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Abstract

We estimate a dynamic oligopoly entry game in the early U.S. local telephone market. We observe the identities of potential entrants into local markets and therefore the waiting time of each potential entrant before it commits actual entry. To capture the feature of the data, we allow firms to be heterogeneous long-run players who have option value of waiting. We find that firm-level heterogeneity in entry costs plays a significant role in determining a firm's entry behavior into a local market. Our model can be used to conduct counterfactual simulations to understand the effectiveness of subsidy policies with different focuses.

Key words: option value of waiting, entry, dynamic oligopoly entry game, telecommunications

JEL: L1, L96

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1 Introduction

This paper seeks to understand the importance of heterogeneity in dynamic, strategic interactions in a high-stake market. Specifically, we study the entry decisions into local telephone markets in the United States. The passage of the Telecommunications Act of 1996 opened the floodgate to competitive entry into this industry. The entrants (known as Competitive Local Exchange Carriers, or CLECs) were aided by implicit subsidies from government and equipment suppliers, and they varied substantially in size, ownership and financial structure, and telecommunications experience.

What would be an effective subsidy policy by the Federal Communication Commission if the goal is to accelerate the arrival of competition in relatively rural, sparsely-populated areas? The current practice is to offer subsidy to every entering CLEC. Would this type of subsidy achieve its goal to alleviate the divide between urban and rural communities? Not necessarily. For example, a subsidy to reduce sunk costs of entry in rural areas, applied to every potential entrant, will also lower the expectation of entry value of every potential entrant because its competitors are also subsidized and thus more likely to enter. Therefore a market may have to wait longer for competition to arrive under this subsidy policy. Ultimately, this is an empirical question which depends on the effects of the subsidy on the entry costs and entry value. Since entry costs can vary substantially across firms and entry value across markets, it is important for researchers and policy markers to take into account firm- and market- level heterogeneity.

In order to answer the above question, we need to directly estimate the distributions of heterogenous sunk costs. We employ a rich panel data which record all CLECs' entry decisions into local telephone markets. What is novel in our data is that we observe the identities of potential entrants into local markets and therefore the waiting time of each potential entrant before entry. This is because one key feature of the industry after the 1996 Act: a firm needs to obtain certification from a state in order to become a potential entrant to cities within the state. From our data we observe a large variation in the waiting time for a CLEC to commit actual entry after it obtains certification. The entry decisions, and more importantly, the timing of entry allow us to infer firms' heterogeneous sunk costs of entry.

To capture the feature of the data, we set up a dynamic entry model based on Pakes, Ostrovsky and Berry (2007) (henceforth, POB) with two deviations. First, a firms is a long-run player and decides on entering or waiting. When making this decision, a firm compares the value of entry subtracted by the entry costs with the option value of waiting. This is in contrast with most entry studies in which a firm enters or perishes and the option value of waiting is set to zero. Second, firms are heterogeneous, which is key to explaining variations in the waiting time. For example, a better-funded firm may have smaller sunk costs of entry and be able to implement its entry plans more quickly. With these two deviations, we are able to incorporate the timing of entry and firm heterogeneity into a dynamic oligopoly entry game.

To estimate this model, we follow the recent development in two-step estimation strategy of dynamic oligopoly entry games. In such two-step procedure of estimation, the econometrician first obtains the conditional choice probability at any given state from data and then matches the empirical conditional choice probability with its counterpart predicted by the model. The first step is nonparametric, which puts a limit on the number of state variables in such a model. In our setting, this means a limit on the number of firms' heterogeneous characteristics that we can incorporate.[1] We solve this problem by mapping multiple dimensions of firm- and firm/market-specific heterogeneity into a single dimension —- a variable which describes the type of a potential entrant in a market. To do this, we incorporate the decision of an established firm to become a potential entrant into a local market. This decision reflects a firm's expectation of future payoffs against its expectation of the entry costs. The observed relationship between firm- and firm/market-specific attributes and a firm's decision in a market therefore gives us a projection from these attributes to a single-dimensional characteristic of this firm. We then categorize firms into high-cost and low-cost types based on this single dimension of firm heterogeneity. The advantage of our approach is that the type of a firm is estimated from firms' decisions on becoming potential entrants instead of imposed on in an ad hoc fashion.

This is still work in progress and we have two key findings from preliminary results. First, market size, measured by the number of business establishments, affects the operating profit of a CLEC positively once it becomes an incumbent. Therefore the larger the market is, the more likely is a CLEC to enter. This befits the traditional wisdom in line of Bresenhan and Reiss (1991) that a larger market size is necessary to support

[1] On the contrary, in the reduced-form framework incorporating firm-level heterogeneity is typically not an issue. For example, Morton (1999) conducts regression analysis to study which attributes of potential entrants contribute to entry decision in the pharmaceutical industry. However, a reduced-form framework cannot allow us to properly address the policy question that we are interested in.

more competitors. Second, firm-level heterogeneity in entry costs plays a significant role in determining what types of CLECs enter and what types do not. We do find that two types of CLECs, one with lower entry costs than the other, display asymmetric entry behavior. These two results are as expected. The estimate of the competition effect, however, is significantly positive. We discuss several potential reasons behind this anti-intuitive result and possible solutions correspondingly. For example, we think that accounting for unobserved heterogeneity and the growing trend of this new industry will improve our estimates. The next step for us is to modify the model and the estimation strategy and produce finalized estimation results.

With estimated structural parameters, we can conduct counterfactual simulations to understand the effectiveness of subsidy policies with different focuses. We are also able to assess the impact of ignoring the option value of waiting upon policy results. Due to lack of data on potential entrants, existing research in the literature has to assume an arbitrary number of potential entrants. Moreover, these potential entrants are assumed to be short-run players, who enter or perish in any given period. The optional value of waiting for any potential entrant is essentially assumed to be 0. This is obviously not the case in many real world entry settings, including the local telephone markets we study. In our data we observe the identities of all potential entrants and the waiting time of each potential entrant before it enters. This feature provides aids for us to show bias in our parameter estimates and policy counterfactual simulations due to these unreasonable assumptions.

This paper proceeds as follows. In section 2 we briefly review literature on entry game and telecommunication industries. In section 3 we introduce data we use. Section 3 and 4 describe in detail our model and estimation strategy. Section 4 reports some preliminary results.

2 Literature Review

Estimating dynamic entry games has been the research front of empirical Industrial Organization for the last decade. The goal is to recover the distribution of sunk cost, which is usually unobserved in data but plays a pivotal role in determining market structure and competition. The distribution of sunk cost is usually recovered from the observed probability of entry conditioning on some market-level attributes determining post-entry payoffs. A handful of papers (Aguirregabiria & Mira, 2007; Bajari, Benkard & Levinson, 2007;

POB; Pesendorfer & Schmidt-Dengler, 2003) have made significant progress after Hotz & Miller (1993) by proposing a two-step estimation strategy, in which it is unnecessary to solve for equilibrium in a complex dynamic model. However, due to lack of data there is an important limitation to applications utilizing this approach (Collard-Wexler, 2008; Ryan, 2009; Dunne, Klimek, Roberts, & Xu, 2009). Researchers do not observe the number and the identities of potential entrants and therefore have to assume that potential entrants are homogenous and short-run players. Without considering the firm- and market- level heterogeneity reflected in the timing of entry, we will lose critical information in the process of recovering the distribution of entry costs. Moreover, paradoxically the players who play these dynamic games face a short-run decision of entering or perishing.[2] In fact, in most industries the decision facing a potential entrant is usually to enter or wait. This paper is trying to make amendment in these regards and, because of this amendment, bridge up the gap between the advancement in empirical dynamic entry game literature and the complex real entry game in which firms are often ex-ante heterogeneous long run players.

This project is also related to the literature on examining competition in the local telephone markets. Closest to this work are Greenstein and Mazzeo (2006), who study CLEC entry decisions into differentiated categories using a "static" entry model, and Goldfarb and Xiao (2010), who emphasize heterogeneity in managerial ability in the first year of this industry. Our paper complements theirs in that all three emphasize the importance of firm-level heterogeneity in understanding the strategic interaction of the US local telephone market.[3]

3 Data

We combine information from the CLEC annual reports by the New Paradigm Resources Group, Inc. (NPRG) and the US Census Bureau's Zip Code Business Pattern to create a panel data set on firms' entry decisions into cities, firm-level characteristics, and market attributes.

A) The NPRG CLEC annual reports

We acquired the 1998-2002 CLEC annual reports from NPRG, which contain information on the uni-

[2] Fan (2006) makes the distinction between a short-run and long-run player in a dynamic entry game.

[3] Other papers on the US telecommunication industry include Economides, Seim, and Viard (2008), Ackerberg et al (2009), Mini (2001), and Alexander and Feinberg (2004).

verse of facilities-based CLECs[4] in the United States since the passage of the1996 Telecommunication Act.[5] NPRG provides a detailed profile for every CLEC on its history, management, ownership and organization, technology, state certification, and the location of its local voice and data networks. From these profiles, we know all local voice markets that a CLEC served and the exact year of entry. We also have firm attributes such as the year the company was founded, the zip code of the headquarter, whether the company is public or private, whether the company is venture capital funded, and whether the company is a wholly owned subsidiary of a larger telecommunications company.

More importantly, a neat feature of the local telephone industry enables us to identify the set of potential entrants in each local market. CLECs must first apply for and receive certification from the state regulators before they can operate in a given city within the state. Once approved, the CLEC can operate anywhere within the state. Therefore, we identify potential entrants in a market as the set of CLECs with certification to operate in the state. It is important to note that while regulatory approval is necessary for entry, it is not sufficient. After receiving state certification, it usually takes months to years for a CLEC to commit actual entry. Some CLECs never entered a state for which they are approved to enter during the years covered in this study.

B) Market definition, characteristics, and selection

To complement CLECs' voice network data, we obtain information on location characteristics in corresponding years. The locations in the NPRG reports, i.e., the cities a CLEC provides services to, are best interpreted as the Census "place", rather than the county or metropolitan statistical area. So we choose a Census place as our market definition and use the popular name "city" henceforth.

As most of these CLECs' cater to small business clientele in the early years of the competitive telecommunication industry, we think the best proxy of market size is the number of business establishments in a city. We decompose a city into a set of Zip Code Tabulation Areas (ZCTAs) and obtain the numbers of

[4]Facilities-based CLECs are CLECs which build their own physical networks instead of resell the existing networks of the Incumbent Local Exchange Carriers (ILECs). Facilities-based CLECs are deemed by the industry experts to represent true competition to ILECs.

[5]The NPRG CLEC annual reports cover 1996 to current. 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 only published for the year of 2001 and 2002. Therefore, we are only able to assemble information on the universe of facilities-based CLECs from 1998 to 2002.

business establishments from the Zip Code Business Patterns.[6]

Lastly, we select medium-sized cities based on the number of business establishments. We only keep cities in which the number of business establishments in 1997 falls between 2,000 and 15,000. Basically, we drop the top 25 large cities because CLECs may not be able to directly compete with each other in large cities[7]; we drop a vast number of very small cities because in the early years of this industry the CLECs may not be realistically considered having true intentions to enter these cities. In the end, we have 398 medium cities remaining in our sample.

C) Summary statistics

This combination of NPRG data and city-level business pattern data has several appealing features. Not only do we have information on all entry by all firms from the effective start of the industry, we can also match this to rich data on firm characteristics and measures affectively profitability in each market.

Tables 1, 2, and 3 report descriptive statistics. Table 1 shows that these firms are generally privately owned (59% to 64% across years) and have a high variance in age (the standard deviation is about twice the mean for all years). A small proportion of these firms are subsidiaries to large corporations (26% to 32% across years) and partially funded by venture capital (17% to 22% across years). On average, these firms aim at regional or national scale, reflected by the large numbers of states which they have certification to enter. The average number of states certified has increased gradually and peaked in year 2001, when the market suffered a valuation crash. In contrast, the average number of cities a firm enters has increased more dramatically from 4 in 1998 to 12 in 2002. Lastly, the number of firms across years reflects the rapid boom and bust in the early years of this industry.

Table 2 describes the 398 mid-size cities that we use for our analysis. The number of business establishments is gradually increasing until year 2001, reflecting the ups and downs of the macro econmy in this period. The number of incumbent CLECs is gradually increasing despite the 2001 crash, reflecting the fact that this is a new and growing industry. A typical city has a large set of potential entrants but only a few incumbents (including ILECs)[8] and entrants. The number of new entrants has gone through a rapid increase

[6]The Business Census only provides this information at the city level every five years while the Zip Code Business Patterns do this annually. We use a cross reference from ZCTAs to cities based on the 2000 Business Census to perform the decomposition.

[7]Atlanta is the cutoff city based on this threshold. We will try different cutoffs to check robustness of results in revisions.

[8]Each market only has one ILEC.

and then a sharp drop, which again echoes the 2001 valuation crash.

Table 3 summarizes the data at the firm-market level to give the readers a sense of the exogenous and endogenous variables we use in our structural model. The firm attributes we select to use are those we think determine whether a firm is a high- or low-cost potential entrant. These attributes include the ownership, organizational and financial structure of the firm, as well as the age of the firm at the beginning of the CLEC industry. We also have a measure of the distance (in kilometers) between a firm's headquarter zip code and the centroid of a state [9], which is a significant determinant of sunk cost of entry as well. The endogenous variable in this study is a potential entrant' decisions on entry each year.

4 Model

The model is based on Pakes, Ostrovsky and Berry (2007) with three differences: (1) We assume that potential entrants, i.e., firms with state certification, are long-run players. Therefore, a potential entrant chooses to enter or to wait in each period, and accordingly the option value of waiting is positive. (2) We allow potential entrants to be ex ante heterogenous. While POB assume that all potential entrants draw their entry costs from the same distribution, our model allows different types of potential entrants to face different distributions. (3) In addition to firms' entry decisions, we also use their decisions on whether to become potential entrants in the model. Specifically, we use them to identify a firm's type. The key idea is that a firm decides to apply for state certification only if its expected value of being a potential entrant is sufficiently high. The expected value depends on the distribution of the entry costs, which is in turn determined by the type of the firm. A firm's decision on certification application therefore reveals its type.

4.1 Decisions on Becoming Potential Entrants

In this Section, we describe a firm's decision on obtaining state certification, that is, a firm's decision to become a potential entrant, and explain how these decisions can help us infer firms' heterogenous entry costs. When a firm decides on whether to obtain certification in a state, it compares the expected value of being a potential entrant in this state with the cost of obtaining certification. The expected value of being

[9] We use a software to calculate the distance between the zip code's and the state's centroid (in latitude and longitude).

a potential entrant, in turn, is determined by the expected payoff and the expected cost of future entry into cities located in the state. The expected payoff of entry can be captured by the state-year fixed effect, which reflects a general expectation of aggregate operating profits for a given state in a given year. We allow entry costs to depend on firm-specific characteristics. Specifically, firm attributes (denoted by z_f) such as whether the firm is private owned, whether it is a subsidiary, whether it is financed by venture capital and the age of the firm in 1998 [10] contribute to the determination of entry costs. We also consider the distance between the headquarter of the firm and the centroid of a state (denoted by d_{fs}) to capture the idea that firms face higher entry costs in more distant geography. As the firm is trading off the expected value and cost of becoming a potential entrant, the mapping between these characteristics and a firm's decisions on obtaining state certification provides information on the unobserved association of these characteristics and a firm's entry costs. To this end, we specify a Logit model:

$$\Pr\left(certification_{fst}|z_f, d_{fs}\right) = \frac{\exp\left(\xi_{st} + \varphi_1 z_f + \varphi_2 d_{fs}\right)}{1 + \exp\left(\xi_{st} + \varphi_1 z_f + \varphi_2 d_{fs}\right)}, \tag{1}$$

where ξ_{st} is the state/year fixed effect capturing the payoff of future entry into cities located in the state. With estimated $\hat{\boldsymbol{\varphi}}$, we project multiple dimensions of firm-level heterogeneity into a single dimension, which we will use in the next stage as a proxy for firms' ex-ante heterogeneity affecting entry costs. In short, $\hat{\varphi}_1 z_f + \hat{\varphi}_2 d_{fs}$ represents a firm f's "type" in state s.

4.2 Decisions on Entry into Local Markets

After obtaining state certification, a firm decides whether to enter a city inside the state in each period. At the beginning of each period, a potential entrant observes its entry cost. The entry costs, distributed independently across firms, is a firm's private information. This distribution of the entry costs depends on the type of a firm, which has been described in the previous section. To restrict the dimensions of the state space, we discretize the type $\hat{\varphi}_1 z_f + \hat{\varphi}_2 d_{fs}$ into two categories: type 1 and type 2. Given that firm attributes are publicly observed by all firms, the number of potential entrants of each type in a city is common knowledge.

A potential entrant then compares its entry costs with its expected present value of the profit stream

[10] 1998 is the first year of our data, when NPGR starts to collection information on the universe of CLECs.

it will receive upon entry. The profit of a firm operating in a city, which is assumed to be identical for all firms, depends on the size of the market and the number of incumbents. Let m_{ct} be the market size and n_{ct} the number of incumbents in city c in year t. The market size m_{ct} evolves exogenously according to a first-order Markov chain process, whereas the number of incumbents n_{ct} evolves endogenously. We assume that the one-period profit function is of the following parametric form:

$$\pi\left(n_{ct}, m_{ct}\right) = m_{ct}^{\alpha} / \left(1 + n_{ct}\right)^{\gamma}, \tag{2}$$

where α (the market size effect) and γ (the competition effect) are parameters to be estimated.

We now describe the dynamic game of entry. At the beginning of each period, a potential entrant observes its own cost of entry and knows the types of other potential entrants. If a potential entrant decides to enter the market, it will start to earn profits next period after paying an up-front cost of entry this period. Therefore, the value of entry is the expected value of being an incumbent next period. 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 the following two reasons. First, in our data an incumbent always stays in the local market until the firm exits as a whole, which is consistent with the observation that the variable cost of maintaining operation is very low. Second, during the time span covered in study the large-scale exiting (often referred to as telecom "meltdown") is largely due to "exogenous" shocks related to macro economy and FCC re-regulation. Therefore, we assume that a firm exits exogenously with probability p_{xt} in year t. All firms face the same probability of exit. However, in line with the data this probability varies across years. We assume that p_{xt} is centered at p_x, which is common knowledge among all firms and will be estimated from the data.

As specified above, the distribution of a firm's entry costs depends on the type of the firm and the profit depends on (m_{ct}, n_{ct}). Therefore, the number of potential entrants of each type (T_{1ct}, T_{2ct}) as well as (m_{ct}, n_{ct}) are relevant state variables for firms' decisions. For notation simplicity, we suppress subscripts c, t hereafter in this section. From now on, we use the phrase "geographic state" for an actual state, such as California, and the word "state" for a state in this dynamic model.

A potential entrant's decision is based on the comparison of the value of entry net of entry costs with the value of waiting. As explained, the value of entry is the expected value of being an incumbent next period.

The value of waiting is the expected value of being a potential entrant 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) and $V^E(m,n,T_1,T_2,\tau,\zeta)$ that of a potential entrant of type τ with entry costs ζ. Therefore,

$$\begin{aligned} V^E(m,n,T_1,T_2,\tau,\zeta) \quad = \quad & \max\Big\{-\zeta+\beta(1-p_x)E^e_{(m',n',T_1',T_2')|(m,n,T_1,T_2,\tau)}V^I(m',n',T_1',T_2'), \\ & \beta(1-p_x)E^w_{(m',n',T_1',T_2')|(m,n,T_1,T_2)}E_{\zeta'|\tau}V^E(m',n',T_1',T_2',\tau,\zeta')\Big\}, \end{aligned} \tag{3}$$

where $E^e_{(m',n',T_1',T_2')|(m,n,T_1,T_2,\tau)}$ is the potential entrant's expectation *conditional* on itself *entering* at state (m,n,T_1,T_2) when its own type is τ, $E^w_{(m',n',T_1',T_2')|(m,n,T_1,T_2,\tau)}$ is that *conditional* on itself *waiting* and $E_{\zeta'|\tau}$ denotes the expectation of its entry cost of next period. Note the difference of the value functions and the expectation operators in the first term and the second term in (3). For a long-run player, the value of waiting is the expected value of being a potential entrant next period. Consequently, the expectation is taken over the distribution of next period's entry costs ζ' conditional on its own type τ and over the state next period (m',n',T_1',T_2') conditional on itself waiting. In (3), β is the discount factor and p_x is the expected probability that a firm will exit. For notation simplicity, we define $\delta=\beta(1-p_x)$ as the discount factor adjusted by the probability of exit. Similarly, the value of incumbent in (3) is

$$V^I(m,n,T_1,T_2)=\pi(m,n)+\delta E_{(m',n',T_1',T_2')|(m,n,T_1,T_2)}V^I(m',n',T_1',T_2'). \tag{4}$$

Following Hotz and Miller (1993) and using the assumption that entry costs and sell-off value are i.i.d., we define the value of entry for potential entrants at state (m,n,T_1,T_2) before they observe their entry costs as

$$VE(m,n,T_1,T_2,\tau)=E^e_{(m',n',T_1',T_2')|(m,n,T_1,T_2)}V^I(m',n',T_1',T_2'); \tag{5}$$

and the value of waiting as

$$VW(m,n,T_1,T_2,\tau)=E^w_{(m',n',T_1',T_2')|(m,n,T_1,T_2)}E_{\zeta'|\tau}V^E(m',n',T_1',T_2',\tau,\zeta'). \tag{6}$$

Note that the value of waiting depends on a firm's own type τ because it determines the distribution of the firm's entry costs next period.

Plugging (3) into (6) gives the value of waiting:

$$
\begin{aligned}
& VW(m,n,T_1,T_2,\tau) && (7)\\
= \;& E^{w}_{\left(m',n',T_1',T_2'\right)|(m,n,T_1,T_2)} E_{\zeta'|\tau} \max\left\{\delta VW\left(m',n',T_1^{'},T_2^{'},\tau\right), \delta VE\left(m',n',T_1^{'},T_2^{'},\tau\right)-\zeta'\right\}\\
= \;& E^{w}_{\left(m',n',T_1',T_2'\right)|(m,n,T_1,T_2)} \Big\{\left(1-p^{e}\left(m',n',T_1',T_2',\tau\right)\right)\delta VW\left(m',n',T_1^{'},T_2^{'},\tau\right)+\\
& p^{e}\left(m',n',T_1',T_2',\tau\right)\left(\delta VE\left(m',n',T_1',T_2',\tau\right)-E\left[\zeta|\zeta<\delta VE\left(m',n',T_1',T_2',\tau\right)-\delta VW\left(m',n',T_1^{'},T_2^{'},\tau\right)\right]\right)\Big\},
\end{aligned}
$$

where $p^e(m',n',T_1',T_2',\tau)$ is the probability of entry at state (m',n',T_1',T_2') for a firm of type τ.

We close the model by specifying the density of the entry cost (ζ). Following POB, we assume that the density for firms of type 1 or 2 is

$$
f_j(\zeta)=\mu_j^2\left(\zeta-\frac{1}{\mu_j}\right)\exp\left(-\mu_j\left(\zeta-\frac{1}{\mu_j}\right)\right),
$$

where μ_j, $j=1,2$ are parameters to be estimated.

5 Estimation

Estimation is carried out in two main steps. In the first step, we "determine" the type of each firm. To this end, we estimate (φ_1,φ_2) in (1). We then compute $\hat{\varphi}_1 z_f+\hat{\varphi}_2 d_{fs}$ for each firm and divide the value into two groups: a firm f is of type 1 in geographic state s if and only if $\hat{\varphi}_1 z_f+\hat{\varphi}_2 d_{fs}$ is below the mean; otherwise, this firm is of type 2 in geographic state s.

In the second step, we estimate the parameters in the profit function (α,γ), the parameters in the entry costs functions (μ_1,μ_2) and the discount factor adjusted by exiting probability $\delta=\beta(1-p_x)$. We first estimate p_x directly from data[11] and we choose β according to the real interest rate. The rest of the parameters are estimated following the procedure in POB with only one modification: we need to consistently estimate the value of waiting as well as the value of entry.

To estimate these values, we rewrite equations (3), (5) and (7) in vector forms. The state in this study is a quadruple (m,n,T_1,T_2). Suppose there are J distinct states. We denote the ith state by (m_i,n_i,T_{1i},T_{2i}). With slight abuse of the notation, let $V^I(\alpha,\gamma,\delta)$ be the vector with $V^I(m_i,n_i,T_{1i},T_{2i})$ as its ith element.

[11] The average exit probability at the firm level is about 23.9% from 1999 to 2002.

Here we add parameters as inputs of $V^I(\alpha,\gamma,\delta)$ explicitly. Similarly, the ith element of the vector $\pi(\alpha,\gamma)$ is $\pi(m_i,n_i)$. With these notations, we can rewrite equation (3) in vector forms as

$$V^I(\alpha,\gamma,\delta)=\pi(\alpha,\gamma)+\delta MV^I(\alpha,\gamma,\delta), \tag{8}$$

where M is the transition probability matrix, i.e., 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 will be estimated directly from data.

Similarly, we define the vectors of the entry value of both types as $VE_1(\alpha,\gamma,\delta)$ and $VE_2(\alpha,\gamma,\delta)$. Their ith elements are $VE(m_i,n_i,T_{1i},T_{2i},\tau=1)$ and $VE(m_i,n_i,T_{1i},T_{2i},\tau=2)$, respectively. Analogously, we define the vectors $VW_1(\alpha,\gamma,\delta)$ and $VW_2(\alpha,\gamma,\delta)$ as the value of waiting for type-1 and type-2 firms. Then, equations (5) and (7) for $\tau=1,2$ can be rewritten as

$$VE_\tau(\alpha,\gamma,\delta)=M_\tau^eV^I(\alpha,\gamma,\delta), \tag{9}$$

and

$$\begin{aligned}VW_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right) &= M_\tau^w\left\{(1-p^e)\,\delta VW_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right)\right.\\ &\quad \left.+p^e\left(\delta VE_\tau(\alpha,\gamma,\delta)-E\left[\zeta|\zeta<\delta VE_\tau(\alpha,\gamma,\delta)-\delta VW_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right);\mu_\tau\right]\right)\right\},\end{aligned} \tag{10}$$

where M_τ^e is a matrix whose 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}) conditional on a type-τ potential entrant *entering* and M_τ^w is the same matrix conditional on *waiting*.

To estimate $VE_\tau(\alpha,\gamma,\delta)$ and $VW_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right)$, we need consistent estimates of the transition probability matrices M, M_τ^e and M_τ^w. We use their empirical counterparts following POB. See Appendix A for the details on $\hat{M}$, $\hat{M}_1^e$, $\hat{M}_2^e$, $\hat{M}_1^w$,and $\hat{M}_2^w$.

With $\hat{M}$, $\hat{M}_1^e$, $\hat{M}_2^e$, $\hat{M}_1^w$,and $\hat{M}_2^w$ estimated, the estimate of the value of entry is given by

$$V\hat{E}_\tau(\alpha,\gamma,\delta)=\hat{M}_\tau^e\hat{V}_\tau^I(\alpha,\gamma,\delta) \tag{11}$$

where

$$\hat{V}_\tau^I(\alpha,\gamma,\delta)=\left(I-\delta\hat{M}\right)^{-1}\pi(\alpha,\gamma),$$

and I is an identity matrix. Similarly, $V\hat{W}_\tau(\alpha,\gamma,\mu_\tau,\delta)$ is the fixed point of (10) when $VE_\tau(\alpha,\gamma,\delta)$, M_τ^w and p^e are replaced by their empirical counterparts. Note that the RHS of equation (10) is a contraction mapping of $VW_\tau\left(\alpha,\gamma,\mu_j,\delta\right)$ because ζ is assumed to be a log concave random variable (with gamma distribution) and therefore, $0\leq\frac{\partial E(\zeta|\zeta<d)}{\partial d}\leq 1$ (see Proposition 1 of Heckman and Honore (2003)).

Once we have consistent estimates of the values of entry and waiting, we can get consistent estimates of the probabilities of entry for given parameters. The probability of entry at state (m_i,n_i,T_{1i},T_{2i}) is the probability that the entry costs are smaller than the difference between the value of entry and the value of waiting at this state. In other words, the vector of probabilities of entry for a type-τ firm is

$$\hat{p}^e_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right)=\Gamma\left(2,\lambda\delta\left[V\hat{E}_\tau\left(\alpha,\gamma,\delta\right)-V\hat{W}_\tau\left(\alpha,\gamma,\mu_\tau,\delta\right)\right]\right)$$

where $\Gamma\left(a,x\right)=\int_0^x t^{a-1}e^{-t}dt$ is the incomplete gamma function.

We estimate four parameters: distribution parameters $\left(\mu_1,\mu_2\right)$ and model parameters (α,γ). These parameters are estimated with the Generalized Methods of Moments. Five moment conditions are used. These conditions posit that prediction errors are uncorrelated the explanatory variables. Specifically, the prediction errors of entry at state (m_i,n_i,T_{1i},T_{2i}) for type-1 firms and type-2 firms are respectively

$$\begin{aligned}\varepsilon_1\left(m_i,n_i,T_{1i},T_{2i}\right) &= \hat{p}^e_{1i}\left(\alpha,\gamma,\mu_\tau,\delta\right)T_{1i}-\frac{1}{\#K\left(m_i,n_i,T_{1i},T_{2i}\right)}\left(\sum\nolimits_{k\in K(m_i,n_i,T_{1i},T_{2i})}e_{1(k)}\right)\\ \varepsilon_2\left(m_i,n_i,T_{1i},T_{2i}\right) &= \hat{p}^e_{2i}\left(\alpha,\gamma,\mu_\tau,\delta\right)T_{2i}-\frac{1}{\#K\left(m_i,n_i,T_{1i},T_{2i}\right)}\left(\sum\nolimits_{k\in K(m_i,n_i,T_{1i},T_{2i})}e_{2(k)}\right),\end{aligned}$$

where $K\left(m_i,n_i,T_{1i},T_{2i}\right)$ is the collection of all market/years whose states are (m_i,n_i,T_{1i},T_{2i}) and $\#K\left(m_i,n_i,T_{1i},T_{2i}\right)$ is the cardinality of the set K. Since the predication errors depend on the state interacting with the type of a firm, we define a quintuple $(m_h,n_h,T_{1h},T_{2h},\tau_h)$. Given that there are J distinct states defined by (m,n,T_1,T_2), there exist $2J$ such quintuples. The moment conditions are

$$\frac{1}{2J}\sum_{i=1}^{2J}\varepsilon_\tau\left(m_i,n_i,T_{1i},T_{2i},\tau_i\right)\cdot x_i=0 \text{ for } x_i=1,\ m_i,\ n_i,\ T_{1i},\ T_{2i},\ \tau_i-1. \tag{12}$$

These conditions hold because the market size, the number of incumbents, the number of potential entrants of either type and the type of a firm are all assumed exogenous. Note that the value of the type τ is either 1 or 2. We use $\tau-1$ to define a dummy according to the firm type.

6 Preliminary Results

We have obtained some preliminary results using the model and estimation strategy described above. In the follows we discuss these results and their potential problems. Note that we drop year 1998 from our data and only use year 1999 to 2002 for our structural estimation. This is because we do not observe the number of CLEC incumbents in 1998, the first year we have data. Consequently, if we were to include 1998, we would have to assume that all firms in 1998 were new entrants. This might lead to an underestimate of incumbents and an overestimate of new entrants because CLECs have started to enter right after the 1996 Telecom Act.

A) State Certification regression results

Table 4 shows results from the reduced-form regression of established CLECs' decisions to obtain state certification for the first time. The first two columns are OLS and Logit regressions and the last two are their counterparts with state-year fixed effects included. Here we use state-year fixed effects to capture a general expectation of aggregate operating profits for a given state in a given year. The results in the last two columns clearly indicate that observed firm attributes are key determinants of firms' decisions to obtain state certification. CLECs that are privately owned or subsidiaries to larger firms are significantly less likely to obtain state certification. However, those funded by venture capital or having a longer operating history show the opposite pattern. Furthermore, CLECs are significantly less likely to obtain state certification in states farther from their headquarters, suggesting that CLECs may have higher entry costs into a more distant geography.

As described in the section on firms' decisions to become potential entrants, we use results from certification regression (Column 4, Table 4)to obtain $\hat{\varphi}_1 z_f + \hat{\varphi}_2 d_{fs}$, which determines a firm$f$'s "type" in state s. We then divide firms into two categories according to this measure: type 1 and type 2 —- CLECs with $\hat{\varphi}_1 z_f + \hat{\varphi}_2 d_{fs}$ smaller than the mean are labeled as type 1 and those with this measure larger than the mean are considered type 2.[12] As explained in the model, any market-year combination can now be characterized by four state variables: market size (log of the number of business establishments), number of incumbents (including both ILECs and CLECs), number of type-1 potential entrants, and number of type-2 potential

[12]Note when we calculate the mean we disregard 1998 data as we do not use 1998 in estimation.

entrants. A potential entrant's decision on entry also depends on its own type. Table 5 summarizes these 5 state variables, which are used to explain entry behavior of CLECs from 1999 to 2002.

B) Estimates of structural parameters

Table 6 report preliminary results on the 4 structural parameters in the model. The left panel reports results with equal weight on all states (i.e. the aforementioned four-dimensional state of a year-market). The right panel reports results with different weights on states, where the weight is the number of year-markets with each state. That is, in the right panel we put more weight on states with more year-market observations. We can see that estimates in the two panels are very close, however, 3 out of 4 parameters are estimated without much precision with equal weight. Moving from the left panel to right we gain much significance in our estimates, suggesting that the number of year-market observations in a state adds relevant information to the estimation. We thus choose the results with weighted states as our baseline results.

From the right panel results, two of them are as expected. First, market size, measured by the number of business establishments, affects the operating profit of a CLEC positively once it becomes an incumbent. Therefore the larger the market is, the more likely is a CLEC to enter. This befits the traditional wisdom in line of Bresenhan and Reiss (1991) that a larger market size is necessary to support more competitors. This also implies that smaller markets may be kept from having a competitive structure, as these markets do not have sufficient demand to attract competitive entry. Second, firm-level heterogeneity in entry costs plays a significant role in determining a firm's entry into a local market. We find that potential entrants labeled as "type 1" have higher entry costs that the "type 2" potential entrants,[13] and are therefore more likely to enter ceteris paribus. Both types' entry cost parameter is estimated with statistical significance.

The competition effect, however, is not easy to be rationalized. We expected γ to be positive to capture the negative competition effect, which suggests more incumbents in a market will erode average profitability. This is the result from traditional wisdom and a vast number of studies on market structure and competition.

There are a few possible reasons for this anti-intuitive result. First, there may exist market-level unobserved heterogeneity which is not captured by our single variable measuring market demand —- the size of the market. For example, the income and education level of the local population may be strong factors in

[13]The mean of entry cost is $\frac{2}{u_j}$.

attracting CLEC entry. If this unobserved market-level heterogeneity is persistent over time, it is positively correlated with the number of incumbents, which may lead to an upward bias in the effect of latter on profitability. We have tried dropping the number of incumbents from instruments in our GMM estimation; but results stay similar and we suspect this is because the number of potential entrants suffers the same endogeneity problem. To deal with this problem, we can either incorporate more market characteristics into our model and estimation or allow an unobserved state affecting firms' decision. The latter solution is computationally intensive. Regarding the former solution, we are currently investigating relevant market attributes. Second, the model we have set up has the exogenous variables following a stationary process while the real market may be a non-stationary process. As the competitive local telephone market is new and growing rapidly in the first decade after the 1996 Telecom Act, a non-stationary entry model may fit the data better. We can potentially set up a model that incorporates the growing feature at the beginning of this industry. The estimation strategy described above can still be used as long as we add the restriction that the process is stationary from certain year on. Both reasons call for a major revision of the model and we are still trying to figure out the best way to proceed. Adding persistent unobserved heterogeneity or adding a non-stationary process are both challenging and in the very front of research on empirical dynamic models.

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A Appendix: Estimation of the Transition Probability

The estimate of the unconditional transition probability is

$$\hat{M}_{ij} = \frac{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} 1\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = (m_j, n_j, T_{1j}, T_{2j})\right\}}{\#K\left(m_i, n_i, T_{1i}, T_{2i}\right)}, \tag{13}$$

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

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

$$\hat{M}^e_{1ij} = \frac{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} e_{1(k)} 1\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = (m_j, n_j, T_{1j}, T_{2j})\right\}}{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} e_{1(k)}}, \tag{14}$$

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 potential entrants do enter over the probability that e_1 out of T_1 potential entrants enter. Similarly,

$$\hat{M}^e_{2ij} = \frac{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} e_{2(k)} 1\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = (m_j, n_j, T_{1j}, T_{2j})\right\}}{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} e_{2(k)}}. \tag{15}$$

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 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

$$\hat{M}^w_{1ij} = \frac{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} 1\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = (m_j, n_j, T_{1j}, T_{2j})\right\}}{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} \left(T_{1i} - e_{1(k)}\right)}, \tag{16}$$

and

$$\hat{M}^w_{2ij} = \frac{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} 1\left\{\left(n_{(k+1)}, m_{(k+1)}, T_{1(k+1)}, T_{2(k+1)}\right) = (m_j, n_j, T_{1j}, T_{2j})\right\}}{\sum_{k\in K(m_i,n_i,T_{1i},T_{2i})} \left(T_{2i} - e_{2(k)}\right)}. \tag{17}$$

Table 1: Summary Statistics by Firm-Year

	1998	1999	2000	2001	2002
Variables			Mean(std. dev.)		
Privately owned	0.639	0.595	0.581	0.598	0.640
	(0.483)	(0.493)	(0.495)	(0.493)	(0.483)
Subsidiary	0.320	0.284	0.282	0.268	0.302
	(0.469)	(0.453)	(0.452)	(0.445)	(0.462)
Financed by venture capital	0.176	0.190	0.222	0.206	0.221
	(0.383)	(0.394)	(0.418)	(0.407)	(0.417)
Firm age	8.897	8.526	11.020	13.330	15.267
	(20.207)	(16.548)	(20.769)	(23.416)	(24.695)
# states certified	9.021	10.810	12.872	13.134	12.814
	(10.941)	(11.859)	(12.103)	(12.271)	(12.326)
# cities to enter	100.381	111.138	127.111	122.196	123.209
	(108.026)	(112.298)	(116.250)	(114.220)	(117.659)
# cities entered	3.938	8.034	10.923	11.330	11.837
	(8.092)	(14.935)	(16.345)	(14.243)	(15.860)
# observations (firm)	97	116	117	97	86

Table 2: Summary Statistics by Market-Year

	1998	1999	2000	2001	2002
Variables			Mean(std. dev.)		
# business establishments	3999.083	4022.314	4020.987	4010.884	4091.384
	(2435.669)	(2438.102)	(2437.705)	(2426.277)	(2468.029)
# incumbents	n.a.	0.819	1.598	1.960	2.224
		(1.429)	(2.185)	(2.603)	(2.776)
# potential entrants	24.465	31.573	35.769	27.822	24.400
	(7.481)	(9.616)	(10.493)	(10.000)	(8.577)
# entrants	0.960	1.523	1.613	0.802	0.334
	(1.618)	(1.798)	(2.268)	(1.150)	(0.599)
# observations (city as market)			98		

Table 3: Summary Statistics by Firm-Market

Variable	Mean	Std. Dev.	Min	Max
Firm attributes				
Privately owned	0.417	0.493	0	1
Subsidiary	0.184	0.387	0	1
Funded by venture capital	0.197	0.398	0	1
Firm age in 1998	10.443	22.000	0	114
Firm-market attribute				
Distance from firm headquarter to market (in 1000 kms)	1.470	1.107	0.0002	4.390
Market-level state variables				
# business establishments (in 1000s)	3.998	2.383	1.648	14.136
# incumbents	1.4	2.305	0	17
# potential entrants	32.217	10.171	5	55
Firm-market endogenous decision				
Enter	0.035	0.183	0	1
# observations (firm-market)		59950		

Table 4: Firms' Decisions to Obtain State Certification

	(1) OLS	(2) Logit	(3) OLS	(4) Logit
dummy: =1 if the CLEC is privately owned	-0.123	-1.326	-0.117	-1.365
	(0.005)***	(0.056)***	(0.005)***	(0.059)***
dummy: =1 if the CLEC is a subsidiary of a firm	-0.085	-1.077	-0.084	-1.128
	(0.005)***	(0.060)***	(0.004)***	(0.063)***
dummy: =1 if the CLEC is funded by venture capital	-0.005	0.195	0.011	0.302
	-0.006	(0.068)***	(0.005)**	(0.071)***
log of firm age in 1998	0.001	0.003	0.006	0.080
	(0.003)	(0.026)	(0.002)***	(0.027)***
distance b/w headquarter zip code and state	-0.052	-0.696	-0.066	-0.884
	(0.002)***	(0.030)***	(0.002)***	(0.036)***
Constant	0.303	-1.105	0.312	
	(0.007)***	(0.080)	(0.008)***	
State-year fixed effects included?	No	No	Yes	Yes
Number of state-year fixed effects			235	227
R-squared	0.07		0.08	
Log Likelihood		-6050.482		-4972.045
Observations	20531	20531	20531	19933

Standard errors in parentheses* significant at 10%; ** significant at 5%; *** significant at 1%

Table 5: Description of Variables Used in Structural Estimation

Variable	Mean	Std. Dev.	Min	Max
State variables				
log(# business establishments)	8.145	0.458	7.407	9.554
# incumbents	2.485	2.217	1	18
# type 1 potential entrants	15.210	8.463	0	31
# type 2 potential entrants	17.419	7.664	0	36
Potential entrant's type	1.541	0.498	1	2
Firm-market endogenous decision				
Enter	0.036	0.186	0	1
# observations (firm-market)		47586		

Table 6: Preliminary Results

	Equal Weights on States		Weighted States	
Structural parameter	Estimate	Std. error	Estimate	Std. error
α (market size effect)	0.303	0.533	0.871	0.360**
γ (competition effect)	-1.321	0.607**	-0.801	0.290***
μ_1 (entry cost parameter for type 1)	0.035	0.040	0.024	0.013*
μ_2 (entry cost parameter for type 2)	0.037	0.042	0.026	0.014*
# observations (firm-market)		47586		
Standard errors in parentheses* significant at 10%; ** significant at 5%; *** significant at 1%				