Does Municipal Supply of Communications Crowd-Out Private Communications Investment? An Empirical Study

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There are 2,007 municipalities across the United States that provide electricity service to their constituents. Of these, over 600 provide some sort of communications services to the community. An important policy question is whether or not public investment in communications crowds out private investment, or whether such investment encourages additional entry by creating wholesale markets and economic growth. We test these two hypotheses – the *crowding out* and *stimulation* hypothesis – using a recent dataset for the state of Florida. We find strong evidence favoring the stimulation hypothesis, since public investment in communications network increases competitive communications firm entry by a sizeable amount.

I. Introduction

There are 2,007 municipalities across the United States that provide electricity service to their constituents.[1] In 2004, 616 of these provided some sort of communications services to the community, and this number has grown by 37% since 2001.[2] As evidence mounts of the substantial positive effects on economic development from municipal construction of communications networks, more municipalities are expected to deploy such networks in the near future.[3] Even large cities that do not provide electricity services are entering the communications business; St. Louis, Atlanta, Philadelphia, Los Angeles, Seattle, and

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1 See *Electric Power Statistics* page at www.appanet.org. These systems serve about 14% of total households in the United States. Also see APPA Annual Directory & Statistical Report (Multiple years) and Gillett et al (2004).

2 Id., years 2003, 2004, and 2005.

3 See, e.g., Lisa Eckelbecker, High-speed services critical for growth; WPI panelists discuss future of broadband, Telegram & Gazette (Jul 27, 2004); McGregor McCance, Broadband Service Seen as Economic Stimulus for Rural Virginia Communities, Knight Ridder Tribune Business News. (Mar 4, 2003); Brad Carlson, Citi Cards to boost Boise staff, build new center in Silverstone, The Idaho Business Review (Mar 15, 2004); John Snow, Broadband scarcity hurts economic development, Business Journal (Jacksonville) (May 16, 2003), Vol.19, Iss. 31; and various archived news releases at http://www.muniwireless.com/.

many more cities are all in the process of deploying city-wide wireless broadband networks.[4] This phenomenon is not restricted to the United States – municipal wireless broadband networks are operating in Hamburg, Brussels, and the British Virgin Islands.[5]

Recently, municipal entry into the communications markets has come under attack by private communications firms – primarily incumbent local phone and cable companies. These politically powerful firms have successfully lobbied for legislation banning or severely restricting municipal communications networks in Arkansas, Missouri, Minnesota, Nebraska, Nevada, Pennsylvania, South Carolina, Texas, and Utah.[6] Anti-muni laws are being debated by legislators in Indiana, Nebraska, Ohio, and Florida.

One the principle arguments against municipal provision of communications services is that this public investment will "crowd out" private investment: if we view that a market is capable of sustaining *N* firms, then the entry of a municipality will displace (at least) one private firm.[7] While simple, the argument is arguably too simple when applied to the communications industry. Entry into the communications industry typically requires large sunk investments in fixed assets that render non-trivial scale economies. In many cases, therefore, the municipality will be the only entrant for some communications services or in particular geographic areas, since the expected return may not be sufficient to warrant the investment by a private firm.[8] Or, the municipality may be the only competitor to a monopoly private firm in cases where additional entry may be precluded absent the positive spillovers available to the municipality. So, in many cases, municipal entry may have no effect on private entry, but it may be an important element of a well-functioning communications market.

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4 Esme Vos, FIRST ANNIVERSARY REPORT, Muniwireless.com (June 2004) and articles on www.muniwireless.com.

5 *Id*.

6 PUBLIC POWER: PROVIDING THE 21ST CENTURY THROUGH COMMUNITY BROADBAND SERVICES, American Public Power Association (December 2004): www.appanet.org.

7 Another criticism of municipal supply of services in that governments are inefficient. However, the vast majority of published empirical studies on the public provision of utility services suggests that, if there is any difference, public provision is slightly more efficient (e.g., Hausman and Neufeld 1991; Foreman-Peck and Waterson 1985; Byrnes, et al 1986; Bruggink 1982; Ohlsson 2003; Renzetti and Dupont 2003; Estache and Rossim 2002).

8 In most cases (particularly in Florida), the municipal provision of communications services arose from a refusal of incumbent phone companies to provide high-capacity telephone and Internet services desired by the community even after a direct request for such services was made. See, e.g., *Scottsburg, Indiana Wireless Network Saves the Community*, MUNIWIRELESS (April 29, 2004): ("Scottsburg, Indiana is a community of 6000 people, 29 miles (47 km) north of Louisville Kentucky. Scottsburg does not have wired broadband and the costs of deploying one are prohibitive. Just to give you an example, it costs $1300 per month to lease a T1 line in Scottsburg; in Louisville, it costs only $300 per month. The town approached Verizon about bringing broadband to their community, but the latter told them that there were not enough residents to make it worth Verizon's trouble"). Also see Phil Davies, Broadband.gov: A growing number of small cities in the district offer their residents high-speed Internet access. Does local government belong in the telecom business? Fedgazette (November 2004).

Further, in many cases, the investments made by the public sector may increase private investment since municipally run communications networks typically provide wholesale access to key components of telecommunications infrastructure.[9] Like the unbundling obligations of the 1996 Telecommunications Act, this wholesale access to fixed and sunk assets promotes entry. So, there is a plausible argument that municipal entry may actually encourage private firm entry and investment. In Leesburg, Florida, for example, the municipality leases strands of its dark fiber network to alternative providers, so that those private firms can utilize Leesburg's network to deploy their own services cost-effectively. Similarly, the municipal communications system serving Gainesville, Florida, provides high-capacity circuits to independent wireless providers, making those networks more cost-effective, and leases itself long-haul capacity from competitive carriers. In New Smyrna Beach, Florida, the city purposely avoided capital expenses by contracting with competitive local exchange carriers to provide the necessary facilities for service provision.

Even the incumbents recognize, to some extend, that municipal participation may increase private firm investment. A spokesperson for Verizon Communications (the largest incumbent local exchange carrier) lauded Erie County's wireless broadband infrastructure, stating that the county's network "makes people more aware of the benefits of broadband" and will eventually help Verizon sell its own broadband services.[10] So, to the extent municipal communications services create awareness and/or promote economic development, then the resulting market expansion may increase the equilibrium number of firms in the market (Sutton 1995; Beard and Ford 2003).

In sum, we have two opposing hypotheses related to municipal entry into telecommunications: a) the *crowding out* hypothesis and b) the *stimulation* hypothesis. As is typical of important policy questions, theory does not provide unambiguous guidance. The relationship between municipal entry and the magnitude of private entry and investment relationship is, in the end, and empirical question, and "an empirical question cannot be settled by non-empirical arguments (Stigler 1968: 115)."

In this paper, we subject the "crowding out" hypothesis to an empirical test using data recently released by the Florida Public Service Commission on the number of competitive local exchange carriers ("CLECs") serving particular markets. Combining this city level data on CLEC entry with demographic and other data, we specify and estimate an empirical model that quantifies the effect of municipal communications on private firm entry. The empirical model performs remarkably well in terms of fit, and specification testing indicates the model is correctly specified. Our empirical model provides no evidence to support the crowding out hypothesis. In fact, we find statistically significant evidence of more private firm entry in markets where municipalities operate communications network (a 63% increase). Thus, this evidence presented here supports the stimulation hypothesis.

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[9] In Wisconsin, for example, state law requires that municipalities can provide communications services only on a wholesale basis.

[10] Fred O. Williams, "Urban Wi-Fi 'hot spots' criticized in report," *Buffalo (N.Y.) News* (Feb. 4, 2005).

II. The Empirical Framework

In the *Annual Report to the Florida Legislature on the Status of Competition in the Telecommunications Industry in Florida* (2004), the Florida Public Service Commission lists the number of CLECs operating in each local rate exchange in the state of Florida.[11] This format is convenient, since it can be merged city-specific demographic and other data to evaluate how municipal provisioning of services affects private firm entry, holding other factors systematically influencing CLEC entry constant.

Using this data, we specify an empirical model of the number of private, competitive communications firms serving a market as:

$$\begin{aligned} N_i = {} & \gamma_1 DMUNI_i + \gamma_2 DCOMM_i + \beta_1 HH_i + \beta_2 INC_i + \beta_3 URBAN_i + \beta_4 URBANC_i \\ & + \beta_5 DENSE_i + \beta_6 LOOP_i + \beta_7 + \beta_8 DVERIZON_i + \beta_9 DSPRINT_i + \varepsilon_i \end{aligned} \tag{1}$$

where N is the number of CLECs serving market i. In addition to variables indicating whether or not the market has municipally-supplied electricity service (*DMUNI* = 1) and whether these particular markets also have a municipally-supplied communications network (*DCOMM* = 1; *DMUNI* = 1), there are a number of other regressors that measure the demographic and ILEC profile of the market. Demographic determinants of CLEC entry include five variables: 1) *HH* is city households (in thousands); 2) *INC* is per-capita income (in thousands); 3) *URBAN* is the percent of population in urban areas; 4) *URBANC* is a dummy variable that equals one if the entire urban population is located inside the urbanized areas (a measure of urban density); and 5) *DENSE* is households per thousand square land miles. CLEC entry is likely to be affected by the characteristics of the incumbent local exchange carrier serving the city, since interconnection among entrants and incumbents is nearly always required. ILEC-specific factors are captured by three variables: 1) *LOOP* is the price of an unbundled loop in the market; 2) *DVERIZON* is the dummy variable for Verizon exchanges; 3) *DSPRINT* is a dummy variable for Sprint exchanges. The constant term of the regression (β_7) includes the mean effect for BellSouth Communications.

We expect all the demographic variables (*HH*, *INC*, *URBAN*, *URBANC*, and *DENSE*) to be positively related to CLEC entry; more firms are expected in larger, richer, and more densely populated markets. Following the law of demand, a negative sign on the loop price (*LOOP*) is expected, since an unbundled loop in an input of production. About half of competitive lines in Florida are serviced using unbundled elements, so we expect loop prices to be a significant determinant of CLEC entry.[12] We expect the dummy variables *DVERIZON* and *DSPRINT* to be negative, indicating less entry in these regions than in the BellSouth region. Of these three ILECs, BellSouth was the only one in Florida required to satisfy the competitive checklist of Section 271 of the 1996 Act prior to offering interLATA long distance services. Under Section 271, interLATA entry was a

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[11] http://www.psc.state.fl.us/.

[12] LOCAL TELEPHONE COMPETITION – STATUS AS OF JUNE 30, 2004, Federal Communications Commission (December 22, 2004), at Table 10: www.fcc.gov.

quid pro quo for opening its markets to competition from CLECs. As a result, Bellsouth was required to do far more to allow competition than its sister ILECs, and consequently there is considerably more competition in the BellSouth region than any of the other ILEC regions in Florida. Random influences on CLEC entry are captured by the econometric disturbance term (ε).

The variables of most interest in the model are the two dummy variables *DMUNI* (municipal electric, no communications) and *DCOMM* (municipal electric, with communications). For expositional purposes, let the term βX be the sum of the estimated β coefficients multiplied by the sample means of the variables.[13] We have three scenarios of interest, with the mean number of CLECs per market being:

i) No Municipal Electricity $N_1 = \beta X$;

ii) Municipal Electricity, No Communications $N_2 = \gamma_1 + \beta X$;

iii) Municipal Electricity, With Communications $N_3 = \gamma_1 + \gamma_2 + \beta X$.

The difference $N_3 - N_2$ measures the mean effect of the provision of communications network by a city with municipally-supplied electricity. Thus, the competing hypotheses on the effect of communications supply are tested directly by the coefficient γ_2. If γ_2 is negative, then there is crowding out. If γ_2 is positive, then the municipal supply of communications network leads to an increase in private firm entry (the stimulation hypothesis). Of course, municipal entry may have no effect, and in that case, γ_2 will be statistically indistinguishable from zero. The null hypothesis of zero-effect can be tested using the t-statistic on γ_2. If the hypothesis is rejected, then the sign on γ_2 serves to distinguish which of the two competing hypotheses is more consistent with the data. In sum, the crowding out hypothesis is supported if $\gamma_2 < 0$, and the stimulation hypothesis is supported if $\gamma_2 > 0$. The sign on γ_2, if the coefficient is statistically different from zero, indicates which hypothesis is more consistent with the data.

1. Estimation Technique

Evaluating the two competing hypotheses on municipally-supplied communications network hinges on hypothesis testing (particularly on γ_2), so the efficiency of our estimated coefficients is critical. The dependent variable N is a non-negative count of CLECs, and linear regression for such data can result in inefficient, inconsistent, and biased estimates (Long 1997: 217). So, we employ more appropriate estimation techniques including Negative Binomial and Poisson regressions. Due to evidence of (mild) overdispersion in the data, the Negative Binomial estimation technique is more appropriate, since the Negative Binomial regression does not require equality of the conditional mean and variance (but does require the conditional variance to exceed the

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[13] We cannot reject the null hypothesis that the means of each of the demographic regressors and the loop price are equal across cities with and without municipal electric operations.

mean).[14] In the presence of overdispersion, the Negative Binomial regression is more efficient than Poisson (Wooldridge 2002: Ch. 19).[15] The Poisson regression, which requires that the conditional mean of the data equal the conditional variance, can produce estimated standard errors that are too small in the presence of overdispersion, thereby leading to a spurious overstatement of statistical significance (Gourieroux et al., 1984). However, it is possible to calculate fully robust standard errors for Poisson regression in the presence of overdispersion, and we do so (Wooldridge 2002: 649-650).[16]

2. DATA

Our dependent variable, *N*, is taken from the *Annual Report to the Florida Legislature on the Status of Competition in the Telecommunications Industry in Florida* (2004), Appendix B. We limit our analysis to the state of Florida since it is the only source of which we are aware that provides exchange specific data on CLEC activity. We also focus on business CLECs, since entrants serving businesses are more likely to be facilities-based, thereby the analysis can be extrapolated to financial investments in the community rather than just a count of sellers. We combine the exchange-level CLEC data presented in the *Annual Report* with city demographic data from the 2000 Census.[17] We use the 2004 *APPA Annual Directory & Statistical Report* to indicate whether cities with municipal electric services also operate some type of communications network. We also match the cities to unbundled loop prices, which apply only to the largest incumbent phone companies in Florida (BellSouth, Verizon, and Sprint). The *Annual Report* (2003) lists data for 277 cities, but only 225 of these are in the large ILEC regions for which loop price data is available. Of these markets, we are able to match up demographic and loop price data to 163 cities. Since our focus is on business CLECs, it is reasonable to limit our sample to cities with at least some urban population, leaving us with 126 exchanges in

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[14] Overdisperson tests include those proposed by Cameron and Trivedi (1988; t = 2.45) and Wooldridge (1997; t = 4.41), which both tests indicate the presence of overdispersion in the dependent variable. Both tests require first estimating the equation using Poisson, then running secondary regressions $(y - \hat{y})^2 - y$ on $\hat{y}^2$ for the first test and $e^2 - 1$ on $\hat{y}$, where e is the standardized residual, for the second. The t-statistic on the sole regressor is a test of the null of "equality of the conditional mean and variance." The statistically significant and positive coefficient α in the Negative Binomial regression (Table 1) indicates mild overdispersion (Cameron and Trivedi 1988, at 79).

[15] Overdispersion occurs when one event makes other within-observation events more likely; underdispersion occurs when one event makes other events less likely. Overdispersion is a somewhat peculiar finding for the number of firms serving a market, since an additional firm should reduce the profitability of further entry. The overdispersion is mild, however.

[16] By robust, we mean robust to overdispersion. The robust standard errors reported in Table 1 are not robust to heteroskedasticity. We estimated Equation (2) using least squares and the natural logarithmic transformation of *N* as the dependent variable, a specification which is close approximation to the Negative Binomial model. From this regression, we were unable to reject the White Test's null hypothesis of homoscedastic disturbances (χ^2 = 14.05, Probability 0.52). So, we do not believe heteroskedasticity is a problem. Linear, least squares with a logarithmic transformation of the dependent variable is a close approximation to Poisson and Negative Binomial regression: for Poisson we have $y = \exp(x\beta)$, and for the transformed least squares regression we have $\ln(y) = x\beta$ (Long 1997: 224-8). However, the least squares estimates remain biased, despite the transformation (King 1988).

[17] The data for "Places" in Florida was extracted from the SF3 data using the *dataferret* interface made available by the Census Bureau (www.census.gov).

our final dataset. In the final sample, there are 24 (19%) municipally operated electric companies, with 15 of these municipalities operating communications plant (12%).

III. Results

Estimates of Equation (1) are summarized in Table 1, along with the descriptive statistics of the variables in the model. Both the traditional (in parenthesis) and robust standard errors (in brackets) are used to compute the t-statistics. Equation (1) is estimated both by the Negative Binomial (Model 1) and Poisson techniques (Model 2). Both of the models fit the data well, with a Psuedo-R^2 of 0.81 for Model 1 and 0.82 for Model 2. Such a good fit to the data is encouraging, particularly for cross sectional data.[18]

Model specification is evaluated using RESET, where Wooldridge's (1991, 1999) robust RESET specification test for conditional mean regressions is employed. The null hypothesis of "no specification error" cannot be rejected for either model at even the 10% level.[19] This result is encouraging, since RESET is a rather general specification error test capable of detecting incorrect functional form, omitted variables, and simultaneity (though it has the most power against incorrect functional form) (Ramsey 1969; Gujarati 1995: 464-6; Godfrey 1988). Both the standard and robust RESET F-statistics are reported for each regression.

The estimated coefficients across the two models are very similar, and all but two (*INC* and *DENSE*) of the estimated coefficients are statistically different from zero at traditional levels, regardless of whether the traditional or robust standard errors are used for hypothesis testing. *DENSE* is statistically significant in Model 2, though not in Model 1. The signs on all statistically-significant demographic variables (*HH*, *URBAN*, *URBANC*, and *DENSE*) are as expected – CLECs tend to enter large, densely populated markets. Also as expected, the signs on *DVERIZON* and *DSPRINT* are negative and statistically different from zero, indicating that CLECs are more likely to enter cities inside the BellSouth region, even after accounting for variations in unbundled loop prices. Higher unbundled loop prices also reduce CLEC entry, as indicated by the negative and statistically significant coefficient on *LOOP*. We compute the elasticity for loop prices to be about -0.30, implying a 10% increase in the price of an unbundled loop reduces the number of CLECs serving a market by 3% (about one CLEC).[20]

Turning to the key policy variables, we observe that both *DMUNI* and *DCOMM* are statistically different from zero. The estimated parameters indicate that CLEC activity

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[18] The Psuedo-R^2 is computed as the squared correlation coefficient between the predicted and actual values of the dependent variable While an imperfect measure of fit (as are all such measures for non-linear regression), the statistic does illustrate that our chosen model explains a large percentage of the variation in CLEC counts (Cameron and Trivedi 1998: 151-8).

[19] In an alternate specification, continuous regressors were log transformed. We were able to reject the null hypothesis of RESET at the 10% level for this specification, suggesting our chosen specification is preferred.

[20] As recommended by Cameron and Trivedi (1998: 80-81), we report the average response over all cities in the sample, rather than compute the responses at the sample means.

generally is lower in cities with municipal electric operations ($\gamma_1 \approx -0.40$). However, the regression also indicates that the provision of communications services by municipal electrics significantly increases CLEC entry ($\gamma_2 \approx 0.50$). Both coefficients are statistically different from zero. Based on the average response over all cities in the sample, the mean predictions from the regression are summarized as follows:

i) No Municipal Electricity $N_1 = 23.18$;

ii) Municipal Electricity, No Communications $N_2 = 16.12$;

iii) Municipal Electricity, With Communications $N_3 = 26.24$.

The model predicts that cities that self-supply electricity have approximately 7 fewer CLECs ($N_3 - N_1$), on average, than do similarly situated cities without municipal electricity operations (a 30% reduction). Within the group of cities self-supplying electricity (*DMUNI* = 1), those cities with communications networks (*DCOMM* = 1) average about 10 more CLECs ($N_3 - N_2$), other things constant (a 63% increase).[21] Relative to cities that do not have municipal electric operations, municipalities operating both electric and communications networks ($N_3 - N_1$) have on average about three more CLECs (a 13% increase) than similarly situated cities without municipally-supplied electricity. This latter increase is statistically different from zero; that is, the hypothesis $\gamma_1 + \gamma_2 = 0$ is rejected for both models. Thus, there are more CLECs in those municipalities operating communications networks than in cities that do not have municipally run communications networks, regardless of whether those cities also supply electricity services.

Our empirical model provides no support for the crowding out hypothesis ($\gamma_2 > 0$; $\gamma_1 + \gamma_2 > 0$), but strong support for the stimulation hypothesis ($\gamma_2 > 0$; $\gamma_1 + \gamma_2 > 0$). Other things constant, the empirical model indicates that municipally operated communications networks lead to a 63% increase in CLEC count relative to other cities supplying their own electricity, and a 13% increase in CLEC count relative to cities with privately-supplied electricity.

IV. Conclusions

The municipal supply of communications services is on the rise. While constituents are generally delighted with the municipal services, incumbent firms that compete (or may do so at some later date) with these systems are unsurprisingly displeased with the rise of municipal communications. The incumbents levy many arguments against municipal entry, one of them being that public investment in communications networks crowds out private investment. In this paper, we subjected this hypothesis to an empirical test, and found no evidence to support the "crowding out" hypothesis. Indeed, the empirical model indicates that municipal communications actually increases private firm entry

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[21] Depending on what services the municipal is providing, it may be a required to be a registered CLEC.

and, presumably as a consequence, private investment (thereby supporting the stimulation hypothesis).

One inference from this study is that legislation restricting or precluding municipal provision of communications services reduces the overall level of competition in the market, and this reduction is not merely limited to the absence of a municipal provider. The empirical model predicts that the absence of municipal provider is accompanied by 13% fewer competitors. Any reduction in competition in communications is troubling, since communications services, particularly broadband Internet services, are generally believed to produce large positive externalities. In the presence of positive externalities, even a perfectly competitive market will undersupply communications. So, reductions in competition will move consumption even further away from the optimal level of communications services.

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Table 1. Regression Results

	Model (1) Neg. Binomial	Model (2) Poisson	Mean (St. Dev)
DMUNI	-0.363 (-2.70)* [-1.92]**	-0.405 (-3.75)* [-1.98]*	0.190 (…)
DCOMM	0.487 (3.17)* [2.49]*	0.530 (4.42)* [2.49]*	0.119 (…)
HH	0.003 (4.49)* [3.65]*	0.003 (7.39)* [5.22]*	18.332 (36.258)
INC	-0.005 (-1.38) [-1.40]	-0.005 (-1.74) [-1.29]	20.706 (9.176)
URBAN	0.882 (2.57)* [5.35]*	0.853 (3.12)* [5.07]*	0.965 (0.092)
URBANC	0.280 (3.42)* [3.79]*	0.269 (4.32)* [3.63]*	0.698 (…)
DENSE	0.181 (1.46) [1.29]	0.217 (2.91)* [2.03]*	0.347 (0.251)
LOOP	-0.014 (-3.48)* [-3.81]*	-0.015 (-4.98)* [-3.95]*	26.320 (10.096)
Constant	2.710 (7.27)* [11.96]*	2.757 (9.38)* [11.63]*	…
DVERIZON	-0.672 (-7.21)* [-10.34]*	-0.657 (-9.59)* [-10.54]*	0.159 (…)
DSPRINT	-0.771 (-11.09)* [-11.19]*	0.461 (-14.90)* [-11.13]*	0.421 (…)
α	0.037 (-11.21)* [-8.02]*	…	… …
N			23.333 (16.39)
(Psuedo) R^2	0.81	0.82	
RESET F	2.04	1.39	
RESET Robust F	2.57	1.34	
χ^2: $\gamma_1 + \gamma_2 = 0$	4.09*	3.77*	
Obs.	126	126	

* Significance at 5% level; ** Significance at 10% level.
Traditional t-stats in parenthesis, robust t-statistics in brackets.