

Helmut-Schmidt-Universität
Universität der Bundeswehr Hamburg
University of the Federal Armed Forces Hamburg

Fächergruppe Volkswirtschaftslehre
Department of Economics

Discussion Paper No.
December 2004

33

Estimating Demand Elasticities for Mobile Telecommunications in Austria

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DISTRIBUTION STATEMENT A
Approved for Public Release
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20060508019

REPORT DOCUMENTATION PAGE

Form Approved OMB No. 0704-0188

Public reporting burden for this collection of information is estimated to average 1 hour per response, including the time for reviewing instructions, searching existing data sources, gathering and maintaining the data needed, and completing and reviewing the collection of information. Send comments regarding this burden estimate or any other aspect of this collection of information, including suggestions for reducing this burden to Washington Headquarters Services, Directorate for Information Operations and Reports, 1215 Jefferson Davis Highway, Suite 1204, Arlington, VA 22202-4302, and to the Office of Management and Budget, Paperwork Reduction Project (0704-0188), Washington, DC 20503.

1. AGENCY USE ONLY (Leave blank)

2. REPORT DATE

December 2004

3. REPORT TYPE AND DATES COVERED

Discussion Paper

4. TITLE AND SUBTITLE

Estimating Demand Elasticities for Mobile Telecommunications in Austria

5. FUNDING NUMBERS

6. AUTHOR(S)

Ralf Dewenter and Justus Haucapf

7. PERFORMING ORGANIZATION NAME(S) AND ADDRESS(ES)

Universitaet fuer der Bundeswehr Hamburg
Holstenhofweg 85
D-22043 Hamburg GERMANY

8. PERFORMING ORGANIZATION
REPORT NUMBER

9. SPONSORING/MONITORING AGENCY NAME(S) AND ADDRESS(ES)

Universitaet fuer der Bundeswehr Hamburg
Holstenhofweg 85
D-22043 Hamburg GERMANY

10. SPONSORING/MONITORING
AGENCY REPORT NUMBER

Discussion Paper No. 33
December 2004

11. SUPPLEMENTARY NOTES

Text in English, 19 pages, 26 references, 5 tables.

12a. DISTRIBUTION/AVAILABILITY STATEMENT

Public Release. Copyrighted.

12b. DISTRIBUTION CODE

ABSTRACT (Maximum 200 words)

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14. SUBJECT TERMS

ISL, German, Mobile telephony, Price elasticities, Unbalanced panel data, Dynamic panel analysis

15. NUMBER OF PAGES

16. PRICE CODE

17. SECURITY CLASSIFICATION
OF REPORT

UNCLASSIFIED

18. SECURITY CLASSIFICATION
OF THIS PAGE

UNCLASSIFIED

19. SECURITY CLASSIFICATION
OF ABSTRACT

UNCLASSIFIED

20. LIMITATION OF ABSTRACT

UL

NSN 7540-01-280-5500

Standard Form 298 (Rev. 2-89)
Prescribed by ANSI Std. Z39-18
298-102

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Abstract

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JEL-Classification: C23, L13, L96.

Keywords: Mobile telephony, Price elasticities, Unbalanced panel data, Dynamic panel analysis.

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AQ F06-07-6418

1 Introduction

While mobile telecommunications markets have largely been left unregulated in Europe until recently, they have started to draw regulators' and policy makers' attention in more recent times (see, e.g., European Commission, 2004). Apart from more narrowly defined issues such as mobile number portability, mobile termination rates, and international roaming, an area of concern has also been the general competitiveness of the mobile telecommunications industry. For example, Ofcom and the UK Competition Commission have argued that the mobile telecommunications industry as a whole is not subject to effective competition, due to the oligopolistic industry configuration (see Competition Commission, 2003). Since there is only a limited amount of radio spectrum available and as the fixed and common costs associated with mobile network investments are relatively high, mobile telecommunications markets have been argued to be natural oligopolies (see Gruber, 2001; Valletti, 2003). Accordingly, concerns have been voiced by various regulatory and competition authorities about competition in mobile telecommunications markets (or, more precisely, the lack thereof), especially with respect to the potential for collusive behavior.

In fact, as oligopolistic industries are often prone to collusion, it is important to analyze the market participants' conduct in these industries in more detail. Apart from factors such as the number of operators, barriers to entry, product differentiation, the firms' cost structures, and market transparency, one indicator for the firms' incentives to engage in collusive behavior is the market's and the firms' demand elasticity (see, e.g., Carlton and Perloff, 2004). If the market demand is relative inelastic, the firms' rewards from engaging in collusive conduct are relatively high, as prices can be increased without losing much custom. In contrast, a relatively elastic demand im-

plies that the additional profit from collusion is relatively low. In addition, a high *firm-specific* elasticity of demand implies that deviating from a collusive agreement is relatively profitable (as a small price decrease generates a relatively high increase in the quantity sold) so that collusion is more likely to break down due to the “cheating problem”.

Moreover, demand elasticities have also been the subject of debate in various hearings on price regulation and the allocation of common costs, for which demand elasticities play an important role (e.g., for Ramsey pricing). Hence, as demand elasticities have become a subject of debate, the number of studies that estimate demand elasticities has also been increasing, some of which are reviewed below. This paper adds to this growing literature. However, in contrast to most other research which is based on aggregate market data we had access to firm-specific data from three different competitors in the Austrian mobile telecommunications market. These three firms who are the three largest mobile operators in Austria account for around 90% of the Austrian market for mobile telecommunications. In our analysis, we will use firm specific data on prices and quantities for these firms and analyze price elasticities for mobile telecommunications services. Our main result is that both static and dynamic panel data approaches lead to consistent results according to which the demand for mobile telecommunications services is relatively elastic in Austria (when compared to studies from other countries).

The remainder of the paper is now organized as follows: The next section provides an overview over empirical studies of demand elasticities in telecommunications markets, before section 3 offers some basic facts on the Austrian mobile telecommunications market and its historical development. In section 4 we describe that data used and present and discuss two different empirical

specifications for the demand equations with respect to panel data analysis. Finally, our main results and conclusions are summarized in section 5.

2 Brief Review of the Empirical Literature

Empirical studies on demand elasticities for mobile markets have, in principle, been using two different approaches. While the first approach is based on highly aggregated data on country or regional level, a second method to measure price elasticities relies on individual or survey data of consumer behavior.

Independently of whether aggregated or individual data has been used most studies have found relatively moderate price elasticities. Hausman (1999) and (2000), for example, finds a price elasticity of access to mobile services of -0.51, using aggregate data on 30 U.S. markets for the period 1988 to 1993. Analyzing the price elasticity of subscription using data on 64 different countries Ahn and Lee (1999) estimate an average elasticity of -0.36.

Summarizing the results from different studies by *DotEcon*, *Frontier Economics* and *Holden Pearmain*, the UK Competition Commission (2003) reports own-price elasticities of mobile subscriptions between -0.08 and -0.54. For mobile calls, own-price elasticities between -0.48 and -0.62 have been measured. In a study on the Australian mobile market *Access Economics* reports a price elasticity of -0.8 (see Competition Commission, 2003).

Rodini et al. (2002) analyze the substitutability between fixed and mobile access in the U.S. and, for this purpose, estimate own and cross-price elasticities. Using survey data on telephony services Rodini et al. (2002) find own-price elasticities of -0.43 for mobile subscription rates. Furthermore, a total elasticity of -0.6 is estimated for the access *and* usage price.

A quite different approach to analyze conduct in mobile markets has been carried out by Parker and Röller (1997) and Grzybowski (2004). Both studies apply structural models in order to examine the competitive behavior of mobile operators. While Parker and Röller (1997) find an own-price elasticity of -2.5 using data on the United States covering the period 1984-1988, Grzybowski (2004) finds rather moderate elasticities for the EU countries in 1998-2002, ranging from -0.2 to -0.9.

In order to analyze the price elasticities of demand for the Austrian mobile telecommunications market, and in contrast to existing studies, we (i) use data on firm specific tariffs and (ii) apply both static and dynamic panel techniques. By these means we are able to distinguish between short- and long-run elasticities and to distinguish between consumer behavior on the firm level.

3 The Austrian Market for Mobile Telecommunications

In contrast to most other European countries, the Austrian market for mobile telecommunications services has only been opened to competition relatively late, namely in 1996. While mobile telecommunications services have been offered since 1979, *Mobilkom Austria*, the former state-owned enterprise, was allowed to operate as a monopoly provider until October 1996 when *max.mobil* (now *T-Mobile Austria*) entered the market. Then two years later, *Connect Austria* (now *One*) was granted a license, and in 2000 a fourth carrier (*tele.ring*) entered the market (for details see Kruse et al., 2004).

However, even though deregulation and liberalization have been introduced rather late, Austria is nowadays one of the few European countries

with four GSM-1800 networks that provide almost full coverage.¹ Moreover, further entry is likely to occur as two potential entrants, *Hutchison 3G* and *3G Mobile (Telefonica)*, were - apart from the incumbents *Mobilkom Austria*, *T-Mobile*, *One* and *tele.ring* - successful in the Austrian UMTS license auction in 2000. Today, the Austrian mobile telecommunications industry is considered to be one of the most competitive ones in Europe (see WIK, 2002; Grzybowski, 2004).

Comparing the market shares of the four "incumbent" carriers, we see that *Mobilkom's* market share has declined significantly, while the other operators' market shares have increased (see Figure 1). In April 2004 the market share of the former state-owned monopolist, *Mobilkom*, was 42.67% (*T-Mobile* 27.45 and *One* 20.16) but, more interestingly, the share of *tele.ring* has increased from 2.65% in 2001 to 9.32% in 2004 (see Figure 2). As can also be seen from Figure 1, the shares of *T-Mobile*, the first competitor, have decreased following the market entry of *One* and *tele.ring*.

*** Insert Figure 1 about here ***

*** Insert Figure 2 about here ***

¹Other European countries with four mobile network operators are Finland, Denmark, Germany, Italy, or the UK.

4 Empirical Analysis

4.1 Data and Empirical Specification

Data

To analyze short- and long-run elasticities, we use monthly data on mobile telephone traffic in Austria over the period January 1998 to March 2002. The data on prices, quantities and networks' subscriber bases has been provided by the three largest Austrian mobile operators: *Mobilkom*, *One* and *T-Mobile*. In total we have information on 37 different tariffs from the three operators mentioned (see Table 1 for a summary of the data). In addition, information on the price index has been gathered from official statistics of Austria.

***** Insert Table 1 about here *****

The variable 'total number of outgoing minutes' measures the monthly traffic (Q) for each tariff. The variable consists of the sum of all outgoing calls, independently of the exact type of service (except for SMS or data services). Hence, the variable represents an aggregate over various services (such as on-net, off-net, mobile to fixed, and international calls) within a specific tariff. Also, to analyze price elasticities we have calculated the average traffic per subscriber (q) by building the ratio of $Q/TNet$, where $TNet$ is the tariff specific subscriber number.

Furthermore, we had to use an average call rate (P), which has been constructed by dividing the total revenue for each tariff by the total number of outgoing minutes for that tariff. While mobile markets are characterized by price differentiation between peak and off-peak times, more detailed data has not been available. To obtain real prices P has been deflated by the Austrian consumer price index. Furthermore, information on the networks' (total)

subscriber bases (*TNet* and *Net*), and time and firm dummies, respectively, have been used as explanatory variables. All variables but the dummies are in logarithmic forms (see Table 2 for descriptive statistics of the variables).

***** Insert Table 2 about here *****

In order to take a closer look on consumer behavior in the Austrian mobile telephone market, we have used four different samples: The first includes information on all of the three providers, whereas samples 2 to 4 only include data on the single firms' tariffs.

Specifications

According to the literature typically two different empirical specifications are used to analyze telecommunications demand. The first is a simple log-linear specification of an iso-elastic demand function of the form $q = p^\eta$, where η denotes the (long-run) own price elasticity of demand. Taking into account the panel structure of our data, one can derive an adequate specification as

$$\ln q_{it} = \alpha_{it} + \beta_1 \ln p_{it} + \sum_{k=2}^K \beta_k x_{it,k} + \varepsilon_{it}, \quad (1)$$

where q_{it} is the average quantity demanded for tariff i at time t , p_{it} is the respective average price, $x_{it,k}$'s are k additional explanatory variables, ε_{it} is an error variable, and α and the β 's are parameters to be estimated. Assuming that α_{it} is fixed over time, but differs with cross-section units, equation (1) can be estimated using the fixed effects. Assuming that α_{it} can be decomposed into a *common* constant (α) and a unit specific random variable (ν_i) so that $\alpha_{it} = \alpha + \nu_i$, equation (1) can be estimated with the random effects model.

However, neglecting a possible endogeneity of one or more of the covariates (in our case prices are, for example, likely to be endogenous) the use of

ordinary fixed effects or random effects models lead to biased results. For this purpose, panel instrumental variable methods are appropriate to account for endogeneity problems. In case of fixed effects, the 'two-stage least-squares within' estimator is applied. Computing a random effects model, we use the two-stage least-squares one-way error component model using feasible generalized least squares (FGLS). Moreover, since our panel is unbalanced, we apply the consistent estimator by Baltagi and Chang (1994) of the variance components.

A second typical approach towards the estimation of demand equations can be derived from the Houthakker-Taylor model, which takes possible path dependencies of consumption into account (see Houthakker and Talor, 1970). In this model, demand at time t can be expressed by $q_t = q_{t-1}^\phi p_t^\eta$ so that the short-run price elasticity is determined by η , whereas the long-run price elasticity is equal to $\eta/(1 - \phi)$. Taking into account the panel structure of the data, the following specification can be derived:

$$\ln q_{it} = \gamma_{it} + \delta_1 \ln q_{it-1} + \delta_2 \ln p_{it} + \sum_{k=3}^K \delta_k x_{it,k} + \varepsilon_t. \quad (2)$$

However, the use of usual panel data techniques leads to biased results in this case, not only because prices are endogenous, but also because of the lagged endogenous variable q_{it-1} . Hence, a dynamic panel analysis is appropriate. Applying a first difference transformation on equation (2) leads to

$$\Delta \ln q_{it} = \delta_1 \Delta \ln q_{it-1} + \delta_2 \Delta \sum_{k=3}^K \delta_k \Delta \ln x_{it,k} + \Delta \varepsilon_t, \quad (3)$$

which can be estimated consistently using a GMM approach suggested by Arellano and Bond (1991). Both lagged dependent variable and endogenous variables can be instrumented by lagged values. Furthermore, Arellano and Bond (1991) provide a heteroscedastic robust estimator. Since the GMM

estimator is not consistent under the existence of second order autocorrelation, Arellano and Bond have derived an adequate test of autocorrelation. Furthermore, a Sargan test of over-identifying restrictions on the number of instruments can be applied.

4.2 Results

To prevent the hazard of spurious regressions we first have applied panel unit root tests for the variables used in this study. While the pooled panel estimator yields consistent estimates even if some of the variables are integrated of order one (or higher) and also independently of the existence of a cointegrating relation (see Phillips and Moon, 1999), in this case only long-run relationships can be analyzed using the integrated variables. Using lagged variables is not appropriate though. Thus, the first step is to test the samples against unit roots using a non-parametric approach suggested by Maddala and Wu (1999).² The authors suggest to choose a test-statistic by Fisher (1932) where the p-values of single unit root tests (π_i) from each cross-sectional unit $i = 1 \dots N$ are used to calculate the test statistic $p_\lambda = -2 \sum_{i=1}^N \ln(\pi_i) \sim \chi_{2N}^2$. In order to account for possible autocorrelated and heteroscedastic errors, Phillips-Perron tests (see Phillips 1987 and Phillips and Perron 1988) have been applied to each time series to calculate respective p-values.³

*** Insert Table 3 about here ***

At first, specification (1) has been tested using data on Austrian mobile telephony, by regressing q_{it} on p_{it} , $TNet$, time and firm dummies, and an

²While there is a set of unit root tests for panel data, tests such as those proposed by Im, Pesaran and Shin (1997) are not appropriate here since the samples used are unbalanced.

³The number of lags used for each series has been calculated by $l = \text{int}(4 \cdot (T/100)^{(2/9)})$ (see Newey and West, 1987), where T is number of observations.

error term, applying fixed effects and random effects IV estimators. Since prices are likely to be endogenous, we have used their first and second lags as instruments to account for this possible endogeneity.⁴

As can be seen from Table 4, price elasticities strongly vary from -0.19 to -3.56, depending on the sample used. However, in nearly each case prices are found to have a negative and statistically significant influence on the average demand for mobile telephony. Regarding the Hausman specification tests, in all samples (except the Firm 2 sample) the random effects model can not be rejected.⁵ Moreover, the results for the network's subscriber bases are ambiguous even if most equations show a negative impact of $\ln TNet$. Since q_t is defined as $q_t \equiv Q_t/TNet$, a negative coefficient for $\ln TNet$ means that the average quantity consumed per subscriber is decreasing with an increasing subscriber bases. Hence, we conclude that there is only weak (if any) evidence for network effects in Austrian mobile telephony.⁶

*** Insert Table 4 about here ***

In order to test the second specification according to equation (3) and to distinguish between short- and long-run elasticities, dynamic panel techniques have been applied using the approach introduced by Arellano and Bond (1991). Again, we have used four different samples to test for the consistency of our results. Since regressions in Table 5 are first difference

⁴Hausman-Wu tests (see Greene, 2003) have been applied to test for possible endogeneity of current prices. However, in any of the results we present below the tests failed to reject the null hypothesis of exogeneity.

⁵Using lagged prices of order 3 and higher leads to statistically significant and higher coefficients in absolute value. However, using lags of higher orders also reduces the sample significantly, therefore, we do not report the results.

⁶Not only $TNet$, but also the variable Net has been used to account for possible network effects. However, neither use of the tariff specific subscriber base nor use of the total network size has produced evidence for positive network effects.

transformations of the Houthakker-Taylor model in logs, the estimated parameters can be also interpreted as elasticities. According to all models, the lagged outgoing traffic is statistically significant and ranges between -0.44 and -0.7. Therefore, demand in the Austrian mobile telecommunications market seems to be path dependent. Interestingly, the short-run price elasticities obtained are consistent over the different models employed, even though our samples differ in size.

The network's subscriber base has a negative impact on the average quantity consumed per customer. Again, there seems to be some evidence that an increase in the number of customers implies a decrease in the average quantity consumed. In fact, considering outgoing call minutes instead of the outgoing traffic per subscriber, the respective parameters are less than unity in each sample. This means, that additional consumers (who are relative late adapters) use their mobile telephone less than the early adapters. According to our analyses, this effect dominates any positive network effect that there may be.

Finally, the constant is statistically insignificant (as expected) as the difference transformation of equation (2) leads to a vanishing γ_{it} .

***** Insert Table 5 about here *****

The long-run elasticities are considerably higher than short-run values and vary between -0.61 and -1.05. In comparison to the other empirical studies (discussed above) these values seem to be appropriate but at the upper end of the range.

***** Insert Table 6 about here *****

The Sargan test of over-identifying restrictions does not reject the null of validity of instruments used. The Arellano-Bond test on second order au-

to correlation does not allow to reject the hypothesis of no-autocorrelation, except for sample "Firm 2", where the null can be rejected at the 8% level of significance.

To summarize, the results provide some evidence that the demand for mobile telephony services is relatively elastic in Austria when compared to other countries.

5 Summary and Conclusions

In this paper we have analyzed the demand for mobile telecommunications services in Austria, using firm level data on different tariffs. Both static and dynamic panel data techniques have been applied in order to estimate short- and long-run price elasticities and the impact of networks's subscriber bases on the demand for mobile calls. As the results indicate, the demand for mobile telecommunications services in Austria appears to be relatively elastic when compared to other countries.

From a competition policy perspective, these results imply that policy makers in Austria have to be less concerned about collusion between mobile operators than in other countries. This result is not only consistent with more qualitative comparative studies that have demonstrated the high degree of competition in the Austrian mobile telecommunications sector (see WIK, 2002), but our results also nicely complement these analyses. Given the regulatory framework that is in place in Austria (involving the regulation of termination fees and mobile number portability), we do not see any obvious reason for any further regulatory intervention or supervision at this point. The intense competition between mobile operators not only makes collusion relatively unlikely, but most of all it benefits consumers.

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A Figures

Figure 1: Mobile Operators' market shares

(April 1998 – April 2004, Source: RTR, 2004)

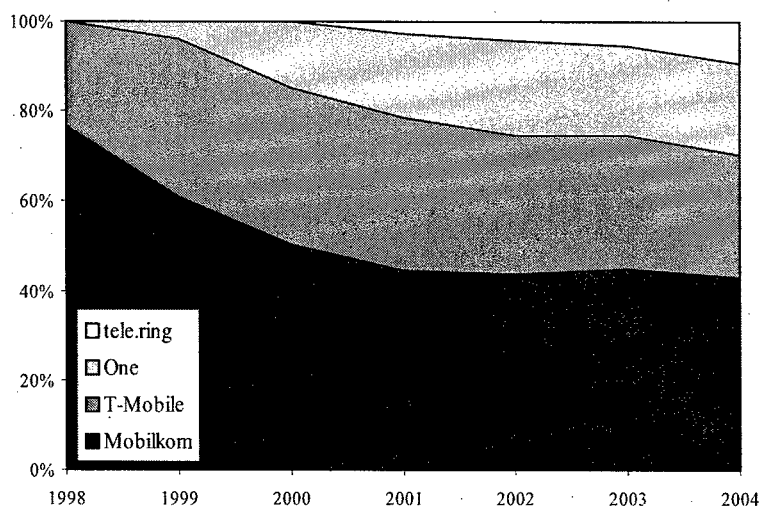
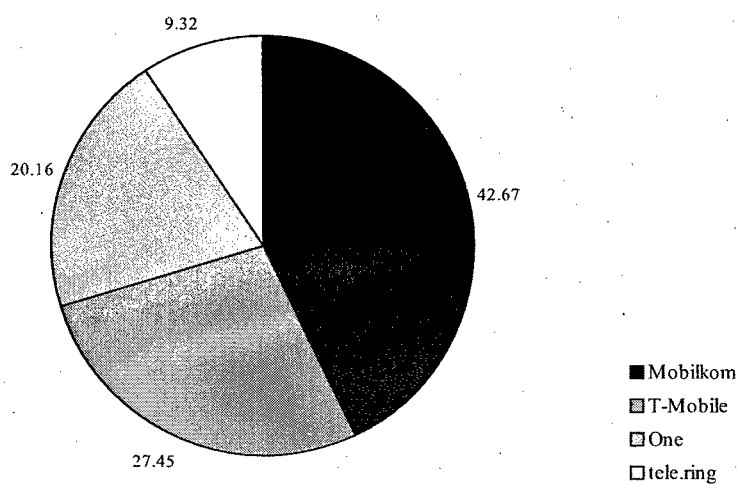


Figure 2: Mobile Operators' market shares

(April 2004, in % Source: RTR, 2004)



B Tables

Table 1: Data

No.	Firm	Tariffs	Min	Max
1	One	17	7/1998	2/2002
2	T-Mobile	6	1/1998	2/2002
3	Mobilkom Austria	14	1/1999	3/2002

Table 2: Descriptive Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
ln Price	960	-0.48	0.82	-3.78	1.78
ln Firm Net	1156	13.17	2.08	4.49	14.79
ln Minutes	1032	14.91	3.02	2.61	20.11
ln TNet	928	9.91	2.88	0.69	13.99

Table 3: Maddala-Wu unit root tests

Sample, $\chi^2_{\alpha, 2N}$	$\ln p_t$	$\ln q_t$	$\ln TNet_t$
All Firms, $\chi^2_{0.001, 80} = 125$	164.87	252.69	242.42
Firm 1, $\chi^2_{0.001, 40} = 73$	95.87	112.62	102.09
Firm 2, $\chi^2_{0.001, 6} = 22.5$	23.07	50.40	84.63
Firm 3, $\chi^2_{0.001, 14} = 36.1$	48.90	89.66	55.70

Table 4: Panel IV estimates of mobile demand

	All Firms		Firm 1		Firm 2		Firm 3	
	FE	RE	FE	RE	FE	RE	FE	RE
$\ln p_t$	-0.5601 (-3.17)	-0.6861 (-4.23)	-0.1934 (-2.28)	-0.3446 (-3.75)	-0.2038 (-0.98)	-2.7815 (-26.23)	-1.3012 (-2.3)	-3.5649 (-2.90)
$\ln TNet$	-0.0711 (-3.61)	-0.0675 (-3.55)	-0.0892 (-6.61)	-0.1057 (-6.63)	0.0815 (1.30)	-0.0872 (-3.80)	-0.9613 (-1.90)	0.2160 (2.98)
Constant	5.0446 (27.22)	4.1525 (11.71)	5.4533 (34.57)	5.4074 (26.34)	3.1225 (5.13)	1.0354 (2.56)	19.1568 (2.57)	2.4495 (2.92)
Time dummies	YES	YES	YES	YES	YES	YES	YES	YES
Wald χ^2 ($H_0 : \alpha_i = 0$)	296288 (0.00)	62.30 (0.00)	367819 (0.00)	158.76 (0.00)	196323 (0.00)	927.37 (0.00)	17.07 (0.19)	43.67 (0.00)
Observations	347	347	239	239	157	157	347	347
No. of Groups	14	14	17	17	6	6	14	14
Hausman Test	17.21 (0.44)		18.19 (0.15)		66.72 (0.00)		14.93 (0.53)	

Note: Heteroscedasticity consistent t-statistics are given in parenthesis.

Table 5: One-step GMM estimates of mobile demand

	All Firms	Firm 1	Firm 2	Firm 3
$\Delta \ln q_{t-1}$	0.4389 (9.92)	0.5943 (10.58)	0.5443 (3.59)	0.6894 (11.97)
$\Delta \ln p_t$	-0.4177 (-4.09)	-0.2489 (-8.83)	-0.3872 (-3.85)	-0.3265 (-2.38)
$\Delta \ln \text{TNet}$	-0.1002 (-2.20)	-0.0237 (-3.28)	-0.0080 (-0.15)	-0.1248 (-1.22)
Constant	-0.0039 (-1.08)	-0.0028 (-1.03)	0.0022 (0.86)	-0.0047 (-1.44)
Time dummies	YES	YES	YES	YES
Sargan Test	590.90	203.20	110.13	334.92
(Prob.)	(0.00)	(0.00)	(0.00)	(0.00)
AR(2) Test	-0.86	-0.99	1.77	-0.77
(Prob.)	(0.39)	(0.32)	(0.08)	(0.44)
Observations	697	205	145	347
No. of Groups	37	17	6	14
long-run elasticity	-0.74	-0.61	-0.85	-1.05

Note: Heteroscedasticity consistent t-statistics are given in parenthesis.

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