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Journal of Asian Economics 16 (2005) 688–703
JOURNAL OF ASIAN ECONOMICS

How does the entry of new firms change demand? An empirical estimation for a Thai telecommunications company

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Received 9 September 2004; received in revised form 31 May 2005; accepted 7 June 2005

Abstract

This paper uses cointegration, error correction mechanism (ECM) and causal effect techniques to estimate the demand for the services supplied by the Telephone Organization of Thailand (TOT). The entry of new competitors reduces access demand but increases usage demand in both long-run and short-run. Access and usage externalities exist and the former has stronger effects than the latter. It is also found that consumers initially overreact to increases in income and telephone subscription but they do not overreact to services price changes. Causality examination shows that the number of telephone subscribers and income cause demands for access and usage, vice verse.

JEL classification: C22; L96

Keywords: Access demand; Usage demand; Cointegration; Error correction model; Causality; Telecommunication

1. Introduction

Demand for telecommunications services as a research topic has attracted wide attention since the pioneering works by Littlechild (1975) and Taylor (1980). It is an

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1049-0078/$ – see front matter
doi:10.1016/j.asieco.2005.06.002

interesting topic because it differs from the demand for other conventional goods. Firstly, telecommunications demand can be divided into two parts: demand for access and demand for usage. But more importantly, telecommunications involve two parties of subscription: the party who makes a phone call and the party who receives the call. Therefore, telephone networks have strong externality effects.

In terms of access demand, Taylor (1980) calculates the price elasticity of demand for residential access in the United States. His result, an elasticity of −0.030, is comparable with more recent study by Cain and MacDonald (1991) who report an elasticity of −0.038 in the United States. Similarly, the study of Canada by Bodnar, Dilworth, and Iacano (1988) shows an elasticity of −0.009 while Madden, Bloch, and Hensher (1993) report an elasticity of −0.003 for the Australian demand. These results demonstrate that the access demand is quite inelastic, i.e. when the cost of installing a telephone line connection rises, the number of new subscribers declines marginally.

For the usage demand, Bidwell, Wang, and Zona (1995) investigate the price elasticity of demand for local calls in the United States and find it equal to −0.04. Park, Wetzel, and Mitchell (1983) study three states in the United States and conclude that the elasticity is equal to −0.008. In contrast, the Australian study by Madden et al. (1993) finds that the price elasticity of demand for local calls is −0.46. So, there is a large difference between the US demand and the Australian demand; the former is much more inelastic. This difference is likely to stem from the different pricing structures between the US and Australia. In the US, the price of a local call is very low and in many cases a flat monthly fee for local calls is applied. Thus, customers are insensitive to a change in the price of local calls. In contrast, the Australian demand for local calls is more sensitive to price because Australian consumers pay for each local call on top of monthly charges.

A weakness of these studies is that they treat demand as a static variable and the issue of unit root is ignored. In fact, most empirical analyses of telecommunications demand use time series data such as income and the price of telecommunications services. Like other empirical studies, there is a problem of unpredictability when time series are used to estimate demand for access and usage because of the change in the trend of these time series. Economists have paid increasing attention to this problem since late 1970. Testing the unit root of a time series model plays a central role in time series studies since Dickey (1976), Dickey and Fuller (1979), and Fuller (1976). The unit root testing model has been widely accepted after Nelson and Plosser (1982) investigated unit roots in the US macroeconomic time series and demonstrated that unit roots existed in 13 out of 14 time series. Empirical studies by Granger and Newbold (1974), Nelson and Kang (1981), and Nelson and Plosser (1982) also suggest that the majority of macroeconomic variables are non-stationary, which cause serious implications and unreliability for econometric models, in particular, the spurious issue.

Applying the cointegration (long-run) and error correction mechanism (ECM) models (short-run), Amaral, Gonzalez, and Jimernez (1995) study the demand for Spanish telecommunications services between 1980 and 1991. They conclude that the price elasticities of local calls are −0.17 in the long-run and −0.19 in the short-run. The price elasticities of long distance calls are −0.34 in the long-run and −0.36 in the short-run, respectively. They also find that the short-run effect is stronger than the long-run effect, because consumers initially overreact to a change in price, but gradually they adapt to the new prices. Andres and Amaral (1998) study long-run and short-run access demands for telephone services in Spain for the period 1980–1993. They report that the elasticities for installation charge in the long-run and short-run are −0.56 and −0.30, respectively, and the income elasticities are 1.25 and 3. These results suggest that the demand for new telephone lines is very sensitive to changes in income (both in the long-run and short-run) while it is insensitive to changes in installation charge.

The purpose of this paper is to use time series data compiled by the Telephone Organization of Thailand (TOT) to estimate the long-run (cointegration) and short-run (error correction mechanism) responsiveness of access demand and usage demand to the changes in price and the number of telephone subscribers, with particular interests in the effects of the introduction of concessions which allows new firms to operate in the industry.[1] The rest of paper is arranged as follows. Section 2 outlines the method of our analysis and presents demand models for TOT services. Section 3 briefly describes the data used in our analysis. Section 4 focuses on empirical analysis of demand for the access to the TOT network and demand for local call services provided by TOT. It also documents the price elasticity, income elasticity, access externality (new subscriber effects) and usage externality (calling effects) and causality. Since the introduction of concessions has a substantial effect on the demand for TOT services, we distinguish the demand for TOT services before and after the concessions were granted. Section 5 summarizes the conclusions from our analysis.

2. Demand for the services of Telephone Organization of Thailand: methods and models

2.1. Methods

In this paper we adopt the techniques of cointegration, error correction mechanism and the causality test developed by Engle and Granger (1987), and Granger (1981, 1986) to study the demand for the TOT's telecommunications services. The test of the long-run and short-run relationships between demand and independent variables can be divided into three steps.

First, we test the stationarity of the time series and determine the orders of integration for each series by using the Dickey–Fuller (DF) and Augmented Dickey–Fuller (ADF) tests.

Second, we estimate the cointegration regression of ordinary least squares by using the same order of variable integrations.

[1] Before 1993, Thai telecommunications services were supplied by two state-owned firms: TOT, which provided domestic and long distance services, and the Communications Authority of Thailand (CAT), which provided international services. But the fast growth of Thai economy in 1980s and early 1990s made demand for telecommunications services much greater than the supply provided by TOT and CAT. To resolve the issue of services shortage, the Thai government made a concession which allowed two private firms, Telecom Asia (TA) and Thai Telephone and Telecommunications (TT&T), to build telephone networks. These two companies were permitted to provide services to the public for 25 years after the networks were built, under the conditions specified by concession agreements between the government and the companies. Thus, the private companies can earn profits from the operation after recovering their initial investment costs although the ownership of network assets belongs to the Thai government. The concession to TA was granted in 1992 and to TT&T in 1993. Both companies began operation in 1993.

Finally, we test the stationarity of cointegration regression residuals. This can be divided into two cases: (i) when the residuals are I(0) stationary (by ADF test), the error correction mechanism can be used to estimate the short-run relationships; (ii) when the residuals are non-stationary but the residuals are stationary in first-order difference, I(1), the short-run relationships can still be tested by the ECM.

After these tests, we investigate, based on ECM models, Granger causality between demand and independent variables. The Microfit Econometric Program version 4 is used to estimate access and usage demand for Thai telecommunications.[2]

2.2. Demand models

Usually the demand for access is defined by the number of new subscribers. However, the supply of telephone access is short in Thailand and clients have to wait for 3 years on average before a telephone line connection is installed. Since there are a substantial number of clients waiting for the access to services, the demand for access in our model equals the number of new subscribers plus the potential subscribers on the waiting list. The demand for (local call) usage in our model is conventional; i.e., the number of local calls.

Most studies of the demand for access include connection fee, monthly charge, income and the number of existing subscribers as explanation variables. For the demand for usage, it is usually considered to be determined by the price of a local call, income and the number of existing subscribers. However, Thailand has a specific feature in the cost of installing telephone connection; i.e., when a telephone line is installed, the subscriber has to pay a bond in addition to conventional connection fee. The bond is refunded when the subscriber terminates the lease of the telephone line but it receives no interests. To certain extent, the bond can be considered as an interest-free loan from clients to the telephone company. Therefore, we add the amount of the bond a client must pay into the model to determine the demand for the access to the TOT telephone network. In addition, the price of posting a letter has been added into the usage demand model because it can be considered as a substitute for local calls. Both models have a time dummy variable distinguishing the pre-concession period from the post-concession period, so we can clearly observe the demand change caused by the entry of new competitors.

We write the demand functions for the access to the TOT network and its local call services as follows:

$$\mathrm{NS} = f(\mathrm{CF}, \mathrm{MC}, \mathrm{GDP}, \mathrm{DEN}, \mathrm{BO}, D), \tag{1}$$

$$\mathrm{LOC} = f(\mathrm{PR}, \mathrm{GDP}, \mathrm{DEN}, \mathrm{LET}, D), \tag{2}$$

where NS is the demand for access to the TOT network (the number of new subscribers plus the number of intended subscribers on waiting list), CF the connection fee, MC the monthly charge, GDP the per capita GDP, DEN the telephone density (the total number of existing telephone subscribers), BO the amount of bond, D the dummy variable (equal to 1 after 1993—the year concessions were granted and 0 otherwise), LOC the total number of local

[2] See Pesaran and Pesaran (1997).

calls made by the subscribers of TOT, PR the price of a local call, LET the price of posting a letter.

All variables in our models except for the dummy variables are transformed to natural logarithms. Since network externalities (access and usage externalities) are particularly important in telecommunications and are distinguishing features of demand for telephone services, the number of existing subscribers can be used to measure access externality and usage externality. If the number of new subscribers (local calls) is positively related to the number of existing subscribers, there is a positive access (usage) externality. Network externality effects in developing countries are likely to be stronger than in developed countries because they have a lower percentage of telephone penetration than developed countries. If the coefficient of number of existing subscribers is positive in (1) or (2), then we can say that an externality exists.

3. Data

The main sources of the data used in this study are the TOT Annual Reports and official documents compiled by the National Statistical Office of Thailand, the Department of Internal Trade and the Bank of Thailand. The sample covers the period 1973–2000. The data for the number of new subscribers, the total number of subscribers, connection fee, monthly fee and bond are obtained from the TOT. The number of local calls is obtained from the National Statistical Office of Thailand while GDP per capita data is collected from the Bank of Thailand. All prices and income are in terms of baht (the Thai currency) and deflated by the consumer price index in 1988. Tables 1 and 2 report the statistics of variables (in their log-transformations) used in access and usage demand models.

4. Empirical results

4.1. Testing for unit root

The formal analysis of unit root is the Dickey–Fuller test and the Augmented Dickey–Fuller test. To begin empirical analysis, the orders of integration for all variables are

Table 1
Means and standard deviation of access demand: 1973–2000

	NS	CF	MC	GDP	DEN	BO
Maximum	13.4053	8.5061	4.6022	10.8556	14.8721	8.0024
Minimum	9.9293	4.8675	2.5649	9.5722	12.0732	6.1995
Mean	11.9297	7.5801	4.1820	10.2169	13.5787	7.8436
Standard deviation	1.0103	1.1459	0.66618	0.43696	0.96018	0.36303
Skewness	−0.31564	−1.7677	−1.8411	0.074745	−0.034637	−3.6143
Kurtosis-3	−0.85814	1.5203	1.7239	−1.4316	−1.4698	13.3413
Coefficient of variation	0.084691	0.15117	0.15929	0.042768	0.070712	0.046284

Table 2
Means and standard deviation of usage demand: 1973–2000

	LOC	PR	GDP	DEN	LET
Maximum	22.9455	1.0956	10.8556	14.8721	1.0956
Minimum	19.7304	−0.026344	9.5722	12.0732	−0.35954
Mean	21.3253	0.69540	10.2177	13.5787	0.41752
Standard deviation	1.0938	0.44176	0.43794	0.96018	0.54248
Skewness	0.18397	−0.55830	0.077047	−0.034637	−0.28854
Kurtosis-3	−1.3581	−1.3159	−1.4326	−1.4698	−1.4208
Coefficient of variation	0.051289	0.63526	0.042861	0.070712	1.2993

examined. For the level test, the model is:

$$Y_t = \phi Y_{t-1} + u_t, \tag{3}$$

where Y_t is the time series under the consideration (Y can be CF, MC, GDP, DEN, BO, PR or LET in our analysis), ϕ is a constant and u_t is white noise. If $\phi = 1$, $Y_t = Y_{t-1} + u_t$ and it is a pure random walk. If $|\phi| < 1$, the value of Y_t swings in sign but converses. If $|\phi| > 1$, the trend of Y_t grows without limit. For the Dickey–Fuller test, the model is:

$$\Delta Y_t = \delta Y_{t-1} + u_t, \tag{4}$$

where $\Delta Y_t = Y_t - Y_{t-1}$ and $\delta = \phi - 1$. For the Augmented Dickey–Fuller test, the model is:

$$\Delta Y_t = \delta Y_{t-1} + \sum_{i=1}^{k} \delta_i \, \Delta Y_{t-i} + u_t. \tag{5}$$

Hypotheses for testing are:

H_0: unit root exists; i.e., $\delta = 0$ (or $\phi = 1$);
H_1: stationary; i.e., $\delta < 0$ (or $|\phi| < 1$).

The results of the tests are shown in Table 3 (access demand) and Table 4 (usage demand). For the level of the time series, none of the variables in both the access and usage

Table 3
Testing stationarity for access demand: 1973–2000

	DF test		ADF test	
	Level	First-order difference	Level	First-order difference
NS	−1.7127	−5.841*	−1.2131	−4.023*
CF	−2.4386	−5.0467*	−2.2551	−3.655**
MC	−2.6489	−5.0768*	−2.4558	−3.2661**
GDP	−0.6944	−2.5693	−2.7827	−3.1849**
DEN	−1.8397	−6.9693*	−0.9427	−3.8842**
BO	−2.641	−5.147*	−3.465	−4.6531*

* Significant at 1% level.
** Significant at 5% level.

Table 4
Testing stationarity for usage demand: 1973–2000

	DF test		ADF test	
	Level	First-order difference	Level	First-order difference
LOC	−1.174	−3.2406**	−1.6348	−2.0933
PR	−1.3117	−5.625*	−0.9971	−3.0713**
GDP	−0.6944	−2.5693	−2.7827	−3.1849**
DEN	−1.8397	−6.9693*	−0.9427	−3.8842**
LET	−2.289	−5.4245*	−2.2151	−4.1482*

* Significant at 1% level.
** Significant at 5% level.

demand models reject the null hypothesis of non-stationarity at 5% statistical significance level. However, after first-order differencing, all variable series rejected the null hypothesis at least at 5% statistical significance level. In other words, the series are stationary in the first-order difference.

4.2. Cointegration test

This subsection adopts the two-step method of Engle and Granger (1987) to test cointegration. The first step estimates the long-run access demand and usage demand by OLS regression and then tests the stationarity of the residuals. If they contain a unit root, the null hypothesis of non-cointegration cannot be rejected. The second step uses the ECM model, if cointegration exists, to estimates the residuals represented by the disequilibrium terms.

An informal test for the cointegration can be conducted by comparing the Durbin–Watson (DW) statistic with adjusted R^2. The cointegration may exist if the DW statistic is greater than the adjusted R^2.[3] Results in Tables 5 and 6 below illustrate that the adjusted R^2 is less than the DW statistic for both demand models which means that the cointegration is likely to exist. For the access demand, the adjusted R^2 is 0.9272 while the DW statistic is 2.9262. For usage demand, the adjusted R^2 is 0.9987 but the DW statistic is 2.6504.

A formal test for the cointegration (long-run) relationship is done by testing the statistics on the residuals. The cointegration equations of access and usage demands are:

$$\mathrm{NS}_t = \hat{a}_1 + \hat{b}_1\mathrm{CF}_t + \hat{b}_2\mathrm{MC}_t + \hat{b}_3\mathrm{GDP}_t + \hat{b}_4\mathrm{DEN}_t + \hat{b}_5\mathrm{BO}_t + e_t, \tag{6}$$

$$\mathrm{LOC}_t = \hat{a}_2 + \hat{c}_1\mathrm{PR}_t + \hat{c}_2\mathrm{GDP}_t + \hat{c}_3\mathrm{DEN}_t + \hat{c}_4\mathrm{LET}_t e_t. \tag{7}$$

Their cointegrating regression equations can be written as:

$$e_t = \mathrm{NS}_t - \hat{b}_1\mathrm{CF}_t + \hat{b}_2\mathrm{MC}_t - \hat{b}_3\mathrm{GDP}_t - \hat{b}_4\mathrm{DEN}_t - \hat{b}_5\mathrm{BO}_t - \hat{a}_1, \tag{8}$$

$$e_t = \mathrm{LOC}_t - \hat{c}_1\mathrm{PR}_t - \hat{c}_2\mathrm{GDP}_t - \hat{c}_3\mathrm{DEN}_t - \hat{c}_4\mathrm{LET}_t - \hat{a}_2. \tag{9}$$

[3] If cointegration exists, the error correction mechanism occurs as well.

Table 5
Cointegrating equation for access demand: 1973–2000

	Coefficient	T-ratio
Constant	−113.13	−4.4924*
CF	−1.6018	−4.2579*
MC	−2.0576	−2.7591**
GDP	−1.2301	−1.0265
DEN	2.6724	4.6158*
BO	15.478	4.8581*
D	−0.9184	−3.4234**
R^2	0.97092	
Adjusted R^2	0.9272	
DW	2.9226	
F-statistic $F(15, 10)$	22.256*	
DF	−5.6579**	
ADF(1)	−4.6328**	

Note: The dependent variable is NS. The independent variables are CF, MC, GDP, DEN, BO, and dummy variable D. The residual test of the Dickey–Fuller (DF) and Augmented Dickey–Fuller (ADF) in cointegration equations are based on the following models: $u_t = \phi u_{t-1} + e_t, \Delta u_t = \delta u_{t-1} + \sum_{i=1}^{k} \Delta u_{t-1} + e_t$.

* Significant at 1% level.

** Significant at 5% level.

We now test stationarity in the residual (e_t) from Eqs. (8) and (9). Stationary residuals imply that dependent and independent variables are cointegrated (long-run relationship). Hypotheses for testing are:

H_0: unit root in residuals;
H_1: stationary in residuals.

Table 6
Cointegration equation for usage demand: 1973–2000

	Coefficient	T-ratio
Constant	4.788	2.6723**
PR	0.0604	0.413
GDP	0.7396	1.9859***
DEN	0.6645	3.3473*
LET	0.0556	0.4727
D	0.2445	2.7343**
R^2	0.999	
Adjusted R^2	0.9987	
DW	2.6504	
F-statistic $F(6, 19)$	3484.8	
DF	−5.496**	
ADF(1)	−2.699***	

Note: The dependent variable is LOC. The independent variables are PR, GDP, DEN, LET and dummy variable D. The residual test of the Dickey–Fuller (DF) and Augmented Dickey–Fuller (ADF) in cointegration equations are based on the following models: $u_t = \phi u_{t-1} + e_t, \Delta u_t = \delta u_{t-1} + \sum_{i=1}^{k} \Delta u_{t-1} + e_t$.

* Significant at 1% level.

** Significant at 5% level.

*** Significant at 10% level.

Tables 5 and 6 present, respectively, the results of the cointegrations of access and usage demands. They show that the DF and ADF statistics reject the null hypothesis (unit root in residuals). The residuals are statistically significant at the 5% and 10% levels in access demand and usage demand, respectively. This means that there exists a long-run linear relationship between demand for access (NS) and independent variables (connection fee, monthly charges, income, telephone density and bond) and between the number of local calls (LOC) and independent variables (local call price, income, telephone density and price of posting a letter).

More specifically, Table 5 shows that all independent variables have correct signs except for per capita GDP and bond. Connection fee and telephone density are statistically significant at 1% level, whereas monthly fee and dummy variable are statistically significant at 5% level. A 1% increase in connection fee decreases access demand by 1.60%. This is significantly higher than the connection fee elasticities in the United States (Cain & MacDonald, 1991; Taylor, 1980), in Canada (Bodnar et al., 1988), in Australia (Madden et al., 1993) and in Spain (Andres & Amaral, 1998). The demand for access in these countries is very inelastic to connection fee. We conjecture that the price elasticity of demand for the access to a telephone network decreases in per capita income and increases in price. Since Thailand has a lower per capita income relative to OECD countries and since the connection fee in Thailand is higher than in OECD countries and other Asian countries, Thai customers are, not surprisingly, more sensitive to connection fee. Table 5 also shows that a 1% increase in telephone density raises access demand by 2.67%. This implies that access externality exists in Thai telecommunications in the long run. Moreover, the data illustrate that monthly fee has a greater elasticity than connection fee. For the dummy variable, it is statistically significant at 5% level. This can be interpreted as that there is a difference between pre- and post-telecommunications concessions in terms of access demand. After the introduction of concessions, new operators, TA and TT&T, entered into the Thai telecommunications market. With stronger competition, the demand for the access to TOT services in the long-run decreases by 0.92%. This is clearly consistent with economic intuition that incumbents lose market share when new firms enter the market.

Although the effect of GDP has a wrong sign,[4] it is statistically insignificant. On the other hand, the effect of bond also has a wrong sign with a statistical significance level of 1%. Since the bond represents an interest-free loan from clients to the TOT, we expect the demand for access declines as bond increases. But the coefficient of BO in Table 5 is positive, opposite to our expectation. The wrong sign is possibly caused by subscribers' expectation that they will get the bond back when they terminate their lease of the telephone line. So even if bond rises, the demand for access does not decline.

Turning to the usage demand, Table 6 shows that the long-run effects of telephone density and per capita GDP on usage demand are statistically significant at 1% and 5% levels, respectively. A 1% increase in telephone density or per capita GDP increases 0.66% or 0.74% in the demand for local calls. Since the coefficient of telephone density is positive, the usage externality exists in the long-run. The dummy variable has a positive effect on demand for usage; that is, after the concessions were granted in 1993, the demand

[4] The coefficient of GDP is negative but we expect it to be positive since demand for access should increase as per capita GDP rises.

for TOT's local call services rises by 0.24%. This is an interesting observation. In contrast to access demand, stronger competition after the introduction of concessions enhances the usage demand for TOT services. The reason behind this observation is the network externalities: a TOT customer can phone a subscriber of the rival companies. So, the demand for a company's local services increases as new competitors enter the market.

The price elasticity of local calls is statistically insignificant. Therefore, we can interpret that the price of local calls has no substantial effect on the demand for local calls. This interpretation is also consistent with the cross-subsidisation between local calls and long distance calls in developing countries, including Thailand. The price of local calls is suppressed to a very low level in these countries, often below marginal cost, and customers are insensitive to the local call price.

4.3. Error correction mechanism testing

The error correction mechanism model is a short-run relationship, which includes lagged dependent and independent variables in a form of difference equation. The ECM equations of access demand and usage demand in the short-run are, respectively,

$$\begin{aligned}\Delta \mathrm{NS}_t = {} & b_1\,\Delta \mathrm{NS}_{t-1} + b_2\,\Delta \mathrm{CF}_{t-1} + b_3\,\Delta \mathrm{MC}_{t-1} + b_4\,\Delta \mathrm{GDP}_{t-1} \\ & + b_5\,\Delta \mathrm{DEN}_{t-1} + b_6\,\Delta \mathrm{BO}_{t-1} + \lambda u_{t-1} + \varepsilon_t,\end{aligned} \tag{10}$$

$$\begin{aligned}\Delta \mathrm{LOC}_t = {} & c_1\,\Delta \mathrm{LOC}_{t-1} + c_2\,\Delta \mathrm{PR}_{t-1} + c_3\,\Delta \mathrm{GDP}_{t-1} \\ & + c_4\,\Delta \mathrm{DEN}_{t-1} + c_5\,\Delta \mathrm{LET}_{t-1} + \lambda u_{t-1} + \varepsilon_t,\end{aligned} \tag{11}$$

where u_{t-1} is the disequilibrium error or short-run relationship, ε_t the white noise and $\lambda \in (0, 1)$ the adjustment parameter.

There is an issue of choosing the lag lengths in an ECM model. Two techniques can be used to decide the lag length. First, determining the lag length by minimizing Akaike Information Criterion (AIC) or Schwarz Bayesian Criterion (SBC). Second, choosing an arbitrary lag length and then eliminating statistically insignificant lags by using Hendry's general specific modelling. This paper adopts the SBC method. The relationship between the cointegration and ECM is conversely affected.[5] The cointegration implies the existence of ECM and vice versa.

In the previous subsection, we found that the long-run relationship between access demand and independent variables exists. Thus, it can be expected that the short-run (ECM) relationship exists as well. Table 7 confirms this and shows that the number of existing subscribers is significant at 1% level while per capita GDP is significant at 5% level. A 1% growth in telephone density increases the demand for access by 3.32% in the short-run. Thus, the short-run access externality exists in Thai telecommunications. A 1% increase in per capita GDP increases access demand by 10.30% in the short-run. This is consistent with a previous study in Spain by Andres and Amaral (1998). On the other hand, connection fee, monthly fee and bond are insignificant at 5% or even 10% level. This implies that the decision to apply for a telephone line in the short-run in Thailand does not

[5] See Granger (1969).

Table 7
ECM equation for access demand: 1973–2000

Dependent variable: ΔNS		
Independent variable	Coefficient	*T*-ratio
Constant	−140.49	−3.9914*
ΔCF	−0.4468	−1.0214
ΔMC	0.3085	0.4040
ΔGDP	10.3027	3.0293**
ΔDEN	3.3187	4.7079*
ΔBO	1.0544	0.3501
D	−1.1412	−3.2271**
ECM(−1)	−0.241	−7.2192*
R^2	0.8992	
Adjusted R^2	0.7481	
DW	2.9226	
F-statistic *F*(11, 14)	8.1158	

Note: The error correction mechanism equations are based on the following as: $\Delta Y_t = b_1\,\Delta Y_{t-1} + b_2\,\Delta X_{t-1} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.
* Significant at 1% level.
** Significant at 5% level.

depend on these variables. Since the average waiting time for access is 3 years, this insensitivity seems plausible. If clients want to have a new telephone line in the next few years, they have to apply for it now. This is in contrast to developed countries where connection takes less than a week for a new telephone line.

The dummy variable is statistically significant at 1% level. As the data in Table 7 show, the demand for the access to TOT network in the short-run decreased by 1.14% after TA and TT&T were allowed to operate in 1993. Therefore, it again confirms that stronger competition dose reduce the demand for the access to a telecommunications company's existing network.

The ECM model can also illustrate how fast variables change or adjust from a short-run equilibrium into a long-run equilibrium. ECM(−1) of access demand in the short-run shown in Table 7 has a correct sign (negative) and is significant at 1% level. The coefficient is −0.241, which means that within a year about 25% has been adjusted towards the long-run equilibrium. In other words, it takes about 4 years for variables in short-equilibrium to be adjusted into the long-run equilibrium.

As we have seen from the previous subsection, the residuals of the usage demand model reject the null hypothesis, implying that there is a long-run relationship between the dependent and independent variables in usage demand. So we can expect that there exists a short-run relationship between these variables. As Table 8 confirms, the price of a local call (PR) and the price of posting a letter (LET) are of positive sign, which means that if these prices increase, the local call usage will increase. The positive sign of local call price is in contrast to economic intuition. But the effects of both prices on the demand for usage are statistically insignificant and therefore are negligible. The per capita GDP variable is statistically significant at 10% level in the short-run. A 1% rise in income increases usage demand by 0.30%. The telephone density has a positive effect on demand for local calls at

Table 8
ECM Equation for usage demand: 1973–2000

Dependent variable: ΔLOC		
Independent variable	Coefficient	*T*-ratio
Constant	1.9041	2.5123**
ΔPR	0.024	0.4151
ΔGDP	0.2941	1.8448***
ΔDEN	0.2642	3.4251*
ΔLET	0.0221	0.4628
D	0.0972	2.4095**
ECM(−1)	−0.3976	−7.6134*
R^2	0.7993	
Adjusted R^2	0.7391	
DW	2.6504	
F-statistic *F*(6, 19)	13.2771	

Note: The error correction mechanism equations are based on the following as: $\Delta Y_t = b_1\,\Delta Y_{t-1} + b_2\,\Delta X_{t-1} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.
* Significant at 1% level.
** Significant at 5% level.
*** Significant at 10% level.

1% significance level. A 1% increase in telephone density increases usage demand by 0.26%. So, the usage externality does exist. Once again, the dummy variable has a positive effect on the demand for usage. One year after the concessions were granted, the demand for usage increases by 0.10%. Thus, the effect of the entry of new competitors on usage demand is opposite to the effect on access demand: the entry increases the demand for the incumbent's local call services.

Table 8 also shows ECM(−1) has a correct sign (negative) and is significant at 1% level. The adjustment coefficient is −0.398; i.e., about 40% is adjusted towards the long-run equilibrium within a year. Comparing usage demand with access demand, the adjustment coefficient (ECM(−1)) of access demand is smaller. Therefore, the adjustment from the short-run to the long-run in usage demand is faster than in access demand.

4.4. Causality

Causality relationship is important as it specifies the direction of the effect of one variable on other. The causality can be tested by the Granger and Modified Sims tests. These tests are very similar.[6] This subsection investigates causality between a pair of variables by adopting the Granger causality method, taking its advantage of being acceptable in small samples (Guilkey & Salemi, 1982). The Granger causality in access demand and usage demand can be formed in unrestricted and restricted models as follows:

- Unrestricted model of access demand

$$\Delta \mathrm{NS}_t = a + b_1\mathrm{NS}_{t-1} + b_2\mathrm{NS}_{t-2} + c_1\mathrm{CF}_{t-1} + c_2\mathrm{CF}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{12}$$

[6] For more details, see Guilkey and Salemi (1982).

If $c_1 = c_2 = \lambda = 0$, then CF does not Granger cause NS. The reverse effects on access demand can be written as:

$$\Delta \mathrm{CF}_t = a + b_1 \mathrm{CF}_{t-1} + b_2 \mathrm{CF}_{t-2} + c_1 \mathrm{NS}_{t-1} + c_2 \mathrm{NS}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{13}$$

If $c_1 = c_2 = \lambda = 0$, then NS does not Granger cause CF. Similar processes are applicable to MC, GDP, DEN or BO.

- Unrestricted model of usage demand

$$\Delta \mathrm{LOC}_t = a + b_1 \mathrm{LOC}_{t-1} + b_2 \mathrm{LOC}_{t-2} + c_1 \mathrm{PR}_{t-1} + c_2 \mathrm{PR}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{14}$$

If $c_1 = c_2 = \lambda = 0$, then PR does not Granger cause LOC. Reverse effects on usage demand can be written as:

$$\Delta \mathrm{PR}_t = a + b_1 \mathrm{PR}_{t-1} + b_2 \mathrm{PR}_{t-2} + c_1 \mathrm{LOC}_{t-1} + c_2 \mathrm{LOC}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{15}$$

If $c_1 = c_2 = \lambda = 0$, then LOC does not Granger cause PR. Similar processes are applicable to GDP or DEN or LET.

The Granger causality hypothesis can be tested by the following statistic:

$$\frac{(\mathrm{SSR_R} - \mathrm{SSR_U})/h}{\mathrm{SSR_U}/(n - k_\mathrm{U})} = F(h, n - k_\mathrm{U}), \tag{16}$$

where n is the sample size, k_U is the number of parameters in the unrestricted equation, k_R is the number of parameters in the restricted equation, h is the number of restrictions being imposed ($h = k_\mathrm{U} - k_{\mathrm{R},}$), $\mathrm{SSR_U}$ is the sum of squared residuals of the unrestricted equation and $\mathrm{SSR_R}$ is the sum of squared residuals of the restricted equation.

The restricted demand model can be written similarly to (12)–(15) but eliminating the independent variables.

- Restricted model of access demand

$$\Delta \mathrm{NS}_t = a + b_1 \mathrm{NS}_{t-1} + b_2 \mathrm{NS}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{17}$$

Reverse effects on access demand is:

$$\Delta \mathrm{CF}_t = a + b_1 \mathrm{CF}_{t-1} + b_2 \mathrm{CF}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{18}$$

Table 9
The relationship between new subscribers and connection fee (monthly charged, income, telephone density and bond): access demand

1 year lag	F-statistic(1, 25)
CF causes NS	2.6526
MC causes NS	1.688
GDP causes NS	10.7146*
DEN causes NS	11.14*
BO causes NS	6.7**

Note: The Granger causality can be formed as follows: $\Delta Y_t = a + b_1 \Delta Y_{t-1} + b_2 Y_{t-2} + c_1 X_{t-1} + c_2 X_{t-2} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.

* Significant at 1% level.

** Significant at 5% level.

Table 10
The relationship between new subscribers and connection fee (monthly charged, income, telephone density and bond): reverse effects of access demand

1 year lag	F-statistic(1, 25)
NS causes CF	4.9082**
NS causes MC	5.457**
NS causes GDP	6.7857**
NS causes DEN	5.497**
NS causes BO	2.7621

Note: The Granger causality can be formed as follows: $\Delta Y_t = a + b_1\,\Delta Y_{t-1} + b_2\,Y_{t-2} + c_1\,X_{t-1} + c_2\,X_{t-2} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.
** Significant at 5% level.

Table 11
The relationship between local calls and price of local calls (income, telephone density and price of posting a letter): usage demand

1 year lag	F-statistic(1, 25)
PR causes LOC	0.6343
GDP causes LOC	10.577*
DEN causes LOC	11.459*
LET causes LOC	0

Note: The Granger causality can be formed as follows: $\Delta Y_t = a + b_1\,\Delta Y_{t-1} + b_2\,Y_{t-2} + c_1\,X_{t-1} + c_2\,X_{t-2} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.
* Significant at 1% level.

- Restricted model of usage demand

$$\Delta \mathrm{LOC}_t = a + b_1 \mathrm{LOC}_{t-1} + b_2 \mathrm{LOC}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{19}$$

Reverse effects on usage demand is:

$$\Delta \mathrm{PR}_t = a + b_1 \mathrm{PR}_{t-1} + b_2 \mathrm{PR}_{t-2} + \lambda u_{t-1} + \varepsilon_t. \tag{20}$$

For access demand, the results in Tables 9 and 10 show that only income and the number of new subscribers, and the number of existing telephone subscribers and the number of new subscribers have a two-way relationship (Granger causality and reverse Granger

Table 12
The relationship between local calls and price of local calls (income, telephone density and price of posting a letter): reverse of usage demand

1 year lag	F-statistic(1, 25)
LOC causes PR	2.0242
LOC causes GDP	8.5216*
LOC causes DEN	10.5528*
LOC causes LET	0.7903

Note: The Granger causality can be formed as follows: $\Delta Y_t = a + b_1\,\Delta Y_{t-1} + b_2\,Y_{t-2} + c_1\,X_{t-1} + c_2\,X_{t-2} + \lambda u_{t-1} + \varepsilon_t$, where Δ denotes the first-order difference of the variables.
* Significant at 1% level.

causality). For other variables, we find that bond causes the number of new subscribers, and the number of new subscribers causes connection fee and monthly charge.

For usage demand, Tables 11 and 12 show Granger causality and reverse Granger causality between income (GDP) and the number of local calls (LOC), and between the number of existing telephone subscribers (DEN) and LOC. LOC causes GDP and LOC causes DEN. All coefficients are statistically significant at 1% level.

5. Concluding remarks

How does the entry of new firms change demand? Our empirical findings demonstrate that the entry of new competitors reduces the demand for the access to the incumbent's telecommunications network but increases the demand for its local call services in the long-run and in the short-run as well. After the concession year (1993), the demand for the access to TOT's network decreases almost 1% in the long-run and 0.5% in the short-run, due to two new firms TA and TT&T entering into the market. However, the entries have increased the demand for TOT's local call services both in the long-run and in the short-run because the subscribers of TOT can make (receive) a phone call to (from) TA and TT&T subscribers.

The number of existing subscriber has a statistically significant effect on access demand and usage demand in both long-run and short-run. This demonstrates that there are access externality and usage externality in Thai telecommunications. In other words, the larger number of existing subscribers attracts more new subscribers and makes them use more local call services.

Our empirical analysis also shows that the majority of estimates of the short-run coefficients in absolute terms are smaller than the long-run coefficients even though they are likely to be statistically insignificant. This implies that customers do not overreact to a change in the short-run. On the other hand, for access demand, the coefficients of income and the number of existing telephone subscribers in the short-run are greater than their counterparts in the long-run. These results suggest that consumers initially overreact to the changes in their income and to the increase in other subscribers. Usage demand moves from the short-run towards the long-run faster than access demand. There is about one half and one quarter of adjustment within 1 year in usage demand and access demand, respectively.

Income has a statistically significant effect on both long-run and short-run usage demands and on short-run access demand but its effect on long-run access demand is statistically insignificant. The effects of connection fee are statistically significant only in the long-run. The price elasticity of local calls is statistically insignificant in both long-run and short-run, due to the cross-subsidisation between local calls and long distance calls in Thai telecommunications. In causality analysis, we also find that both telephone density and income cause access demand and usage demand, vice versa.

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