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Is Fixed-Mobile Substitution strong enough to de-regulate Fixed Voice Telephony? Evidence from the Austrian Markets.

Paper

Is Fixed-Mobile Substitution strong enough to de-regulate Fixed Voice Telephony? Evidence from the Austrian Markets

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Abstract:

We estimate own-price elasticities for fixed network voice telephony access and (national) calls services for private users and cross-price elasticities to mobile using time series data from 2002-2007 from the Austrian market. Using instrumental variable estimates and taking into account the possibility of cointegration we find that access is inelastic while calls are elastic. We conclude that the retail market for national calls of private users can probably be deregulated due to sufficient competitive pressure from mobile. Access-substitution on the other hand does not seem to be strong enough to justify de-regulation.

Keywords:

Fixed-Mobile Substitution, Market Definition, Hypothetical Monopolist Test, (De-)Regulation, Fixed networks, Voice Telephony

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The views expressed are entirely those of the author and do not necessarily represent those of RTR or Telekom-Control-Kommission (TKK) and WIFO. The usual caveat applies.

1 Motivation

With regard to the ex ante regulation of communications markets, the European Commission publishes a list of relevant markets every few years. This is the so called 'Recommendation on Relevant Markets', which has to be considered by each national regulatory authority (NRA) and therefore is the starting point of the market analyses process. In 2007, the European Commission issued a new Recommendation replacing the Recommendation from 2003.[1] The most substantial change that came along with the 2007 Recommendation concerns the retail markets for voice telephony: According to the new Recommendation, markets for national and international calls (markets number 3-6 in the 2003 Recommendation, see footnote 1) should no longer be subject to ex ante regulation.

The European Commission substantiates the 'non-relevance' of these markets in its new recommendation with reference to the increasing importance of broadband connections and associated technological innovations (most notably, IP-based telephony) on the one hand, and (in part only recently) imposed regulatory instruments on the wholesale level (such as Unbundling, Naked DSL, Wholesale Line Rental, Carrier Selection) on the other hand. However, it is questionable whether these 'intramodal' developments really justify any

[1] See European Commission (2003) for the 'old' recommendation as well as European Commission (2007) for the 'new' recommendation. According to the 'old' recommendation the following fixed voice telephony markets were declared to be relevant (although disprovable by national regulators) at the retail level:

1. Access to public telephone network at fixed locations for residential customers
2. Access to public telephone network at fixed locations for non-residential customers
3. Publicly available local and/or national telephone services provided at a fixed location for residential customers
4. Publicly available international telephone services provided at a fixed location for residential customers
5. Publicly available local and/or national telephone services provided at a fixed location for non-residential customers
6. Publicly available international telephone services provided at a fixed location for non-residential customers

changes of the Recommendation on empirical grounds.[2] On the other hand, in many member states the mobile sector has recently begun to exert increasingly competitive pressures (‘intermodal’) on fixed voice telephony markets (‘Fixed-to-Mobile Substitution’, FMS). In our paper we use time series data for the years 2002-2007 from the Austrian markets to empirically examine the extent of intermodal competition, since this will not only be crucial for the initial analysis stage of market definition but will also determine whether retail regulation is still necessary or not.

FMS became a subject of interest for policy makers and economists as mobile telephony became more widely spread. The New Zealand Commerce Commission (2003) reviews a number of studies primarily from the 1980s and 1990s estimating elasticities for fixed network services for different countries, finding almost consistently inelastic demand for fixed access and usage.[3] The evidence from more recent studies with regard to FMS is as follows: Barros/Cadima (2000) report a negative effect of mobile phone diffusion on the fixed network penetration rate based on 1981-1998 time series data from Portugal. This can be interpreted as evidence of fixed-mobile access substitution. Sung/Kim/Lee (2000) find that a 1% increase in the number of mobile phones reduced fixed connections by 0.1-0.2% in Korea in the 1990s. Gruber/Verboven (2001), on the other hand, conclude that countries with a large fixed network penetration tend to be more advanced in adopting mobile phones based on data from 140 countries from 1981-1995 (but without using price variables). They interpret this as evidence of complementary use. Hamilton (2003) finds that evidence on fixed mobile substitution in Africa (based on 1985-1997 data) is mixed with some evidence of complementarity and some for substitutability (also without using price variables). Sugolov

[2] See European Commission (2008), volume 2, figure 41 for competition on access markets and figures 35-36 for competition on national calls markets. The market share of most incumbents was still very high in 2006 in access as well as in calls markets.

[3] See New Zealand Commerce Commission (2003) and statement of Vodafone New Zealand limited (available at: http://www.comcom.govt.nz/IndustryRegulation/Telecommunications/TelecommunicationsServiceObligations/ContentFiles/Documents/Vodafoneweightedrevenue0.PDF).

(2005) analyses transition economies in the years 1993-2001 and finds that fixed and mobile access might be complements in the beginning but become substitutes as the mobile market matures. Madden/Grant (2004) analyse data from 58 countries for 1995-2000 and report a significant substitution effect between fixed and mobile subscriptions. Based on 2000/2001 survey data from the US, Rodini/Ward/Woroch (2003) find that *second* fixed lines and mobile services are substitutes for one another. Even for second fixed lines, the elasticity is in the inelastic range. Vagliasindini/Güney/Taubman (2006) analyse the relation between fixed and mobile access lines used by businesses in transition economies in 2002. They find evidence for some substitution effects at the country level. Garbacz/Thompson (2007) analyse 1996-2003 data from developing countries and find asymmetric substitution effects: although fixed connections are substitutes in the mobile market, mobile phones may be considered compliments in the wireline market.

Regarding usage substitution, Sung (2003) finds that fixed toll calls and mobile calls are economic substitutes (i.e., ε_{ij}>0) based on 1993-1997 panel data for Korea. The own price elasticity of toll calls is in the inelastic range, however. Ingraham/Sidak (2004), focusing on the elasticity of mobile services in the US for the years 1999-2001, also find that mobile and long-distance calls are substitutes with an own-price elasticity of long distance calls in the elastic range (-1.3). Horvath/Maldoom (2002) find that the ownership of a mobile phone ignificantly reduces the use of fixed lines using UK survey data from 1999, 2000 and 2001. Ward/Woroch (2004) analyse fixed-mobile usage substitution in the US around the year 2000 and find that mobile service is a moderate substitute for wireline usage. They expect that the services will become closer substitutes over time.

While most of the literature differentiates between access and usage substitution, several market segments are usually aggregated. As will be discussed in section 2.1, however, market data and empirical evidence indicate that FMS differs with regard to different market segments, e.g., residential vs. non-residential users, and national vs. international calls. Therefore, it is worth to focus on specific segments when analysing FMS. We focus on a segment, where FMS can be expected to be high: The segment of national calls of private users. Since demand for access and usage is closely linked, we also consider access for private users. As far as we are aware, up-to-date empirical evidence with such focus is still missing. Closing this gap seems to be of primary importance for NRAs on account of the increasing regulatory importance of FMS.

The paper is structured as follows: Section 2 outlines the main hypotheses as regards FMS, for empirical evidence we generally refer to the data of European member states. Section 4 describes the data used in the estimation. The main estimation results are presented and discussed in section 5. Section 6 concludes. The Appendix contains unit root-, cointegration-, and instrument tests.

2 Hypotheses and Conceptual Framework

2.1 Basic Hypotheses and Empirical Evidence

FMS basically suggests an opposing development of volumes in both sectors. Indeed, with regard to the mobile sector (2G and 3G) we observe persistent growth in penetration levels/access lines and call minutes[4] whereas access lines and usage in the fixed line sector have been decreasing steadily for many years.[5] It can be assumed that substitution of traditional fixed line voice telephony services by mobile telephony is the main driver of these developments. However, the influence of mobile competition varies significantly among (EU/OECD) countries.[6] The regulatory relevance of FMS is therefore an issue that has to be dealt with on a country specific level by each NRA, so market definition should be expected to differ across member states. As regards the Austrian market situation we have a specific

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4 European Commission (2008), Volume 1, p. 10.

5 See European Commission (2008), Volume 1, p. 17, Schäfer/Schöbel (2006), pp.6-87, for international case studies or RTR (2008a), p. 141-142, for the Austrian communications market development.

6 See OECD (2009), Figure 3.8, which shows that the share of mobile revenues in total revenues differs among EU countries between 65% in the Slovak Republic, 62% in Austria (ranked second) and 26% in the Netherlands in 2007.

case at hand, since mobile competition is particularly advanced in Austria, where about 75% of voice traffic originated from mobile networks in 2007.[7] In addition, Austria is not only particularly advanced not only with regard to voice FMS bus also with regard to mobile broadband.[8]

On average the decline of fixed line telephony is evident, yet, concerning substitution patterns we have to distinguish between market segments. From the old recommendation – as listed in footnote 1 – we infer that market definition, in principle, has to be analysed in terms of the following dimensions: (i) access versus calls (ii) national calls versus international calls and (iii) residential versus non-residential consumers.

Demand for fixed narrowband access is likely to be significantly much less elastic than demand for calls. This is suggested by past estimates as well as the fact that substitution of access is related to a discrete decision (access or no access at all),[9] whereas calls can be substituted continuously, on a minute-by-minute (or call-by-call) basis if the user has both a fixed and a mobile connection. Of course, the complementarities between demands for access and calls imply that access substitution also triggers complete substitution of calls. As markets mature, both forms of FMS typically tend to increase: Because relative price differences have become smaller during liberalisation, due to technological advances in mobile services and the availability of number portability and, last but not least, because of the recent advent of mobile broadband connections lessening the attractiveness of fixed lines for the purpose of access to the internet. Although these factors contribute to both, FMS relative to access and to calls, the rate of FMS still appears to be much stronger with respect to calls.

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[7] See Kruse (2007), chapter 3.2.3, or Analysys Research (2007).

[8] See European Commission (2009) p. 12 where Austria ranks first among EU countries in mobile broadband penetration.

[9] Two sub-types of access substitution might be distinguished: 'Cut the cord' means that users, who previously had access to both fixed and mobile services, have chosen to give up the fixed line entirely. 'Straight to mobile' means that some mobile-only users (such as students or young household people) never used fixed access before; see ACCC (2008), p. 9. ''

In general, FMS can be expected to be strongest with regard to national calls. It is usually much less pronounced with regard to international calls services simply because per minute prices still remain much more expensive if the call originates from a mobile.[10]

As far as consumer groups are concerned, we typically observe that FMS is most evident among a subset of predominantly residential customers: Those who do not require a fixed line for access to the internet, who are part of a single-person household and therefore do not share the costs of a fixed line with other family members or exhibit a low usage intensity and therefore benefit most from pre-paid mobile offers.[11] Non-residential/business customers, in turn, are usually much more reluctant to FMS because they tend to assign higher importance to remaining fixed network characteristics, such as quality, access to internet and other data services. Office-based enterprises, in particular, may be concerned about reliability and coverage within buildings. Finally, business customers usually have a proportionally higher demand for international calls and some consumers seem to attach less confidence to companies which are available only via mobile numbers.[12] Market data from Austria show that the decline in the number of fixed access lines and fixed network minutes is much smaller for business compared to residential customers (see RTR, 2008b).

These observations motivate our focus on national calls of residential users. This is the segment where FMS is likely to be strongest. Since demand for access and usage is closely linked, we also consider access for residential users.

2.2 Conceptual Framework of Market Definition[13]

According to the current regulatory framework for telecommunications regulation in the EU, the underlying methodology of market delineation is to be based on economic principles.[14]

[10] See Stumpf (2007), p. 4.

[11] See Stumpf (2007), pp. 6-7.

[12] See ACMA (2008), p. 26.

[13] This subsection draws on the presentation in Briglauer (2008), p. 316-317.

With respect to methodological aspects, the hypothetical monopolist test (HMT) has become standard, and as such, also part of the ex ante communications framework.[15]

The HMT asks whether a small but significant non-transitory price increase from the competitive level would be profitable for a hypothetical monopolist of a particular product. 'Small but significant' has been interpreted as 5-10%[16] in the past while 'non-transitory' should be interpreted as a period of 1-2 years.[17]

Traditionally, market definition was thought to be based on the degree of cross-price elasticities of demand, measuring price and quantity reactions of two (potential) substitute products in relative terms, i.e.

$$(1) \qquad \varepsilon_{ij} = \frac{\partial D_j * p_i}{\partial p_i * D_j} ,$$

with demand $D(p)$, and i, j denoting the initial product and substitute candidate, respectively. However, there is no theoretical threshold value, indicating how high cross-price elasticity of demand needs to be, which allows us to conclude that two goods belong to the same market. Secondly, cross-price elasticities focus by definition only on the potential effect of individual substitute products, whereas the collective competitive restriction of all potential substitutes matters in determining whether a 5-10% increase of the HM is profitable or not. Generally, the elasticity of residual demand (ε_i) completely summarizes a firm's market power, i.e. the ability to raise prices above competitive levels and pertain monopoly profits. The higher the

14 The legal basis for this is the Framework Directive (2002/21/EC) Art. 15 (3).

15 Long time before the HMT became part of the sector specific regulations framework this approach to market definition was introduced by the US Department of Justice (1982 Merger Guidelines, revised in 1992, 1997) which is currently being used by antitrust authorities worldwide, see for instance Bishop/Walker (1999), table 3.1. The HMT is sometimes also called SSNIP test since it explores the consequences of a (hypothetical) Small but Significant Non-transitory Increase in Price.

16 The US Department of Justice refers to a 5% increase whereas the SMP-Guidelines (§ 40) refer to a 5-10% increase in price.

17 The US Department of Justice suggest a response period of one year. In ex ante framework the typical minimum period of market analyses suggest a time frame of two years.

elasticity of residual demand, the lower is the potential market power of the firm under consideration. However, for market definition purposes one does not refer to an existing or specific firm, but to the HM firm.

In carrying out the iterative HMT procedure cross-price elasticities constitute the most adequate method of ranking the closest substitute products. $\varepsilon_{ij} = \varepsilon_{ji}$ does not hold in general, so in practice the particular case under consideration to which the initial price increase is applied determines which cross-price elasticity is meant. For instance, if we examine whether mobile services constrain a HM of fixed network services, we would ask for $\varepsilon_{f,m}$. In (regulatory) practice the initial set of products as well as potential substitute candidates will always have to be derived on prior industry knowledge. In case of voice telephony, mobile services are by far the most obvious substitute candidate. Therefore conducting the HMT routine is basically tantamount to estimating ε_f and $\varepsilon_{f,m}$.[18]

Finally, in order to give a meaningful interpretation of estimated demand elasiticities, one needs a benchmark to which estimated values of the own-price elasticity can be compared. The so called critical demand elasticity (ε_c) can be calculated a priori by making specific assumptions about cost and demand functions. Assuming the standard case with constant marginal cost and linear demand we get ε_c = 1/(PCM + t), whereby PCM stands for the competitive price-cost margin prior to the price increase (t).[19] Elasticities less than ε_c imply that the price increase will increase overall profits. Elasticities greater than ε_c mean that a price increase will reduce profits.

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[18] Since retail competition in the mobile sector is usually sufficiently intense so that ex ante regulation is not needed, $\varepsilon_{m,f}$ is typically not of relevance for ex ante market definition purposes. See however Dewenter/Haucap (2007) for a study on mobile demand elasticities in Austria.

[19] See Church/Ware (2000), pp. 607-610.

2.3 Relation between access and calls

Fixed as well as mobile operators tend to set two-part tariffs for their services with a fixed fee and a per-minute calls price.[20] In our estimation we therefore have to allow for the influence of four prices on FMS: The fixed network access price, the mobile network access price, the fixed network per minute price and the mobile network per minute price. In principle, fixed-mobile access as well as call substitution can be expected to depend on all four prices. In the short run (within a few months) and given the decision to subscribe to fixed, mobile or fixed and mobile access, usage substitution may depend on calls prices only. But over a longer period of time (within several months or years), the subscription decision can also be changed. While previous studies usually only include access or call prices or a single price which is a combination of both ('total revenues'),[21] we include all four prices in our model. Since access prices turn out to be insignificant in the usage model and the other way round, we eliminate them from the estimation in a second specification. We also make an additional robustness check with variables incorporating both the fixed and the variable price component (total revenues divided by total minutes).

3 Empirical Specification

We specify a (reduced form) demand model, where demand depends on the own price, the price of a substitute and income. Since cointegration tests (see Appendix) show that (almost) all our variables are integrated of order 1, we cannot estimate in levels, but have to use first differences (which are stationary). Differencing may, however, eliminate valuable information about the (long run) relationship among integrated series. We therefore also consider the

[20] Pre-paid contracts and flat-rates are extreme forms of two-part tariffs where either the access or the calls price is equal to zero.

[21] An exception is Rodini/Ward/Woroch (2002) who include all four prices, i.e., access and calls prices for fixed and mobile telephony. Access and usage prices are not directly observed but estimated from monthly (total) bills, however.

possibility of cointegration and take into accont levels and differences in our estimate. At the same time we have to solve the well known endogeneity problem associated with the estimation of demand equations.

As the empirical specification of the demand for telephone services we propose an error correction model (ECM).[22] Steen and Salvanes (1999), for example, proposed a dynamic formulation of the oligopoly model of Bresnahan (1982) and Lau (1982) within an ECM and applied it to the French market for fresh salmon. In contrast to them, we concentrate on the demand side, only. An ECM allows for short-run departures from long-run equilibrium in the data, and with this approach Steen and Salvanes (1999) address not only statistical problems generated by short run dynamics in the data, but also incorporate dynamic factors such as habit formation of consumers and adjustment costs for producers. We concentrate on the dynamics on the demand side and neglect the supply side.

We specify our basic specification for the demand for telephone services with an ECM such that

$$(2) \qquad \Delta Q(t) = \beta(0) + \beta(1)*\Delta P(t) + \beta(2)*\Delta W(t) + \beta(3)*\Delta Y(t) + \beta(4)*D(1) + \ldots + \beta(n)*D(m) + \gamma*[Q(t-1) - \alpha(1)*P(t-1) - \alpha(2)*W(t-1) - \alpha(3)*Y(t-1) - \alpha(4)*Trend] + \varepsilon(t),$$

where $Q(t)$ denotes quantity, $P(t)$ the price, $W(t)$ the price of the substitute, i.e. the price of mobile phone, $Y(t)$ income and $D(1)$ to $D(m)$ are dummy variables that account for seasonal effects.

We estimate equation (2) with a two-step procedure following the so-called Bardsen (1989) transformation. First, we estimate the equation with OLS to obtain a consistent estimate of γ, i.e. $\hat{\gamma}$. Then, we construct $\Delta Q(t) - \hat{\gamma}\, Q(t-1)$ and regress it on the remaining explanatory variables in equation (2). Instead of getting the long-run parameters directly from a non-linear

[22] See Davidson/MacKinnon (1993) for a detailed description of the error correction model.

estimation procedure, with this procedure we obtain the long-run parameters by dividing all the estimated level parameters by $\hat{\gamma}$.

To account for the endogeneity of output and prices, we estimate equation (2) with two-stage least squares (TSLS) using instruments for $P(t\text{-}1)$ and $\Delta P(t)$. The instrumental variables are used in their levels and their first differences.

4 Data

Since the HMT refers to the aggregate product level we do not analyse firm-specific data. As outlined before, we analyse the market segments of national calls of residential (private) customers and access of private customers. All data have been collected by the Austrian Regulatory Authority for Broadcasting and Telecommunictaions (RTR GmbH) directly from all the operators in the market.

To analyse short- and long-run elasticities, we used monthly data[23] on prices and quantities in Austria over the period from January 2002 to December 2007. As quantities we use the number of subscribers in the fixed network and the traffic volume (total number of national call minutes fixed to fixed and fixed to mobile). Prices were constructed by calculating the average revenue per subscriber and per minute as a proxy for the access and the calls price in fixed and mobile networks. To obtain real prices, we deflated our price data using the Austrian consumer price index.

In order to control for changes in income we use a production index which is highly correlated with GDP (the index is available on a monthly basis while GDP is only available on a quarterly basis). To control for seasonal patterns in demand, we use seasonal dummies.

[23] Unfortunately, quantities and revenues for the mobile sector are only available on a quarterly basis from July 2003 to December 2007. In order not to lose too many observations we linearly interpolated the 'missing' monthly values. This will affect the mobile price variables which are calculated as revenues divided by volumes.

Since prices and quantities are determined simultaneously in equilibrium, we need instruments for the fixed network price(s) in order to identify the demand equation. Ideally, an instrument would be highly correlated with the variable which is about to be instrumented, and simultaneously, it should be uncorrelated with the error term of the equation to be estimated. In the context of demand estimation, cost shifters are likely to be such instruments.

We use the following instruments:

For the fixed network calls price, we use two instruments:

(i) A basket of fixed and mobile termination charges: Termination constitutes a main (wholesale) input for call services (offered at the retail level). Changes in termination charges therefore are changes in variable (marginal) costs – at least for off-net calls. Termination charges are determined by the regulator and therefore can be viewed as exogenous.[24] Since reduction in termination rates are not passed on to consumers immediately, we lag the basket of termination rates for 10 months.

(ii) The number of fixed network access lines: A change in the number of access lines will not directly affect the variable costs of a call, but will – due to the high share of fixed costs – effect average costs.

As regards retail access prices, the numbers of broadband and voice over broadband subscribers appear to be reasonable cost variables. An increasing broadband penetration results in economies of scope (the access line is used for voice and broadband) and can

[24] Of course this assumption can be criticized since the termination charges set by the regulator may be affected e.g. by changes in volumes. However, termination rates are the only factor which directly influences variable costs (at least of off-net calls) and which can be observed by the econometrician. Within the empirical analysis we then also test for the exogeneity of our instruments using exclusionary restrictions. Our results indicate that we cannot reject the exogeneity of our instruments.

result in significant technological cost savings for voice telephony if ‘voice over broadband’ (an access based form of IP telephony) is used.

In addition to these instruments we also considered lagged endogenous variables. A test for exogeneity showed, however, that these variables are not exogenous and therefore could not be used as instruments.

Table 1 and 2 contain descriptions and summary statistics for all the variables used. In the estimation all variables (but the dummies) were used in logarithmic forms. All variables refer to access or national calls of residential customers and are available on a monthly basis (except for mobile prices, see footnote 23) from January 2002 to December 2007. Price variables are deflated by the Austrian consumer price index.

Table 1: Description of variables

Variable	**Description**
q_fn_use	minutes of use in the fixed network (in Mio.)
q_fn_acc	number of access lines in the fixed network (end of month, in Mio.)
p_fn_use	price of fixed network calls (revenues from calls divided by minutes)
p_fn_acc	price of fixed network access (revenues from access divided by users)
p_fn	price of fixed network calls (revenues from *calls and access* divided by minutes)
p_mn_use	price of mobile network calls (revenues from calls divided by minutes)
p_mn_acc	price of mobile network access (revenues from access divided by users)
p_mn	price of mobile network calls (revenues from *calls and access* divided by minutes)
prod	production index (proxy for GDP), seasonally adjusted
p_term	price basket of fixed an mobile termination rates (instrument, lagged for ten months in the TSLS estimation)
q_bb_acc	number of fixed network broadband connections (instrument, in Mio.)
q_vb_acc	number of voice over broadband connections (instrument, in Tsd.)
D(1), D(2)	dummy variables accounting for seasonal effects (in the calls equations)

Table 2: Summary statistics

Variable	observations	mean	std. dev.	Min	max.
q_fn_use	72	405.24	72.62	257.32	528.92
q_fn_acc	72	2.33	0.14	2.04	2.51
p_fn_use	70*	0.052	0.003	0.046	0.061
p_fn_acc	72	12.349	0.365	11.388	13.150
p_fn	70*	0.125	0.008	0.110	0.146
p_mn_use	72	0.099	0.035	0.040	0.145
p_mn_acc	72	5.541	0.214	4.921	6.042
p_mn	72	0.150	0.043	0.072	0.209
prod	72	85.9	5.0	78.4	94.4
p_term	72	0.096	0.017	0.056	0.112
q_bb_acc	72	0.87	0.39	0.32	1.55
q_vb_acc	72	35.44	45.90	0	166.38

* Two observations which were obviously implausible (sudden and strong increase in prices in Nov 2006 followed by a sudden and strong decrease in Dec 2006 before going back to the “normal” level) have been eliminated.

5 Estimation Results

We describe the results with regard to calls in section 5.1 and the results with regard to access in section 5.2. Section 5.3 discusses the implications for market definition and regulation.

5.1 Calls

As discussed in section 2.3, we estimate three models in order to deal with the two-part tariff nature of fixed and mobile telephony prices:

- model 1: with all four prices (*p_fn_use, p_fn_acc, p_mn_use, p_mn_acc*)
- model 2: only with the calls prices (*p_fn_use, p_mn_use*)
- model 3: with ‘average’ call prices (*p_fn, p_mn*)

Since the coefficients on the access prices in model 1 are insignificant, we report the details of the estimation only for model 2. The estimation results are depicted in Table 3.

We report the results from steps one and two, i.e. the OLS estimates and TSLS estimates. To obtain the coefficients $\alpha(1)$, $\alpha(2)$, $\alpha(3)$, and $\alpha(4)$, the respective estimated coefficients have been divided by $\hat{\gamma}$. The standard errors are calculated using the delta method.[25]

As Table 2 shows, the estimated coefficient on γ is equal to -0.3969 and significantly different from zero. This means that it takes the demand two and a half months (1/0.4) to go back to the long run equilibrium path after a shock had occurred. Besides obtaining an estimate for γ, we also obtain long run and short run estimates for the own price elasticity and the cross price elasticity as well as the income elasticity. However, with the first step (OLS) estimations (columns (1) and (2)) we do not account for endogeneity of demand and supply. We do so with the second step (TSLS) estimations (columns (3) and (4)).

[25] The delta method calculates standard errors using a Taylor series expansion. See for example, Ramanathan (1993).

Table 3: Regression results calls (model 2)					
		(1)	(2)	(3)	(4)
		First step (OLS)		Second step (TSLS)	
Dependent variable		*Δq_fn_use(t)*		*Δq_fn_use(t)* - $\hat{\gamma}$ *q_fn_use(t)*	
Variable		Estimated coefficient	Standard error	Estimated coefficient	Standard error
q_fn_use(t-1)	*γ*	-0.3969***	0.098		
p_fn_use(t)	*α(1)*	-1.9528***	0.193	-1.3663***	0.368
p_mn_use(t)	*α(2)*	0.4610***	0.005	0.4995***	0.003
prod(t)	*α(3)*	0.6310***	0.102	0.4124*	0.231
trend	*α(4)*	-0.0072***	0.000	-0.0043***	0.000
Δp_fn_use(t)	*β(1)*	-0.7409***	0.182	-0.7494	0.555
Δp_mn_use(t)	*β(2)*	0.1456	0.243	0.0614	0.283
Δprod(t)	*β(3)*	0.4638***	0.152	0.4009**	0.189
D(1)	*β(4)*	-0.1180***	0.013	-0.1087***	0.015
D(2)	*β(5)*	-0.0623***	0.008	-0.0593***	0.012
constant	*β(0)*	5.1394***	1.771	6.1633***	1.617
Number of observations*		69		69	
Adjusted R^2		0.8308		0.9168	

Significance at the 1%, 5% and 10% level is indicated by ***, ** and *.

The estimated long run own price elasticity in the OLS estimation is equal to -1.95 and the estimated long run cross price elasticity is equal to 0.46. Once we account for the endogeneity of output and prices, the estimated long run own price elasticity becomes smaller in absolute values. It is then equal to -1.37. The estimated long run cross price elasticity is equal to 0.50. The values of the OLS and the TSLS estimations lie in a similar range and indicate that national fixed calls of private users are own price elastic. The long run cross price elasticity suggests mobile phone to be a substitute to fixed telephone.

The short-run elasticties are less elastic than the long run elasticities. This is as one would have expected. The estimated short run demand elasticity is equal to -0.74 (-0.75 with TSLS, not significantly different from zero) and the estimated short run cross price elasticity is equal to 0.14 (0.06) (both not significantly different from zero). These values indicate that national fixed calls of private users are inelastic in the short run. This, however, has to be considered with respect to the monthly data we use. It indicates that some consumers cannot switch or cancel contracts due to minimum contract durations and/or that some consumers are sluggish or not well informed.

The elasticities obtained by estimating models 1 and 3 show that the results obtained with model 2 are fairly robust (see Table 4).

Table 4: Summary of elasticities for calls

Sort run elasticities		**own price (fixed)**	**cross price (to mobile)**
model 1 (4 prices)	OLS	-0.84***	0.15
	TSLS	-1.76**	0.28
model 2 (calls prices)	OLS	-0.74***	0.15
	TSLS	-0.75	0.06
model 3 (‘avg.’ calls prices)	OLS	-1.06***	0.09
	TSLS	-1.41***	0.11
Long run elasticities			
model 1 (4 prices)	OLS	-2.01***	0.34**
	TSLS	-3.57***	0.02
model 2 (calls prices)	OLS	-1.95***	0.46***
	TSLS	-1.37***	0.50***
model 3 (‘avg.’ calls prices)	OLS	-1.20***	0.44***
	TSLS	-1.15***	0.50***

Significance at the 1%, 5% and 10% level is indicated by ***, ** and *.

Short run own price elasticities are always smaller (in absolute value) than long run elasticities (with the only exception being the TSLS estimate in model 3) and range from -0.74 to -1.76, i.e., from the inelastic to the elastic range. The long run own-price elasticities range from -1.15 to -3.57 and are consistently in the elastic range.

Cross price elasticities to mobile are insignificant in the short run, but significantly positive in the long run (with the only exception being the TSLS estimate in model 1). They are in the inelastic range and go from 0.34 to 0.50. This suggests that national mobile calls are a substitute to national fixed calls for private users.

To test the validity of the instruments, we test for the joint significance of the instruments in the first stage. We also use a Hausman test to test for the endogeneity of quantities and prices. Finally, we test for overidentifying restrictions. The detailed results of these tests can be found in the Appendix. We conclude from these tests that our instruments are exogenous and correlated with the endogenous variable on the right hand side. We further conclude that there are no significant differences between the OLS and the TSLS estimates.

5.2 Access

Prices and quantities of access are also integrated of order one. So we estimated the same error correction model as for calls. However in the first stage the coefficient on *q_fn_acc(t-1)*, γ, was insignificant. This means that there is no (significant) cointegration relation between the variables. We therefore simply estimate the model in differences:

$$(3) \qquad \Delta Q(t) = \beta(0) + \beta(1)^* \Delta P(t) + \beta(2)^* \Delta W(t) + \beta(3)^* \Delta Y(t) + \varepsilon(t),$$

Similar to calls, we estimate three model specifications:

- model 1: with all four prices (*p_fn_use, p_fn_acc, p_mn_use, p_mn_acc*)
- model 2: only with the access prices (*p_fn_acc, p_mn_acc*)
- model 3: with 'average' call prices (*p_fn, p_mn*)[26]

Since the coefficients on the usage prices are insignificant in specification 1, we report details only for specification 2.

The results of the estimation are depicted in Table 5. We included 2 AR-terms to deal with autocorrelation in the residuals. The coefficient on *Δp_fn_acc(t)* can be interpreted as the short run own-price elasticity. It is significant but very small/inelastic. This counts for the OLS as well as for the TSLS estimation. The short-run cross-price elasticity ($\beta(2)$) is negative, but very small and insignificant in both the OLS and the TSLS estimate.

[26] A fourth specification with price variables dividing total revenues by the number of subscribers (instead of minutes) was also calculated but did not deliver (qualitatively) different results compared to model 3.

Table 5: Regression results access					
		(1)	(2)	(3)	(4)
		OLS		TSLS	
Dependent variable		*Δq_fn_acc(t)*		*Δq_fn_acc(t)*	
Variable		Estimated coefficient	Standard error	Estimated coefficient	Standard error
Δp_fn_acc(t)	*β(1)*	-0.0584***	0.0124	-0.1025***	0.021
Δp_mn_acc(t)	*β (2)*	-0.0026	0.0122	-0.0028	0.012
Δprod(t)	*β (3)*	-0.0282***	0.0067	-0.0257***	0.006
AR(1)	*β (5)*	0.1751*	0.0963	0.1854**	0.092
AR(2)	*β (6)*	0.5227***	0.1028	0.5052***	0.103
constant	*β (0)*	-0.0031***	0.0007	-0.0032***	0.001
Number of observations		69		69	
Adjusted R-squared		0.3474		0.2823	

Significance at the 1%, 5% and 10% level is indicated by ***, ** and *.

The long-run elasticity can be calculated from the short-run elasticities and the AR-terms as follows:

(4) $$\varepsilon_{LR} = \frac{\varepsilon_{SR}}{(1-\beta(5))(1-\beta(6))}$$

With (4), the long-run own-price elasticity becomes -0.25 in the TSLS estimate and -0.15 in OLS estimate. They are still clearly inelastic.

The elasticities obtained by estimating models 1 and 3 show that the results obtained with model 2 are again fairly robust (see Table 6).

Table 6: Summary of elasticities for access

Sort run elasticities		**own price (fixed)**	**cross price (to mobile)**
model 1 (4 prices)	OLS	-0,06***	-0.00
	TSLS	-0,09***	-0.00
model 2 (access prices)	OLS	-0,06***	-0.00
	TSLS	-0,10***	-0.00
model 3 ('avg.' calls prices)	OLS	-0.01	-0.00
	TSLS	-0.01	-0.00
Long run elasticities			
model 1 (4 prices)	OLS	-0.15**	-0.00
	TSLS	-0.21**	-0.00
model 2 (access prices)	OLS	-0.15**	-0.01
	TSLS	-0.25**	-0.01
model 3 ('avg.' calls prices)	OLS	-0.01	-0.00
	TSLS	-0.02	-0.00

The own price elasticities are significant (except for model 3) but very inelastic even in the long run. Cross price elasticities are insignificant which suggests that mobile access is not a substitute for fixed access for private users.

Again, we test for the validity of our instruments using the same test as for calls. The detailed results of these tests can be found in the Appendix. We conclude from these tests that our instruments are exogenous and correlated with the endogenous variable on the right hand side. We further conclude that there are no significant differences between the OLS and the TSLS estimates (with exception of model 1).

5.3 Implications for market definition and regulation

To arrive at conclusions for market definition, the estimated elasticities have to be compared to the critical elasticity, ε_c.

With regard to access, the estimated elasticity is very inelastic and therefore does not exceed the critical elasticity (the smallest possible value of the critical elasticity being 0.91 in case of PCM=1 and t=10%). The estimate therefore indicates that mobile access is not part of the same market as fixed access for residential customers.

To determine ε_c for national calls of private customers, we have to estimate the share of variable costs in total costs to determine the competitive price-cost margin. As regards the definition of variable costs we have to refer to the relevant time horizon of the HMT, i.e., 1 to 2 years. On this basis we estimate variable costs to be in the proximity of 25% of total cost yielding a competitive price-cost margin of 75%.[27] Depending on t (5% or 10%), we get $\varepsilon_c^{5\%}$ = -1,25 and $\varepsilon_c^{10\%}$ = -1.18 respectively.

The range of the estimated (long run) elasticity for calls (-1,15 to -3,57) for the largest part exceeds the range of the critical elasticity. This indicates that fixed and mobile national calls of private users are part of the same market. This is a result which is also supported by evidence from Austrian consumer surveys and general market developments.[28] Given the existence of four mobile operators, this would also indicate that a retail regulation for fixed network national calls of private users is no longer needed.[29]

6 Discussion and Conclusions

Deregulation of fixed voice telephony seems hardly justifiable solely on grounds of intramodal competition within most EU member states.[30] Therefore, inferences about economic FMS and the state of intermodal competition would appear to be a cornerstone on which future regulatory judgements should be based. On account of this, any reliable evidence from econometric studies should be seen as an important part of the regulatory decision making process. Since such evidence is most helpful if available on the level on

[27] This is also within the range given in Stumpf (2007), 60-90%.

[28] See RTR (2008a) and RTR (2008b)

[29] Although one has to take into account that the fixed network incumbent and the largest mobile operator are part of the same company. Nevertheless, the Austrian Regulatory Authority RTR concluded that fixed network retail market(s) for residential users are no longer relevant for ex ante regulation (see RTR (2008b)).

[30] As it can be inferred from European Commission (2008), volume 2, figure 37 intramodal competition is most pronounced with respect to international calls. Indeed, those markets were the first ones to be deregulated since the first recommendation was introduced (see Cullen International (2008)).

which regulatory decisions are taken, we focus on the segment of private users and estimate elasticities for access and national calls demand.

As expected given the highly intense mobile competition in Austria we find evidence for FMS with respect to calls. Furthermore, we found empirical support for our hypotheses according to which FMS is much less pronounced as regards retail access compared to calls markets. Whereas we found that demand for access services is inelastic and that the cross price elasticity to mobile is insignificant throughout all of our model specifications, estimates for calls show elastic demand and a significant positive cross price effect. As the estimated elasticity of national calls of private users likely exceeds the critical elasticity, we conclude that national fixed and mobile calls of private users are likely to be part of the same market. Fixed and mobile access for private users, on the other hand are unlikely to be part of the same market based on our evidence. In this respect, one has to say, however, that mobile broadband, which might significantly increase the willingness of consumers to switch from fixed to mobile access, only started to take off in Austria in 2007, which is the last year of our sample. An estimate based on more recent data might therefore result in higher elasticities.[31]

What could this mean for other markets? Since elasticities in the business segment can be assumed to be significantly lower than in the residential segment, we would expect that fixed and mobile telephony (be it access or calls) do not form a single market. The same goes for international calls, where significant price differences between fixed and mobile still prevail. Since Austria is among the countries with the highest FMS in the EU, it is even questionable, from our point of view, if FMS in strong enough in other EU countries (even for national calls of residential customers). But these are certainly fields for future research.

Finally, we have to point out, that, although we have rich data available, there are some shortcomings: data for mobile prices had to be interpolated from quarterly to monthly date for a significant part of the sample. With regard to the instruments we use, we find that they work

[31] The coefficient on an interaction term (fixed network access price times a dummy variable for 2007) was insignificant in model 2, however.

quite well in levels, but not as well in differences. The results should therefore be interpreted with caution an in the context of other evidence (e.g. consumer surveys) available. For Austria, such additional evidence points in the same direction as our estimates, however.

Appendix – Instrumental variable and cointegration tests

Unit root and cointegration test - calls

We test for a cointegrating relation by conducting unit root tests on all series in the ECM demand model and the estimated residuals. We use Dickey-Fuller's augmented unit root test including an intercept and a time trend. As Table 7 shows, we cannot reject the null hypothesis of a unit root for the series *q_fn_use(t)*, *p_mn_use(t)*, *p_mn* and *prod(t)* at the 10% level, and for the series *p_fn_use(t)* at the 1% level. We also find that these variables are stationary in first differences.

The estimated residuals are stationary in all three models. We use the adjusted asymptotic critical value from Table 20.2 in Davidson and MacKinnon (1993). In all cases we can reject the null hypothesis of a unit root. Therefore, we conclude that the residuals are stationary and that there is a cointegrating relation, i.e. a long run demand relation, between quantities, prices and income.

Table 7: Results from unit root tests – calls

Variable	Test statistic*	Critical values 1%	5%	10%
q_fn_use(t)	-0.77	-4.12	-3.49	-3.17
p_fn_use(t)	-4.02	-4.09	-3.48	-3.17
p_mn_use(t)	-2.34	-4.06	-3.46	-3.15
p_fn(t)	-4.46	-4.09	-3.48	-3.17
p_mn(t)	-0.24	-4.05	-3.45	-3.15
prod(t)	-1.87	-4.08	-3.45	-3.15
Estimated OLS residuals model 1	-10.59	-5.52	-4.98	-4.70
Estimated TSLS residuals model 1	-10,23	-5.52	-4.98	-4.70
Estimated OLS residuals model 2	-10.61	-4.97	-4.43	-4.15
Estimated TSLS residuals model 2	-10.42	-4.97	-4.43	-4.15
Estimated OLS residuals model 3	-9.26	-4.97	-4.43	-4.15
Estimated TSLS residuals model 3	-9.32	-4.97	-4.43	-4.15

*including intercept and time trend

Unit root tests - access

The same test as for calls are also made for the variables used in the access estimation.

Table 8: Results from unit root tests – access

Variable	Test statistic*	Critical values 1%	5%	10%
q_fn_acc(t)	-0.97	-4.09	-3.47	-3.16
p_fn_acc(t)	-1.82	-4.09	-3.47	-3.16
p_mn_acc(t)	-4.21	-4.05	-3.45	-3.15

*including intercept and time trend

We cannot reject the null hypothesis of a unit root for the series *q_fn_acc(t)*, and *p_fn_acc(t)* at the 10% level. We also find that these variables are stationary in first differences.

Analysis of instrumental variables

To test the validity of the instruments, we test for the joint significance of the instruments in the first stage (see column '1^{st} stage: F-statistic (p-value)' in Table 9).

Additionally, we use a Hausman test to test for the endogeneity of quantities and prices. We use the residuals from the first stages as explanatory variables in the second stage and test for joint significance with an F-test (see column 'Hausman test: F-/t-statistic (p-value)' in Table 9).

Table 9: Instrument tests

model	endogenous variable(s)	instruments	1st stage: F-statistic (p-value)	Hausman test: F-/t-statistic (p-value)
calls, model 1	*p_fn_use(-1)*	*p_term(t-11), Δp_term(t-10), q_fn_acc(t-1), Δq_fn_acc(t), q_bb_acc(t-1), Δq_bb_acc(t), q_vb_acc(t-1), Δq_vb_acc(t)*	3.72 (0.00)	1.98 (0.11)
	Δp_fn_use		1.12 (0.37)	
	p_fn_acc(-1)		37,89 (0.00)	
	Δp_fn_acc		2,33 (0.03)	
calls, model 2	*p_fn_use(-1)*	*p_term(t-11), Δp_term(t-10), q_fn_acc(t-1), Δq_fn_acc(t)*	3.42 (0.01)	0.94 (0.40)
	Δp_fn_use		1.44 (0.23)	
calls, model 3	*p_fn(-1)*	*p_term(t-11), Δp_term(t-10), q_fn_acc(t-1), Δq_fn_acc(t), q_bb_acc(t-1), Δq_bb_acc(t), q_vb_acc(t-1), Δq_vb_acc(t)*	6.41 (0.00)	0.55 (0.58)
	Δp_fn		0.70 (0.69)	
access, model 1	*Δp_fn_acc*	*Δq_bb_acc(t), Δq_vb_acc(t), Δp_term(t-10)*	0.61 (0.61)	3.01 (0.06)
	Δp_fn_use		5.07 (0.00)	
access, model 2	*Δp_fn_acc*	*Δq_bb_acc(t), Δq_vb_acc(t),*	3.57 (0.03)	-0.83 (0.41)
access, model 3	*Δp_fn*	*Δq_bb_acc(t), Δq_vb_acc(t), Δp_term(t-10)*	9.46 (0.00)	0.05 (0.96)

Finally, we test the calls model 2 for overidentifying restrictions. We estimate the structural equation (2) by TSLS using only one instrument. We regress the residuals from this regression on all exogenous variables, i.e. all exogenous variables in equation (2) and all instruments, and obtain a value of 0.13 for the R^2. This gives a value of 8.97 for the test statistic $n*R^2$, which is χ^2 distributed. With two degrees of freedom and a significance level of 1%, the critical value of the χ^2 distribution is 9.21. We therefore cannot reject the null hypothesis that all instruments are uncorrelated with the estimated residuals at the 1% level. We conclude that our instruments are exogenous.

Similarly, we obtain a value of 0.05 for the R^2 of access model 2. This gives a value of 3.06 for the test statistic $n*R^2$, which is Chi-square distributed. With two degrees of freedom and a significance level of 1%, the critical value of the Chi-square distribution is 9.21. We conclude that our instruments are exogenous.

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