Statewide Cable Franchising and Broadband Connections

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INTRODUCTION

Licenses for cable television operators have increasingly been granted at a state-wide level. According to the most recent data collected by Holmes and Mowrey (2009) as of October 2009 25 states have adopted such provisions. State-wide franchising, it is argued, lowers to cost of entry into the cable television market. Without a statewide franchising law a potential statewide cable TV competitor is required to negotiate a license for operation in each and every locality in the state. A statewide franchise allows such a firm to operate throughout the state subject a uniform set of rules and with a single application facilitating their entry into the cable TV market.

Advocates for statewide franchising have generally been large telecommunications firms wishing to offer their cable TV services at a statewide level. Opponents have included local cable incumbents. Advocates of statewide franchising have argued that its adoption would increase telecommunications investment and lead to more competitive cable television services. Opponents have denied such claims.

To the best of our knowledge there is no evidence in the refereed academic literature of the impact of statewide cable franchise laws on either the quantity of investment in telecommunication infrastructure or on cable television rates. This is not surprising as the both cable television rates and telecommunications infrastructure investment is proprietary information.

However, since 1999 the FCC has kept data on the number of broadband connections by state. This FCC data offers an avenue to assess the impact of statewide franchising on an important telecommunications metric: broadband connections. Telecommunications providers have increasingly offered bundled services, blurring the line between a cable provider, a phone provider and an internet provider. A statewide cable franchise encourages a traditional land-line telephone provider not only to enter the cable TV market but also the market for broadband service. Although broadband service could be offered in a local market by a land line telephone provider in the absence of a statewide franchise, a statewide franchise “sweetens’ the potential returns to the capital investments necessary to facilitate the provision of both cable and broadband services.

There is, therefore, reason to suspect that entry into a cable TV market will be accompanied by entry into the broadband market. Increased competition in broadband should be consistent with higher take rates for broadband, holding all other factors constant. The empirical issue we pose is straightforward: do states adopt statewide cable franchising have higher growth rates in household and firm broadband connections than states that have not adopted such provisions—controlling for all other relevant factors? Not only does this offer to provide indirect evidence as to the initial claims of statewide franchise advocates—that such laws increase telecommunications investment—but also offers to potentially quantify another benefit of a statewide cable franchise law—increased internet access.

A Cournot Model with Sub-additive Costs

A stylized version of broadband access in the United States would appear thusly: Local Cable Access Television (CATV) and Incumbent Local Exchange Carriers (ILEC) provide two separate services under state level regional regulation. Under the traditional regulatory arrangement, CATV providers are unable to offer wireline telephony, while ILEC's are unable to offer cable television services. However, the regulatory schema does not preclude either provider from bundling their regulated service with broadband. Indeed, both do. However, we posit a change to the output decision on offering of broadband for each provider under statewide franchising. To develop this scenario we offer a model of output decisions using the traditional Cournot framework, but include an assumption of subadditivity in costs of production (Baumol, Panzar and Willig, 1982). We begin with the Cournot profit function for firm i:

$$\Pi_i(q_i, q_i) = q_i P(q_i + q_i) - c_i(q_i)$$

Which is the well known single product Cournot profit function, which we extend to a two product (A,B) model.

$$\Pi_i\ (q_i^A, q_i^B, q_j^A, q_j^B) = q_i^A P^A(q_i^A + q_j^A) + q_i^B P^B(q_i^B + q_j^B) - c_i(q_i^A) - c_i(q_i^B)$$

Whose first order conditions are:

$$\partial\Pi_i\ (q_i^A, q_i^B, q_j^A, q_j^B) \Big/ \partial q_i^A = P^A(q_i^A + q_j^A) - c_i(q_i^A) = 0$$

And

$$\partial\Pi_i\ (q_i^A, q_i^B, q_j^A, q_j^B) \Big/ \partial q_i^B = P^B(q_i^B + q_j^B) - c_i(q_i^B) = 0$$

To which we introduce sub-additivity of costs (Baumol, Panzar and Willig, 1982). Sub additive costs represent a case of joint production that was in BPW introduced as a departure point for understanding the least cost provider case. Here we adapt this observation to understand the bundling of telecommunications services so that:

$$c_i\left(q_i^A\right) + c_i(q_i^B) > \hat{c}_i\left(q_i^A, q_i^B\right)$$

Which is applicable to *j* firms, resulting in a two product output decision for a quantity setting oligopolist:

$$\Pi_i \ \left(q_i^A, q_i^B, q_j^A, q_j^B\right) = \ q_i^A P^A\left(q_i^A + q_j^A\right) + q_i^B P^B\left(q_i^B + q_j^B\right) - \hat{c}_i\left(q_i^A, q_i^B\right)$$

From which we can solve the reaction function *R* for each firm.

$$q_i = R_i(q_i) = {1 - q_i - \frac{\alpha}{n}} \Big/_2$$

Where $\frac{\alpha}{n}$ is the market share of firm *i*, and *n* is the elasticity of demand. Solving for equilibrium conditions for both the separate and subadditive cost case we have:

$$q_i = {1 - 2c_i + c_j}/_3$$ and

$$\widehat{q_i} = {1 - 2\hat{c}_i + \hat{c}_i} \ /_3$$

where the observation that the subadditive costs are less than the separated costs for each firms leads to the result $\widehat{q_i} > \ q_i$.

The result of which offers the following prediction, under conditions of subadditive costs, output under a two form quantity setting oligopolist will be higher than under conditions of separate production costs. Other results are also derived from this observation. The equilibrium conditions derived from R, under the case of subadditive costs also yields a higher elasticity of demand than under the separate production schema. That is necessary to equilibrate the reaction function under the differing cost functions.

Within the context of this derivation, some comments on the institutional arrangement of statewide franchising is in order. First, we treat firms as enjoying homogeneous costs, a condition which is invisible to the econometrician in this case. Further, we will employ empirics that permit multi-firm entrance, extending the two firm model to an *n* firm reality. We think these two concerns have modest quantity effects on the output decisions which we leave as an empirical matter in this case.

From this rather simple model we are able to conclude theoretically that the relaxation of regulations preventing entrance into cable access TV markets through statewide franchising brings with it two testable predictions. The first is that the output decision is effected with unambiguous increases in quantity resulting from the lower sub-additive costs. Second, the elasticity of demand at the new equilibrium would be higher than under the separate cost conditions. As a consequence we can also predict (but not test) that price for telecommunications services would decline.

Table 1

	Statewide Franchising	**Legislation**	**Last Action**
Alabama	No	N/A	N/A
Alaska	Yes	N/A	N/A
Arizona	No	HB 2812	Enacted March 2006
Arkansas	No	N/A	N/A
California	Yes	AB 2987	Enacted September 2006
Colorado	No	HB 1222	Dead as of June 2007
Connecticut	Yes	HB 7182	Enacted July 2007
Delaware	No	N/A	N/A
Florida	Yes	HB 529	Enacted May 2007
Georgia	Yes	HB 227	Enacted July 2007
Hawaii	Yes	N/A	N/A
Idaho	Pending	S1100/In House	Passed Senate February 2009
Illinois	Yes	SB 0678	Enacted June 2007
Indiana	Yes	HR 1279	Enacted March 2006
Iowa	Yes	SF 554	Enacted March 2007
Kansas	Yes	SB 449	Enacted April 2006
Kentucky	No	N/A	N/A
Louisiana	Yes	SB 807	Enacted June 2008
Maine	Yes	HB 1515	Enacted April 2008
Maryland	Pending	HB 1182/ In Senate	As of February 2009
Massachusetts	Pending	S2649	As of January 2009
Michigan	Yes	HB 6456	Enacted December 2006
Minnesota	No	SB 3337	Enacted May 2008
Mississippi	No	N/A	N/A

Missouri	Yes	SB 284	Enacted March 2007
Montana	No	N/A	N/A
Nebraska	No	N/A	N/A
Nevada	Yes	AB 526	Enacted June 2007
New Hampshire	No	N/A	N/A
New Jersey	Yes	ACS 804	Enacted August 2006
New Mexico	No	HB 675/SB 522	Legislation exhausted as of April 2009
New York	Pending	AB 4469	As of February 2009
North Carolina	Yes	H 2047	Enacted July 2006
North Dakota	No	N/A	N/A
Ohio	Yes	SB 117	Enacted July 2007
Oklahoma	No	N/A	N/A
Oregon	No	N/A	N/A
Pennsylvania	No	HB 1490	As of May 2009
Rhode Island	Yes	N/A	N/A
South Carolina	Yes	HB 4428/HB 3396	Enacted May 2006 & March 2007 resp.
South Dakota++	No	HB 1160	Enacted March 2005
Tennessee	Yes	HB 1421/SB 1933	Enacted May 2008
Texas	Yes	SB 5	Enacted August 2005
Utah	No	SB 209	Exhausted as of February 2007
Vermont	Yes	N/A	N/A
Virginia	Yes	HB 568/HB1404	March & July 2006
Washington	Pending	SB 5421	Senate Hearing February 2009
West Virginia	No	HB 3161	Legislation Exhausted as of 2003
Wisconsin	Yes	AB 207/SB 107	Enacted April 2007
Wyoming	No	N/A	N/A

Data and Empirical Modeling

Our empirical strategy is straightforward. We seek to test the relationship between statewide franchising legislation – the relaxing of geographic market constraints on the degree of competition within cable networks. To do so we must construct models which account for the presence or introduction of statewide franchise legislation as well as indications of competition in broadband and cable services in each U.S. state.

To begin this process we obtained semi-annual, state level data on subscribers from the Federal Communications Commission, Form 477 reports. This data provides administrative subscriber accounts as of June and December each year, beginning in June, 1999. The data lag is roughly 16 months, so as of this writing we possess June, 2008 data.

We also collected data on the presence of statewide franchising through a census of states. See Table 1. From this data we crafted a panel of dummy variables which accounted for the presence of statewide franchising, by state, in the semi-annual period which corresponds to the FCC data. In order to be conservative in our estimate, we imposed no restrictions on the duration of implementation. For example Illinois' Senate Bill 0678 was implemented in June of 2007, which dictated our coding Illinois as possessing statewide cable franchising during the period January-June 2007. For a practical matter, this would not likely be sufficient time to observe a competitive response to this change in regulation. However, we have adopted this convention because insofar as it imposes any bias in the treatment of de-regulation it would tend to reduce the magnitude of the impacts. This is a conservative assumption.

Further, we collected data on subscribers, by type from the FCC Form 477 reports. However, these data cover a far shorter duration, with annual observations of no more than four years. While this is a richer data set with respect to the share of subscribers by provider type, the time frame is not really sufficient for dynamic analysis. This data contains nine different types of broadband providers, albeit with considerable data suppression in smaller states. To this data we were able to add a variable for total years of statewide cable franchise availability, and demographic data on population, population density, per capita personal income and the share of population less than 65 years of age.

As a consequence, we have two data sets. The first is a semi-annual panel from 1999:S2 through 2008:S1 comprising broadband subscribers (in aggregate) and the presence of statewide cable franchising legislation. The second is a cross sectional model with detailed information on demographic, geographic, economic and regulatory information on broadband subscribers by state. Thus we have two potential families of competitive models to test.

Statewide Franchising and Subscriber Dynamics

A fundamental consideration in the context of statewide cable franchising was the extension of broadband subscribers as a consequence of the price effect of statewide competition. Historical data on prices for internet services are unavailable. As a consequence we must rely upon other data to estimate this effect.[1] Estimating this on statewide data provides us the following relationship:

$$Subscribers = f(X, Cable\ Franchise, Trends)$$

[1] We must also rely upon a relatively straightforward derivation of the demand curve from the expenditure function where the underlying indirect utility function contains internet or cable services.

where a measure of broadband subscribers are a function of regional specific conditions (X), the presence of a statewide cable franchising and trend dynamics. To test this we create a basic mode:

$$d\left(\log S_t^i\right)\Big/_{dt} = \alpha + \alpha_i + \beta\left(F_t^i\right) + \varphi R_t + \gamma_i T + \delta\theta_{t-1} + e_i$$

Where the change in the log of Subscribers, in each state during each semi-annual time period t, (Jan-Jun and July to Dec) is a function of a common and cross sectional fixed effects α, the presence dummy of the statewide cable franchising deregulation (F) in state i, in time period t. We also report a recession dummy for the last two business cycles. To that we introduce a cross section specific time trend and first order autocorrelation variable. The error term is assumed iid → $N(0,\sigma^2)$.

An alternative specification includes a first order spatial autocorrelation term$\rho(\widehat{W}S_t^j)$, where the coefficient ρ is estimated for a first order spatial weight matrix W, with row normalized weights, estimated for the dependent variable $d\left(\log S_t^i\right)/dt$ in adjacent state j, in time t. All tests are on the 48 conterminous states.

A brief explanation of these terms is helpful. We are interested in detecting a year to year variation in the number of subscribers in each state as a consequence of statewide cable franchising changes and other factors which may influence broadband subscriber growth. By estimating the dependent variable as a percent change, we abstract from state level population differences in the estimate. The use of state specific fixed effects over such a brief temporal period allows us to control for all non-stochastic (e) variance which is common to each state for the duration of the sample period. Thus, this variable accounts for such things as relative population density, regional age differences, other demographic characteristics and incomes. The recession dummy accounts for business cycle specific changes to broadband adoption. By permitting the time trend to vary by state we are attempting to isolate the differential growth in take rates by states which absorbed different technologies at different times. The common AR specification accounts for national dynamics. The spatial autocorrelation factor corrects for a problems associated with similarity across regions which are not otherwise accounted for within the model specification.

Both specifications were estimated in a panel setting, with error terms which have been treated with White's [1980] heteroscedasticity invariant, variance, covariance matrix. The use of a panel with spatial autocorrelation terms argues for an EGLS estimator to correct for potential endogeneity of the contemporaneous spatial term.[2]

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[2] We employed the EGLS estimate with a Cochrane-Orcutt iterative algorithm (employing the OLS coefficients as initial coefficient values to avoid convergence on a local maximum).

We estimation results speak primarily to the effect of statewide cable franchising deregulation. While the effect of recessions, broad regional influences and state trends are also of interest, these variables are primarily designed to control for other influences, hence isolating the effect of statewide franchising changes. The results are displayed in Table 2.

Table 2, Results of Panel Estimate
(change in the Log of Subscribers as dependent variable)

Coefficient	Description	Specification 1	Specification 2
		0.3546***	0.070832***
α	Common Intercept	(21.04)	(5.38)
		0.0623***	0.019623**
β	Statewide Franchise	(3.15)	(1.99)
		-0.00024	-0.0011
φ	Recession Dummy	(-0.02)	(-0.18)
		…	0.887695***
ρ	Spatial Autocorrelation		(25.38)
		0.1522***	-0.12998***
δ	Temporal Autocorrelation	(2.86)	(-3.04)
	Adjusted R-squared	0.315879	0.685926
	S.E. of regression	0.126407	0.099819
	F-statistic	4.547788***	16.68481***
	Durbin-Watson stat	2.017825	1.951055
	Observations	N=754	N=712

Note: *** denotes statistical significance of .01 percent, ** of statistical significance of .05 percent. T-statistics in parentheses are asymptotic. Fixed effects and cross sectional specific trends are not reported.

We have not reported statewide trends or fixed effects. These are of tangential interest. However, the regression results are of significant interest. Treating them in order of importance, we find that there are strong spatial and temporal autocorrelation apparent in both models. The magnitude of the effect, were it statistically valid is likewise quite small. We think that is because the trend of increasing take rates in broadband persisted across the business cycle. Simply, the most two recent recessions have not imposed statistically measurable reductions in broadband deployment rates.

The important result of this model for this research is the role statewide cable franchise de-regulation ahs changes the number of broadband subscribers in the state, all things held constant. For that we turn to the coefficient results in both models. In both instances, the results have strong statistical significance, a result which is robust to alternative model specifications. In unreported estimates we estimated a common time trend, which yielded coefficient estimates similar to those reported in both significance and magnitude. The two results vary by roughly an order of magnitude in their point estimate. Despite the fact that statewide franchising does not extend across borders, we prefer the model which accounts for spatial autocorrelation.

These results suggest that for each observed period (six months) of statewide franchising, a state will experience a roughly four percent increase in subscribers. Within our sample of 48 conterminous states, twenty-two have adopted statewide franchising (with four pending in 2010). The mean duration of statewide franchising is just under two years and four months. It is possible then to provide a point estimate of additional broadband connections for each state with a deregulated cable franchising. See table 3.

Table 3, Additional Broadband Connections attributable to Statewide Cable Franchising

State	Total attributable to SW franchising	% of total new subscribers attributable to SW franchising
California	1,489,551	2.41%
Connecticut	110,085	2.04%
Florida	444,977	2.03%
Georgia	149,513	1.93%
Illinois	305,114	2.05%
Indiana	226,719	2.47%
Iowa	59,469	2.04%
Kansas	98,983	2.33%
Louisiana	25,730	1.66%
Maine	7,925	1.85%
Michigan	284,587	2.23%
Missouri	111,962	2.03%
Nevada	69,556	1.99%
New Jersey	393,890	2.21%
North Carolina	278,784	2.22%
Ohio	184,494	1.91%
Rhode Island	176,634	5.32%
South Carolina	158,608	2.49%
Tennessee	50,385	1.82%
Vermont	86,493	5.88%
Virginia	327,981	2.42%
Wisconsin	105,987	2.04%
Total	5,147,425	

These findings are prime evidence of increased competition in broadband services which resulted from enactment of statewide cable franchise legislation in a few states. Another important facet of the debate is the change in competition resulting from changes to statewide franchising of cable services.