-run average costs. Aided by these policies, CLECs are able to avoid negotiating interconnection agreements or building overlapping networks with ILECs.[28] In the simulations that follow, we study several explicit subsidy policies *on top of* the existing policies and examine their effects on promoting competitiveness in local markets. All of the policies studied subsidize the entry cost. Throughout the analysis, we focus on comparing the impact of these policies on reducing the number of monopoly markets.[29] In the simulations below, we keep the set of CLECs holding a certification from a state the same as that in the data.

□    **Subsidy to every new entrant in all markets.**    Table 8 shows how applying a subsidy to every entrant could encourage entry into local markets. The first row shows the status quo—the model-simulated distribution of market structures with no subsidies.[30] Row (2) and Row (3) show,

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[28] The subsidies imposed by the Act are implicit and thus difficult to quantify. One may worry that these implicit subsidies, which lower average variable costs by allowing CLECs to rent networks and facilities from ILECs, vary across markets and thus impact our estimates. However, the main component of CLECs' average variable cost is maintaining and servicing the networks and facilities. It is thus reasonable to assume that these subsidies are only a function of the size of the network, which depends on market size alone. Therefore, most of the unobserved market-level policy heterogeneity is captured by market size, which is already included in the model.

[29] Although a market could theoretically become more competitive even without entry, as the mere threat of entry after the Act could make the incumbent act more competitively, there is little evidence on such effects. On the other hand, some studies have found a positive welfare effect of an increase in the actual number of competitors in the local phone industry (see, e.g., Economides, Seim, and Viard, 2008).

[30] We use a model-simulated market structure because we do not want realizations of unobservables in the data to affect the comparison between results with and without subsidies. Furthermore, we have shown above that the simulated market structure is close to the observed market structure.

TABLE 9    Change in Percentage of Monopoly Market under a 5% Subsidy: Our Model vs. POB

| | Percentage of Monopoly Markets (%) | | | |
|---|---|---|---|---|
| | 1999 | 2000 | 2001 | 2002 |
| Our model | −20.00 | −19.65 | −17.92 | −15.27 |
| Hybrid | −8.67 | −7.63 | −7.79 | −6.66 |
| POB, 20 potential entrants | −10.05 | −7.49 | −7.54 | −6.77 |
| POB, 30 potential entrants | −8.78 | −6.71 | −6.53 | −5.42 |
| POB, 40 potential entrants | −7.54 | −4.98 | −5.64 | −4.26 |

respectively, the effect of a subsidy equaling 5% and 10% of the entry cost mean averaged across the two types (i.e. 5% and 10% of $(\hat{\mu}_1 + \hat{\mu}_2)/2$) to every entrant in every local market. From Table 8, we can see that the 5% subsidy reduces the share of monopoly markets to 32% by the beginning of 1999 (compared to 52% without any subsidy), whereas the 10% subsidy further reduces this share to 14% over the same period. By the beginning of 2002, the 5% subsidy reduces the percentage of monopoly markets to less than 10% (compared to 23% without any subsidy), whereas the 10% subsidy nearly eliminates monopoly markets.

We have shown that our model fits the data better than the POB model and the hybrid model. To understand whether this difference in fitting the data and the different estimates across the two models also have economic significance, we next investigate whether they have different implications for evaluating the effect of subsidy policies. To this end, we simulate the effect of a 5% subsidy to every entrant in all markets using the estimated POB models with different assumptions on the number of potential entrants. We do so using the estimated hybrid model as well, where the number of potential entrants is taken from the data. However, we still maintain the assumption that potential entrants either enter or exit and that they make their entry decisions based on the assumption that a fixed number of potential entrants are born every period. We compare the simulated results in Table 9. For example, our model predicts a drop in the percentage of monopoly markets by 20% (from 52%–32%) in 1999 under a 5% subsidy to every entrant in all markets. The predicted change using the POB model or the hybrid model varies between 7.5% and 10%, depending on the assumption about the number of potential entrants. Overall, we can see from Table 9 that the estimated POB model underestimates the subsidy effect by more than 50%. Note that in the POB model and the hybrid model, firms think that there is a fixed number of potential entrants born every period, and thus an entry subsidy does not affect the number of potential entrants. However, in fact, such a subsidy can lead to more entries now and hence decreases the number of potential entrants in the future, which increases the value of entry for a firm. Thus, ignoring the effect of a subsidy on the number of potential entrants leads to an underestimation of the subsidy's effect on entry.

Returning to the discussion on the effect of subsidies, though applying a subsidy to every entrant in every market is effective, it may also be costly. In our model, the number of monopoly markets under a subsidy equaling 5% (10%) of the average entry cost would be the same as in a scenario where there is a 5.1% (10%) exogenous increase in market size of every market and in every period. Recall that an increase in the market size leads to higher post entry profit and hence attracts more entry. Another way to understand the magnitude of the subsidy is to use information on the annual capital expenditures that we observe for the majority of the CLECs. Recall from Section 2 that the average entry cost per market is calculated to be $6.5 million. This translates into roughly $325,000 per firm for a 5% subsidy and $650,000 for a 10% subsidy. The question that arises next is: can alternative subsidy designs that target selected markets or selected CLECs be more cost effective?

☐    **Subsidy in small markets only.** To answer the above question, we study in this section whether offering a subsidy in small markets only, that is, in cities with fewer than 5000 business

establishments in 1998, is more effective at reducing monopoly markets.[31] In other words, for the same amount of total subsidy paid, does a subsidy in small markets only lead to fewer monopoly markets than a subsidy applied to all markets? Intuitively, on the one hand, a small market is less attractive to potential entrants. The same amount of subsidy per firm may be less effective at encouraging entry into small markets than at encouraging entry into larger markets. Thus, a higher subsidy for each entrant would be needed. On the other hand, small markets are more likely to be monopoly markets before a subsidy. Therefore, a subsidy encouraging entry into a small market may immediately eliminate a monopoly, whereas a subsidy in a larger market may only help to add another competitor to an already relatively competitive market.

To study the overall effect of a small market subsidy, we use the simulation results in the previous subsection as the benchmark. Doing so, we find that, in terms of costs, a 5% subsidy in all markets is equivalent to a 7.3% subsidy in only small markets. In other words, under these two subsidy schemes, the total amount of subsidy paid in 1998 to 2001 is the same. Similarly, we see that a 10% subsidy in all markets is equivalent to a 12.6% subsidy in small markets only. Rows (4) and (5) of Table 8 show the percentage of monopoly markets under these equivalent subsidies in small markets only. The comparison of Row (2) (5% subsidy in all markets) and Row (4) (the corresponding equivalent subsidy in small markets only) shows that providing a subsidy in small markets only is more effective than providing a subsidy in all markets. The same amount of money spent leads to a larger reduction in the share of monopoly markets in all years of our study. The comparison of Row (3) and Row (5) for the effect of a 10% subsidy in all markets and the effect of its equivalent subsidy in small markets yields the same result. This result indicates that the first effect (a higher subsidy per entrant is needed to attract entry into a smaller market) is dominated by the second effect (the subsidy to small markets only is more likely to be right on target at reducing monopoly markets).

We have shown that a small-market-only subsidy policy is more cost effective at reducing the number of monopoly markets. However, is it more desirable from a welfare point of view? On the one hand, it leads to a larger reduction in monopoly markets. On the other hand, it affects fewer customers per market. This trade-off is similar to a trade-off between a positive extensive margin effect and a negative intensive margin effect. To investigate the overall effect, we now examine the effect of different subsidy policies on the percentage of business establishments located in monopoly markets (in addition to the percentage of monopoly markets as studied above). According to our simulation results, under the all-market 5% subsidy, the percentages of business establishments located in monopoly markets are 24.43%, 11.07%, 9.25%, and 5.67% in 1999, 2000, 2001, and 2002, respectively. Under the equivalent small-market subsidy, they become 22.37%, 7.31%, 6.25%, and 3.33%. In other words, the equivalent small-market-only subsidy is more effective at reducing not only the number of monopoly markets, but also the number of business establishments located in monopoly markets. Thus, even though this comparison is not a full welfare analysis, it suggests that the small-market-only subsidy is likely to be more efficient, given that the dollar amount paid under these two subsidies is the same.

The result is, however, different when we compare the 10% all-market subsidy to its equivalent small-market-only subsidy: the percentages of business establishments located in monopoly markets under these two subsidies are, respectively (10.32%, 2.21%, 1.79%, 0.97%) and (11.00%, 2.15%, 1.96%, 1.03%) between 1999 and 2002. The latter is slightly larger (except in year 2000), implying that the 10% all-market subsidy benefits more business consumers. This change in the results is not surprising because compared to a 5% all-market subsidy, a 10% all-market subsidy eliminates more monopoly markets, small or large. When a subsidy is already successful at eliminating monopoly markets, restricting it to only small markets can be less efficient. For example, at the extreme, if a subsidy eliminates all monopoly markets, restricting it to only small markets may leave some large markets monopolistic. In conclusion, for a large subsidy that can greatly reduce the number of monopoly markets, switching to a small-market-only subsidy may not be

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[31] In our sample, 310 of 398 cities fall into this category.

efficient. In contrast, for a moderate subsidy program, focusing on small markets only may be more efficient.

□ **Subsidy to low-cost CLECs only.** Another option to improve the effectiveness of a subsidy policy is to provide a subsidy to only low-cost type potential entrants. Intuitively, a subsidy to a low-cost potential entrant would be more effective than the same subsidy offered to a high-cost potential entrant because it would be more likely to help the former to overcome entry costs. So, a subsidy to only low-cost potential entrants may be more cost effective. This intuition is illustrated in Rows (6) and (7) of Table 8, where we compare the percentage of monopoly markets when a 10% subsidy is applied to type-1 (high-cost type) potential entrants only (Row (6)) to the percentage of monopoly markets when a 10% subsidy is applied to type-2 (low-cost type) potential entrants only (Row (7)).[32] We can see from this comparison that the same amount of subsidy per firm is more likely to help a low-cost potential entrant to enter than a high-cost potential entrant. This comparison also shows that the estimated difference between the entry cost means of the two types (10.082 vs. 9.671) has significant implications for these two types' entry behavior.

However, at the same time, such a discriminatory policy is applied to fewer potential entrants. Thus, it is also possible that it encourages less entry. To study the overall effect, we compare the percentage of monopoly markets when a 10% subsidy is given to all new entrants (Row (3) in Table 8) to the percentage when an equivalent subsidy is given to low-cost potential entrants only (Row (8) in Table 8). Again, two subsidy policies are "equivalent" if the total amount of subsidy paid under the two policies is the same. The comparison of Row (3) and Row (8) shows that a subsidy to low-cost potential entrants only is less effective than a subsidy to both types. Unlike the policy that exploits market heterogeneity, a policy exploiting firm heterogeneity is less effective at reducing the number of monopoly markets than a nondiscriminatory policy.

This result indicates that the latter "fewer firms" effect of the discriminatory policy dominates the former "entry cost heterogeneity" effect. A subsidy to low-cost firms may be more cost effective because the same amount of subsidy is more likely to help a low-cost firm to overcome entry costs than it would a high-cost firm. The magnitude of this effect is governed by the difference between the entry costs for the two types of firms. Even though the estimated difference has significant implications for the two types of firms' entry behavior, as shown by the comparison of Rows (6) and (7) in Table 8, it is not large enough to dominate the fact that the discriminatory policy applies to fewer potential entrants, which leads to less entry.

□ **Subsidy in 1998 only.** Last, we study the effect of option values for the arrival speed of competition in local markets and its implication for the design of entry subsidies. Specifically, we consider the market structure implications of changing the option value of waiting for potential entrants by offering a one-shot subsidy in 1998. This modification affects the timing of competition arrival through two channels. First, a potential entrant can receive a subsidy only if entering in 1998, not in subsequent years, which decreases the value of waiting in 1998. The second channel is through the indirect competition effect. Potential entrants know their competitors will not be subsidized to enter in years other than in 1998. So, there might be less competition in the future compared to when the subsidy is offered in all years, which increases the value of entry in 1998. To illustrate how these two effects would impact the timing of competition arrival, we simulate the effects of offering a 10% subsidy in 1998 only and present the results in Row (9) of Table 8. This table shows that the subsidy in 1998 only reduces the share of monopoly markets to 8.72% by the beginning of 1999, compared to 13.77% when a subsidy is offered in all years (Row (3) in Table 8). This result suggests that a one-shot subsidy speeds up the arrival of competition and thus reduces the number of monopoly markets. This is indeed because of the two effects explained above. Specifically, we find that the average value of waiting (averaged over different values of

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[32] A similar comparison when a 5% subsidy is used yields qualitatively similar results.

the state variables) in 1998 for type-1 potential entrants decreases by 42% when we move from a 10% subsidy offered every year to a 10% subsidy offered in 1998 only. Similarly, we find that the average value of waiting for type-2 potential entrants decreases by 33%. At the same time, the average value of entry increases by 0.4%.

To further understand the effects of the above two channels, we conduct two decompositions. As explained, the overall effect of these two channels is that the percentage of monopoly markets at the beginning of 1999 drops from 13.77% when a 10% subsidy per entrant is applied in all years to 8.72% when it is applied in 1998 only. In other words, there is a decline of 5.05%. For the first decomposition, we keep the value of entry the same as that under the all-year subsidy while allowing the value of waiting to be that under the 1998-only subsidy. Doing so, we find that with only the decrease in the value of waiting, the percentage of monopoly markets drops to 11.30%. In other words, of the total decline of 5.05%, 2.47% is due to the decrease in the value of waiting. The remaining 2.58% of the decline can be attributed to the increase in the value of entry. In a second decomposition, we keep the value of waiting the same as that under the all-year subsidy while allowing the value of entry to change. The simulation results show that the percentage of monopoly markets at the beginning of 1999 decreases from 13.77% to 10.28%. This decomposition shows that the effect of the increase in the value of entry is 3.49%. Together, these two decompositions indicate that both channels contribute to a decline in the number of monopoly markets. They also show that the indirect competition effect through the increase in the value of entry is slightly higher than the direct effect of reducing the value of waiting. This latter result suggests that it is important to consider the competition effect of entry when designing a subsidy policy.

Returning to the comparison of the all-year subsidy policy and the 1998-only subsidy policy, the long-run effect of the 1998-only subsidy policy may be different from the short-run effect. For example, under the 1998-only subsidy, the percentage of monopoly markets gradually grows over time as shown in Row (9) of Table 8. We attribute this result to a high exit rate. Without continuous subsidies in all years, the industry loses competitive markets later on. In other words, the long-run effect of this subsidy policy on market structure is influenced by the exit rate. When the exit rate is high, the long-run effect may be small.

As in previous counterfactuals in which we keep the total amount of subsidy spent the same, we also apply a 16.7% subsidy in 1998 only, which is equivalent to a 10% subsidy in all years. The results, presented in Row (10) of Table 8, show that this equivalent 1998-only subsidy drastically reduces the percentage of monopoly markets, nearly eliminating them by the beginning of 1999. However, due to the high exit rate, the percentage of monopoly markets at the beginning of 2001 under the equivalent 1998-only subsidy policy becomes higher than that under the 10% subsidy offered in all years.

## 8.  Conclusion

■    Before 1996, decades of regulation left the US local telephone industry with a monopolistic market structure. The 1996 Telecommunications Act opened the telecommunications markets to new entrants. However, due to substantial entry costs, many local markets remained monopolistic, leaving deregulated incumbents with the freedom to exercise market power. In this study, we explore the effect of subsidizing entry costs on new firm entry. We do so by combining economic theory with data on the entry decisions of CLECs from 1998 to 2002. We estimate a dynamic oligopolistic entry game, in which potential entrants are heterogeneous long-run players with an option value of waiting. Through counterfactual experiments, we obtain results that suggest that policy makers should exploit market heterogeneity but not firm heterogeneity when designing subsidy policies. Our results also indicate that policy makers should consider the dynamic, oligopolistic nature of local competition. In particular, we find that subsidies in only early periods speed up the initial arrival of competition, due to a direct effect that reduces the value of waiting and an indirect competition effect that increases the value of entry. These results shed new light on a

critical policy area in a fundamental infrastructure industry. Overall, our policy recommendations exploit information that regulators can readily access (e.g., size of the market) and actions that regulators can easily control (e.g., timing of the subsidy).

In this study, we focus specifically on the local telephone market. The issues about encouraging competitive entry are common in many telecommunications industries, especially the wireless telephone and Internet industries. Moreover, local telephone companies have now become major players in the Internet market.[33] CLEC entry thus brings competition to local telephone market services and Internet services. Our findings in the local telephone markets are therefore likely to have implications for policy design in these related telecommunications markets, where competitive entry can alleviate the discrepancies in the availability and quality of services across different markets.

Several limitations of our work need to be acknowledged. First, we capture firm-level heterogeneity in entry costs by allowing firms to draw entry costs from two different distributions. If we had a longer panel, and in turn more data points, we might be able to discretize firms into more types and thus better capture firm heterogeneity. Second, our model does not incorporate post entry firm level heterogeneity. In the real world, CLECs may cater to different clienteles and offer differentiated value-added services. Without data on post entry competition, however, we are not able to provide insight on this issue. Third, we assume that entry decisions are independent across markets. This is a standard assumption in the dynamic entry literature, as the state space would increase substantially otherwise. One may be concerned about the possibility of entry clustering due to spillover effects.[34] Despite these limitations, we find that our model fits the data reasonably well. We believe that we have made a first step that we hope will encourage future research in this area.

## Appendix

This Appendix provides some details on the estimation procedure.

▢ **Discretization of the state space.** We discretize the state space and estimate the transition probabilities directly from the data. Specifically, we discretize the market size into two bins: small markets (market size $\leq 5000$) and large markets (market size $> 5000$). We assign the average market size (averaged across all market-years within the same bin) as the value for a bin. Similarly, we discretize the number of incumbents into three bins using the following cutoffs: $n = 1$, $n = 2$, $n \geq 3$; and assign, for the last bin, the average number of incumbents within the bin as the value for the bin. We also discretize the number of type-1 and type-2 potential entrants into two bins, each using their respective average across market-years as the cutoffs (16.243 and 12.660, respectively) and again assign the average within each bin as the bin value. This discretization is fairly coarse. In web Appendix C, we show that our estimates are robust to various finer discretizations.

▢ **Estimation of the transition probability.** After the discretization, we estimate the transition probabilities directly from the data. The estimate of the unconditional transition probability is given by

$$\hat{M}_{ij} = \frac{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} \mathbb{1}\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = \left(m_j, n_j, T_{1j}, T_{2j}\right)\right\}}{\#K(m_i, n_i, T_{1i}, T_{2i})}, \quad (A1)$$

where $K(m_i, n_i, T_{1i}, T_{2i})$ is the collection of all market-years whose states are $(m_i, n_i, T_{1i}, T_{2i})$, $k$ represents such a market-year, and $k + 1$ represents the same market one year later. The cardinality of the set $K(m_i, n_i, T_{1i}, T_{2i})$ is $\#K(m_i, n_i, T_{1i}, T_{2i})$.

The estimate of the transition probability conditional on a type-1 firm entering is given by

$$\hat{M}^e_{1ij} = \frac{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} e_{1(k)} \mathbb{1}\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = \left(m_j, n_j, T_{1j}, T_{2j}\right)\right\}}{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} e_{1(k)}}, \quad (A2)$$

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[33] The three main types of Internet service providers are ILECs, CLECs, and cable TV companies. As of December 2003, ILECs and CLECs together account for about one third of the high-speed Internet lines in the United States (Xiao and Orazem, 2011).

[34] In web Appendix D, we estimate a model allowing the entry cost to depend on whether a firm is already operating in a close-by market but still treat entry into different markets as independent. We find that a firm with neighboring presence does have a slightly lower entry cost on average, but the estimates of other parameters stay fairly robust.

where $e_{1(k)}$ is the number of type-1 entrants in market-year $k$. Note that the transition of the states is weighted by the ratio of the probability that $(e_1 - 1)$ out of $T_1 - 1$ potential entrants enter over the probability that $e_1$ out of $T_1$ potential entrants enter. Similarly,

$$\hat{M}^e_{2ij} = \frac{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} e_{2(k)} \mathbb{1}\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = \left(m_j, n_j, T_{1j}, T_{2j}\right)\right\}}{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} e_{2(k)}}. \tag{A3}$$

The same argument gives the estimate of the transition probability conditional on waiting. The transition of the states conditional on waiting for type-1 firms is weighted by the ratio of the probability that $e_1$ out of $T_1 - 1$ potential entrants enter over the probability that $e_1$ out of $T_1$ potential entrants enter. The transition for type-2 firms is analogously weighted. In other words, the estimates are

$$\hat{M}^w_{1ij} = \frac{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} \left(T_{1i} - e_{1(k)}\right) \mathbb{1}\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = \left(m_j, n_j, T_{1j}, T_{2j}\right)\right\}}{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} \left(T_{1i} - e_{1(k)}\right)}, \tag{A4}$$

and

$$\hat{M}^w_{2ij} = \frac{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} \left(T_{2i} - e_{2(k)}\right) \mathbb{1}\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = \left(m_j, n_j, T_{1j}, T_{2j}\right)\right\}}{\sum_{k \in K(m_i, n_i, T_{1i}, T_{2i})} \left(T_{2i} - e_{2(k)}\right)}. \tag{A5}$$

□ **Computing the Markov perfect equilibrium.** We compute the Markov perfect equilibrium (MPE) to show the fit of the model and to perform counterfactual simulations. As explained in Section 4, a MPE is a triple $(p^e, V^I, V^E)$ that solves equations (3), (4), and (5). Therefore, to compute the MPE, we start with a policy function, $p_0^e$, which is a vector with each element representing the probability of entry in a state. We then compute the three transition probability matrices (the unconditional transition probability of the state, the transition probability conditional on entering, and that conditional on waiting) corresponding to this policy function. With these transition probability matrices computed, we compute the fixed point (denoted by $V_0^I$) of equation (3). We plug $V_0^I$ into equation (4) and solve for $V^E$ from equation (4) (denoted by $V_0^E$). We compute a new policy function ($p_1^e$) according to equation (5) by plugging in $V_0^I$ and $V_0^E$. We continue this process until the vector $p^e$ converges. We "rule out" multiple equilibria by starting with different initial guesses of the policy function.

## References

ACKERBERG, D.A., RIORDAN, M.H., ROSSTON, G.L., AND WIMMER, B.S. "Low-Income Demand for Local Telephone Service: Effects of Lifeline and Linkup." No 08-047, Discussion Papers, Stanford Institute for Economic Policy Research, 2009.

AGUIRREGABIRIA, V. AND MIRA, P. "Sequential Estimation of Dynamic Discrete Games." *Econometrica*, Vol. 75 (2007), pp. 1–53.

ALEXANDER, D.L. AND FEINBERG, R.M. "Entry in Local Telecommunications Markets." *Review of Industrial Organization*, Vol. 25 (2004), pp. 107–127.

BAJARI, P., BENKARD, L., AND LEVIN, J. "Estimating Dynamic Models of Imperfect Competition." *Econometrica*, Vol. 75 (2007), pp. 1331–1370.

BASKER, E. "Selling a Cheaper Mousetrap: Wal-Mart's Effect on Retail Prices." *Journal of Urban Economics*, Vol. 58 (2005), pp. 203–229.

BRESNAHAN, T.F. AND REISS, P.C. "Entry and Competition in Concentrated Markets." *Journal of Political Economy*, Vol. 99 (1991), pp. 977–1009.

COLLARD-WEXLER, A. "Demand Fluctuations in the Ready-Mix Concrete Industry." *Econometrica*, Vol. 81 (2013), pp. 1003–1037.

CRANDALL, R.W. "An Assessment of the Competitive Local Exchange Carriers Five Years After the Passage of the Telecommunications Act." Unpublished Report, June 2001.

———. "Competition and Chaos: US Telecommunications since the 1996 Telecom Act." Washington, DC: Brookings Institution Press, 2005.

DORASZELSKI, U. AND SATTERTHWAITE, M. "Computable Markov–Perfect Industry Dynamics." *RAND Journal of Economics*, Vol. 41 (2010), pp. 215–243.

DUNNE, T., KLIMEK, S., ROBERTS, M., AND XU, D.Y. "Entry, Exit and the Determinants of Market Structure." *RAND Journal of Economics*, Vol. 44 (2013), pp. 462–487.

ECONOMIDES, N. "The Telecommunications Act of 1996 and Its Impact." *Japan and the World Economy*, Vol. 11 (1999), pp. 455–483.

ECONOMIDES, N., SEIM, K., AND VIARD, V.B. "Quantifying the Benefits of Entry into Local Telephone Service." *RAND Journal of Economics*, Vol. 39 (2008), pp. 699–730.

FAN, Y. "Estimating the Costs of Market Entry and Exit for Video Rental Stores." Working Paper, University of Michigan, 2006.

——— "Ownership Consolidation and Product Characteristics: A Study of the US Daily Newspaper Market." *American Economic Review*, Vol. 103 (2013), pp. 1598–1628.

GOLDFARB, A. AND XIAO, M. "Who Thinks About the Competition? Managerial Ability and Strategic Entry in US Local Telephone Markets." *American Economic Review*, Vol. 101 (2011), pp. 3130–3161.

GREENSTEIN, S. AND MAZZEO, M.J. "The Role of Differentiation Strategy in Local Telecommunication Entry and Market Evolution: 1999–2002." *Journal of Industrial Economics*, Vol. 54 (2006), pp. 323–350.

HAUSMAN, J.A. AND TAYLOR, W.E. "Telecommunications Deregulation." *American Economic Review: Papers and Proceedings*, Vol. 102 (2012), pp. 382–390.

HECKMAN, J.J. AND HONORÉ, B.E. "The Empirical Content of the Roy Model." *Econometrica*, Vol. 58 (1990), pp. 1121–1149.

HOTZ, V.J. AND MILLER, R.A. "Conditional Choice Probabilities and the Estimation of Dynamic Models." *Review of Economic Studies*, Vol. 60 (1993), pp. 497–529.

KELLOGG, R. "The Effect of Uncertainty on Investment: Evidence from Texas Oil Drilling." *American Economic Review*, Vol. 104 (2014), pp. 1698–1734.

KENNEDY, C.H. *An Introduction to U.S. Telecommunications Law*. 2d ed. Nerwood, MA: Artech House Telecommunications Library, 2001.

MATSA, D. "Competition and Product Quality in the Supermarket Industry." *Quarterly Journal of Economics*, Vol. 126 (2011), pp. 1539–1591.

MAZZEO, M.J. "Competition and Service Quality in the US Airline Industry." *Review of Industrial Organization*, Vol. 22 (2003), pp. 275–296.

MINI, F. "The Role of Incentives for Opening Monopoly Markets: Comparing GTE and BOC Cooperation with Local Entrants." *Journal of Industrial Economics*, Vol. 49 (2001), pp. 379–414.

MIRAVETE, E. "Estimating Demand for Local Telephone Service with Asymmetric Information and Optional Calling Plans." *Review of Economic Studies*, Vol. 69 (2002), pp. 943–971.

NARDOTTO, M., VALLETTI, T., AND VERBOVEN, F. "Unbundling the Incumbent: Evidence from UK Broadband." CEPR Discussion Papers no. 9194, CEPR Discussion Papers, 2012.

NEVO, A. "Mergers with Differentiated Products: The Case of the Ready-to-Eat Cereal Industry." *RAND Journal of Economics*, Vol. 31 (2000), pp. 395–421.

New Paradigm Resources Group. 1999–2003, 1998–2002 CLEC Reports, Chicago, IL.

PAKES, A., OSTROVSKY, M., AND BERRY, S. "Simple Estimators for the Parameters of Discrete Dynamic Games, with Entry/Exit Examples." *RAND Journal of Economics*, Vol. 38 (2007), pp. 373–399.

PESENDORFER, M. AND SCHMIDT-DENGLER, P. "Asymptotic Least Squares Estimators for Dynamic Games." *Review of Economic Studies*, Vol. 75 (2008), pp. 901–928.

PINDYCK, R. "Irreversibility, Uncertainty, and Investment." *Journal of Economic Literature*, Vol. 29 (1991), pp. 1110–1148.

——— "Sunk Costs and Real Options in Antitrust." In W. Collins, ed., *Issues in Competition Law and Policy*. Chicago, IL: American Bar Association, 2008.

RYAN, S. "The Costs of Environmental Regulation in a Concentrated Industry." *Econometrica*, Vol. 80 (2012), pp. 1019–1061.

XIAO, M. AND ORAZEM, P. "Does the Fourth Entrant Make Any Difference? Entry and Competition in the Early U.S. Broadband Market." *International Journal of Industrial Organization*, Vol. 29 (2011), pp. 547–561.

YOUNG, S., DREAZEN, Y., AND BLUMENSTEIN, R. "On Hold: Local Phone Giants Beat Back Rivals, as Competition Stalls." *Wall Street Journal*, February 11, 2002, pp. 547–561.

## Supporting information

Additional supporting information may be found in the online version of this article at the publisher's website:

**Appendix A:** List of Markets in the Sample.
**Appendix B:** Robustness Analysis: Different Discount Factors.
**Appendix C:** Robustness Analyses: Different Discretizations.
**Appendix D:** Synergy in Entering Neighboring Markets.