Entry and Exit in the United States Local Daily Newspaper Industry

by
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1 Introduction

Your local newspaper has been in trouble for a long time. Over the past two decades, the popularization of internet media has seen the proliferation of inexpensive or free substitutes for the information provided in daily newspapers. However, this was not the first time technological change challenged the profitability of print journalism. During the 1930s, radio's rapid rise to prominence offered consumers an alternative source of daily news. Television's ascent to ubiquity in the 1950s and early 1960s further intensified the competition that other forms of media posed to newspapers. These new forms of media competed with print newspapers on two fronts: on one hand, they provided substitutes for the informational and entertainment content of print newspapers; and on the other, they offered alternative for retailers looking to purchase advertising space.

A priori, one might reasonably expect the effect on local newspaper of the emergence of these technologies to have been devastating. In some respects, the data bear out this expectation, as show in Figure 1.[1] The number of active daily newspapers peaked in the 1910s, and has gradually declined since that time. In spite of the decline in the number of active newspapers after 1910, total circulation continued to climb until the 1960s, after which time it remaining mostly stagnant until beginning to fall in the 1990s. Changes in the per capita circulation of daily newspapers over time suggest that population growth was the driving force behind the growth in newspaper circulation after 1950. Although per capita circulation continued to grow even after the number of active newspapers began to decline, it reached a peak around 1950 and has fallen at a nearly constant rate since that time. A similar pattern is observable in the average per capita circulation of active newspapers.

[1]The construction of these measures and the collection of these data are discussed in depth in Section 3.

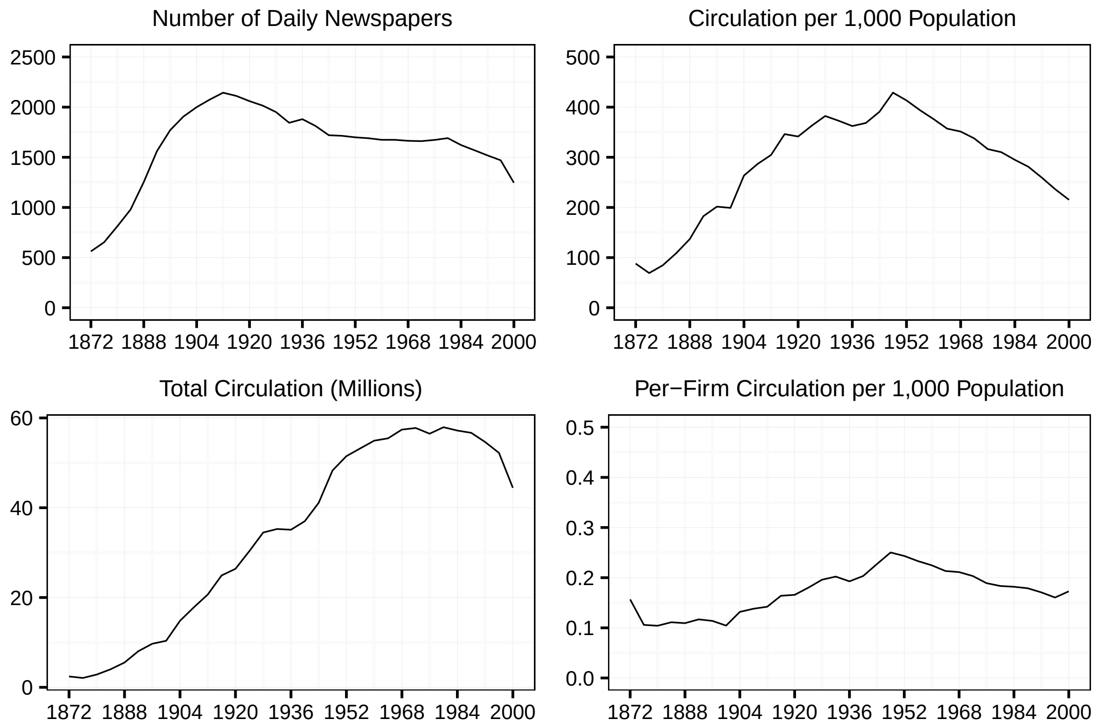

Figure 1: Changes in the Local Daily Newspaper Industry, 1872 - 2000

Local newspaper markets are oligopolistic by nature. Newspapers have large fixed costs of entry, and established newspapers are able to attract advertising revenue more easily compared to relatively smaller entrants. This limits the ability of entrants to compete with incumbent newspapers on subscription prices, as incumbents are able to 'subsidize' subscription prices using ad revenues. The tendency toward concentration in local newspaper markets has only increased over time, as declining demand for newspapers has coincided with an increased prevalence of monopolized local newspaper industries. The total number of counties with local newspapers consistently increased throughout the period between 1872 to 2000, driven exclusively by growth in the number of counties with a monopolistic local newspaper. Growth in the number of monopolistic local newspapers, in turn, is the result both of the exit of newspapers from counties with two or more newspapers and of the entry of newspapers into counties that previously did not have local newspapers.

I study patterns of entry and their competitive effects in local newspaper markets using an adaptation of the empirical models of endogenous market structure initially developed in Bresnahan and Reiss (1991) and Bresnahan and Reiss (1994). First, I estimate a static version of the model that predicts the number of newspapers active in a given local market at a given time by embedding a reduced-form estimate of per firm variable profits within a game-theoretic model of the equilibrium conditions characterizing a given market structure. Using the coefficient estimates from this model, I estimate what Bresnahan and Reiss termed the 'entry threshold': the minimum market size necessary to justify the entry of a given number of firms into a market.

The estimates from the static framework suggest that the average minimum market size required to support the entry of a single newspaper has fallen slightly over time, while the thresholds for two or more newspapers have increased. I find that between 1872 and 1908, a population of 25,000 was required to support the entry of a monopolist, while 49,000 and 116,000 people were required for a second and third newspaper, respectively, to find entry profitable. Between 1976 and 2000, the monopoly entry threshold was roughly 15,000 persons, while the population sizes needed for duopolists and triopolists to enter profitably were 240,000 and 837,000, respectively. Over time, the per firm population required to support the entry of two newspapers has increased considerably relative to the monopoly entry threshold. Conversely, the per firm market size required to support three or more newspapers has consistently remained at about double the level required to support two newspapers. This suggests a sizable competitive effect of entry, as firms' per capita variable profits continue to decline with the entry of the third newspaper.

Next, I estimate a pseudo-dynamic version of the model that predicts the number of newspapers in a local market conditional on the number of firms active in that market during the previous period. Conditioning the model's prediction of the current-period market structure on the market structure observed during previous periods makes it possible to allow incumbents and entrants to earn different profits in the empirical model. Using this information, it is possible to estimate two types of demand thresholds: the entry threshold and the exit threshold. In this context, the entry threshold is defined as the minimum market size required to justify the entry of an N^{th} firm into a market with N incumbents. The exit threshold is defined as the minimum market size required for N to profitably remain in the market. Below this market size, we would expect one or more incumbents to shut down.

When the profits of entrants are incumbents are allowed to differ in the model, the estimated results suggest that entry thresholds increased for any number of firms over the course of the sample period. Entry thresholds for duopolistic and triopolistic local newspaper markets increased by a relatively larger amount than those for local monopolies. The monopoly exit threshold appears to have declined over time, and has essentially remained constant since the period between 1912 and 1940. Duopoly and triopoly exit thresholds have increased over time, although the increase is smaller than that seen in entry thresholds.

An extension of the model that allows sunk costs to increment with the number of incumbents in market provides evidence that sunk costs are increasing in the order of the entry. However, I find that the size of the gap between entry and exit thresholds is approximately constant in the log of county population. As such, the estimates of entry and exit thresholds yielded by the baseline sunk cost specification appear to be appropriate, in spite of the finding that the sunk cost of entry is increasing in the number of incumbents.

The pseudo-dynamic variant of the model tends to predict higher entry thresholds compared to the model without sunk costs. This is because the entry threshold estimates yielded by the sunk cost framework are actually weighted averages of entry and exit thresholds, resulting in a downward-biased estimate of the entry threshold. The magnitude of the bias increases in the size of the sunk costs of entry. During the 1872 to 1908 period, this bias is relatively small, since entry rates in the local daily newspaper industry were high during this time. When entry rates slow during later periods, incumbents become relatively more prevalent, increasing the size of the downward bias in the entry threshold estimates. This result suggests the static Bresnahan-Reiss framework is not well suited to measuring entry thresholds in industries in which entry rates are low and there are sizable sunk costs of entry.

The results of the sunk cost specification show that the gap between entry and exit thresholds has widened over time. I find evidence suggesting that the popularization of television during the 1950s contributed to the widening of this gap, as the popularization of television increased the size of newspapers' sunk costs of entry (increasing entry thresholds) and decreased their variable profits (increasing entry and exit thresholds). Between 1952 and 1960, the gap between entry thresholds at the 100^{th} and 0^{th} percentiles of television diffusion in terms of log population was 3, while the gap between exit thresholds was 2.7. I also find that the popularization of radio in the 1930s and

1940s decreased newspapers' variable profits, thus increasing entry and exit thresholds, but I find little evidence to suggest that radio's popularization affected newspapers' sunk costs separately from variable profits, unlike television.

The paper proceeds as follows. Section 2 reviews the literature concerning the industrial organization of the newspaper industry and empirical models of entry. Section 3 describes the data used in the study. Section 4 provides a descriptive analysis of changes in the market structure of local newspaper markets, and their effect on competition between local newspapers. Section 5 reviews the principles underlying the of static and pseudo-dynamic models of entry developed in Bresnahan and Reiss (1991) and Bresnahan and Reiss (1994), respectively, and explains how I adapt them for use in the context of local daily newspaper markets. The results of estimating these models are reported and discussed in Section 6. Section 7 concludes.

2 Literature Review

Empirical models of market structure typically rely on cross-sectional data containing information on the number of firms in a given geographic market and other market-specific observations, such as demographic characteristics and land prices. Most research in this area focuses on the development of discrete-choice game-theoretic models that can feasibly be taken to the data. These models typically embed reduced-form estimates of profits and costs in a game-theoretic framework. This semi-structural approach allows researchers to circumvent the limitations imposed by the relative paucity of firm-level data available to them, especially concerning key variables such as profits and costs. Using these estimates, the researcher can then make inferences regarding the importance of fixed costs and the rate at which variable profits fall as the intensity of competition increases.

The seminal empirical work on entry in markets with homogeneous firms is Bresnahan and Reiss (1991). This paper, along with its precedents Bresnahan and Reiss (1988) and Bresnahan and Reiss (1990), introduces the concepts of entry thresholds: the smallest market size capable of supporting a given number firms. Using information on the observed market structure in geographically isolated markets along with some demographic and geographic characteristics of these markets, Bresnahan and Reiss parameterize an ordered probit model of the number of firms active in the market. An important advantage of this approach is that it does not require the econometrician to have detailed information concerning price, cost, and demand in the markets under study: it is only necessary to know the number of firms and some market characteristics with which to estimate a reduced-form profit function. The estimates of entry thresholds for different numbers of firms provide information about how the degree of competition in a market changes with the number of firms. If entry thresholds increase more than proportionately with the number of firms, this suggests that firms earn less profits per consumer as the number of competitors increase, which in turn implies that entry increases the degree of competition in the market. They estimate this model for a variety of retail and professional industries, and find that the competitive effects of entry appear to dissipate once three firms are active in the market.

Bresnahan and Reiss (1994) extends the concept of demand thresholds by introducing the notion of the exit threshold, defined as the minimum market size necessary to support a given number of incumbent firms in a market. Entry and exit thresholds differ in cases where firms pay per-period fixed costs in addition to their sunk costs of entry. Hence, entry thresholds must be at least as large as exit thresholds, as the former requires demand to be sufficiently large for a firm to justify the expenditure of the per-period fixed cost and the sunk cost of entry, while the latter only needs to be large enough for incumbents to recover their per-period fixed costs. Bresnahan and Reiss use this concept to formulate an empirical model of market structure based on a two-period model of the entry decision, in which the profits of entrants and incumbents differ due to the presence of sunk costs of entry. Applying this model to a dataset of dentists in geographically isolated markets, they find that sunk costs are increasing in market size and exit thresholds are significantly lower than entry thresholds in this industry.

An important limitation of the empirical framework established in Bresnahan and Reiss (1991) is the assumption of homogeneity among firms. Beside the obviously counter-factual nature of this assumption for many industries, it has the side effect of making the model unable to rationalize cases in which entry and exit occurs simultaneously. In the presence of symmetry and the absence of firm-specific shocks, it does not make sense for identical firms to make different decisions. Unfortunately, it is quite difficult to relax this assumption in practice, even in a static framework. The main reason for this difficulty is that the relaxation of the homogeneity assumption leads to models with either multiple pure-strategy equilibria or no pure-strategy equilibrium at all. This significantly complicates the estimation process. Maximum likelihood estimation, for example, requires each parameter vector to generate a single prediction, which precludes the estimation of models with multiple equilibria.

The literature offers two main methods for relaxing the assumption of homogeneous products in empirical models of endogenous market structure. One approach imposes further structure on the entry game in order to ensure uniqueness, while the other allows for the existence of information asymmetries between firms, which can allow the researcher to avoid the complications resulting from multiple equilibria under certain circumstances. Berry (1992) provides an influential and early example the former approach by allowing for both observed and unobserved firm heterogeneity in his study of airlines' decisions of whether to serve a specific flight route using a sequential-move entry model. In this model, firms offer homogeneous products but have heterogeneous fixed costs of entry. To ensure the existence of a unique equilibrium, entry is assumed to proceed sequentially, with the firms moving in order of profitability (determined by their idiosyncratic fixed costs). The data requirements of this model can be difficult to meet in many industries, as it requires the econometrician to have information about potential entrants. Information of this sort may be available for industries in which only a few firms compete in many separate geographic markets, as is the case in the airline industry, but it is likely impossible to obtain such information for industries with large numbers of firms.

The empirical model of oligopoly market structure model developed in Mazzeo (2002), can be thought of as taking the reverse of the approach used in Berry (1992). While Berry (1992) assumes firms are heterogeneous pre-entry and homogeneous post-entry, Mazzeo (2002) assumes firms are homogeneous pre-entry (i.e., there are no firm-specific shocks) and heterogeneous post-entry. In Mazzeo's framework, firms participate in a two-stage game. In the first stage, firms decide their product type (in this case, motels can choose to offer either a low- or a high-quality product) and whether to enter the market. In the second stage, entrants compete and payoffs are determined. To ensure the existence of a unique equilibrium, Mazzeo proposes two equilibrium concepts, either of which will result in a unique outcome. The first is simply to assume that firms engage in Stackelberg competition, in which firms sequentially determine their product type and whether to enter. In the other equilibrium concept, firms sequentially decide whether to enter and then simultaneously determine product types.[2] These solution concepts are then empirically implemented in a demand threshold-type framework similar to that used in Bresnahan and Reiss (1991), and then applied to a dataset of highway motels (which are classified as either low- or high-quality according to their star rating). In this framework, firms of each type experience difference market-level shocks drawn from a multivariate distribution. The results of the estimation suggest that the choice of equilibrium concept is empirically unimportant, and that firms in the highway motel industry face a strong incentive to differentiate.

Another example of a static empirical model of entry with endogenous product-differentiation can be found in Seim (2006). Seim's model allows firms to face idiosyncratic shocks to their profitability that are unobservable to other firms. This asymmetric environment reduces the computational burden of estimating the model relative to the full information case. In this model, firms choose whether to enter at some point on a discrete set of locations. When making the entry decision, firms must balance the benefits of choosing a location with inherently higher levels of demand with the adverse effects of heightened competition that these locations are liable to attract. Although the profitability of each location is identical for each firm, firms also privately observe idiosyncratic shocks to their own profitability. Due to the assumptions listed above, firms have symmetric expectations of the probability that another firm will enter at a given location. This makes it possible to to estimate the model using a relatively straightforward multinomial logit procedure. Similar to the findings of Mazzeo (2002), Seim (2006) finds that firms in the video retail industry have significant incentives to differentiate through location choice.

Some recent research has sought to deepen the structural foundation of demand threshold-type models. Abbring and Campbell (2010) imposes a "last-in, first-out" (LIFO) rule on a sequential entry decision game. Under LIFO, incumbent firms decide whether to remain in the market in order of the firms' ages, reflecting the empirical tendency for newer firms to have much higher exit rates than their older competitors.[3]

[2]Note that both of these equilibrium concepts require the imposition of further assumptions on firms' profit functions: namely, that profits are strictly decreasing in the number of entrants and that firms' profits decrease by a relatively larger amount in response to the entry of another firm of the same type.

[3]For examples from multiple industries, see Dunne, Roberts, and Samuelson (1989) and Geroski (1995).

They show that under relatively weak assumptions governing the distribution of the market-level demand shocks, the assumption that firms use threshold rules to determine their entry and exit decisions in an infinite-horizon game, such as the one used in Bresnahan and Reiss (1994), is valid. The use of the LIFO framework presented in Abbring and Campbell (2010) has the advantage of considerably simplifying both equilibrium analysis and the testing of counter-factual policies.

Schaumans and Verboven (2015) extend the static entry model of Bresnahan and Reiss (1991) to allow for product differentiation. They derive specific conditions under which entry thresholds provide a meaningful measure of the presence and size of the competitive market effects. They also introduce a method for obtaining unbiased estimates of the competitive effects of entry by augmenting the product-differentiated demand threshold model with a revenue equation.

Several studies have applied static models of endogenous market structure to estimate demand thresholds and analyze the relationship between market structure and competitive conduct in specific industries. Asplund and Sandin (1999) use the Bresnahan and Reiss (1991) framework to analyze regional markets for driving schools in Sweden, and find that the number of firms scales less than proportionately with market size. Similarly, Manuszak (2002) uses the static Bresnahan-Reiss model to study a historical dataset of beer brewing firms in American frontier states in the late 19th century. He finds evidence that competitive conduct intensified in local brewing markets as the number of firms increased.

Relatively recently, variants of the pseudo-dynamic model of changes in market structure established in Bresnahan and Reiss (1994) have seen some use in empirical applications. Xiao and Orazem (2011) use panel data to study the importance of sunk costs in the local markets for broadband internet services in the United States as the industry evolved between 1999 and 2003. They find little competitive effect of entry in this industry after the establishment of three incumbent firms. Collard-Wexler (2014) uses a similar model to estimate entry and exit thresholds in the ready-mix-concrete industry, which he combines with estimates of the demand process for this industry to simulate the duration of the competitive effects of a merger from duopoly to monopoly in a regional market.

The relationship between market structure and competition in the newspaper industry is of particular interest due to the politically and socially important roles newspapers serve in the dissemination of information. In this respect, several studies have investigated the political importance of newspapers and other news media technologies.

Gentzkow, Shapiro, and Sinkinson (2011) study the effect of entry and exit in the newspaper industry on voter turnout, particularly during the period between 1869 and 1928. During this time, newspapers faced little competition from other news media technologies and thus were at the height of their political importance relative to other sources of information. They find that, on average, the entry of a newspaper into a county increased county-level turnout per eligible voter to presidential elections by 0.3 percentage points during this period. The magnitude of this effect attenuates over time as radio and television become relatively more important as tools for disseminating of political information, but it remains robustly positive. Notably, they find little evidence to support the hypothesis that competition in the newspaper industry has a practically large effect on voter turnout. Instead, they find that the relationship between entry in a local newspaper market and voter turnout is driven almost entirely by the entry of the first newspaper: the marginal effect of the first entrant on turnout appears to be around 1.0 percentage points per eligible voter, while the marginal effects of subsequent entrants are neither statistically nor practically different from zero.

In a related work, Gentzkow, Shapiro, and Sinkinson (2014) study how competition in newspaper markets affects the incentive for newspapers to differentiate ideologically. They study this relationship using a structural model of newspaper competition with endogenous entry, political affiliation (represented as a discrete choice between Democratic and Republican), and subscription and advertising prices. The model is estimated using cross-sectional data of newspaper location choices and subscription revenues in 1924, augmented with calibrated parameters representing marginal costs, readership overlap, and per-consumer advertising revenues. They find a positive relationship between the intensity of competition and ideological diversity in local newspaper markets. Policy simulations conducted using the model suggest that allowing newspapers to collude on advertising prices functions as a subsidy to subscription prices, increasing consumer welfare by enhancing entry and the incentive to differentiate along ideological dimensions. Another policy simulation suggests that subsidizing newspapers' marginal printing costs could enhance welfare by inducing entry in markets without newspapers and increasing readership per firm in markets with newspapers.

Some related research exists for other forms of news media. Stromberg (2004) finds that, all else equal, counties with relatively higher radio ownership rates received more unemployment relief per capita from New Deal relief efforts between 1933 and 1935. In addition, the study shows a statistically and practically significant effect of radio diffusion on turnout, estimating that a 10 percentage point increase in the proportion of households with radios was associated with a 1.2 percentage point increase in voter turnout.

Gentzkow (2006) exploits exogenous variation in the diffusion of television during the 1950s and finds a negative relationship between the availability of television and voter turnout in presidential and congressional elections. One reason for television's negative relationship with turnout is that it appears to be a poor source of political information relative to other news sources. Survey evidence shows that voters in counties with active television stations in 1952 were less likely to be able to name the candidate for whom they intended to vote compared to voters in counties without television access. Television also appears to cause voters to pay less attention to local politics, as Gentzkow (2006) finds evidence that television had a larger negative effect on turnout in local congressional elections compared to presidential elections.

Oberholzer-Gee and Waldfogel (2009) analyze the relationship between the availability of local news media and local civic behaviour by measuring the effect of the establishment of a television station broadcasting Spanish-language news and turnout rates for Hispanic voters. They estimate that the presence of a Spanish-language local news program in a metropolitan area increases Hispanic voter turnout by between 5 and 10 percentage points for both local and presidential elections, with a relatively larger effect in local elections.

Some research has investigated the nature and extent of product differentiation in the newspaper industry. George and Waldfogel (2006) exploit the New York Times' implementation of an expanded distribution strategy in the 1990s to determine whether the introduction of national newspapers alters locals preferences. They find that the circulation of national newspapers diminishes the readership of local newspapers among the college-educated readers and increases readership among those without college degrees. George (2007) finds that an increasing degree of ownership concentration is associated with increased product differentiation between newspapers, as joint-owners seek to minimize the negative externality associated with products being located too closely together in the product space.

3 Data

3.1 Newspaper Data

This paper uses the panel of English-language daily newspapers in the United States that was developed for and first used in Gentzkow, Shapiro, and Sinkinson (2011). It is available to academics via ICPSR Study 30261. The dataset contains information about US daily newspapers from 1869 to 2004, on a presidential year basis from 1872 onward (i.e., every four years). The data were compiled from a combination of Rowell's *American Newspaper Directory* (1869-1876), Ayer's *American Newspaper Annual* (1880-1928), and the *Editor and Publisher Yearbook* (1932-2004). These newspaper directories are close to comprehensive. Newspapers are considered 'dailies' if they are published on at least four weekdays each week. National newspapers (e.g., *The New York Times*, *Wall Street Journal*) are not included in the data. A detailed discussion of the process by which this data was collected and audited can be found in Appendix A of Gentzkow, Shapiro, and Sinkinson (2011).

The newspapers in the dataset are matched to the city that accounts for the largest share of their circulation. These cities are then matched to their census place, and then matched on that basis to the county representing the largest share of the population of the census place. The tendency of news markets to overlap complicates the identification of the relevant geographic market. A newspaper that circulates predominantly in one large city may also have a readership in surrounding cities, which in turn have their own sets of locally headquartered newspapers which have some additional readership in surrounding cities, and so on. This places all daily newspapers in some degree of indirect competition with one another, with the intensity of inter-newspaper competition diminishing with distance. This aspect of competition in the newspaper industry makes it necessary to make some assumptions concerning the definition of the relevant geographic market.

In this respect, I follow Gentzkow and Shapiro (2010) in treating the county as the relevant 'news market', given their finding that in 2005 the county in which a newspaper was headquartered accounted for over 80 percent of its circulation. Similarly, Fan (2013) defines the relevant market as the set of counties accounting for at least 85 percent of the newspapers' circulation. The Audit Bureau of Circulation, a non-profit organization that verifies circulation figures for newspapers in the US, also uses this standard. The limitations of the data set make it impossible to mirror this standard in this paper, although in most cases treating the county as the relevant geographic market serves as a reasonable approximation of this standard. Chandra and Collard-Wexler (2009), in their study of ownership consolidation in the Canadian newspaper market, also take the county to be the relevant geographic market for a non-national daily newspaper. This definition of the relevant market is also used in Seamans and Zhu (2013). Of the counties identified in the dataset, 77.1 percent contain only one news market, while 14.3 percent contain two news markets, and 5.0 percent contain three news markets. Overall, the assumption that the county is relevant geographic market for non-national daily newspapers is a reasonable approximation of reality.

For each year, I define county-level variables reflecting the number of daily newspapers active in the county and the number of newspapers of each self-declared political affiliation (Republican, Democratic, Independent, or undeclared). For county-years in which the circulation figures for all newspapers active in the county are in the data set, I calculate the total and average circulation of each newspaper in the county. For these county-years, I also obtain the one-firm concentration ratio by calculating the proportion of total circulation in the county accounted for by the newspaper with the largest circulation.

3.2 Demographic Data

I obtain county-level demographic data from the US Census. From the US Census and County Data Books made available by ICPSR Study 02896, I obtain measures of the total population, the white population, and employment in the manufacturing sector for each county for the 1870 to 2000 period. For the 1870 to 1950 period, I also obtain the foreign-born white population, the population of males aged 21 and older, the population living in urban areas with a population above 2,500, the surface area of the county, and the surface area of the county occupied by farmland from the ICSPR data. For 1930, 1940 and 1950, I collect county-level data on the number of households and the number of households owning a radio. I supplement the ICPSR data with census data from the National Historical Geographic Information System for the 1950 to 2000 period. From this source, I obtain county-level figures for the foreign-born white population, the population of males aged 21 and older, and the population living in urban areas.

Although the US Census is only collected decennially, to my knowledge it is the best source for county-level demographic data, especially for the late 19th to early 20th century. To obtain demographic data for election years that do not coincide with census years, I follow Gentzkow, Shapiro, and Sinkinson (2011) in interpolating intercensal presidential years using a natural cubic spline.[4]

I obtain county-level data on television ownership from ICPSR Study 22720. This data set provides information on the number of households in each county that own televisions for the years 1950 and 1953 to 1960. The observations from 1950 and 1960 are drawn from the decennial census for those years. Observations from 1953 and 1955 to 1959 are drawn from the *TV Factbook* for each year, while observations for 1954 come from *TV Magazine*. Using this data, I calculate the share of households owning televisions using data on the number of households per county drawn the decennial census. I interpolate the number of households owning televisions for the presidential year of 1952 for this period using the same approach outlined earlier in this subsection.

3.3 Geographic Definitions and Sample Selection

Some complications arise from the fact that county definitions are not always consistent across time, especially over a period as long as the one used in this study. Over the course of the period, some counties change names, some counties are incorporated into other counties, and still others are dissolved into unincorporated territory. To account for this fact, I have attempted, where possible, to retroactively apply the geographic definitions of counties used by the US Census Bureau in the 1990 census to all earlier years. I treat counties that changed names at some point during their history as a single continuous entity. Similarly, I treat counties that merge into a larger county as though

[4] For variables that are calculated as a ratio of two other variables (e.g., the proportion of county population consisting of males age 21 and older), I interpolate both the numerator and the denominator and then take their ratio, instead of interpolating the ratio itself.

they had always been a part of that larger county by summing the merging counties along the various demographic and geographic measures listed above. I drop cases in which a county is eventually broken up and then merged with two or more pre-existing counties.

To simplify the process of retroactively applying 1990s geographic definitions, I do not include counties that eventually become unincorporated in the sample. I also drop county-years in which total population is less than one thousand. This eliminates the many instances in the late 19th century of extremely small counties being absorbed by larger counties or otherwise dissolved into unorganized territory. It also eliminates outlying cases of population growth, as one is much more likely to observe a 200 percent growth in population from one presidential year to the next in a county with a population of 300 than in a county with a population of 3,000.

Finally, I restrict the sample of counties to include only counties that had a local daily newspaper for during at least one presidential year throughout the period. After restricting the sample along the lines described above, at total of 1,438 counties are included in the sample, out of 3,146 counties in existence in 1990. Over the course of the period, these counties represent an average of 75.7 percent of the US population.

4 A First Look at the Data

The local daily newspaper industry was established through a large initial wave of entry that lasted from at least the beginning of the observed period in 1872 until 1920, after which the industry entered a period of stagnation. Throughout the stagnant period, net entry (the number of entering firms minus the number of exiting firms during presidential year t) has been small in absolute terms and typically slightly negative. Figure 2 plots the industry-wide entry and exit rates over the course of the period. Following Dunne, Roberts, and Samuelson (1989), the entry rate is defined as the number of entrants at time t divided by the number of active firms at time t-1. Conversely, the exit rate is defined as the number of firms that exited between periods t-1 and t divided by the number of firms active at time t.

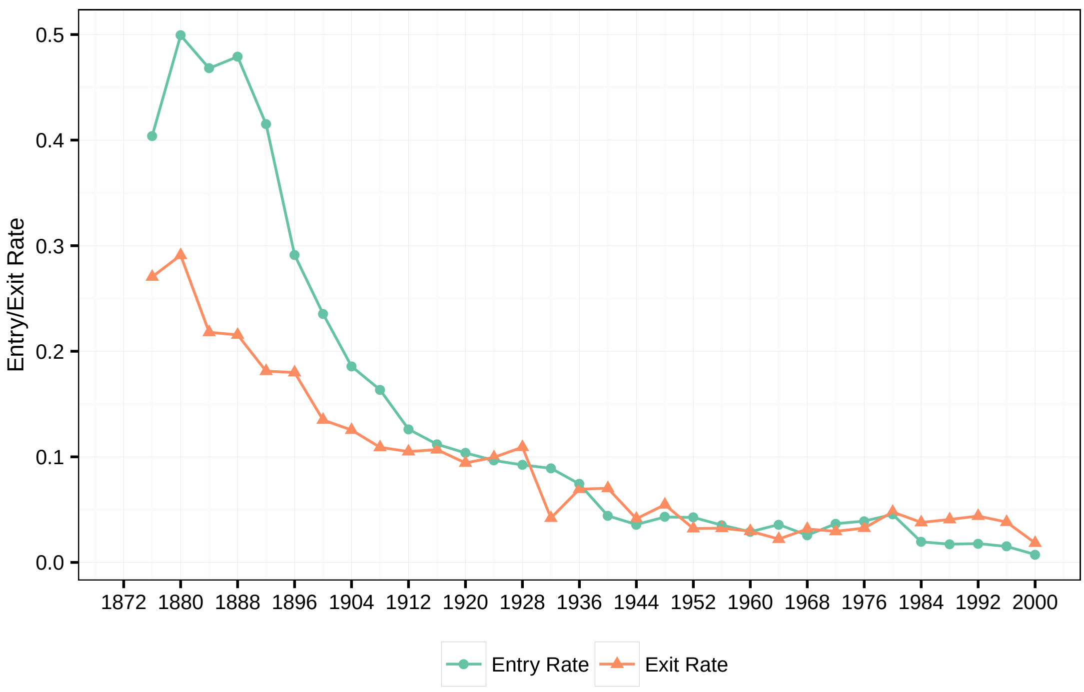

FIGURE 2: Entry and Exit Rates in the Newspaper Industry, 1872 - 2000

From Figure 2, it is clear that entry and exit rates in the newspaper industry have followed similar patterns across time. Initially, both entry and exit rates are quite high. Before 1900, the number of fresh entrants was over 30 percent of the number of existing firms in the previous presidential year. During the same period, exiting firms accounted for between 17.5 and 30 percent of existing firms in the prior presidential year. After 1916, entry and exit rates were roughly equal during all years. This pattern persists until the end of the sample period, although the exit rate slightly exceeds the entry rate after 1980.

For the most part, the observed entry and exit rates in the newspaper industry are consistent with the stylized facts concerning entry and exit outlined in Geroski (1995) - entry and exit rates are positively correlated, vary over time, and tend to be highest at the beginning of an industry's lifespan. In this case, there are no obvious waves of entry and exit after the initial wave of entry that populated the industry in the late 19$^{\text{th}}$ and early 20$^{\text{th}}$ centuries. The lack of visibly obvious waves of exit after 1920 is particularly surprising: in spite of the popularization of radio in the 1930s, and television in the 1950s, exit rates in the local daily newspaper industry appear to have followed a reasonably consistent trajectory since at least 1912. In fact, the largest movement in the exit rate occurred between 1928 and 1932, during which time the exit rate actually *decreased* from 0.11 to 0.04. As we will see, most of the entry that occurred after 1920 was in counties that previously did not have newspapers, while most exit occurred in counties with two or more local newspapers.

Today, most daily newspapers are monopolists. However, this has not always been the case. From 1872 until around 1912, newspaper markets were split roughly evenly across three categories: monopolies, duopolies, and markets with three or more newspapers. The left panel of Figure 3 plots the proportion of counties with active newspapers belonging to each category. In some presidential years during this period, monopoly was actually the least commonly observed market structure in the local daily newspaper industry. After 1912, however, this situation changed rapidly, and one-newspaper counties became the norm. This shift coincided with steady declines in the proportion of counties with two or more newspapers.

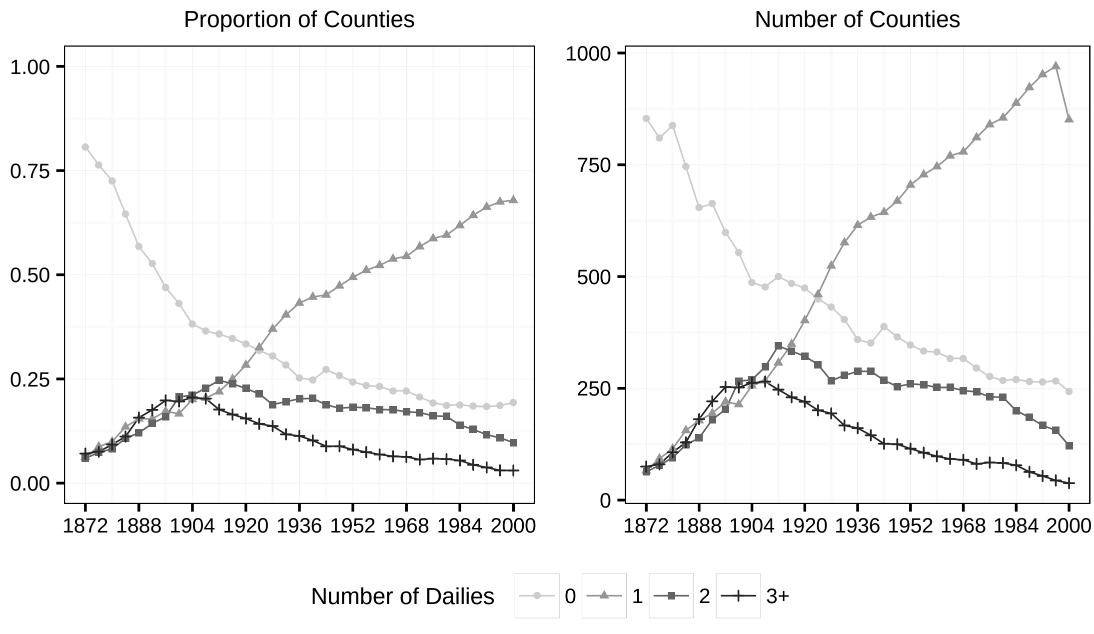

Figure 3: Counties by Number of Papers, 1872 - 2000

The right panel of Figure 3 plots the absolute number of counties of each market structure. As shown in the chart, the increasing prevalence toward monopoly occurred in both relative and absolute terms. Throughout the sample period, the total number of counties with active local newspapers consistently increases, with the exception of the 2000 presidential year. After 1912, increases in the number of one-newspaper counties explain all of this growth. Population growth is presumably partly responsible for the increasing number of counties with at least one newspaper: as the potential market size of a given county grows, the expected profitability of potential entrants increases. Entry occurs once a market becomes sufficiently large to support the profitable entry of a monopolist. Hence, the increasing number of counties with monopolistic local newspaper markets is driven in part by entry into counties that were previously too small to support a local newspaper. At the same time, the absolute number of counties with two or more newspapers declines consistently after 1912. Presumably, it is more likely that a county with two newspapers will transition to having one newspaper rather than to having no newspapers, with a similar logic applying to counties with three or more newspapers. This suggests another source of the increasing number of one-newspaper counties.

Figure 4 explores the relationship between population and the number of newspapers in greater detail. Even without estimating a model of market structure, it should be possible to form a rough estimate of entry thresholds simply by sorting counties by population and market structure. As the figure shows, there is a reasonably clear separation between the various market structures observed in counties of different sizes, although the distinctness of this separation varies over time.

From 1872 to 1908, the minimum market size required to support at least one newspaper appears to be between 22,000 and 36,000 persons. In this early stage of the newspaper industry's development, there is no clear between the monopoly and duopoly entry thresholds. The minimum market size required to support at least three newspapers appears more clearly defined, lying between 60,000 and 100,000 persons.

Between 1912 and 1940, the entry threshold for a monopolist looks to be between 17,000 and 36,000. The threshold for a duopolist is less easy to eyeball, as there is significant overlap between the populations of one- and two-newspaper counties during this year, but it appears to lie between 36,000 and 98,000. As before, the triopoly threshold is easier to see, and likely lies between 98,000 and 162,000.

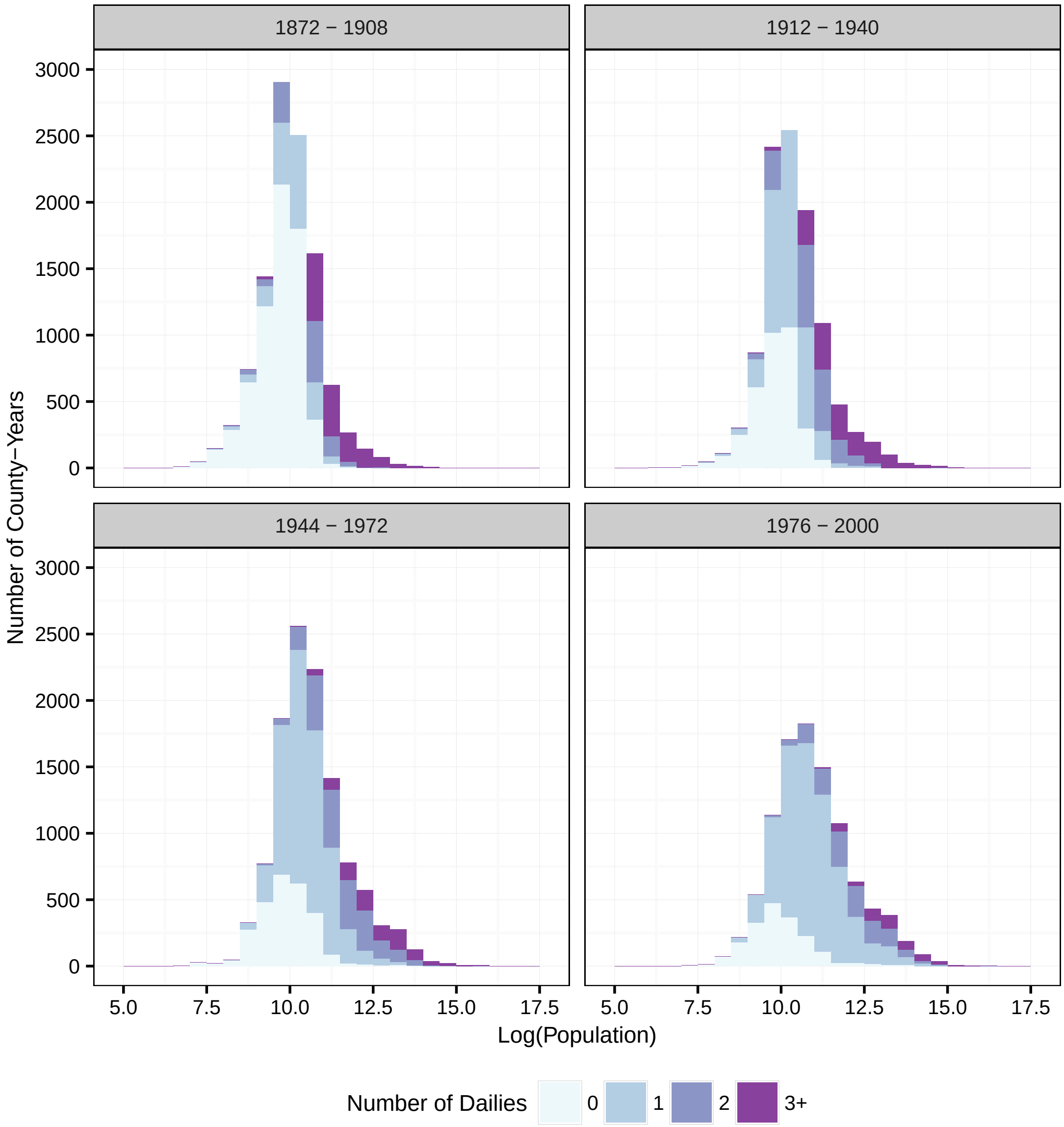

FIGURE 4: County-Years by Population and Number of Dailies (Binwidth = 0.5)

In the early stages of the television period, from 1944 to 1972, the monopoly entry threshold becomes more easily observable. It appears to lie between 13,000 and 22,000. During this period, oligopolistic market structures become relatively less common. The duopoly entry threshold appears to be between 98,000 and 162,000, while the triopoly threshold is above 440,000. A similar pattern is observable in the latter portion of the television period, from 1976 to 2000. The monopoly entry threshold remains around 22,000, while the duopoly threshold appears to have increased to somewhere between 268,000 and 442,000. The triopoly entry threshold is not very visually apparent during this period, but it likely lies above 1.2 million. In any case, it is far beyond the population level of the average county.

In order to gain a more detailed understanding of the increasing tendency toward monopoly in the local daily newspaper industry, it is useful to look at measures of the probability that a county will transition from one market structure to another over time. Figure 5 plots the probability that a county will have a given number of active newspapers in the next presidential year conditional on the number of active newspapers it has in the current presidential year. Each panel corresponds to the number of newspapers active at time t and each line corresponds to the probability of transitioning to a given number of newspapers at $t+1$. The observed probability that a county with a given market structure at time t will transition to a given market structure at time $t+1$ is calculated as the proportion of counties with a given number of newspapers at time $t+1$ that had a given number of newspapers at time t.[5]

At any time during the period, for any market structure, stasis is the most likely transition type from one presidential year to the next: counties with no newspapers during one presidential year will most likely have none the during the next, a county with one newspaper during one presidential year will most likely still have one the next, and so on. The probability of stasis in counties without newspapers is close to one and changes relatively little over the course of the period. In fact, the plot actually understates the probability of stasis, since the sample is restricted to counties that have newspapers during at least one presidential year within the sample period.[6] Nevertheless, the probability of stasis for zero-newspaper counties falls slightly early on, and then gradually rises over the remainder of the period. Changes in the probability of transition to a one-newspaper county essentially mirror the probability of stasis for counties without newspapers. It is exceptionally unlikely for two or more newspapers to find it profitable to enter the same county simultaneously.

[5] For example, suppose there are 100 counties with one newspaper at time t. Of these counties, 15 have no newspapers, 70 have one newspaper, 10 have two newspapers, and 5 have three newspapers at at time $t+1$. The observed transition probabilities from t to $t+1$ for counties with one newspaper at time t are 0.15, 0.70, 0.10, and 0.05 for each respective transition.

[6] In practice, including only counties that have at least one newspaper at some point during the sample period excludes roughly half of all counties in the United States. When all counties are included in the sample, the probability of stasis in counties without newspapers is very near to one, making it difficult to observe any patterns in the transition probabilities for counties without newspapers.

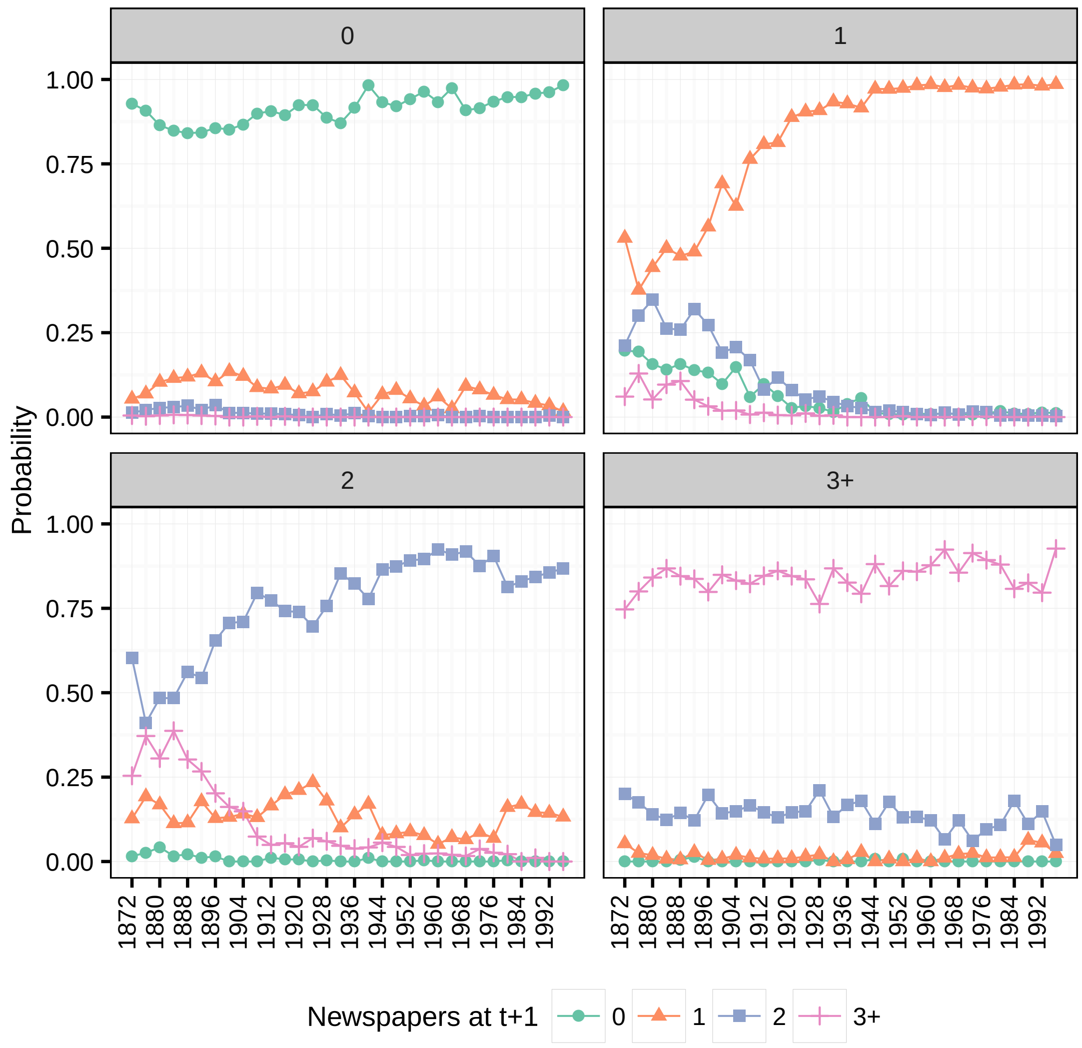

FIGURE 5: Transition Probabilities, Presidential Year to Presidential Year, 1872 - 1996

One- and two-newspaper counties experienced substantial changes in their transition probabilities over the course of the period. In one-newspaper counties, the probability of stasis was around 0.5 prior to 1900, after which it began to increase. By 1944, the likelihood of a one-newspaper county having one-newspaper in the next presidential year was close to 1.0. A decline in the probability of entry, and to a lesser extent the probability of exit, accompanied the increasing probability of stasis in local monopolies. Throughout much of the period, a one-newspaper county is more likely to transition to being a two- or a three-newspaper county than to a zero-newspaper county. Clearly, local monopolists tend to be secure in their positions, especially in the latter half of the sample period. Even after the popularization of television in the 1950s, the probability of a one-newspaper county becoming a zero-newspaper county was close to zero from one presidential year to the next.

Two-newspaper counties saw a similar increase in the probability of stasis over the course of the period, although these markets tend to be less stable than local monopolies. Before 1904, two-newspaper counties were more likely to see entry than exit. Most of the increase in the probability of stasis seen among two-newspaper counties is the result of a decline in the probability of entry, as the probability of exit in two-newspaper counties remained roughly constant throughout the period.

It is difficult to discern the presence of any trends in transition rates of counties with three or more newspapers from one presidential year to the next. The probability of stasis is quite high, remaining above 0.75 for the entirety of the sample period. As was the case with two-newspaper counties, the probability of exit stays nearly constant over the course of the period. A county with three or more newspapers is approximately as likely to transition to having two newspapers as a two-newspaper county is to transition to a monopoly.

The declining absolute number of counties with three or more newspapers is thus explained by the declining probability of entry in two-newspaper counties, as the probability of exit in three-newspaper counties has remained nearly constant. Similarly, the diminishing number of two-newspaper counties is largely explained by the declining probability of entry in one-newspaper markets, as exit rates in counties with three or more newspapers have fluctuated relatively little over time. Finally, the increasing tendency toward stasis in one-newspaper counties was driven by simultaneous declines in the probability of both entry and exit in local monopolies, as the probability of entry in zero-newspaper counties and exit in two-newspaper counties changed little over the course of the period.

Figures 6 and 7 plot transition probabilities over horizons of three and five presidential years, respectively. Naturally, extending the time horizon decreases the probability of stasis for any given market structure. This effect is more apparent in counties with two or more local newspapers than it is in counties with zero or one newspapers. After 1994, two-newspaper counties typically had a greater than 85 percent chance of experiencing neither entry nor exit over a one-presidential year horizon. Over a three-presidential-year horizon, however, the probability of stasis is typically below 65 percent during this period. A similar pattern is observable in three-newspaper counties over the same time horizon.

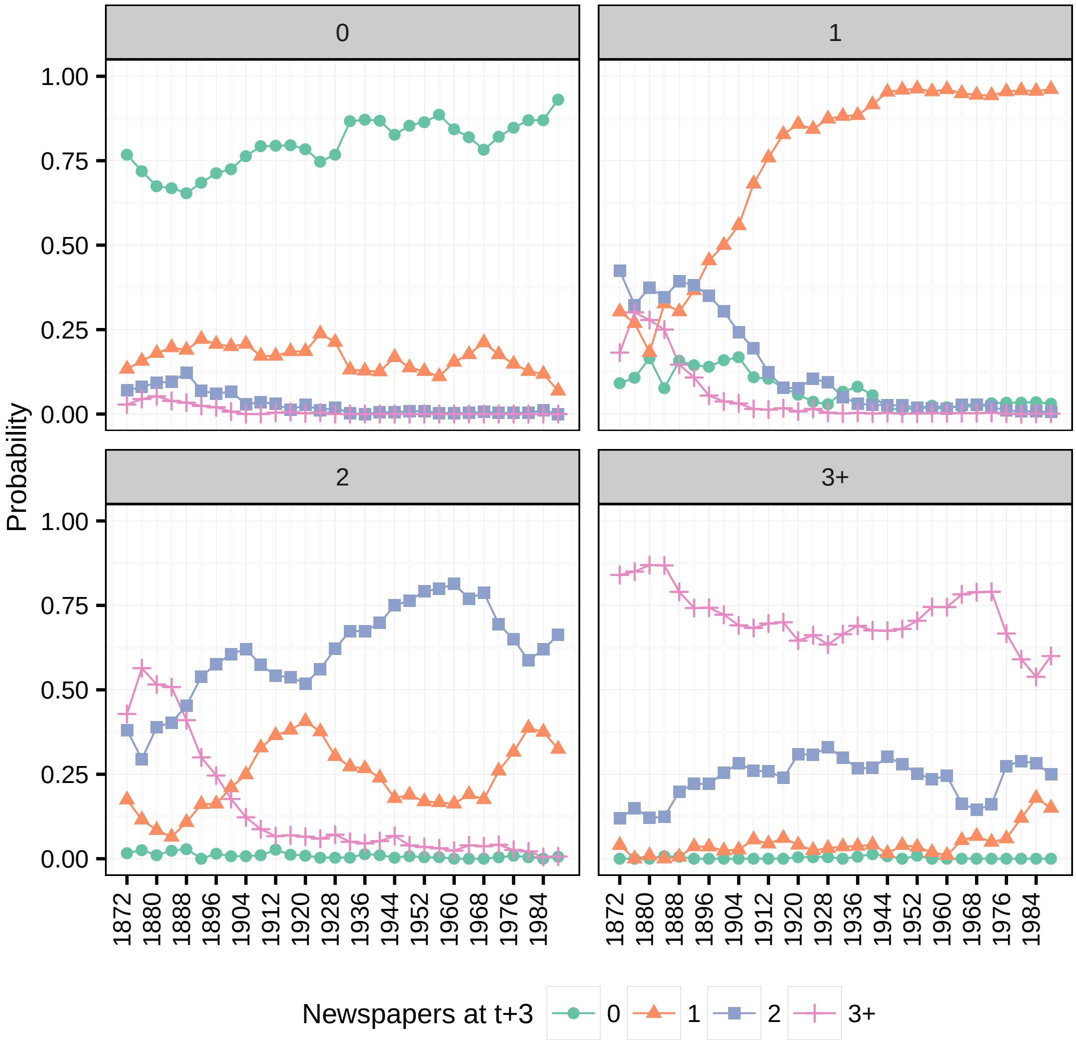

Figure 6: Transition Probabilities Over 3 Presidential Years (12 Years)

Extending the time horizon to five presidential years makes this tendency even more apparent. By the end of the sample period, two- and three-newspaper counties were almost as likely to experience exit over a horizon of five presidential years as they were to retain all of their incumbent newspapers. Interestingly, two-newspaper counties were most likely to experience stasis between 1940 and 1970, the period during which we might expect television to be having a large negative effect on newspapers' profits. Moreover, the increasing probability of stasis for two-newspaper counties during this period is driven largely by a declining probability of exit, rather than a declining probability of entry.

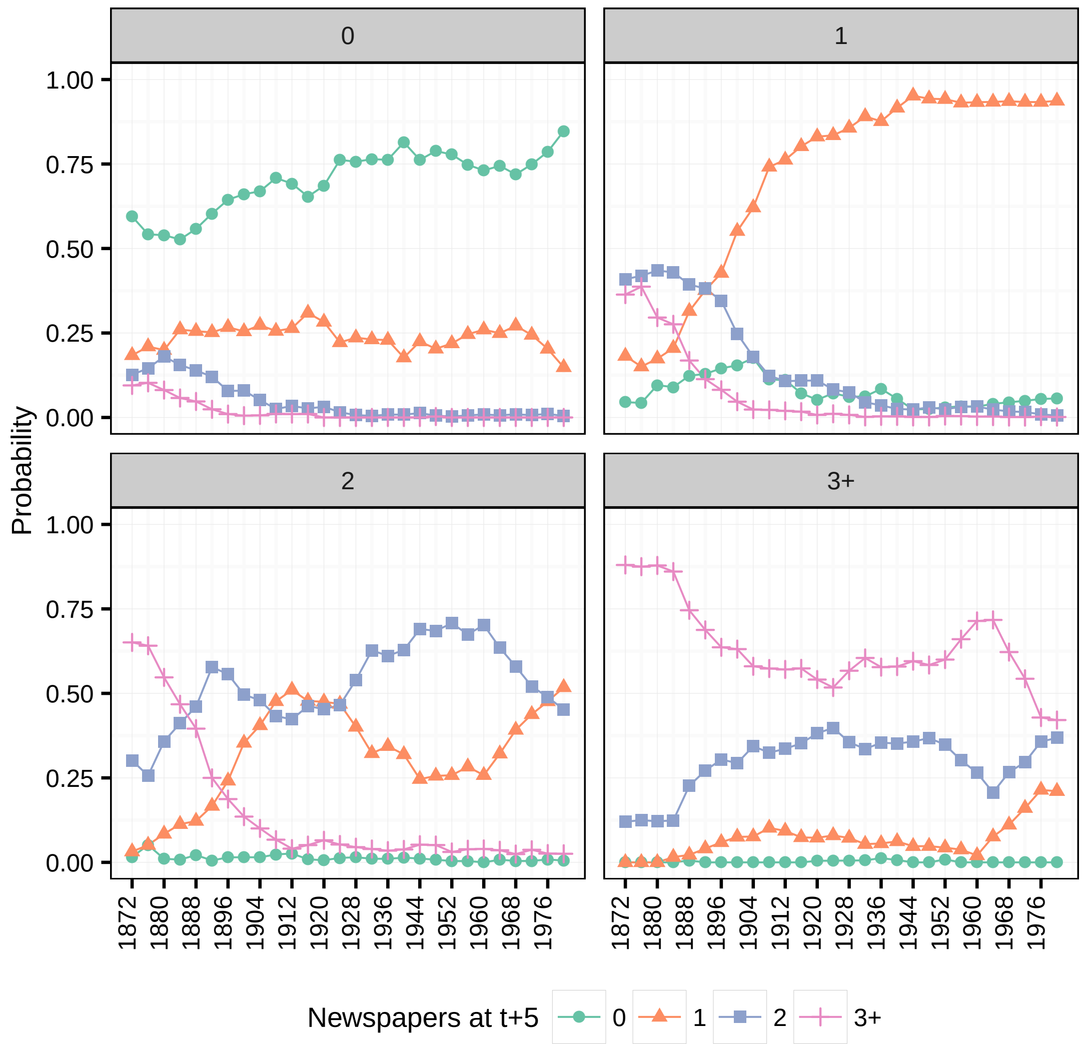

FIGURE 7: Transition Probabilities Over 5 Presidential Years (20 Years)

By contrast, the probability of stasis changes relatively little for one-newspaper counties when we consider a time horizon of three or even five presidential years, especially after 1944. Over relatively longer time horizons, counties without newspapers become increasingly likely to transition to being a local monopoly, while the probability of transition to a duopoly or triopoly remains close to zero, especially during the 20$^{\text{th}}$ century.

Based on this information, we can form some expectations about how entry and exit thresholds for various market structures have changed over time. Clearly, one-newspaper counties exhibit a strong tendency toward stasis, in spite of the presumably negative effects on the demand for newspapers resulting from the popularization of radio and television during the period under study. At the same time, extending the time horizon increases the chance that a county without a newspaper will transition into a local monopoly. The high probability of stasis in the newspaper industry suggests exit thresholds are low relative to entry thresholds. Entry thresholds for local monopolists appear to have increased, but not by a large amount relative to overall population growth. In two- and three-newspaper counties, the declining stasis probabilities and declining probability that counties with fewer newspapers will experience entry suggest that both entry and exit thresholds have increased by a sizable amount over the course of the period for duopolies and triopolies.

Since firms in local newspaper markets are not necessarily symmetric, it is worth considering how market concentration varies with market size and the number of firms. Figure 8 plots the circulation share of the largest newspaper (i.e., the one-firm concentration ratio) against the log of total population for each county-year for which complete circulation data are available. In order to see how the relationship between market size and market concentration has changed over time, I separate the sample period into four subsections: 1872 to 1908, 1912 to 1940, 1944 to 1972, and 1976 to 2000. I chose these periods such that they approximately align with changing patterns in entry and exit rates in the newspaper industry as a whole.

Returning to Figure 2, we can see that the period 1872 to 1912 corresponds to the growth stage of the newspaper industry, characterized by large net entry and high turnover rates. The period of 1912 to 1940 represents a more stable phase of the newspaper industry's development, during which time entry and exit rates were generally below 0.1 and trending downward. Between 1944 and 1972, entry and exit rates were relatively stagnant, in spite of the presumably disruptive intervention of the popularization of television that took place during this period. From 1976 onward, the newspaper industry was in the declining stage of its development, as exit rates began to exceed entry rates, with the former trending upward and the latter trending downward. Each of these periods contain a roughly equal number of county-year observations, with the 1872 to 1908, 1912 to 1940, 1944 to 1972, and 1976 to 2000 periods comprising 11,981, 11,311, 11,409, and 9,867 county-years, respectively.

To provide a sense of how market concentration changes with the number of newspapers active in the market, I colour county-years according to the number of active newspapers in that county-year. In addition, I plot lines of best fit for each grouping of years using LOESS.

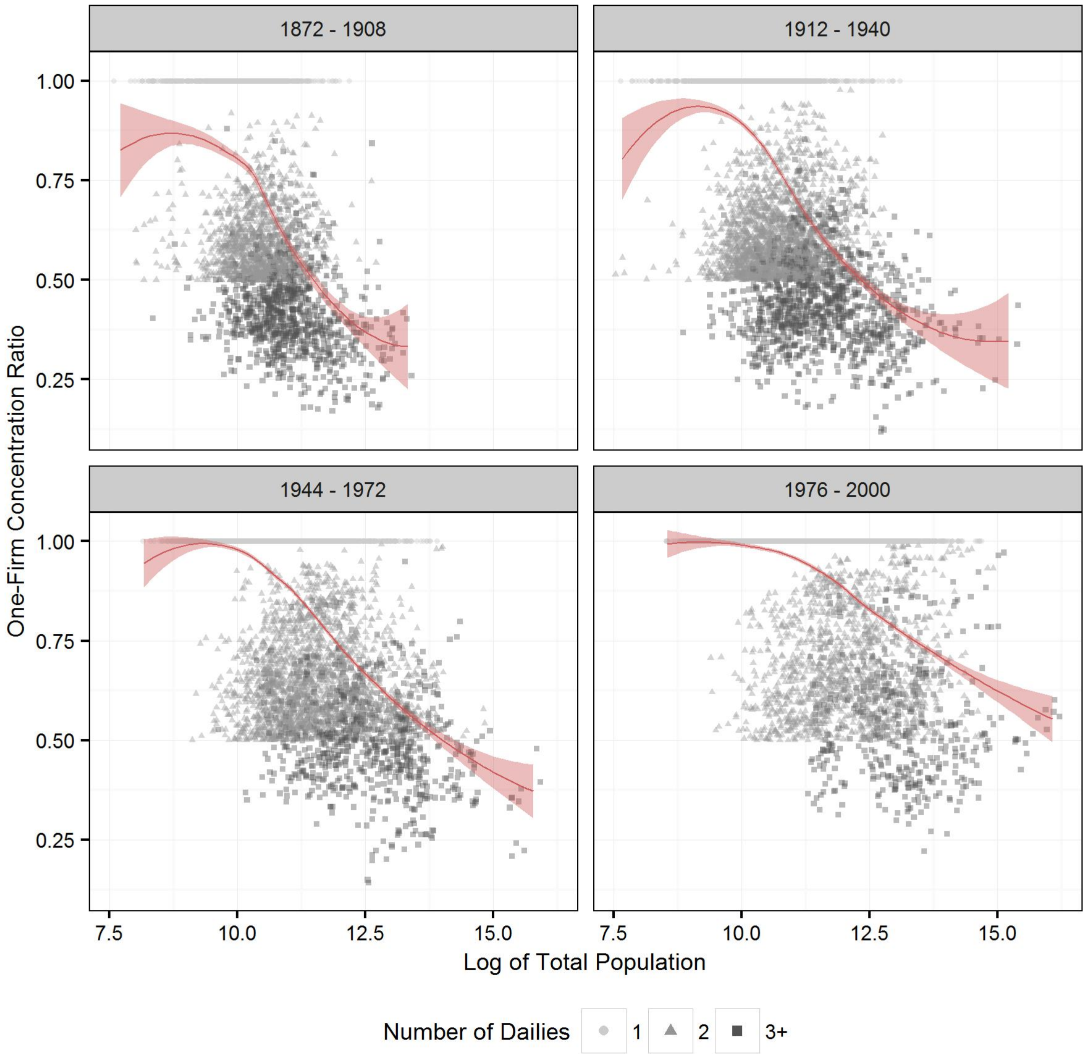

FIGURE 8: One-Firm Concentration Ratios by County Population

At relatively low levels of population, the lines of best fit show a positive relationship between market size and market concentration. This is attributable to the presence of some two-newspaper counties with relatively low populations - so low that even one-newspaper counties are uncommon at that level of population, resulting in the upward sloping segment of the LOESS curve. This upward-sloping segment of the LOESS curve becomes much less prominent from 1944 onward. It is also worth noting that the minimum level of population required to support at least one newspaper appears to increase across the three periods. The LOESS lines of best fit shift upward over time, reflecting both population growth and increasing concentration in the newspaper industry.

Figure 8 also shows that the negative relationship between market size and market concentration has weakened over time. Between 1872 and 1908, and to a slightly lesser extent between 1912 and 1940, markets with two or more newspapers appear to be reasonably competitive. Most two-newspaper county-years have one-firm concentration ratios between 0.5 and 0.75, and the majority of county-years with three or more newspapers have concentration ratios between 0.35 and 0.5. This suggests that these counties are nearly as unconcentrated as they can be without the entry of additional competitors.

For the most part, this pattern is still apparent between 1944 and 1972, although two-newspaper counties with concentration ratios above 0.7 and three-newspaper counties with concentration ratios above 0.5 become relatively more common during this period. In addition, three-newspaper counties begin to disappear during this period, shifting the line of best fit upward.

The negative relationship between population and market concentration further attenuates during the period between 1976 and 2000. Two factors explain this change: the exit of newspapers from three-newspaper counties, and the continued upward drift of one-firm concentration ratios in two-newspaper counties. A possible explanation for the upward trend in one-firm concentration ratios is that newspapers may have exited from markets in which competition between duopolists was relatively more intense, leading to an increase in average concentration in two-firm markets.

The extent to which consumers care about product differentiation may be an important determinant of market structure in the local newspaper industry. If newspapers are differentiated, then an entrant can expand the size of a given market by locating itself away from incumbent newspapers in the product space. I present some descriptive evidence reflecting this aspect of entry in the local newspaper industry in Table 1, which shows the results of simple regressions of growth in county-level circulation per capita and the average circulation per capita of incumbents on the change in the number of newspapers and various demographic controls. I include presidential year-specific fixed effects to account for changing preferences for newspapers resulting from the introduction of radio and television.[7]

[7] County-specific fixed effects were found to be jointly statistically insignificant in the regressions for both per capita circulation and the average per capita circulation of incumbents.

	All newspapers		Incumbents only	
	(1)	(2)	(3)	(4)
Number of newspapers$_t$	0.138***	0.141***	-0.077***	-0.081***
	(0.015)	(0.016)	(0.007)	(0.007)
Number of newspapers$_{t-1}$		0.022***		-0.019***
		(0.004)		(0.005)
Log(Population)$_t$	-0.306***	-0.316***	-0.387***	-0.376***
	(0.028)	(0.028)	(0.023)	(0.023)
Urban$_t$	0.051***	0.052***	0.031*	0.032**
	(0.013)	(0.013)	(0.012)	(0.012)
Manufacturing employment$_t$	0.167***	0.167***	0.064	0.070*
	(0.038)	(0.038)	(0.036)	(0.035)
Foreign-born white$_t$	-0.046	-0.124*	-0.048	-0.093
	(0.063)	(0.060)	(0.055)	(0.054)
White$_t$	0.282*	0.334**	0.266*	0.322**
	(0.126)	(0.106)	(0.103)	(0.105)
Observations	27233	26929	26096	25849
Adjusted R^2	0.189	0.192	0.160	0.160

Table entries represent regression results of the listed regressands on county-level newspaper circulation per capita for all newspapers active in a county (including new entrants) and all incumbent newspapers in a county. Standard errors clustered by county in parentheses. Regressands, number of newspapers, and demographic controls are first-differenced. All specifications include presidential year-specific fixed effects.

Table 1: Effect of Newspaper Entry on Circulation per Capita

Entry has a positive relationship with growth in per capita circulation. The entry of one newspaper being associated with an increase of approximately 13.8 percent in per capita circulation in the county. This is somewhat persistent over time, with the entry of a newspaper in the previous presidential year being associated with a growth of approximately 2 percent in per capita circulation in the current presidential year. Although relatively small, this lagged effect is statistically significant. This finding provides some evidence that there is a market-expanding effect of entry at the county level: if new entrants only 'stole' readers from existing newspapers and did not attract any new consumers into the market, then we would not expect per capita circulation to change as the result of the entry of a new newspaper.

On the other hand, entry is negatively related with growth in the average per capita readership of incumbents in the market. The entry of a new newspaper is associated with a decrease of approximately 7.7 percent in the average per capita circulation of newspapers active in the county. The inclusion of the lag of the change in the number of newspapers shows that this effect persists across presidential years, with entry in the previous period being associated with a 1.9 percent decrease in the average circulation per capita of incumbents. These results suggest that entry into a local newspaper market has both market-expanding and business-stealing effects.

A necessary consequence of the increasing prevalence of one-newspaper counties has been the reduction of the variety of products available to consumers. One manifestation of this homogenization is the increasing rarity of counties with at least one newspaper affiliated with both of the major political parties. As the top panel of Figure 9 shows, the number of counties with newspapers affiliated with both major political parties fell from a peak of 93 in 1908 to 23 in 2000, the lowest level since 1876. This trend would be of less relevance if these exits were concentrated in relatively small counties. In general, we would expect exit to occur primarily in relatively smaller counties whose market size no longer justifies remaining in the market in the wake of declining demand for newspapers. However, as the lower panel of Figure 9 shows, the proportion of the population living in counties with newspapers affiliated with both political parties has also fallen considerably over time, from a peak of 31.3 percent of total population in 1908 to 8.8 percent in 2000, the lowest proportion observed during the sample period. Although this decline is less than proportional to the decline in the number of counties with papers affiliated with both parties, it is a large change relative to the peak.

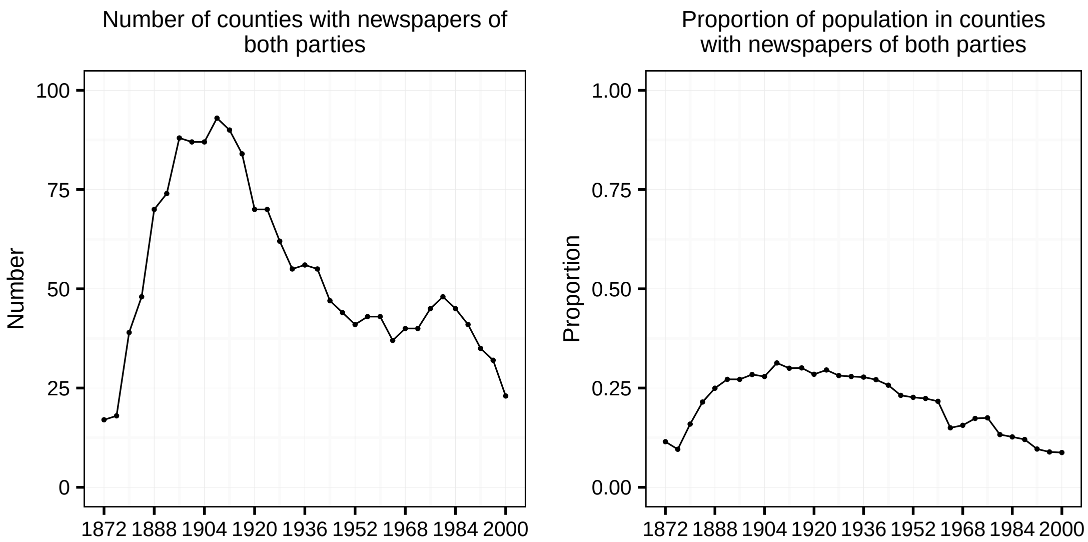

Figure 9: Political Affiliation of Newspapers

5 The Empirical Framework

5.1 Static Bresnahan-Reiss Entry Threshold Model

As a benchmark, I estimate the empirical model of entry developed in Bresnahan and Reiss (1991). In this framework, N firms engage in oligopolistic competition in a local market.[8] By assumption, firms produce homogeneous products, have constant marginal costs c and fixed costs $f \geq 0$. Market demand has the form:

$$Q(p, N) = q(p, N)NS \tag{1}$$

where $q(p, N)$ gives the per capita demand per firm for the product as a function of price and the number of firms active in the market.[9] Market demand is assumed to scale linearly with market size. It is also assumed that $q(p, N)$ is weakly decreasing in price and the number of competitors.

Assuming price competition in addition to the above assumptions, and indexing firms by $i = 1, ..., N$, per firm profits are given by:

$$\pi_i = (p_i - c)q_i(p_i, N)S - f,$$

with the corresponding first-order condition:

$$\frac{\partial \pi_i}{\partial p_i} = q_i(p_i, N) + (p_i - c)\frac{\partial q_i(p_i, N)}{\partial p_i}S = 0.$$

By symmetry, let $p^*(N)$ give the equilibrium price, where $\frac{\partial p^*}{\partial N} \leq 0$. Per firm equilibrium profits, $\pi^*(N)$, are given by:

$$\pi^*(N) = (p^*(N) - c)q(p^*(N), N)S - f. \tag{2}$$

Bresnahan and Reiss define the entry threshold as the minimum market size necessary for N firms to earn non-negative profits. Due to symmetry, this is given by the market size that satisfies $\pi^*(N) = 0$, defined as S_N. Setting (2) equal to 0 and solving for S, we have:

$$S_N = \frac{f}{v(N)}. \tag{3}$$

[8]In reality, the number of firms N is an integer. For ease of exposition, I set this issue aside for now as it is easily dealt with in the empirical implementation.

[9]Note that for the time being, I leave aside the demographic variables included in the original Bresnahan and Reiss (1991) model, as they are not needed at this stage of the exposition.

Hence, the entry threshold for N firms is given by the ratio of fixed costs to per-capita variable profits, where per-capita variable profits are denoted $v(N) \equiv (p^*(N) - c)q(p^*(N), N)$.

The entry threshold increases with fixed costs, since firms must earn relatively more revenue in order to cover these costs, and decreases with variable profits, since firms are able to cover their fixed costs at relatively lower levels of total quantity demanded. To study the competitive effects of entry, Bresnahan and Reiss consider the ratio of the per firm entry thresholds for N and $N-1$ firms:

$$ETR(N) \equiv \frac{S_N}{S_N - 1}\frac{N-1}{N} = \frac{v(N-1)}{v(N)}\frac{N-1}{N}, \tag{4}$$

where we have used expression (3) to establish the rightmost equality.[10]

If the entry threshold increases in exact proportion to the ratio of the pre- and post-entry number of firms, $ETR(N) = 1$. For example, $ETR(2) = 1$ if the number of consumers required for two competing firms to earn non-negative profits is twice the number of consumers required for a monopolist to just break even. From the rightmost expression, we know that this is only possible if $v(N-1)(N-1) = v(N)N$, that is, if per capita industry profits are unaffected by entry. Under the assumptions laid out thus far, this would imply that entry does not induce firms to compete more aggressively. Markets in which this is the case are perfectly competitive: incumbents cannot lower prices without earning negative profits.[11] Conversely, if the per firm number of consumers must more than double to support two firms, we can infer that entry intensifies competition, reducing per firm variable profit.

In the econometric implementation of the static Bresnahan-Reiss model, we assume markets are in equilibrium. In equilibrium, any entrant must earn non-negative profits (i.e., $\pi(p^*(N), N) \geq 0$ for $N \geq 1$). Equilibrium also implies that the entry of one additional firm would be unprofitable: $\pi(p^*(N+1), N+1) < 0$. Recalling the equilibrium profit function (2), these conditions imply:

$$v(N)S \geq f > v(N+1)S. \tag{5}$$

[10] I follow Schaumans and Verboven (2015) in referring to this ratio as the entry threshold ratio (ETR).

[11] Alternatively, an ETR of 1 could indicate that firms are colluding on prices, thereby keeping industry profits constant in N.

Suppose per firm profits in a given market can be observed up to a multiplicative market-level unobservable ξ, such that the true value of per firm profits in a market is given by $v(N)S\xi$. Replacing the per firm profits term in (5) with the version including an unobservable component, we can take logs and rearrange to obtain:

$$\ln v(N) + \ln S - \ln f \geq \ln \xi > \ln v(N+1) + \ln S - \ln f. \tag{6}$$

Since the log of per firm per-capita variable profits cannot be directly observed, I estimate it via the reduced-form equation:

$$\ln v(N) = X\lambda - \sum_{n=2}^{N} \theta^n, \tag{7}$$

where λ is a vector of parameters corresponding to a vector of market-specific demographic characteristics X, and θ^n represents the effect of the entry of an additional competitor on the per capita variable profits earned by each firm. I parameterize the remaining terms in (6) according to $\ln S = \sigma \ln S^*$, $\ln f = \gamma$, and $\ln \xi = \varepsilon$. The true market size, S, is not directly observable. I estimate its log via $\sigma \ln S^*$. Substituting these definitions into (6) and rearranging:

$$X\lambda + \sigma \ln S^* - \gamma - \textstyle\sum_{n=2}^{N} \theta^n \geq \varepsilon > X\lambda + \sigma \ln S^* - \gamma - \sum_{n=2}^{N+1} \theta^n.$$

Provided the incremental effect of an additional entrant on per firm variable profits is negative ($\theta^n > 0$), and imposing the further assumption that the market-level unobservable component of per firm variable profits, ε, is normally distributed according to $N(0,1)$, the probability of observing N firms in a market conditional on the market characteristics X and market size variable S^* is equal to:

$$\Pr(N|X,S^*) = \Phi[-(X\lambda + \sigma \ln S^* - \gamma - \sum_{n=2}^{N} \theta^n)] - \Phi[-(X\lambda + \sigma \ln S^* - \gamma - \sum_{n=2}^{N+1} \theta^n)] \cdot \mathbf{1}(N \geq 1), \tag{8}$$

where $\Phi(\cdot)$ is the cumulative distribution function of the normal distribution and $\mathbf{1}(N \geq 1)$ is an indicator function equal to 1 for $N \geq 1$. These probabilities describe an ordered probit regression on the discrete ordered variable N. The θ^n parameters represent the increments between successive cut points in an ordered probit.

Using the probabilities defined in (8), we can estimate the static model by choosing parameters $\boldsymbol{\theta} \equiv (\lambda, \sigma, \gamma, \theta^n)$ to maximize the log-likelihood function:

$$\mathcal{L}(\boldsymbol{\theta}) = \sum_{m=1}^{M} \Pr(N_m|X_m, S_m^*; \boldsymbol{\theta}), \tag{9}$$

where m indicates market, defined as the county in our context. Since the variance of the error term is normalized to 1, the coefficients are only estimated up to an unknown scale factor. This limitation does not affect the estimates of demand thresholds: since these are calculated as ratios of linear combinations of the coefficient estimates, the unknown scale factor cancels out.

Using the coefficient estimates obtained by maximizing (9), I estimate the entry thresholds $ET(N)$ by solving for the value of S^* for which per firm profits equal 0, evaluating market characteristics X at their mean and the unobservable component of per firm variable profits at zero:

$$ET(N) = S_N^* = \exp\left(\frac{\gamma + \sum_{n=2}^{N} \theta^N - \bar{X}\lambda}{\sigma}\right). \tag{10}$$

Entry threshold ratios can then be estimated by substituting S_N^* for S_N in expression (4).

In the case of local daily newspapers, the assumption that firms produce homogeneous products is not likely to hold. As such, the entry threshold ratios obtained using the empirical model outlined in this section are not usable for making inferences about the competitive effects of entry. Nevertheless, the ETRs calculated using estimates from this model can be used to compare the market sizes necessary to support a given number of newspapers. It is also possible to compare entry thresholds over time.

5.2 Pseudo-Dynamic Bresnahan-Reiss Model

When we talk about 'entry', we usually mean 'there is a firm in the market that was not there before.' In the static Bresnahan-Reiss framework described in the last section, we do not actually observe 'entry' in this sense. Instead, we observe the number of firms active in a market at a given time, and attempt to explain this number using a model of competition that endogenizes the number of firms. However, not all of these firms may be 'entrants' in the sense described above, as some may have been active in the market prior to the period during which we draw our observation of the number of firms. In reality, entrants and incumbents face different decisions. A potential entrant must decide whether to enter the market, taking into account its expected profits stream and any sunk costs of entry. An incumbent, on the other hand, must decide whether to exit, taking into account its expected profits stream (which may be negative) and any scrap value it earns or exit cost it incurs by shutting down and leaving the market.

Bresnahan and Reiss (1994) present a model of entry that distinguishes between entrants and incumbents. As before, firms are assumed to produce homogeneous products and have identical marginal costs. However, the profits of entrants and incumbents may differ at a given time due to the presence of sunk costs: while incumbents are assumed to pay per-period fixed costs of operation, entrants are assumed to pay both the per-period fixed cost, f, and a one-time sunk cost of entry, F. In addition, exiting firms recover some scrap value φ.[12] A firm enters when its expected future profits exceed the fixed cost of entry, remains in the market when expected future profits exceed its scrap value, and exits when expected future profits fall below the scrap value.

Assuming the same demand and cost structures outlined in the discussion of the static framework, a firm's equilibrium profits in a market with N_t firms at time t are given by:

$$\pi(p^*(N_t), N_t) = (1 - \chi_t)[(p^*(N_t) - c)q(p^*(N_t), N_t)S_t - f] - \chi_t^E F + \chi_t \varphi, \tag{11}$$

where χ_t is an indicator variable equal to 1 if a firm decides to exit the market, and χ_t^E is an indicator variable equal to 1 if a firm decides to enter the market.[13] Defining variable profits at time t as $v(N_t) \equiv (p^*(N_t) - c)q(p^*(N_t), N_t)$, profits at t can be expressed more succinctly as:

$$\pi(p^*(N_t), N_t) = (1 - \chi_t)[v(N_t)S_t - f] - \chi_t^E F + \chi_t \varphi. \tag{12}$$

Using the profit function defined in (12), we can infer the conditions on profits that characterize equilibrium in a market with N_t firms. Recall that in the static case equilibrium means all N firms in the market found it profitable to enter, while the $(N + 1)^{th}$ potential entrant did not. With the introduction of per period fixed costs distinct from the sunk cost of entry and the scrap value of exit for incumbent firms, this characterization becomes insufficiently detailed. When firms incur sunk costs of entry separate from per period fixed costs, the definition of an equilibrium at time t depends

[12] Bresnahan and Reiss (1994) assumes firms must pay some amount to exit (i.e., the scrap value is negative). In practice, there is no need to restrict the scrap value φ to be either strictly positive or negative. Hence, depending on its sign, φ could be interpreted as either a scrap value or a cost of exit.

[13] I borrow the practice of indicating entry and exit by χ^E and χ, respectively, from Collard-Wexler (2014).

on the number of firms that were active in the market at $t-1$. In other words, we can infer different information about firms' profits and costs depending on whether a market saw entry, exit, or no change in the number of firms from $t-1$ to t. The following three cases delineate the equilibrium conditions associated with each transition type:

i. Stasis: $N_t = N_{t-1}$

If the number of firms active in a market is unchanged from one period to another, then all N_t firms active at t must be earning variable profits in excess of their continuation value:

$$v(N_t)S_t > f + \varphi.$$

At the same time, profits must be insufficiently large to support the entry of an additional firm. Hence, an $(N_t+1)^{th}$ entrant must not expect to recover its opportunity cost of entry:

$$v(N_t+1)S_t \leq f + \varphi + F.$$

ii. Entry: $N_t > N_{t-1}$

An N^{th} entrant must earn variable profits in excess of its opportunity cost of of entry:

$$v(N_t)S_t > f + \varphi + F.$$

For N_t to be an equilibrium, entry at time t must be unprofitable for the $(N_t+1)^{th}$ firm:

$$v(N_t+1)S_t \leq f + \varphi + F.$$

iii. Exit: $N_t < N_{t-1}$

If one or more firms exit during the transition from $t-1$ to t, all firms remaining in the market must recover their opportunity costs of continuation:

$$v(N_t)S_t > f + \varphi.$$

Given that exit occurred from $t-1$ to t, per firm profits must also fall short of the level justifying the continuation of N_t+1 incumbents:

$$v(N_t + 1)S_t \leq f + \varphi.$$

Taken together, the conditions that must be met given the presence of N_t firms imply:

$$v(N_t)S_t > (f + \varphi) \cdot \mathbf{1}(N_t \leq N_{t-1}) + (f + \varphi + F) \cdot \mathbf{1}(N_t > N_{t-1}), \tag{13}$$

where $\mathbf{1}(\cdot)$ is an indicator function equal to one when its parenthetical condition is satisfied. Similarly, combining the conditions that rule out the possibility of the presence of $N_t + 1$ firms in equilibrium yields:

$$v(N_t + 1)S_t \leq (f + \varphi) \cdot \mathbf{1}(N_t < N_{t-1}) + (f + \varphi + F) \cdot \mathbf{1}(N_t \geq N_{t-1}). \tag{14}$$

From this point, all that remains is to parameterize the profit function and define an error structure that will allow for the econometric estimation of entry and exit thresholds. To do this, I employ a similar approach to that used earlier in the static model, although there is now the added complication of differentiating the opportunity cost of entry from the opportunity cost of continuation.

Assume per firm variable profits at t are known up to a market-specific unobservable component, ξ_t, such that per firm profits in a market are given by $v(N_t)S_t\xi_t$. Taking logs and rearranging, the inequalities given in (13) and (14) can then be restated:

$$\varepsilon_t > -\ln v(N_t) - \ln S_t + \ln(f + \varphi) \cdot \mathbf{1}(N_t \leq N_{t-1}) + \ln(f + \varphi + F) \cdot \mathbf{1}(N_t > N_{t-1})$$

$$\varepsilon_t \leq -\ln v(N_t + 1) - \ln S_t + \ln(f + \varphi) \cdot \mathbf{1}(N_t < N_{t-1}) + \ln(f + \varphi + F) \cdot \mathbf{1}(N_t \geq N_{t-1}).$$

We can now parameterize these conditions using an approach similar to the one taken with the static framework. To this end, define $\ln v(N_t) = X_t\lambda - \sum_{n=2}^{N} \theta^n$, $\ln S_t = \sigma \ln S_t^*$, $\ln(f + \varphi) = \gamma$, and $\ln(f + \varphi + F) = \gamma + \gamma^S$. Under these definitions, the parameter γ^S describes the difference between the continuation and entry thresholds.

Assuming ε_t is normally distributed according to $N(0, 1)$, the probability of observing N_t firms in a market at time t conditional on the market characteristics X_t, N_{t-1}, and S_t^*, is given by:

$$\begin{aligned}\Pr(N_t|N_{t-1}, X_t, S_t^*) = {} & \Phi[-(X_t\lambda + \sigma \ln S_t^* - \sum_{n=2}^{N_t} \theta^n - \gamma - \gamma^S \cdot \mathbf{1}(N_t > N_{t-1})] \\ & -\Phi[-(X_t\lambda + \sigma \ln S_t^* - \sum_{n=2}^{N_t+1} \theta^n - \gamma - \gamma^S \cdot \mathbf{1}(N_t \geq N_{t-1})] \cdot \mathbf{1}(N_t \geq 1).\end{aligned} \tag{15}$$

Using these probabilities, we can estimate the model via maximum likelihood, choosing parameters $\boldsymbol{\theta} \equiv (\lambda, \sigma, \theta^n, \gamma, \gamma^S)$ to maximize the log-likelihood function:

$$\mathcal{L}(\boldsymbol{\theta}) = \sum_{m=1}^{M} \sum_{t=1}^{T} \Pr(N_{mt} | N_{mt-1}, X_{mt}, S^*_{mt}; \boldsymbol{\theta}), \tag{16}$$

where the subscript m indicates county, and t indicates presidential year.

For this version of the model, I estimate entry thresholds $ET(N)$ by solving for the value of S^* that sets profits equal 0 when market characteristics X are evaluated at their means, the unobservable component of variable profit is evaluated at zero, and entry has occurred from $t-1$ to t (i.e., $N_t > N_{t-1}$):

$$ET(N) = S_N^{*E} = \exp\left(\frac{\gamma + \gamma^S + \sum_{n=2}^{N} \theta^n - \bar{X}\lambda}{\sigma}\right). \tag{17}$$

As before, the entry threshold ratio $ETR(N)$ is estimated by substituting (17) into expression (4).

The exit threshold $EX(N)$ (i.e., the minimum market size required to justify the continuation of N incumbents) is estimated using the same approach taken with the entry threshold, except it is assumed $N_t \leq N_{t-1}$:

$$EX(N) = S_N^{*X} = \exp\left(\frac{\gamma + \sum_{n=2}^{N} \theta^n - \bar{X}\lambda}{\sigma}\right). \tag{18}$$

6 Results

In all specifications, I use county population as the market size variable S^*. Table 2 describes the demographic variables and county characteristics used as controls in the reduced-form profit equation in all specifications of the empirical model. I estimate the model separately for four subsets of the sample period: 1872 to 1908, 1912 to 1940, 1944 to 1972, and 1976 to 2000.[14] This makes it possible to see how entry and exit thresholds have changed over time. It also offers some evidence about whether the relationships between the various demographic controls and market structure have changed over time.

[14] The rationale behind this choice of subgroups is discussed in Section 3. In short, they are chosen to reflect the advent of new technologies that compete with newspapers as sources of news and entertainment. The period from 1872 to 1908 reflects a time in which newspapers faced relatively little competition from other media, the period from 1912 to 1940 contains the introduction of radio, while the periods from 1944 to 1972 and 1976 to 2000 reflect the newspaper industry following the introduction and subsequent universalization of television, respectively.

Variable	1872 - 1908 Mean	1872 - 1908 SD	1912 - 1940 Mean	1912 - 1940 SD	1944 - 1972 Mean	1944 - 1972 SD	1976 - 2000 Mean	1976 - 2000 SD
Proportion of surface area of county occupied by farmland	0.69	0.29	0.67	0.28	0.65	0.29	0.53	0.30
Proportion of white population born outside the United States	0.11	0.10	0.07	0.08	0.05	0.07	0.03	0.05
Log of county population per square mile	3.56	1.19	3.80	1.28	4.09	1.39	4.39	1.43
Propoprtion of county population consisting of males aged 21 and older	0.27	0.06	0.30	0.04	0.30	0.03	0.32	0.02
Population employed in manufacturing as a share of males aged 21 and older	0.14	0.16	0.16	0.17	0.23	0.18	0.23	0.17
Proportion of county population living in urban areas (cities with $\geq$ 2,500 residents)	0.22	0.23	0.36	0.24	0.48	0.22	0.55	0.23
Proportion of county population that is white	0.88	0.20	0.89	0.16	0.91	0.13	0.88	0.13

Table 2: Description of Variables

6.1 Results from the Static Model

Parameter estimates from the static Bresnahan-Reiss model described by equation (8) are presented in Table 3. Note that the parameter estimates reflect the relationship between the explanatory variable and factors affecting both demand and costs in the reduced-form profit equation. As such, they do not have any meaningful structural interpretation. Still, we can make some inferences about the relationship between the demographic variables by considering their signs and by comparing the relative size of their estimated coefficients. It is also worth noting that the coefficient estimates yielded by the model represent the partial effect of a change in their associated explanatory variable up to some scale factor. As such, we must be cautious when comparing the magnitudes of the coefficient estimates for the same regressor across years, since the sample means of the regressors and the variance of per firm profits are not necessarily constant across periods.

	1872 - 1908	1912 - 1940	1944 - 1972	1976 - 2000
Log population	1.172^{***}	1.079^{***}	1.156^{***}	0.939^{***}
	(0.071)	(0.074)	(0.059)	(0.065)
Proportion of population white	0.184	0.422^{**}	1.007^{***}	1.038^{***}
	(0.137)	(0.167)	(0.238)	(0.238)
Proportion of white population foreign-born	-2.314^{***}	-1.248^{***}	-0.939^{***}	-0.361
	(0.327)	(0.404)	(0.332)	(0.669)
Proportion of adult population male	6.848^{***}	6.163^{***}	7.674^{***}	-2.342^{**}
	(0.622)	(0.837)	(0.979)	(1.089)
Farmland share of county surface area	0.534^{***}	0.474^{***}	0.520^{***}	0.280^{**}
	(0.138)	(0.118)	(0.104)	(0.111)
Proportion of population in urban areas	4.597^{***}	2.669^{***}	1.286^{***}	1.038^{***}
	(0.201)	(0.155)	(0.091)	(0.167)
Manufacturing empl. share of adult male population	-0.684^{***}	-0.102	0.370^{**}	1.265^{***}
	(0.221)	(0.175)	(0.183)	(0.194)
Log population per square mile	-0.316^{***}	-0.228^{***}	-0.153^{***}	-0.208^{***}
	(0.066)	(0.052)	(0.042)	(0.045)
γ (Fixed costs)	13.801^{***}	12.869^{***}	14.789^{***}	9.286^{***}
	(0.605)	(0.656)	(0.647)	(0.711)
θ^2 (First competitor)	0.772^{***}	1.404^{***}	2.091^{***}	2.584^{***}
	(0.025)	(0.040)	(0.060)	(0.077)
θ^3 (Second competitor)	1.016^{***}	1.216^{***}	1.374^{***}	1.174^{***}
	(0.034)	(0.044)	(0.067)	(0.061)
Observations	11981	11311	11409	9867
Counties	1310	1427	1432	1437
Log-likelihood	-8884.44	-10642.47	-9498.69	-7268.73

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Standard errors clustered by county in parentheses.

Table 3: Estimation Results of Static Bresnahan-Reiss Model

Generally, the signs of the demographic controls are consistent across time. In most cases, the estimated parameters are statistically different from zero at the 95 percent significance level. The log of population, the white proportion of the population, the proportion of the white population born outside the United States, the share of the population living in urban areas, and the proportion of county surface area occupied by farmland are all positively related with the number of active newspapers during each period. The log of population per square mile is negatively associated with the number of active newspapers in all periods. With the exception of the white share of population during the 1876 to 1908, as well as the proportion of the white population foreign-born between 1976 and 2000, the estimated coefficients for these controls are statistically different from zero at the 95 percent significance level.

The share of the overall population accounted for by males over the age of 21 is positively associated with the number of newspapers during the first three periods. The estimated coefficient for this control is statistically different from zero at the 95 percent significance level during all periods. The direction of the relationship between a county's level of manufacturing employment relative to its adult male population and the number of active newspapers appears to change over time. Between 1872 and 1908, the estimated coefficient for this control is negative and statistically significant. Conversely, after 1944 the estimated coefficient is positive and statistically significant. Between 1912 and 1940, manufacturing employment does not appear to have a statistically significant relationship with the number of newspapers in a county.

During all periods, the entry parameter γ is large relative to the competitive effects of subsequent entry. In essence, the parameter γ represents the opportunity cost incurred by entry into the newspaper market when there are no other firms in the market. Under the static framework outlined in Section 5, a firm will not enter into a county in which its expected variable profits fall short of this value. In order to capture the erosion of per firm profits resulting from the competitive effects of entry, I allow this cost to increment with the number of firms. The incremental effects of the second and third entrants are captured by the parameters θ^2 and θ^3, respectively. These increments are equivalent to the differences between the cut points yielded by estimating an ordered probit model. Note that these parameters reflect changes in both variable profits and fixed costs resulting from the entry of additional competitors.

Notably, during the period between 1872 and 1908, the estimated effect of the entry of the first competitor is smaller than that associated with the entry of a second competitor. A possible explanation for the relatively larger competitive effect of the third entrant may be that there are diminishing returns to product differentiation in the newspaper industry, even with relatively low numbers of products. In that case, the market-expanding effect of the entrance of a product-differentiated competitor declines relative to the business-stealing effect of entry as the number of competitors increases. Unfortunately, the counter-factual assumption of product homogeneity required by the Bresnahan-Reiss framework makes it impossible to distinguish between these effects in the empirical model.

This pattern is reversed during all subsequent periods, in which the competitive effects of entry diminish as the number of newspapers in a market increases. Likelihood ratio tests reject the null hypothesis that the competitive effects of subsequent entrants are equal (i.e., $\theta^2 = \theta^3$) at the 99 percent significance level for all four periods. The p-values of likelihood ratio tests of this null hypothesis against a two-sided alternative during each period are reported in Table 4.

Period	$\theta^2 = \theta^3$
1872 - 1908	0.0000
1912 - 1940	0.0000
1944 - 1972	0.0000
1976 - 2000	0.0000

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Table entries represent p-values from the likelihood ratio test of the null hypothesis specified in the column header against a two-sided alternative. Let θ^N denote the competitive effect of entry of the N^{th} entrant.

Table 4: Likelihood Ratio Tests for Static Bresnahan-Reiss Model

Figure 10 presents the entry thresholds predicted by the model for each period, expressed in terms of log population.[15] Between 1872 and 1908, a county population of roughly 25,000 was required to support the entry of a monopolist. Slightly less than twice as many people were needed to support a duopolistic local newspaper market, with the entry threshold lying around 49,000. The triopoly entry threshold was less than three times this amount, at 116,000. The duopoly and triopoly entry thresholds predicted by the static framework increased over time, while the monopoly entry threshold declined. Between 1976 and 2000, the monopoly entry threshold was 15,000 - 60 percent of the level seen between 1872 and 1908. By contrast, the population levels required to support two and three firms were roughly 240,000 and 837,000, respectively, during this period - far greater the levels seen between 1872 and 1908.

[15]Entry thresholds in levels are reported in Appendix Table A.1.1.

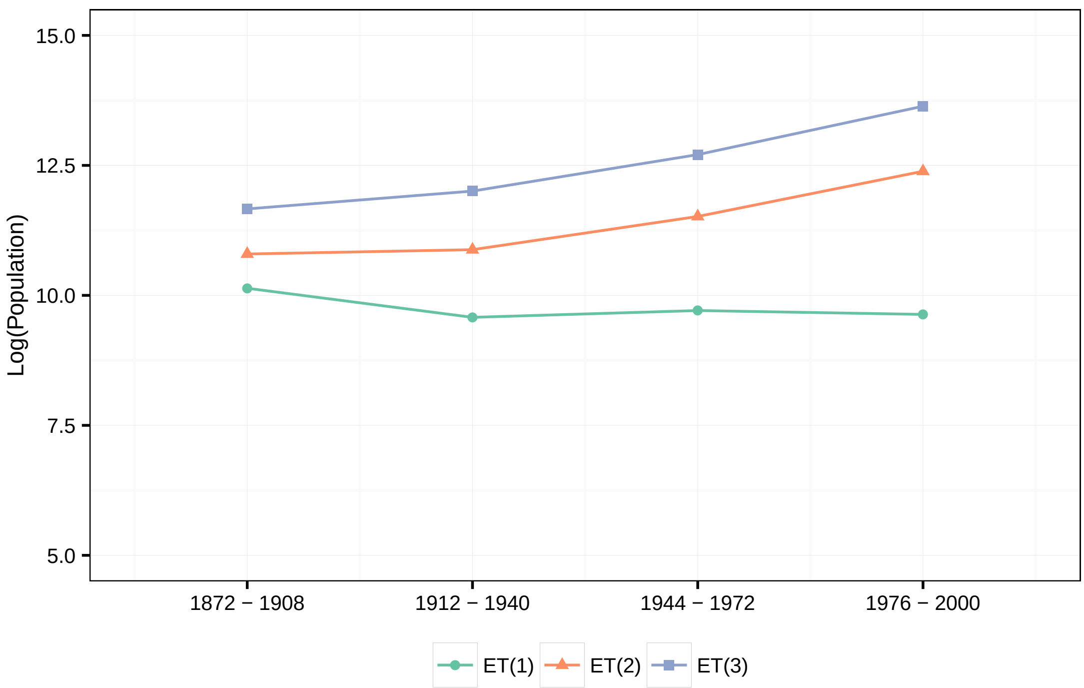

Figure 10: Estimated Entry Thresholds from Static Specification

The entry threshold ratios reflect the divergence in the monopoly and duopoly entry thresholds. Figure 11 plots the entry threshold ratios (ETR) implied by the point estimates of the entry thresholds yielded by the static Bresnahan-Reiss framework. As discussed in Section 5, these ratios compare the number of consumers required per firm to support a given number of entrants.

Notably, the number of consumers per firm required to support two newspapers has increased considerably relative to the number of consumers required to support a monopoly. During the period between 1872 and 1908, when entry rates in the local newspaper were highest, the ETR for a duopoly was near unity. In other words, during this period, the number of consumers per firm required to support a duopolistic local newspaper market was approximately the same as that required to support a monopoly. This ratio was considerably higher between 1976 and 2000: the number of consumers per firm required to support the entrance of a duopolist was nearly eight times greater than the number required to support the entrance of a monopolist. By comparison, the ETR for a triopoly relative to a duopoly appears to have remained stable across time. From 1912 onward, slightly more than double the number of per firm consumers were required to support the entrance of a third newspaper into a local newspaper market relative to the number required for the entrance of a second newspaper.

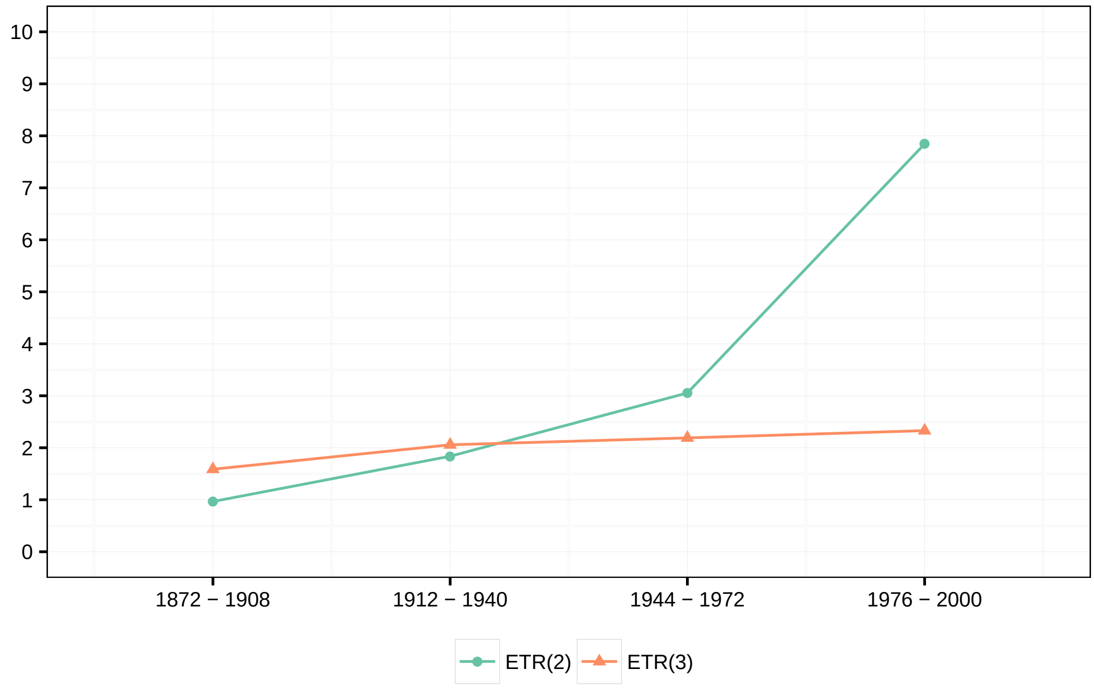

FIGURE 11: Entry Threshold Ratios from Static Specification

6.2 Results from the Pseudo-Dynamic Model

Baseline Sunk Cost Specification

The results of estimating the pseudo-dynamic Bresnahan-Reiss model described by the log likelihood function (16) is given in Table 5. As before, I estimate the model separately for county-years falling between 1872 to 1908, 1912 to 1940, 1944 to 1972, and 1976 to 2000. County-years in 1872 cannot be included in the estimation for the first period because the likelihood function is conditioned on the number of active newspapers in a given county during the previous presidential election year; this information is not available for 1868. In a handful of other cases, county-years in later periods have to be dropped due to the absence of information about the number of active newspapers in a given county during the election year prior. The results of the estimation are reported in Table 5.

	1872 - 1908	1912 - 1940	1944 - 1972	1976 - 2000
Log population	1.010^{***}	0.840^{***}	0.855^{***}	0.802^{***}
	(0.060)	(0.057)	(0.049)	(0.053)
Proportion of population white	0.256^{**}	0.392^{***}	1.383^{***}	-0.150
	(0.118)	(0.125)	(0.178)	(0.192)
Proportion of white population foreign-born	-2.333^{***}	-2.674^{***}	-0.122	0.349
	(0.281)	(0.300)	(0.321)	(0.682)
Proportion of adult population male	6.004^{***}	5.148^{***}	-4.913^{***}	15.243^{***}
	(0.551)	(0.634)	(0.816)	(1.173)
Farmland share of county surface area	0.515^{***}	0.124	0.182^{**}	0.381^{***}
	(0.124)	(0.087)	(0.082)	(0.090)
Proportion of population in urban areas	3.746^{***}	2.261^{***}	1.617^{***}	0.440^{***}
	(0.173)	(0.118)	(0.092)	(0.150)
Manufacturing empl. share of adult male population	-0.488^{***}	-0.279^{**}	0.576^{***}	1.075^{***}
	(0.184)	(0.137)	(0.147)	(0.159)
Log population per square mile	-0.271^{***}	-0.196^{***}	-0.165^{***}	-0.196^{***}
	(0.055)	(0.039)	(0.032)	(0.037)
θ^2 (First competitor)	0.647^{***}	1.051^{***}	1.606^{***}	2.024^{***}
	(0.020)	(0.030)	(0.053)	(0.067)
θ^3 (Second competitor)	0.858^{***}	0.990^{***}	1.032^{***}	0.976^{***}
	(0.029)	(0.033)	(0.053)	(0.058)
γ (Continuation value)	11.329^{***}	8.951^{***}	7.305^{***}	11.412^{***}
	(0.510)	(0.517)	(0.537)	(0.628)
γ^S (Sunk costs)	1.184^{***}	2.195^{***}	2.917^{***}	3.162^{***}
	(0.027)	(0.033)	(0.050)	(0.073)
Observations	10626	11180	11376	9859
Counties	1294	1426	1432	1437
Log-likelihood	-7009.60	-6255.72	-4240.34	-3329.34

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Standard errors clustered by county in parentheses.

Table 5: Estimation Results of Pseudo-Dynamic Bresnahan-Reiss Model

In general, the estimated coefficients for the demographic and other county characteristic variables yielded by this variant of the model are similar to those of the static framework. The coefficient on log population remains positive and statistically significant above the 99 percent level across all periods. Notably, the magnitude of the relationship between the log of population and the number of newspapers is smaller in the specification that includes sunk costs relative to that found in the baseline static specification.

The white share of the population is positively related with the number of active newspapers during the first three periods. The estimated coefficient is statistically different from zero at the 95 percent significance level during each of these periods. During the 1972 to 2000 period, the parameter estimate for this control is negative, although it is not statistically different from zero. The proportion of the white population born outside of the United States is negatively associated with the number of newspapers during the first two periods and positively associated during the last two. After 1944, however, the relationship between the foreign-born share of the white population and the number of newspapers in a county is not statistically different from zero.

During the 1872 to 1908 and 1912 to 1940 periods, the estimated coefficients on the proportion of county population made up of males age 21 and older and the number of newspapers are similar. Both are positive, large in magnitude relative to the other coefficient estimates, and statistically significant at the 99 percent level. During the latter two periods, the static and pseudo-dynamic models show a different relationship between this control and the number of active newspapers. Between 1944 and 1972, the static model finds a statistically significant positive relationship between the adult male share of the population and the number of newspapers, while the sunk cost version of the model finds a statistically significant negative relationship. The converse is true for the 1972 to 2000 period, during which time the static model predicts a negative relationship while the sunk cost model predicts a positive relationship.

As far as I can tell, this inconsistency does not lend itself to any straightforward explanation. The high sensitivity of this parameter to the choice of sample period and the inclusion of sunk costs suggests there is a correlation between the male share of the adult population and unobserved factors unaccounted for by the estimated specification. This lack of robustness serves as a reminder that the parameter estimates in the reduced-form profit equation do not have a meaningful structural interpretation. The principal motivation for their inclusion is simply to reduce the amount of variation left in the market-year-level residuals in order to improve our estimates of the entry and exit thresholds.

Unlike in the static model, the share of county surface area occupied by farmland does not have a statistically significant relationship with the number of newspapers during the 1912 to 1940 period when sunk costs are included in the model. In the other periods, however, the estimated coefficients resemble those yielded by the model without sunk costs. The estimated coefficients on the proportion of county population living in urban areas are mostly unchanged in the sunk cost version of the model. The relationship remains positive and statistically significant at the 95 percent level in all periods. In addition, the magnitudes of the estimated coefficients are reasonably similar under either version of the model.

As was the case with the static model, the ratio of manufacturing employment to the adult male population is negatively associated with the number of newspapers during the first two periods and positively associated during the last two. In the sunk cost version of the model, this relationship is statistically different from zero at the 95 percent confidence level throughout all four periods.

In this specification, the parameter γ^S reflects the difference between the entry and continuation thresholds for a given number of firms. The results of the estimation show that this gap has widened over time. When sunk costs are included in the model, γ reflects the continuation value of a monopolist, while the parameters θ^2 and θ^3 reflect the effects of the entry of the second and third competitors, respectively, on the continuation value. As with the version of the model without sunk costs, the estimated results suggest the entrance of a third newspaper has a relatively larger competitive effect of entry than the entrance of a second competitor during the period between 1872 and 1908. This pattern reverses in later periods, during which the competitive effect of entry of the second entrant is relatively larger than that of the third. A likelihood-ratio test of the null hypothesis that the effect of entry does not change with the number of firms (i.e., $\theta^2 = \theta^3$) fails to reject the null at the 95 percent significance level for the 1912 to 1940 period. During all other periods, the test rejects the null hypothesis at the 99 percent significance level (Table 6). Additionally, likelihood ratio tests reject the null hypothesis that entry and exit thresholds are equal (i.e., $\gamma^S = 0$) in the newspaper industry at any reasonable level of significance.

Period	$\gamma^S = 0$	$\theta^2 = \theta^3$
1872 - 1908	0.0000	0.0000
1912 - 1940	0.0000	0.1293
1944 - 1972	0.0000	0.0000
1976 - 2000	0.0000	0.0000

Table entries represent p-values from the likelihood ratio test of the null hypothesis specified in the column header against a two-sided alternative. Let γ^S denote the sunk cost of entry parameter. Let θ^N represent the effect of the entrance of the N^{th} firm on firms' continuation values.

TABLE 6: Likelihood Ratio Tests for Pseudo-Dynamic Bresnahan-Reiss Model

Once sunk costs are incorporated into the model, we no longer find a decline in the monopoly entry threshold over time (Figure 12), contrary to the findings of the baseline static Bresnahan-Reiss specification.[16] Instead, we find that the monopoly entry threshold increased steadily over time, although at a slower pace than the duopoly and triopoly entry thresholds. The monopoly exit threshold estimated by the sunk cost framework appears to have declined over time, while the duopoly and triopoly exit thresholds have increased over time.[17]

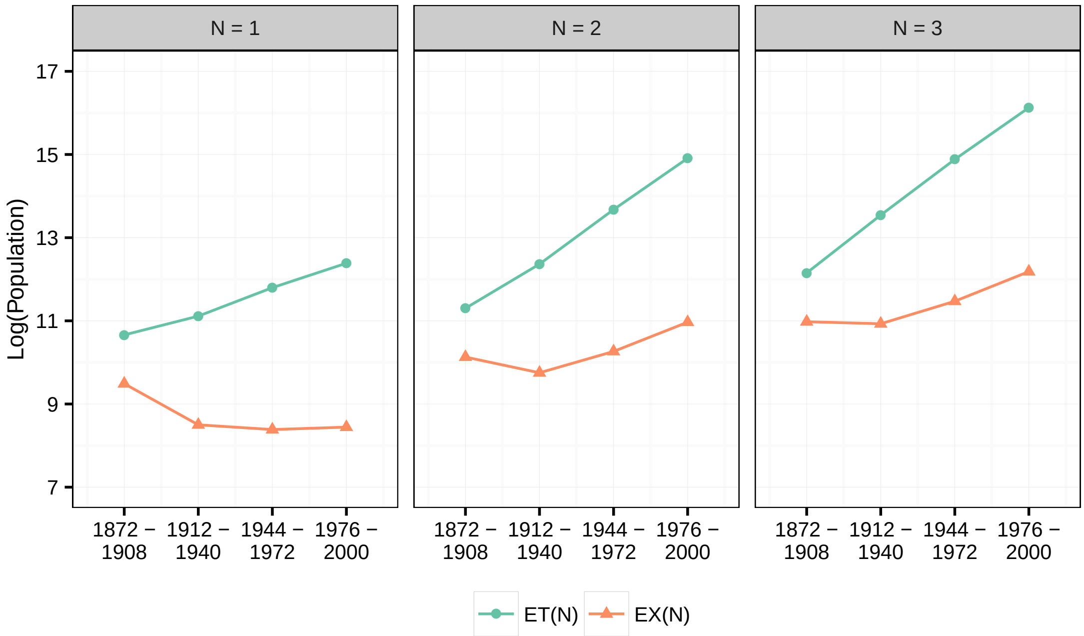

FIGURE 12: Entry and Exit Thresholds, Baseline Sunk Cost Specification

Between 1872 and 1908, a county population of around 43,000 was required to support the entry of a monopolistic local newspaper, while a population of 13,000 sufficed to support the continuation of an incumbent monopolist. The duopoly and triopoly entry thresholds during this period were approximately 81,000 and 189,000, respectively. The exit thresholds for incumbent duopolists and tripolists were 25,000 and 59,000. By the 1972 to 2000 period, the monopoly entry threshold had increased significantly, with a county population of roughly 240,000 required to support the entry of a monopolist - far in excess of the estimated entry threshold of 10,000 yielded by the static framework. By contrast, the estimated monopoly exit threshold for this period is below 5,000, while

[16] Demand thresholds in levels are reported in Appendix Table A.1.2.

[17] Recall that the exit threshold is defined as the minimum market size required to support a given number of incumbent firms.

the duopoly and triopoly exit thresholds are 58,000 and 196,000, respectively.

The large disparity between entry and exit thresholds found here shows the importance of accounting for the existence of sunk costs when modeling market structure in the local daily newspaper industry. In practice, the application of the term 'entry threshold' to the threshold estimate yielded by the static Bresnahan-Reiss framework is somewhat misleading. The threshold estimate produced by that framework is actually a weighted average of the entry and exit thresholds for a given industry. In an industry in which entry rates are low, the threshold estimate yielded by the static framework will tend to approximate the exit threshold more closely than it does the entry threshold.

The ETRs predicted by the sunk cost version of the model exhibit broadly similar patterns to those obtained using the model without sunk costs. As Figure 13 shows, the number of consumers per firm required to support the entrance of a second local newspaper has increased considerably relative to the number required to support the entrance of a monopolist, while the triopoly-to-duopoly ETR has remained essentially constant across time. It is worth noting that the sunk cost specification tends to imply lower ETRs compared to the version of the model without sunk costs. This is especially noticeable with the duopoly-to-monopoly ETR later on in the sample period. For example, between 1976 and 2000 the sunk cost model estimates $ETR(2)$ to be slightly over six, while the static specification places the ratio around eight. This discrepancy appears to be largely the result of the static framework's under-estimation of the monopoly entry threshold. Over time, the magnitude of this downward bias increases, which in turn results in a relatively larger upward bias in the static Bresnahan-Reiss model's estimate of $ETR(2)$.[18]

[18] Note that although the static framework's estimates of the duopoly and triopoly entry thresholds are also biased downward, this should only result in a biased estimate of the ETR if the degree of bias in the estimated entry threshold varies across market structures. As Figure 16 and its accompanying discussion later in this paper shows, it appears that this is indeed the case for the local daily newspaper industry.

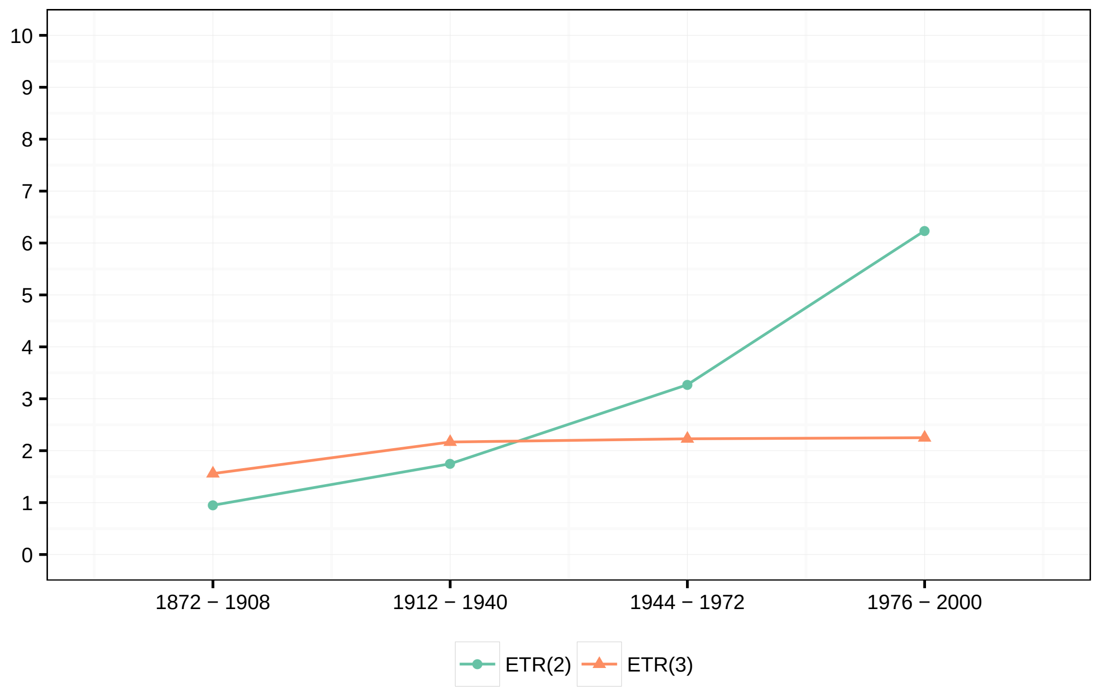

FIGURE 13: Entry Threshold Ratios from Model with Sunk Costs

Model with Flexible Sunk Costs

I also estimate a more flexible version of the sunk cost model by allowing sunk costs to vary with the order of entry. I implement this flexibility by assuming the relationship between variable profits and the number of newspapers has the form $\gamma + \sum_{n=2}^{N} \theta_E^N \cdot \mathbf{1}(N = n)$ and the relationship between sunk costs and the number of newspapers has the form $\gamma^S + \sum_{n=2}^{N} \theta_S^N \cdot \mathbf{1}(N = n)$. This relaxes the assumption implicit in the baseline specification that $\theta_E^N = \theta_S^N = \theta^N$ for all N. In other words, this specification allows entry and exit thresholds to increment by different amounts with the entry of each additional competitor.

The estimated results for this specification are reported in Table 7. Allowing sunk costs to vary with the order of entry has little effect on the estimated coefficients of the demographic controls and county characteristics included in the reduced-form profit function. In most cases, the point estimates of the parameters do not differ even at the second decimal place when compared to the baseline sunk cost model.

	1872 - 1908	1912 - 1940	1944 - 1972	1976 - 2000
Log population	1.009***	0.843***	0.854***	0.806***
	(0.060)	(0.057)	(0.048)	(0.053)
Proportion of population white	0.258**	0.405***	1.393***	−0.165
	(0.118)	(0.124)	(0.179)	(0.189)
Proportion of white population foreign-born	−2.326***	−2.637***	−0.202	0.347
	(0.281)	(0.301)	(0.317)	(0.664)
Proportion of adult population male	5.994***	5.113***	−4.797***	15.359***
	(0.551)	(0.632)	(0.818)	(1.152)
Farmland share of county surface area	0.514***	0.127	0.150*	0.378***
	(0.123)	(0.087)	(0.081)	(0.090)
Proportion of population in urban areas	3.741***	2.228***	1.604***	0.336**
	(0.173)	(0.119)	(0.090)	(0.152)
Manufacturing empl. share of adult male population	−0.486***	−0.259*	0.599***	1.082***
	(0.184)	(0.137)	(0.145)	(0.159)
Log population per square mile	−0.270***	−0.197***	−0.168***	−0.195***
	(0.055)	(0.039)	(0.032)	(0.036)
$\theta_{\mathrm{E}}{}^2$	0.609***	1.095***	1.668***	2.096***
	(0.032)	(0.037)	(0.056)	(0.066)
$\theta_{\mathrm{E}}{}^3$	0.891***	1.046***	1.140***	1.016***
	(0.046)	(0.042)	(0.064)	(0.061)
γ (Continuation value)	11.333***	8.938***	7.243***	11.376***
	(0.511)	(0.513)	(0.527)	(0.624)
γ^S (Sunk costs)	1.165***	2.308***	3.124***	3.440***
	(0.031)	(0.042)	(0.067)	(0.092)
$\theta_{\mathrm{S}}{}^2$	−0.062	0.131**	0.251***	0.512***
	(0.038)	(0.056)	(0.088)	(0.119)
$\theta_{\mathrm{S}}{}^3$	0.059	0.179***	0.454***	0.367***
	(0.053)	(0.065)	(0.095)	(0.116)
Observations	10626	11180	11376	9859
Counties	1294	1426	1432	1437
Log-likelihood	−7007.84	−6243.67	−4208.85	−3304.72

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Standard errors clustered by county in parentheses. Let θ_E^N denote the effect of the N^{th} entrant on firms' continuation values. Let θ_S^N denote the effect of the N^{th} entrant on firms' sunk costs.

Table 7: Estimation Results of Pseudo-Dynamic Bresnahan-Reiss Model, $\theta_E^N \neq \theta_S^N$

Allowing sunk costs to vary with the order of entry separately from firms' continuation values produces an estimate of the monopoly exit threshold parameter γ comparable to the estimates yielded by the specifications used earlier in this paper. As before, the monopoly exit parameter is large relative to the incremental effects of subsequent entrants. The flexible sunk cost specification again finds the competitive effect of the first entrant to exceed that of the second during the period between 1872 and 1908. Between

1912 and 1940, as in the baseline sunk cost specification, the flexible sunk cost specification finds that the competitive effects of the first and second competitors are essentially equal. After this period, the competitive effect of the third entrant exceeds that of the second, consistent with the findings of the other specifications. Likelihood ratio tests of the null hypothesis that the competitive effects of entry do not vary with the order of entry ($\theta_E^2 = \theta_E^3$) fail to reject the null hypothesis at the 95 percent significance level for the period between 1912 and 1940. However, they reject the null at the 99 percent significance level for all other periods (Table 8).

Period	$\theta_E^2 = \theta_E^3$	$\theta_S^2 = \theta_S^3 = 0$	$\theta_S^2 = \theta_S^3$
1872 - 1908	0.0000	0.1727	0.0750
1912 - 1940	0.3229	0.0000	0.5687
1944 - 1972	0.0000	0.0000	0.0986
1976 - 2000	0.0000	0.0000	0.3613

Table entries represent p-values from the likelihood ratio test of the null hypothesis specified in the column header against a two-sided alternative. Let θ_E^N denote the effect of the N^{th} entrant on firms' continuation values. Let θ_S^N denote the effect of the N^{th} entrant on firms' sunk costs.

Table 8: Likelihood Ratio Tests for Flexible Sunk-Cost Model

The model provides little evidence to suggest that sunk costs vary with the order of entry in the nonlinear manner permitted by the flexible sunk cost specification. Likelihood ratio tests of the null hypothesis that $\theta_S^2 = \theta_S^3$ fail to reject the null at the 95 percent significance level during all four periods. Still, the model does provide some evidence to suggest that sunk costs are higher for the second and third entrants than for an initial entrant. I verify this by conducting likelihood ratio tests of the null hypothesis that $\theta_S^2 = \theta_S^3 = 0$. If this equality holds, the model collapses to the baseline sunk cost specification estimated in the previous subsection. For the 1872 to 1908 period, the likelihood ratio test fails to reject this null hypothesis at the 95 percent significance level. However, during all subsequent periods the test rejects the null hypothesis at the 99 percent significance level. This suggests that from 1912 onward, sunk costs increased approximately linearly with the order of entry.

Allowing sunk costs and continuation values to vary separately with the order of entry does not change the trends in entry and exit thresholds estimated by the inflexible sunk cost specification (Figure 14). As implied by the estimates from the baseline sunk cost specification, entry thresholds for all market structures increased over time.[19]

[19]Demand thresholds in levels are reported in Appendix Table A.1.3.

The duopoly and triopoly entry thresholds increased by a larger degree than monopoly thresholds. The estimated monopoly exit threshold decreased over time, while duopoly and triopoly exit thresholds increased slightly over time. In the next section, I compare the findings of the various specifications in more detail.

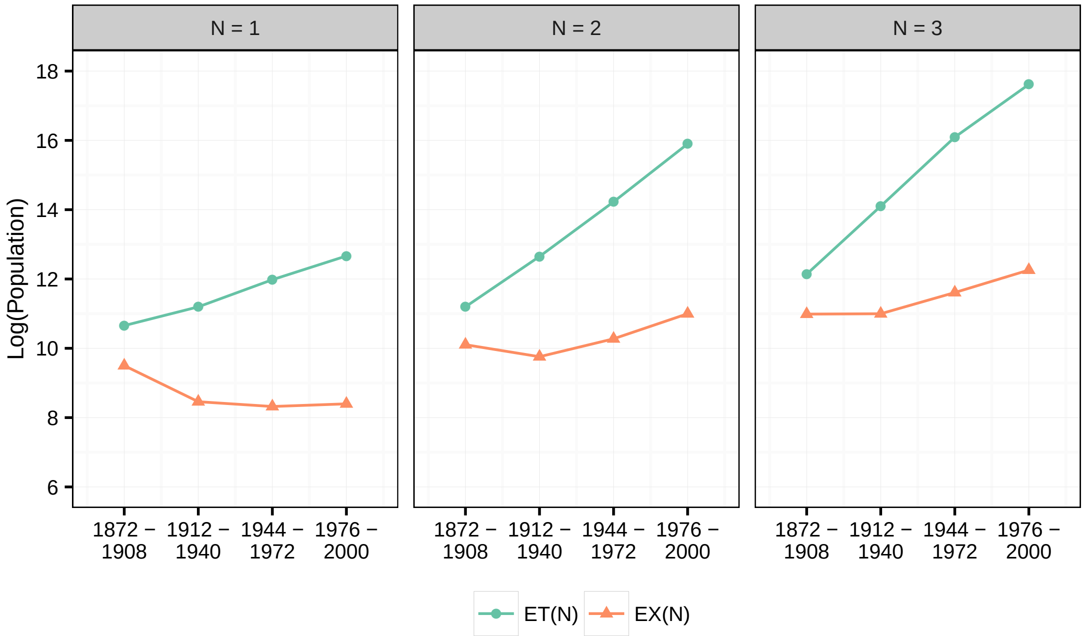

Figure 14: Entry and Exit Thresholds by Market Structure, Flexible Sunk Cost Specification

The flexible sunk cost specification tends to yield higher estimated ETRs relative to the baseline sunk cost specification. Since sunk costs increase with the number of incumbents according the estimated results of the flexible specification, entry becomes relatively more costly as the number of incumbents increases. This tendency is most pronounced during the last two periods. Between 1944 and 1972, the flexible specification estimates that roughly five times as many consumers per firm are required to support the entry of a second entrant relative to the number required for the first entrant (Figure 15). For the 1976 to 2000 period, the disparity is even larger, with nearly 13 times more consumers per firm required to support the entry of the second newspaper relative to the first.

In addition, when sunk costs are allowed to vary with the number of firms separately from continuation values, we no longer find that the estimate of $ETR(3)$ is constant over time. After 1944, the entry thresholds estimated by the flexible sunk cost model imply that roughly four times more consumers per firm are required to support the entry of a third newspaper than were required for the entry of the second firm. By contrast, $ETR(2)$ is estimated to be around 2.2 after 1944 according the both the baseline sunk cost and static specifications.

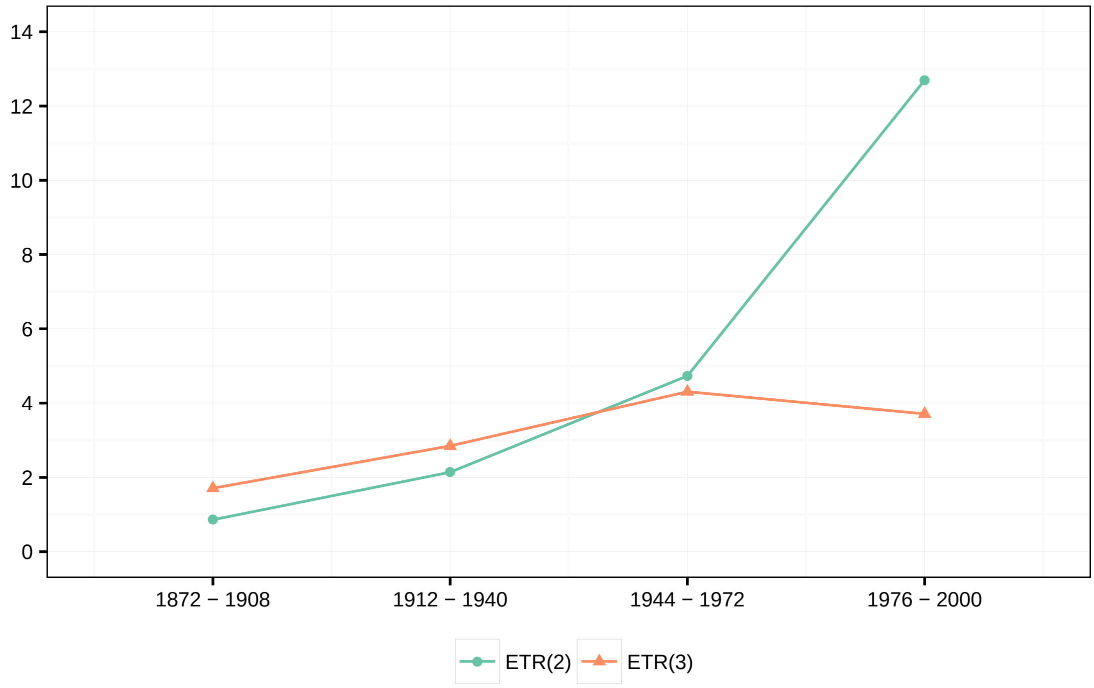

FIGURE 15: Entry Threshold Ratios from Model with Flexible Sunk Costs

6.3 Comparison of Results across Specifications

Figure 16 plots the entry thresholds for each period estimated using the static Bresnahan-Reiss model, the extension of the model allowing for sunk costs, and the flexible sunk cost version of the model. The plot emphasizes the extent to which the static framework underestimates entry thresholds when entry rates are low. The gap between the entry thresholds predicted by the static and sunk cost versions of the model are relatively small across all market structures between 1872 and 1908, when entry rates in the newspaper industry were highest.[20] During subsequent periods, as entry rates fall, the gap between the entry thresholds predicted by the static framework and the sunk cost specifications widens considerably. For monopolistic local newspaper markets, entry thresholds move in opposite directions over time depending on whether sunk costs are explicitly accounted

[20]See Figure 2.

for in the model. This inaccuracy suggests that the static Bresnahan-Reiss framework is not well suited to estimating entry thresholds in industries that have reached a relatively stable long-run equilibrium.

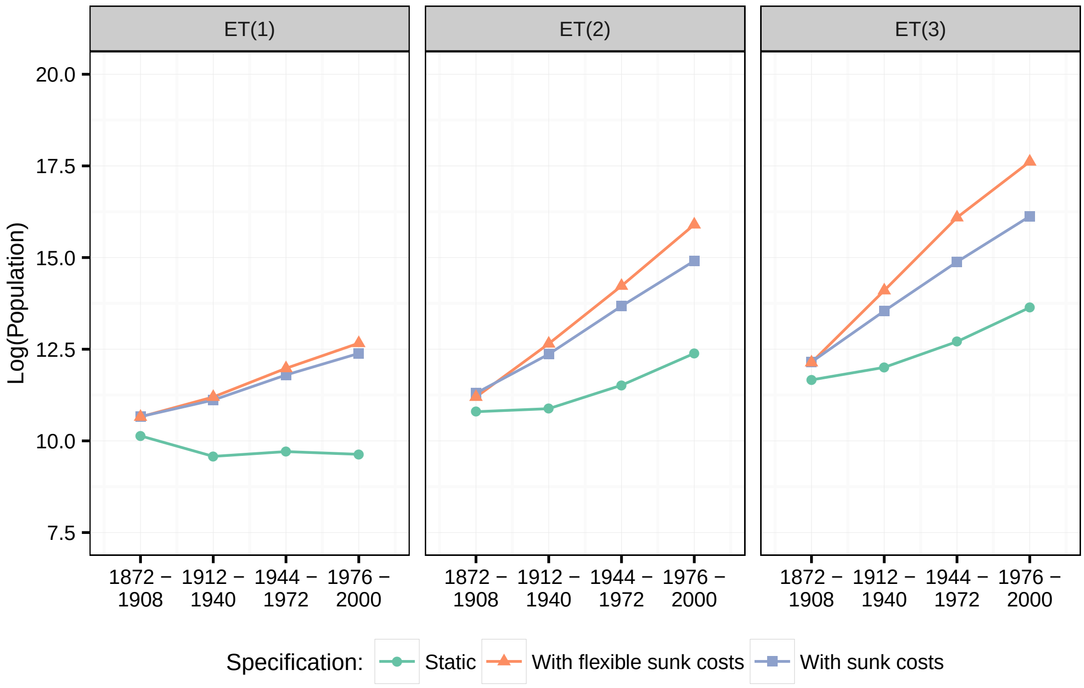

FIGURE 16: Entry Thresholds, Various Specifications

As Figure 16 shows, allowing sunk costs to increment separately from continuation values as the number of firms increases tends to result in higher point estimates of the entry thresholds relative to the baseline sunk cost specification, particularly for two- and three-newspaper counties. However, the differences between these estimates are generally not statistically significant. Table 9 reports p-values from tests of the null hypothesis that the entry threshold for a given period in the flexible sunk cost specification is equal to the point estimate of the entry threshold yielded by the baseline sunk cost specification for the same period. For most periods and market structures, the gap between the point estimates of the two specifications is not statistically different from zero at the 95 percent significance level. There are exceptions to this similarity, however: for triopolistic markets during the 1944 to 1972 and 1972 to 2000 periods, we reject the null hypothesis of equality between the entry threshold estimates of the two specifications at the 95 percent significance level.

Period	$\gamma^{1\{Entry\}}_{Baseline} = \gamma^{1\{Entry\}}_{Flexible}$	$\gamma^{2\{Entry\}}_{Baseline} = \gamma^{2\{Entry\}}_{Flexible}$	$\gamma^{3\{Entry\}}_{Baseline} = \gamma^{3\{Entry\}}_{Flexible}$
1872 - 1908	0.9628	0.7295	0.9492
1912 - 1940	0.7571	0.4206	0.1489
1944 - 1972	0.7029	0.2643	0.0170
1976 - 2000	0.5953	0.0906	0.0154

Table entries represent p-values from the likelihood ratio test of the null hypothesis specified in the column header against a two-sided alternative. Let $\gamma^{N\{Entry\}}_{Baseline} = \gamma + \gamma^S + \sum_{n=2}^{N} \theta^n$ denote the point estimate of the sum of the entry threshold parameters for N firms yielded by the baseline specification. Let $\gamma^{N\{Entry\}}_{Flexible} = \gamma + \sum_{n=2}^{N} \theta^n_E + \gamma^S + \sum_{n=2}^{N} \theta^n_S$ denote the sum of the entry threshold parameters in the flexible sunk cost specification for N firms.

Table 9: Likelihood Ratio Tests for Equality of Entry Thresholds Across Baseline and Flexible Sunk Cost Specifications

The baseline and flexible versions of the sunk cost model yield similar estimates of exit thresholds for the various market structures (Figure 17). In monopolistic and duopolistic local markets, the two specifications produce essentially identical estimates of exit thresholds during all four periods. In triopolistic markets, the flexible sunk cost specification tends to produce a slightly higher point estimate of the exit threshold from 1912 onward. However, likelihood ratio tests fail to reject the null hypothesis that the point estimates of the exit thresholds are the same across the two specifications at the 95 percent significance level for each individual period and market structure (Table 10). As such, there is little evidence to suggest that allowing sunk costs to vary separately from continuation values with the order of entry has any effect on the estimated exit thresholds.

Period	$\gamma^{1\{Exit\}}_{Baseline} = \gamma^{1\{Exit\}}_{Flexible}$	$\gamma^{2\{Exit\}}_{Baseline} = \gamma^{2\{Exit\}}_{Flexible}$	$\gamma^{3\{Exit\}}_{Baseline} = \gamma^{3\{Exit\}}_{Flexible}$
1872 - 1908	0.9902	0.9174	0.9995
1912 - 1940	0.9681	0.9248	0.7967
1944 - 1972	0.8716	0.9986	0.7902
1976 - 2000	0.9342	0.9379	0.8691

Table entries represent p-values from the likelihood ratio test of the null hypothesis specified in the column header against a two-sided alternative. Let $\gamma^{N\{Exit\}}_{Baseline} = \gamma + \sum_{n=2}^{N} \theta^n$ denote the point estimate of the sum of the exit threshold parameters for N firms yielded by the baseline sunk cost specification. Let $\gamma^{N\{Exit\}}_{Flexible} = \gamma + \sum_{n=2}^{N} \theta^n_E$ denote the sum of the exit threshold parameters for N firms estimated in the flexible sunk cost specification.

Table 10: Likelihood Ratio Tests for Equality of Exit Thresholds Across Baseline and Flexible Sunk Cost Specifications

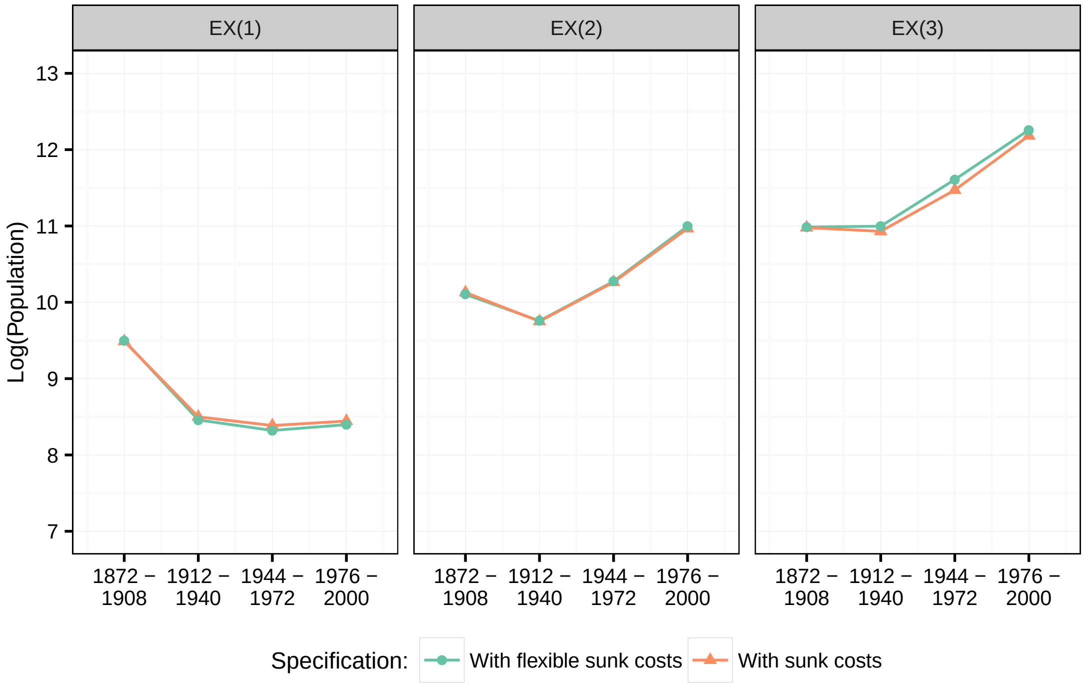

FIGURE 17: Exit Thresholds, Various Specifications

As we have seen, the flexible sunk cost specification has little to recommend it over the baseline sunk cost model in terms of producing sensible estimates of entry and exit thresholds in the local daily newspaper industry. In general, the implicit assumption of the baseline sunk cost specification that the gap between the entry and exit thresholds is constant in the number of firms appears reasonable. Still, the flexible specification is useful for attempting to verify whether sunk costs vary with the order of entry (as Table 8 shows, this indeed appears to be the case in the newspaper industry). Without allowing sunk costs to increment separately from continuation values with the number of firms, we could not test whether increases in sunk costs contribute to rising entry thresholds as the number of firms increases.

6.4 Effect of Competing Technologies on Local Newspaper Markets

During the 20$^{\text{th}}$ century, the newspaper industry faced substantial competition from substitute forms of media. The first disruption resulted from the emergence of radio as an alternative source of news media and entertainment in the United States during the 1930s and 1940s. The subsequent popularization of television during the 1950s further intensified the competitive pressure facing the newspaper industry from other media technologies. Radio and television competed with the newspaper industry along two dimensions: first, they acted as substitute sources of news and entertainment, and second, they offered alternative outlets for retailers and other firms seeking to purchase advertising space. In this subsection, I incorporate data reflecting the diffusion of radio and television across counties into the models of market structure outlined in Section 5 in order to estimate the effect these technologies had on demand thresholds in the local daily newspaper industry.

Radio's Effect on Demand Thresholds

County-level data on household radio ownership are available from the decennial US Census for 1930, 1940, and 1950. As with the demographic controls used throughout this paper, I interpolate the proportion of households owning radios during intercensal presidential election years by fitting a natural cubic spline for the number of households owning radios and the number of households in a county, then taking the ratio of these interpolated values.

During the early 1930s, radio ownership was somewhat uncommon: in most counties in our sample, less than half of all households owned radios (Figure 18). By 1940, radio had become considerably more popular. The modal county in 1940 had a 90 percent household radio ownership rate, and most counties had household radio ownership rates above 80 percent. This expansion continued throughout the 1940s, and by the end of the decade the modal county had a household radio ownership rate above 95 percent.

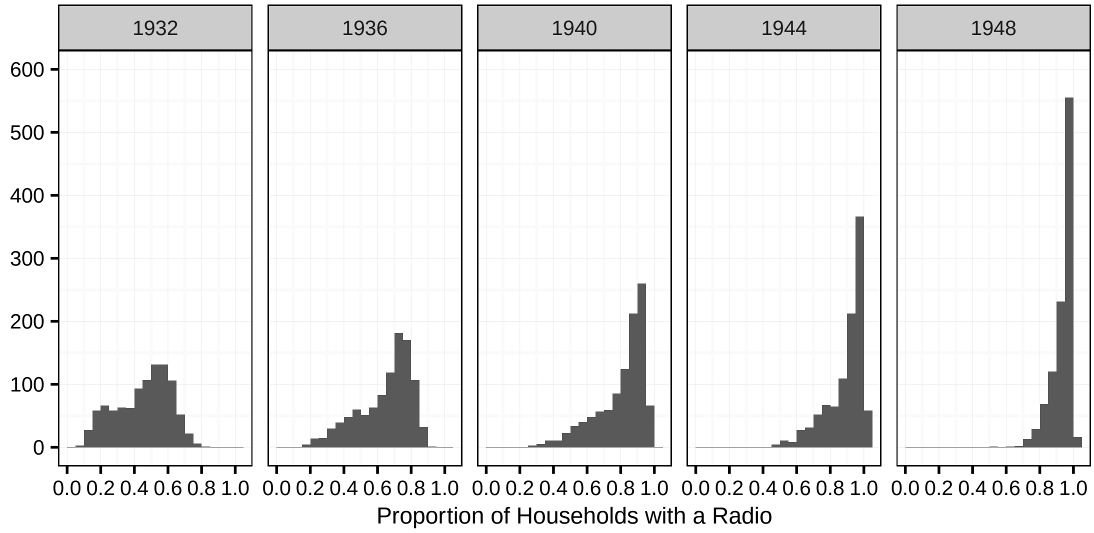

Figure 18: Histogram of Counties by Proportion of Households Owning a Radio, 1932 - 1948 (Binwidth = 0.05)

I estimate the effect of radio's popularization on demand thresholds in the newspaper industry by including the proportion of households with radios as a regressor in the static model and sunk cost models of market structure in local newspaper markets described by the likelihood functions stated in (9) and (16). To this end, I include the share of households owning radios in the per firm variable profit functions in the static and sunk cost models of entry. In the sunk cost model, I also allow radio diffusion to affect variable profits and sunk costs separately by assuming sunk costs γ^S have the form: $\gamma^S = \gamma_0^S + \gamma_1^S(RADIOHH_{mt})$, where $RADIOHH_{mt}$ gives the share of households owning radios in a given county m during presidential year t. Each model is estimated for the subset of county-years for which radio diffusion data are available (1932 - 1948). Because these regressions involve comparing the same counties at different points in time, I detrend the proportion of households owning radios by presidential year. The results of these regressions are reported in Table 11.

	Without sunk costs		With sunk costs		
Log population	1.074***	1.095***	0.945***	0.818***	0.820***
	(0.065)	(0.066)	(0.055)	(0.053)	(0.053)
Proportion of population white	0.623***	0.322	0.439**	1.445***	1.480***
	(0.215)	(0.214)	(0.186)	(0.179)	(0.176)
Proportion of white population foreign-born	-1.263**	-2.376***	-4.440***	-0.380	-0.422
	(0.620)	(0.677)	(0.524)	(0.576)	(0.580)
Proportion of adult population male	5.213***	5.225***	8.439***	8.015***	8.213***
	(1.092)	(1.087)	(0.886)	(0.858)	(0.883)
Farmland share of county surface area	0.214*	0.158	0.171*	0.397***	0.399***
	(0.123)	(0.124)	(0.100)	(0.101)	(0.101)
Proportion of population in urban areas	2.675***	2.655***	2.197***	2.190***	2.204***
	(0.167)	(0.168)	(0.140)	(0.135)	(0.136)
Manufacturing empl. share of adult male population	-0.879***	-0.903***	-0.499***	-0.332*	-0.330*
	(0.226)	(0.228)	(0.193)	(0.183)	(0.184)
Log population per square mile	-0.167***	-0.186***	-0.176***	-0.104***	-0.102***
	(0.045)	(0.045)	(0.036)	(0.034)	(0.034)
Proportion of households with a radio		0.659***		-2.345***	-2.318***
		(0.075)		(0.094)	(0.095)
γ (Continuation value)	12.790***	12.619***	11.340***	11.029***	11.152***
	(0.651)	(0.651)	(0.560)	(0.538)	(0.546)
θ^2 (Effect of first competitor)	1.829***	1.843***	1.462***	1.362***	1.362***
	(0.054)	(0.054)	(0.047)	(0.041)	(0.041)
θ^3 (Effect of second competitor)	1.301***	1.311***	1.051***	1.001***	1.003***
	(0.061)	(0.062)	(0.051)	(0.045)	(0.045)
${\gamma_1}^S$ (Effect of radio on sunk costs)					0.332
					(0.294)
${\gamma_0}^S$ (Sunk costs)			2.269***	2.724***	2.724***
			(0.049)	(0.058)	(0.058)
Observations	7105	7105	7082	7082	7082
Counties	1427	1427	1426	1426	1426
Log-likelihood	-6219.42	-6182.76	-3741.20	-3435.41	-3434.53

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Standard errors clustered by county in parentheses. Each specification is estimated using county-years between 1932 and 1948, the intercensal presidential years for which county-level data on household radio ownership are available.

Table 11: Estimation Results for Static and Sunk Cost Models, Radio Diffusion Included as Regressor, 1932 - 1948

The signs of all of the estimated parameters are robust to the inclusion of radio diffusion as an explanatory variable. Unexpectedly, radio diffusion is positively associated with the number of newspapers active in a county in the static version of the model. Moreover, its estimated coefficient is statistically different from zero at the 95 percent significance level. This finding is likely explained by radio diffusion's correlation with market-level variables that are positively related with the variable profits of local newspapers but cannot be included in the estimation due to data limitations (e.g., education and household income levels).

In the sunk cost specification, conversely, radio diffusion exhibits the expected negative association with the number of newspapers. This relationship is statistically different from zero at the 95 percent significance level. Radio diffusion appears to have little relationship with newspapers' sunk costs of entry: its effect on the sunk cost parameter is not statistically different from zero.

To measure the effect of radio diffusion on demand thresholds, I estimate entry and exit thresholds at varying percentiles of household radio ownership rates. The results of these calculations for the sunk cost specification are presented in Figure 19. Since radio diffusion does not appear to affect firms' sunk costs separately from variable profits, radio diffusion has the same effect on both entry and exit thresholds in terms of log population. In terms of population levels, radio diffusion can contribute to sizable differences in demand thresholds across counties, even within the middle 50 percentiles of household radio ownership rates.

Between 1932 and 1948, nearly 30,000 more people are needed to support the entrance of a monopolist at the 75^{th} percentile of radio diffusion relative to the 50^{th} percentile, while roughly 22,000 more people are required to support the entry of a monopolist at the 50^{th} percentile of radio diffusion relative to the 25^{th}. The monopoly exit threshold is roughly 1,000 persons higher and lower, respectively, at the 75^{th} and 25^{th} percentiles of radio diffusion relative to the 50^{th}. The absolute size of the gap increases for markets with more newspapers: the entry threshold for a third newspaper is 390,000 persons higher at the 75^{th} percentile relative to the 50^{th}, and 520,000 persons higher at the 50^{th} percentile relative to the 25^{th}. The effect of radio diffusion on exit thresholds for three-newspaper counties is less substantial in absolute terms: roughly 14,400 more people are required to support the continuation of three incumbent newspapers at the 75^{th} percentile of radio diffusion relative to the 50^{th}, and 19,000 more people are required at the 50^{th} percentile relative to the 25^{th}.

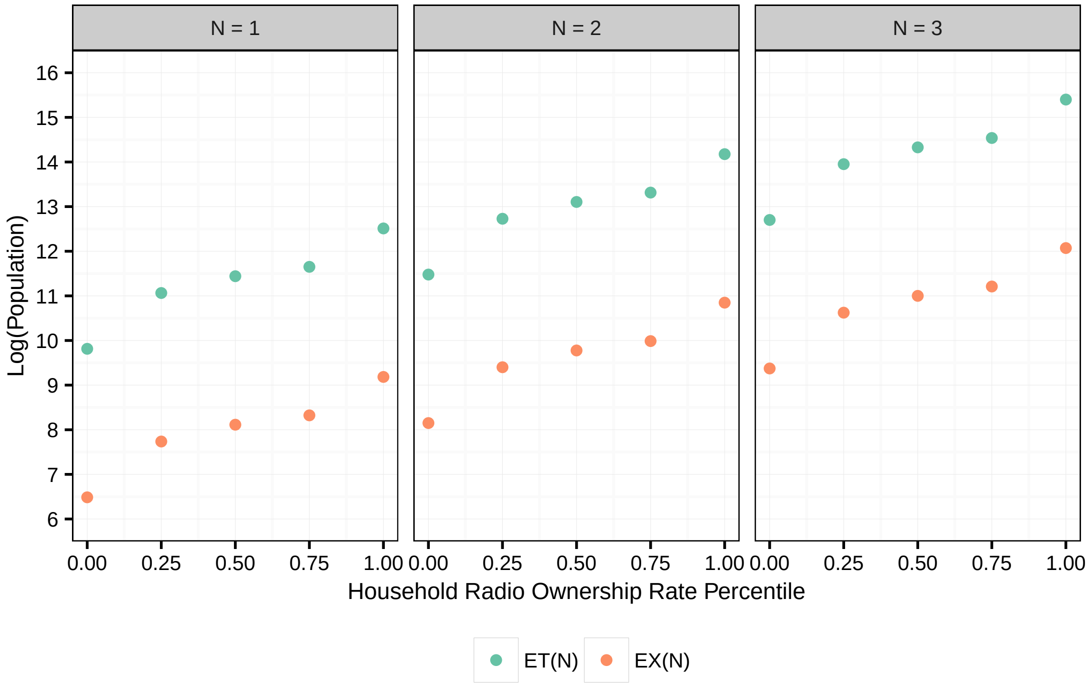

FIGURE 19: Entry and Exit Thresholds by Percentile Rank of Household Radio Ownership Rates, Estimates from Baseline Sunk Cost Specification

Based on these estimates, the popularization of radio during the 1930s and 1940s appears to have contributed to a large increase in the market size required to support the entrance of a newspaper, especially in counties in which incumbent newspapers were already active. Conversely, the effect on exit thresholds was relatively small, even in counties with three incumbent newspapers.

Television's Effect on Demand Thresholds

The popularization of television occurred even more quickly than that of radio (Figure 20). In 1952, the modal county had a household television ownership rate of 0 percent. Outside of the large group of counties without access to television in 1952, the remaining counties are distributed roughly evenly across household television ownership rates, although the distribution is somewhat right-skewed. By 1956, the vast majority of counties had access to television, and television ownership rates quickly came to reflect this fact: the modal county in this year had a household television ownership rate above 90 percent. Outside this modal group, a clear majority of counties having ownership rates above 50 percent. By 1960, most households in nearly all counties included in the sample owned televisions. During this year, the majority of counties had household television ownership rates above 80 percent, and almost no counties had ownership rates below 50 percent.

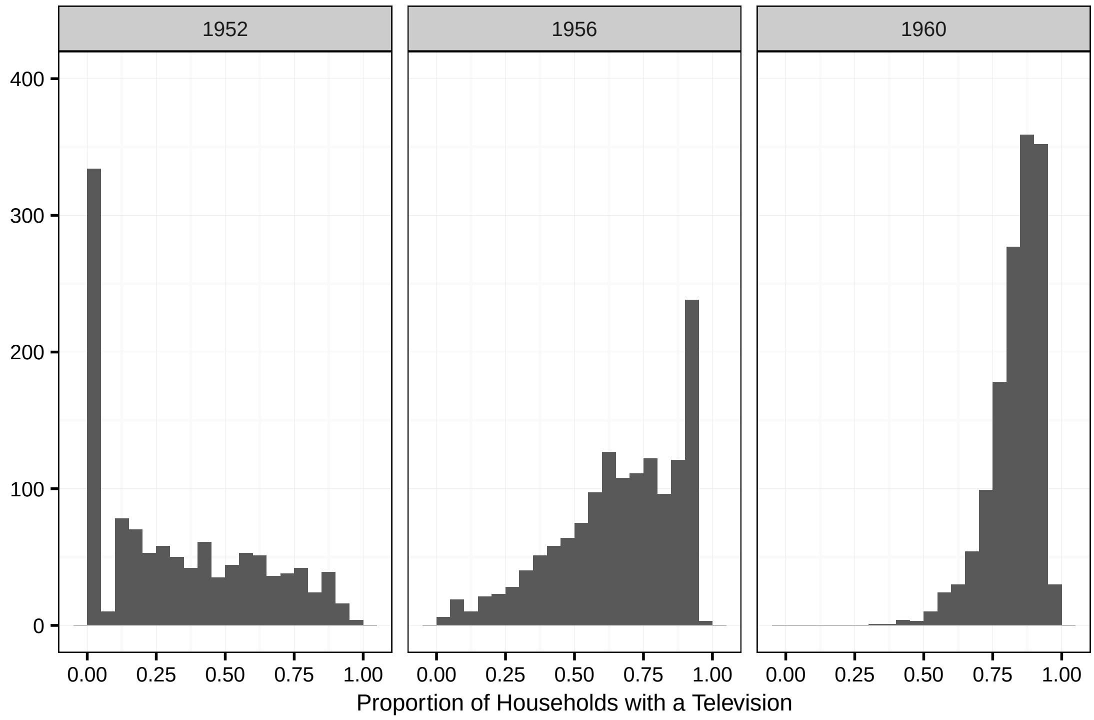

FIGURE 20: Histogram of Counties by Proportion of Households Owning a Television, 1952 - 1960 (Binwidth = 0.05)

As before, I estimate the effect of the popularization of television on demand thresholds by including the household television ownership rate, detrended by year, as a regressor in the variable profits function. I also interact television ownership rates with sunk costs using the same approach taken with radio in the previous subsection. I estimate each specification using county-years from 1952, 1956, and 1960. Table 12 reports the results of estimation.

All coefficient estimates from the static specification are robust to the inclusion of the household television ownership rates in the variable profits equation. In this specification, television diffusion is negatively associated with the number of newspapers active in a given county. This relationship is statistically significant at the 99 percent level.

	Without sunk costs		With sunk costs		
Log population	1.040***	1.062***	0.856***	0.938***	0.935***
	(0.068)	(0.069)	(0.064)	(0.065)	(0.065)
Proportion of population white	1.024***	1.115***	1.571***	1.804***	1.831***
	(0.314)	(0.315)	(0.273)	(0.281)	(0.282)
Proportion of white population foreign-born	-1.364***	-1.527***	0.842**	0.274	0.358
	(0.388)	(0.399)	(0.413)	(0.430)	(0.439)
Proportion of adult population male	7.804***	9.363***	-4.519***	1.023	1.007
	(1.482)	(1.637)	(1.382)	(1.460)	(1.462)
Farmland share of county surface area	0.425***	0.464***	0.329***	0.460***	0.462***
	(0.124)	(0.125)	(0.116)	(0.120)	(0.120)
Proportion of population in urban areas	1.768***	1.706***	1.608***	1.337***	1.345***
	(0.181)	(0.184)	(0.169)	(0.174)	(0.174)
Manufacturing empl. share of adult male population	0.560**	0.668***	0.154	0.568***	0.560***
	(0.223)	(0.228)	(0.205)	(0.213)	(0.214)
Log population per square mile	-0.187***	-0.160***	-0.113***	-0.003	0.005
	(0.045)	(0.045)	(0.042)	(0.044)	(0.045)
Proportion of households with a television		-0.406***		-1.305***	-1.287***
		(0.104)		(0.114)	(0.113)
γ (Continuation value)	13.619***	14.489***	8.277***	11.288***	11.325***
	(0.766)	(0.825)	(0.719)	(0.767)	(0.767)
θ^2 (First competitor)	2.130***	2.138***	1.900***	1.935***	1.935***
	(0.071)	(0.071)	(0.071)	(0.071)	(0.071)
θ^3 (Second competitor)	1.408***	1.413***	1.210***	1.231***	1.235***
	(0.079)	(0.079)	(0.075)	(0.075)	(0.075)
${\gamma_1}^S$ (Effect of TV on sunk costs)					0.922**
					(0.457)
${\gamma_0}^S$ (Sunk costs)			2.568***	2.697***	2.682***
			(0.098)	(0.103)	(0.101)
Observations	3978	3978	3968	3968	3968
Counties	1426	1426	1426	1426	1426
Log-likelihood	-3267.71	-3258.45	-2162.36	-2085.17	-2082.90

$^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.
Standard errors clustered by county in parentheses. Each specification is estimated using county-years in 1952, 1956, and 1960, the presidential years for which county-level data on household television ownership are available.

Table 12: Estimation Results for Static and Sunk Cost Models, Television Diffusion Included as Regressor, 1952 - 1960

In the sunk cost specification, the household television ownership rate enters negatively into the estimated variable profits function. It also enters positively into the size of the sunk costs of entry faced by entrants to local newspaper markets. The estimated coefficients for television diffusion are statistically different from zero at the 95 percent significance level in both the variable profit and sunk cost equations.

Most coefficient estimates from the sunk cost specification are robust to the inclusion of television diffusion, although there are some exceptions. After controlling for the proportion of households that own a television, the log of population per square mile no longer has a statistically significant relationship with the number of newspapers in a county. The coefficient estimate for adult male share of county population is negative, large in magnitude, and statistically significant prior to the inclusion of television diffusion in the model. After accounting for television diffusion, however, the estimated effect of this variable becomes positive and statistically insignificant. The correlation between the adult male share of the population and household television ownership rates appears to explain our earlier finding that the sign of the coefficient estimate for this variable differed in the 1944 to 1972 period from its sign during all other periods.

Figure 21 plots the estimated entry and exit thresholds implied by the sunk cost specification in which television diffusion affects entrants' sunk costs. The positive relationship between household television ownership rates and sunk costs has the effect of widening the gap between entry and exit thresholds as television becomes more popular in a county. In other words, the range of population levels over which newspapers will neither exit nor enter a county expands as television becomes more popular.

The gap between entry and exit thresholds at the extreme ends of household television ownership rates during the 1950s was large. In terms of log population, the difference between entry thresholds at the 100th and 0th percentiles of ownership rates was 3.0, while the difference between exit thresholds was 2.7. In levels, this translates into a gap in entry thresholds of roughly 300,000 persons for an entrant into a county with no newspapers, and a gap of 5.5 million persons across percentiles for the third entrant in a local newspaper market. For exit thresholds, a log population gap of 2.7 between the extreme ends of television diffusion rates implies a difference of roughly 9,000 persons required to support a local monopolist, and a gap of 160,000 in the number of persons required to sustain three incumbents. Notably, the popularization of television appears to have had a relatively larger effect on demand thresholds in the newspaper industry than radio did.

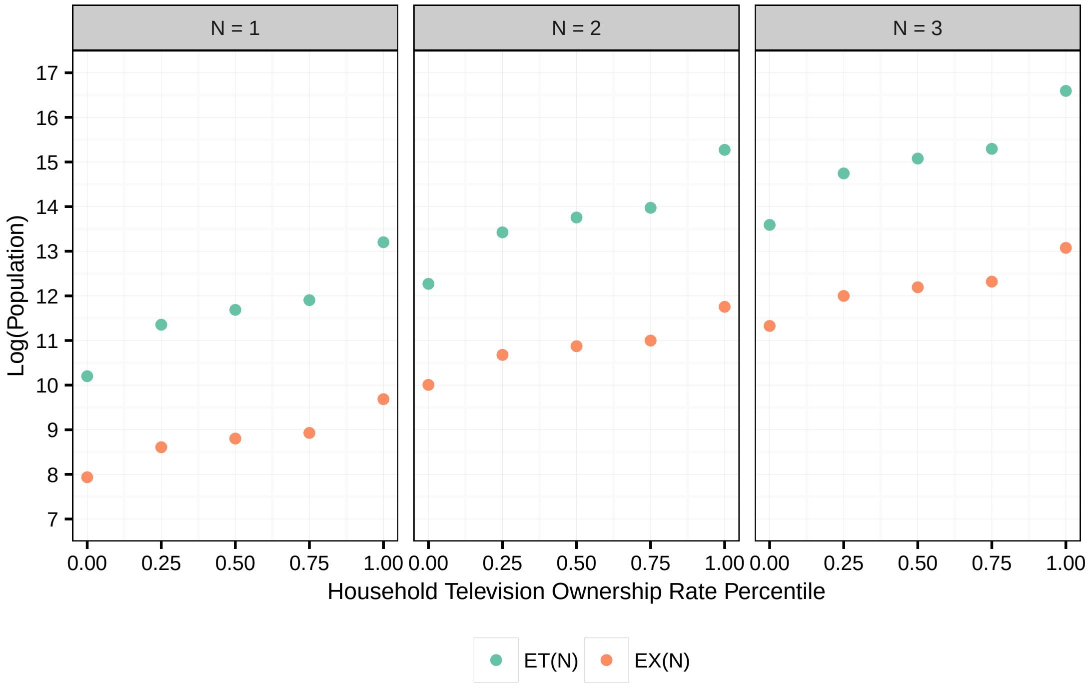

Figure 21: Entry and Exit Thresholds by Percentile Rank of Household Television Ownership Rates, Estimates from Sunk Cost Specification Allowing Television Diffusion to Affect Sunk Costs

The disproportionately large effect that the popularization of television had on entry thresholds relative to exit thresholds during the 1950s explains part of the stabilization in transition probabilities observed in the newspaper industry after 1944 (Figure 5). The introduction of television considerably increased the already large sunk costs of entry facing prospective entrants into local newspaper markets. The negative effect of television diffusion on newspapers' continuation values, while substantial, had a comparatively small effect on exit thresholds. As a result, instead of a mass exodus from the newspaper industry following the universalization of radio and television - two substitute sources of news and entertainment - we instead see a stagnation in the local newspaper industry, with few firms exiting and almost no firms entering.

7 Conclusion

This paper uses variants of models of endogenous market structure developed in Bresnahan and Reiss (1991) and Bresnahan and Reiss (1994) to analyze the importance of sunk costs of entry and entry's effect on competitive behavior in the U.S. local daily newspaper industry. Results from the static model show that the average market size required to support a local monopolist has remained nearly constant since the period between 1912 and 1940, while the average market size required to support two or more newspapers has increased substantially. Evidence from the model suggests that the magnitude of the effect of entry on per firm variable profits increases with the number of firms, at least over the range of market structures most commonly observed in local newspaper markets.

The pseudo-dynamic version of the model, which relaxes (to some degree) the assumption implicit in the static model that the profits earned by entrants and incumbents in a given period are identical, makes it possible to identify the size of barriers to entry in the newspaper industry. The estimated results suggest that the minimum market size required to support the entrance of a newspaper into a county has increased over time for any number of incumbents. The divergence of this result from the finding from the static model that monopoly entry thresholds have decreased over time is explained by the finding from the pseudo-dynamic model that the minimum market size required to support the continuation of an incumbent monopolist has fallen over time. Exit thresholds for counties with two or more incumbents, conversely, have increased over time, although they have increased at slower pace than entry thresholds. An extension of the model allowing sunk costs to change with the order of entry provides some evidence to suggest that sunk costs increase with the number of incumbents. The extended version of the model also suggests that the implicit assumption of the baseline sunk cost specification that the gap between entry and exit thresholds is constant in log demand is reasonably innocuous.

The range of market sizes over which no firms will enter or exit is quite large, and appears to have increased over time for all market structures. The widening of this range has been most apparent in counties with a monopolistic local newspaper. Between 1872 and 1908, the average county population required to induce the entry of a monopolist was around 43,000, and the continuation of a local monopolist required a population of 13,000. By contrast, between 1976 and 2000, a population of roughly 240,000 was needed to induce the entry of a monopolist, and fewer than 5,000 people were sufficient to justify the continuation of an incumbent monopolist, on average.

Examination of the impact of the emergence of radio and television on demand thresholds in the local daily newspaper industry shows that radio and television increased the market size needed to the entry of new firms and the continuation of incumbents. I find little evidence to support the hypothesis that the diffusion of radio affected newspapers' sunk costs separately from variable profits. However, I find a statistically significant positive relationship between the diffusion of television and the size of entering newspapers' sunk costs. This finding suggests that the popularization of television during the 1950s was an important contributor to the widening of the gap between entry and exit thresholds after 1944.

The evidence found in this paper suggests that for local daily newspapers, it is much more difficult to enter a market than it is to remain in one. Given the social importance of newspapers - studies have shown that newspapers increase voter turnout (Gentzkow, Shapiro, and Sinkinson (2011)) and are better sources of political information than other forms of news media (Gentzkow (2006)) - it is possible that high barriers to entry in the newspaper industry have resulted in an inefficiently low number of local newspapers from a social welfare perspective. In the past, policymakers in the United States have taken measures to protect local newspapers. The Newspaper Preservation Act of 1970, for example, allows newspapers to collude on circulation and advertising prices while preserving independent editorial boards in cases in which at least one local newspaper would otherwise have to exit the market (Patkus (1983)). The findings of this paper suggest that a better approach may be to promote entry into local newspaper markets, rather than try to preserve failing incumbent newspapers. According to my estimates of demand thresholds in the local newspaper industry, there are many one-newspaper counties that would be large enough to profitably sustain two newspapers but are too small to justify the entry of a second newspaper due to the presence of sunk costs of entry.

The simulated policy experiments in Gentzkow, Shapiro, and Sinkinson (2014) point to two possible approaches to inducing entry in the newspaper industry in a manner that is welfare-increasing. The first would be to allow newspapers to collude on advertising prices. This increases newspapers' advertising revenues, inducing them to reduce circulation prices in order to attract more consumers through the two-sided market mechanism described in Rochet and Tirole (2006). Gentzkow *et al.* find that allowing newspapers to collude along this dimension significantly increases entry in local newspaper markets.

This approach differs from that taken by the Newspaper Preservation, which allows newspapers to collude on both subscription and advertising prices, and only permits collusion in cases in which a newspaper would otherwise shut down. By contrast, the approach suggested above only allows collusion on advertising prices, but it extends this privilege to all newspapers irrespective of their profitability. An alternative to this approach would be to simply subsidize newspapers' marginal costs for each paper delivered. It is worth noting, however, that the empirical model estimated in Gentzkow, Shapiro, and Sinkinson (2014) uses data from the 1920s. Given the large increases in entry barriers in the newspaper industry since this time, it is not clear how effective these policies would be in the 21st century. In future research, it may be of interest to investigate whether the social welfare benefits of the presence of an additional newspaper outweigh the costs of subsidizing entry given the current conditions facing the newspaper industry.

The estimates of entry thresholds presented in this paper are calculated using models that impose some strong assumptions on the competitive behavior of local newspapers. In particular, both the static and pseudo-dynamic versions of the model assume products are homogeneous, a clearly counterfactual assumption in the case of newspapers. Relatedly, the assumption that the unobservable component of newspapers' profits is determined at the market rather than the firm level is not especially plausible, since market-level shocks are unlikely to have an identical effect on firms offering differentiated profits. Unfortunately, the range of product characteristics that could conceivably be observed for as broad a sample of local newspaper markets over as long a period as that used in this study is rather limited.[21] A possible extension to this paper could incorporate the measures of newspaper political affiliation developed in Gentzkow, Shapiro, and Sinkinson (2011) into a model of discrete product differentiation, similar to the one developed by Mazzeo (2002). With that said, even estimating a model in the spirit of Mazzeo (2002) can easily become computationally expensive: the three product model estimated in that paper requires the use of simulation methods, and that is without the added complication of differentiating between entrants and incumbents.

[21] A similar caveat applies to other measures that could enable the estimation of a more realistic version of the model, such as direct information about newspapers' marginal costs or advertising revenues.

Bibliography

Abbring, Jaap H. and Jeffrey R. Campbell. 2010. "Last-in First-Out Oligopoly Dynamics." *Econometrica* 78 (5):1491–1527.

Asplund, Marcus and Rickard Sandin. 1999. "The Number of Firms and Production Capacity in Relation to Market Size." *Journal of Industrial Economics* 47 (1):69 – 85.

Berry, Steven T. 1992. "Estimation of a Model of Entry in the Airline Industry." *Econometrica* 60 (4):889–917.

Bresnahan, Timothy F. and Peter C. Reiss. 1988. "Do Entry Conditions Vary Across Markets?" *Brookings Papers on Economic Activity* (3):833–871.

———. 1990. "Entry into Monopoly Markets." *Review of Economic Studies* 57 (4):531–553.

———. 1991. "Entry and Competition in Concentrated Markets." *Journal of Political Economy* 99 (5):977–1009.

———. 1994. "Measuring the Importance of Sunk Costs." *Annales D'Economie et de Statistique* (34):181–217.

Chandra, Ambarish and Allan Collard-Wexler. 2009. "Mergers in Two-Sided Markets: An Application to the Canadian Newspaper Industry." *Journal of Economics & Management Strategy* 18 (4):1045–1070.

Collard-Wexler, Allan. 2014. "Mergers and Sunk Costs: An Application to the Ready-Mix Concrete Industry." *American Economic Journal: Microeconomics* 6 (4):407–447.

Dunne, Timothy, Mark J. Roberts, and Larry Samuelson. 1989. "The Growth and Failure of U.S. Manufacturing Plants." *The Quarterly Journal of Economics* 104 (4):671–698.

Fan, Ying. 2013. "Ownership Consolidation and Product Characteristics: A Study of the US Daily Newspaper Market." *American Economic Review* 103 (5):1598–1628.

Gentzkow, Matthew. 2006. "Television and Voter Turnout." *Quarterly Journal of Economics* 121 (3):931–972.

Gentzkow, Matthew and Jesse Shapiro. 2010. "What Drives Media Slant? Evidence from U.S. Daily Newspapers." *Econometrica* 78 (1).

Gentzkow, Matthew, Jesse Shapiro, and Michael Sinkinson. 2011. "The Effect of Newspaper Entry and Exit on Electoral Politics." *American Economic Review* 101 (7).

Gentzkow, Matthew, Jesse M. Shapiro, and Michael Sinkinson. 2014. "Competition and Ideological Diversity: Historical Evidence from US Newspapers." *American Economic Review* 104 (10):3073 – 3114.

George, Lisa. 2007. "Effect of Ownership Concentration on Product Variety in Daily Newspaper Markets." *Information Economics and Policy* 19 (3):285–303.

George, Lisa and Joel Waldfogel. 2006. "The New York Times and the Market for Local Newspapers." *American Economic Review* 96 (1):435–447.

Geroski, P. A. 1995. "What Do We Know About Entry?" *International Journal of Industrial Organization* 13 (4):421–440.

Manuszak, Mark D. 2002. "Endogenous Market Structure and Competition in the 19th Century American Brewing Industry." *International Journal of Industrial Organization* 20 (5):673 – 692.

Mazzeo, Michael J. 2002. "Product Choice and Oligopoly Market Structure." *RAND Journal of Economics* 33 (2):1–22.

Oberholzer-Gee, Felix and Joel Waldfogel. 2009. "Media Markets and Localism: Does Local News en Espanol Boost Hispanic Voter Turnout?" *American Economic Review* 99 (5):2120 – 2128.

Patkus, John P. 1983. "The Newspaper Preservation Act: Why It Fails to Preserve Newspapers." *Akron Law Review* 17 (3):435 – 452.

Rochet, Jean-Charles and Jean Tirole. 2006. "Two-Sided Markets: A Progress Report." *RAND Journal of Economics* 37 (3):645–667.

Schaumans, Catherine and Frank Verboven. 2015. "Entry and Competition in Differentiated Products Markets." *The Review of Economics and Statistics* 97 (1):195–209.

Seamans, Robert and Feng Zhu. 2013. "Responses to Entry in Multi-Sided Markets: The Impact of Craiglist on Local Newspapers." *NET Institute Working Paper No. 10-11* .

Seim, Katja. 2006. "An Empirical Model of Firm Entry with Endogenous Product-Type Choices." *The RAND Journal of Economics* 37 (3):619–640.

Stromberg, David. 2004. "Radio's Impact on Public Spending." *The Quarterly Journal of Economics* 119 (1):189–221.

Xiao, Mo and Peter F. Orazem. 2011. "Does the Fourth Entrant Make any Difference? Entry and Competition in the Early U.S. Broadband Market." *International Journal of Industrial Organization* 29 (5):547 – 561.

A Appendix

A.1 Additional Tables

Period	ET(1)	ET(2)	ET(3)	ETR(2)	ETR(3)
1872 - 1908	25252	48795	116159	0.97	1.59
1912 - 1940	14434	53018	163586	1.84	2.06
1944 - 1972	16447	100451	329954	3.05	2.19
1976 - 2000	15269	239531	836923	7.84	2.33

Entry thresholds for a given number of firms ($ET(N)$) are given in number of persons. Entry threshold ratios ($ETR(N)$) are unitless.

Table A.1.1: Estimated Entry Thresholds, Static Specification

Period	ET(1)	ET(2)	ET(3)	EX(1)	EX(2)	EX(3)	ETR(2)	ETR(3)
1872 - 1908	42668	80937	189249	13212	25062	58601	0.95	1.56
1912 - 1940	67051	234327	761651	4913	17170	55809	1.75	2.17
1944 - 1972	132928	869576	2907522	4384	28681	95898	3.27	2.23
1976 - 2000	239307	2984768	10073203	4647	57960	195608	6.24	2.25

Entry and exit thresholds for a given number of firms ($ET(N)$ and $EX(N)$, respectively) are given in number of persons. Entry threshold ratios ($ETR(N)$) are unitless.

Table A.1.2: Estimated Entry and Exit Thresholds, Baseline Sunk Cost Specification

Period	ET(1)	ET(2)	ET(3)	EX(1)	EX(2)	EX(3)	ETR(2)	ETR(3)
1872 - 1908	42382	72886	186823	13359	24432	59094	0.86	1.71
1912 - 1940	72867	311776	1331953	4715	17277	59715	2.14	2.85
1944 - 1972	159415	1508554	9745315	4108	28978	110058	4.73	4.31
1976 - 2000	316276	8031201	44697090	4435	59699	210683	12.70	3.71

Entry and exit thresholds for a given number of firms ($ET(N)$ and $EX(N)$, respectively) are given in number of persons. Entry threshold ratios ($ETR(N)$) are unitless.

Table A.1.3: Estimated Entry and Exit Thresholds, Flexible Sunk Cost Specification