Measuring the Benefits of Mobile Number Portability

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Abstract

Increasing numbers of countries require mobile telephone networks to offer mobile number portability (MNP). MNP allows customers who wish to switch mobile operator to keep their mobile numbers, avoiding the costs of switching to new numbers. *Ex ante* assessments suggest that MNP should reduce switching costs and strengthen competition. In this paper, we test MNP's impact on market outcomes using international time-series cross-section data. We find that MNP reduces average prices and encourages churn (a proxy for switching) when the switching process is rapid (e.g. less than 5 days) but not when it is slower.

JEL classifications: L96, L51

Key words: Impact of regulation, mobile telecommunications, cost-benefit analysis, competition, switching costs.

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1.   Introduction

Increasing numbers of countries require mobile telephone network operators to offer mobile number portability (MNP). This facility allows customers who wish to switch mobile operator to keep the mobile numbers originally assigned to them, avoiding the costs of switching to new numbers.

Since MNP regulation was first mooted, policymakers have asked whether it can produce positive net benefits. *Ex ante* evaluations of MNP carried out in several countries have produced detailed estimates of expected costs and direct benefits (e.g. the savings accruing to customers from lower switching costs). While researchers have suggested MNP should have a range of potentially important effects, such as strengthened competition and reduced prices (see Buehler, Dewenter and Haucap (2006) for a recent discussion), few attempts have been made to quantify them *ex post*.

The staggered introduction of MNP internationally provides a useful natural experiment. In this paper, we use econometric analysis of international time-series cross-section data to estimate the average treatment effects of MNP on retail prices and switching by customers. The dataset constructed for this purpose includes information from up to 38 countries for 22 quarters (1Q 1999 through 2Q 2004).

We find that the quality of MNP, as proxied by the target maximum porting time, helps explain its impact on switching and average prices. For countries in our sample that required porting to be completed in five or fewer days, MNP was associated with increased customer switching and lower prices. The sub-sample of countries with less stringent porting time standards experienced no significant churn or revenue effects.

The costs associated with the MNP service depend upon the technology used to deliver it (Buehler, Dewenter and Haucap, 2006). The technology, in turn, determines the "quality" of MNP, including dimensions such as porting time and reliability. Previous research, e.g. Gans, King and Woodbridge (2001), has emphasised the importance that the choice of number portability technology has in determining the likely effects of the measure. Our results provide empirical support for this view. Jurisdictions conducting *ex ante* assessments of MNP in the future should consider the likely trade-off between achieving positive market outcomes and cost of implementation.

Section 2 of the paper provides a brief classification of the potential benefits of MNP and refers to some previous research, including both *ex ante* cost-benefit studies and other empirical research. In Section 3, we ask what effects MNP should be expected to have on consumer switching behaviour and prices. The dataset constructed for this study is described in Section 4, along with some descriptive statistics. Sections 5 and 6 set out econometric models of switching and retail prices, respectively, and Section 7 discusses our conclusions and suggestions for future research.

2.   Potential Benefits of Mobile Number Portability

To provide context for the empirical analysis that follows, in this section we briefly review some relevant empirical research. This consists of *ex ante* cost benefit analyses conducted on MNP by regulators and a modest number of *ex post* empirical studies. Existing theoretical research on mobile number portability was recently surveyed in Buehler and Haucap (2004), but to clarify terminology used in the remainder of the section, it is worth restating the standard classification of number portability benefits.

2.1   Classification of benefits

A commonly-used approach to analysing the likely costs and benefits of MNP divides the measure's potential benefits into three types:[1] **Type 1** benefits obtained directly by customers who switch, **Type 2** benefits obtained by all mobile telephony customers (e.g. efficiency gains and price reductions due to strengthening of competition) and **Type 3** benefits obtained by those making calls to ported numbers.

Past *ex ante* evaluations have proceeded on the basis that MNP should be expected to provide net welfare gains if the sum of these benefits exceeds the cost of network investments, process changes and operating expenses incurred to make mobile numbers portable. However, they have tended to focus on the more empirically tractable Type 1 and Type 3 benefits, giving less emphasis to Type 2 benefits. In Section 2.2 we review some of the results of these *ex ante* evaluations.

2.2   *Ex-ante* Cost-benefit Analyses

Full mobile number portability (MNP) was first employed in Singapore in 1997, and since then many countries have introduced this form of regulation. Several cost-benefit analyses (CBAs) are available in published form, notably Oftel (1997) for the UK, NERA/Smith (1998) for Hong Kong, and Ovum (2000) for Ireland. In Table 1 below, we summarise the estimated benefits per customer by type from each of these studies.

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[1] This framework was originally devised by NERA for the UK regulator OFTEL in a study of geographical number portability: Monopolies and Mergers Commission (1995), pp.58-59. See Oftel (1997) for an early application to mobile number portability.

Table 1: Predictions from three *ex ante* assessments of MNP			
Country	**UK**	**Hong Kong**	**Ireland**
Base year	1997	1998	2000
Expected benefits per subscriber *Present value (in USD) of ten year impact divided by subscribers in base year*			
Type 1	28 - 81	39 - 71	78
Type 2	n/a	1	26
Type 3	1 - 5	1 - 3	5

Sources: analysis of estimates in Oftel (1997), NERA/Smith (1998) and Ovum (2000). Exchange rates are base year figures from IMF International Financial Statistics.

Type 2 benefits were viewed as difficult to estimate, and since Type 1 benefits were by themselves expected to be sufficiently high to justify the intervention, Type 2 benefits were either not quantified or subject to only simple scenario analysis. For example, in the CBA for the Irish market, Ovum assumed that MNP would lead to a 3% fall in retail post-pay mobile telephony prices.[2] Sensitivity analysis was carried out for reductions of 1% and 5%. Ovum also noted that there might be benefits from cost efficiencies or greater innovation, but these were not modelled.

2.3   Other empirical research on the effects of MNP

The main *ex post* empirical work on MNP to date has focused on the propensity of those switching mobile provider to use MNP. This is particularly relevant to the size of Type 1 benefits as discussed above.

As part of a wider study of switching costs for the UK Office of Fair Trading, NERA (2003) examined the usage of MNP for inter-operator switching in UK mobile telephony markets. They found that in the first two years after MNP was introduced, the usage of MNP was very limited for residential customers, with only 12% of customers that switched operator taking up the portability option. This is far lower than the rate predicted in *ex ante* assessments. However, half of businesses who

changed numbers in this period ported at least some of their numbers. NERA suggested that the difficulty of using MNP during the first years after implementation may explain its unpopularity: porting a number originally took an average of 25 days. When the delivery time was reduced to five days on average, take-up increased to about 18% for residential customers and 80% for businesses.[3]

Looking beyond the propensity of switchers to use MNP, there has been little previous empirical work on the broader effects of MNP regulation. Ovum (2005) examined the experience of MNP in six countries that have implemented it: Australia, Germany, Hong Kong, Ireland, the Netherlands and the UK. Several of their findings are relevant to this study:

- Usage of MNP can fall significantly if the time it takes to change operator ("porting time") is too long. The authors suggest that two days is a practical upper limit. However, very short porting times do not necessarily increase demand for MNP.

- High end-user charges for MNP can also deter usage of the facility. Lower charges, which the authors suggest are levels of less than 20% of monthly average revenue per user, do not seem to be a "major deterrent to usage".[4] However, zero charges do not seem to increase demand beyond the levels associated with low charges.

- In jurisdictions with MNP, the extent to which switching customers use it varies widely and tends to increase over time.

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[2] Ovum (2000), pp.12-13.

[3] NERA (2003), pp.37-39.

[4] Ovum (2005), p.1.

There has also been a limited amount of academic research on individual markets. Below we cite two concerning MNP and one on number portability in a related market.

Lee, Kim and Park (2004) used contingent valuation techniques to estimate the prospective demand for MNP in South Korea. They found that the average South Korean mobile user was willing to pay an average of 3.24% of his or her monthly bill for a mobile number portability option. Willingness to pay (WTP) showed a strong positive association with income, awareness of MNP, and intention to switch. The authors also found that WTP varied significantly depending upon a user's network operator: the figure was lower for customers of the incumbent operator than those using either of the alternative operators. Other demographic variables such as age, gender and occupation were not found to be significant.

A recent *ex post* study of MNP's effects also focuses on South Korea. Kim (2005) estimated switching costs for customers of two of the country's mobile network operators by applying a random utility model to cross-sectional subscriber-level microdata. The paper compared switching costs calculated using samples before and after MNP, and differences between these estimates were attributed to MNP. Controls included firm-specific dummy variables, prices, non-price network attributes and customer characteristics. The paper estimated that MNP reduced average switching costs in South Korea by more than 35%.[5] Data reported in the paper indicates that there was significantly more switching after MNP was introduced, at least among customers of the largest operators.[6] Service fees maintained a downward trend of about 7% per annum from 2002-2005, with no

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[5] Kim (2005), p.16.

obvious change in relative or absolute prices at the point MNP was introduced for the two largest operators (July 2003).[7] Per-minute prices remained broadly unchanged over the period.[8]

Viard (forthcoming) examined the effect of number portability on prices in the US market for toll-free calls. This service is different from mobile telephony, but it is similar in some respects (e.g. high rates of growth).[9] Estimating price regressions on data from 219 AT&T virtual private network contracts, he found that introduction of number portability was associated with price reductions of 4.4%. A control group of contracts containing no toll-free services showed no relationship between prices and the introduction of number portability. Viard interpreted the results as evidence of an inverse relationship between switching costs and competition in this market: "despite rapid growth in the market, the firms' incentive to exploit their existing 'locked in' users was greater than their incentive to 'lock in' new customers."[10]

3.   Likely effects of MNP on switching and prices

In this section, we outline the main effects that economic theory suggests MNP should have on switching propensity and retail prices.

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[6] *Ibid*, Table 2.

[7] *Ibid*, p.11.

[8] *Ibid*, Figure 5.

[9] Note, however, that there are also important differences between mobile telephony and toll-free calls markets; in particular, mobile operators may be able to price discriminate between new and existing users. NERA (2003) noted that handset subsidies in effect involve lower prices for new customers than for existing ones; pp.30-31.

[10] Viard (2004), p.25.

3.1   MNP and consumer switching

Significant numbers of customers switch operators at some point after their initial acquisition of a mobile subscription. There are likely to be many reasons for such switching, e.g. changes in individual demand patterns, service innovation, learning by customers about the fit between their pattern of demand and operator offerings, and changing price and quality propositions.

To the extent that the component of switching cost associated with changing one's telephone number is high enough to deter some customers from switching operator when they might otherwise have done so, MNP should yield a positive change in the conditional probability of switching (holding other variables constant). This effect might be offset in whole or in part by operators' reactions, e.g. if operators respond to MNP by reducing price dispersion. Nevertheless, MNP should have at least a weakly positive effect on switching.

3.2   MNP and retail prices

The net effect of MNP on retail prices is in principle indeterminate. Empirically, it is likely to depend upon the interplay of three groups of effects:

- Pass-through of costs associated with the facility (increase in prices);

- Effects on competition (probably a decrease in prices); and

- Loss of customer information (increase in prices).

First, and most obviously, the implementation of MNP imposes costs on all operators employing it. Depending upon the extent of competition in a given national market, these costs are likely to be (at least partly) passed on to consumers and thereby lead to increased prices. Some argue that this is likely to be the main effect of number portability, and hence that mandating it through regulation will lead to a net

reduction in welfare; see, for example, Ellig (2005).[11] Aoki and Small (1999) also address the welfare impact of switching cost reductions due to number portability. They identify cases in which switching costs reductions provided by number portability (e.g. reducing the need to purchase complementary goods such as stationery) could be offset by higher marginal costs of providing call services, leaving consumers with lower surplus.

Beyond the simple effect of increased direct costs from implementation of MNP, theory is less definite about the effect of decreased switching costs on prices. A survey by Klemperer (1995) on the effects of consumer switching costs on competition concludes that "switching costs generally raise prices and create deadweight losses of the usual kind in a closed oligopoly."[12] Buehler and Haucap (2004) present a model focussing specifically on MNP that yields an overall reduction in prices for customers but implies that increases for entrants' customers will be more than offset by decreases for incumbents' customers. The switching cost literature also raises the possibility that a fall in switching costs could make it easier to sustain tacit collusion , e.g. Padilla (1995).

The third group of effects concerns an informational channel through which MNP may lead to increases in at least one component of mobile telephony prices. Depending upon how MNP is implemented, it may reduce the tariff information available to both fixed and mobile customers wishing to make calls to mobile numbers. This effect is discussed in Buehler and Haucap (2004) and Gans and King (2000). Particularly if mobile termination rates are unregulated and there is no

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[11] Ellig (2005), p.29.

[12] Klemperer (1995), p.536.

mechanism identifying the terminating operator to each caller, such a decrease in transparency could lead to higher prices for call termination.

4.   Data employed

We have constructed an unbalanced time-series cross-section dataset that includes most of the OECD and a selection of developing countries. It is based principally on the Merrill Lynch Global Wireless Matrix (Merrill Lynch, 2004).

Although this source provides some data on 46 countries, there are many gaps. Also, we found that data for three countries, China, the Czech Republic and South Korea, contained implausibly large fluctuations in reported subscriber numbers. As a result, these countries were excluded from the dataset. The available panel includes data on 38 countries (for churn modelling) and 37 countries (for price modelling). See Table 10 in the annex for details of the countries and the sample coverage.

The data are quarterly, running for up to 22 quarters from 1Q 1999 through 2Q 2004, and we omit the first two quarters to allow use of differenced and lagged variables. Table 2 below lists the variables and provides summary statistics. Figures in this table and elsewhere in the paper are rounded to three significant digits. Further information on some of the variables is provided in the annex.

Table 2: Variable descriptions, sources and summary statistics (individual observations are for country *i* and quarter *t* in each case)

Variable	Description	Source	Churn model Mean	Churn model St Dev	Price model Mean	Price model St Dev
MNP_{it}	= 1 if mobile number portability in place at any time in quarter *t*	See Table 10 in the annex	0.285	0.452	0.240	0.428
$MNPtime_{it}$	Target maximum single line porting period (days)	*Ibid.*	1.81	4.18	1.69	4.22
$MNP5d_{it}$	If MNP = 1 and *MNPtime* <=5 then 1, else 0	*Ibid.*	0.175	0.381	0.128	0.334
$MNP6p_{it}$	If MNP = 1 and *MNPtime* >5 then 1, else 0	*Ibid.*	0.109	0.312	0.112	0.316
$RMNP_{it}$	If MNP = 1, then (1/*MNPTime*), else 0	*Ibid.*	0.390	1.94	0.382	1.96
$CHURN_{it}$	Monthly number of disconnections from a network expressed as % of MNO's avg. subscriber base in the same month. Quarterly avg. of monthly rates.	Weighted avg. of individual MNOs' data from ML	0.0205	0.0102		
DEN_{it}	Cellular density: mobile users as a share of population	Analysis of ML	0.534	0.298		
OPS_{it}	Number of MNOs in country *i*	Analysis of ML	3.76	1.23	3.72	1.21
$RGDPPC_{it}$	Real GDP per capita (US$)	See the annex	17,400	12,300	17,100	12,100
RPM_{it}	Average real revenue per minute for MNOs in country *i* (US$)[13]	Weighted avg. of individual MNOs' data from ML			0.198	0.0794
$TOTMIN_{it}$	Monthly average minutes of mobile telephony traffic in country *i* (millions)	Analysis of ML			3,710	10,000
$PDNST_{it}$	Population density: population per Km²	World Bank WDI (2004)			126	144
HHI_{it}	Herfindahl Hirshman Index: Sum of the squares of the market shares (users) of all MNOs in country *i*	Analysis of ML			3,790	976
$CR1_{it}$	The top MNO's share of total users	Analysis of ML			0.477	0.116

Notes: MNO is an abbreviation for "mobile network operator". Merrill Lynch (2004) is referred to as ML.

5.   Modelling the effect of MNP on switching

In this section, we define and estimate two econometric models of switching

frequency, including proxy variables to capture the effect of MNP.

The switching variable

The ideal measure of switching for our purposes would directly identify flows of

customers between operators, but such data are generally not put in the public

domain. The best available proxy is **churn**, a metric based on the number of

disconnections from each network as a proportion of the average number of network

users in a given period. While inter-operator switching does feed into churn, the

churn rate is not a pure measure of switching. Subscribers that leave a network

without joining another one, for whatever reason, also appear as churn, as do

customers on pre-paid tariff packages that do not use their phones for a specified

period.

Because churn is a proportion, we apply a logistic transformation to the data before

using it as a dependent variable:

$$LGTCHURN_{it} = \ln\left(\frac{CHURN_{it}}{1 - CHURN_{it}}\right) \tag{2}$$

Explanatory variables

Switching propensity should be positively related to the presence or absence of MNP

and to the quality of the MNP service, insofar as the service reduces consumer

switching costs. However, we have no theoretical prior as to the functional form of

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[13] This is rebased to year 2000 prices using GDP deflators and it excludes revenue from data services.

the relationship. To allow for a range of possibilities, we test two alternative proxies

for MNP, both based on the target maximum porting time (*MNPTM*) in force in a

given country.[14]

The first is a threshold approach, distinguishing between countries with a *MNPTM* of

5 days or less (for which *MNP5D* is set to 1) and those with 6 days or more (for

which *MNP6P* is set to 1). Both variables are set to zero for all other cases. This

divides the observations where MNP was in place into two roughly equal parts along

the quality dimension. The second MNP proxy, *MNPR* is equal to the reciprocal of

MNPTM for observations with MNP and to zero for those without the service.

In the remainder of this section, we include some descriptive statistics to illustrate

the key bi-variate relationships in our data.

A comparison of averages suggests that countries with "high quality" MNP had

slightly higher churn than those without MNP, but those with "low quality" MNP

had slightly lower churn (see Table 3).

Table 3: Relationship between churn rates and mobile number portability	
Case	Sample mean quarterly churn
No MNP	0.0203
MNP delivery time <= 5 days	0.0218
MNP delivery time 6+ days	0.0198

Source: see Table 2 above.

In a regression analysis, we expect coefficients on both *MNP5D* and *MNP6P* to be

positive, but the former should be larger than the latter to the extent that MNP quality

is important to consumers. *MNPR* is also expected to have a positive coefficient, but

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[14] Data on actual, rather than target, porting times would probably be a better measure of quality. Unfortunately, these data are not made public in most countries.

its success in explaining churn will depend upon how well its specific function form fits the data.

We also note that the decision to enact MNP regulation may be affected by market conditions, including churn levels. The econometric model will need to take this possible endogeneity into account.

The number of operators, *Ops*, should have a positive coefficient reflecting increased switching options and promotional activity as the number of operators rises.

A proxy for real incomes, *RGDPPC*, is included to a capture possible reduction in disconnections as income rises. This is expected to have a negative coefficient, because we expect that, in line with previous research, customers' demand for mobile network access is positively related to income. If this is the case, it also seems likely that users with higher income are *ceteris paribus* less likely to stop using mobile telephony once they have started than those of lower income. Since the churn figures include those who disconnect from one network without connecting to another, it is likely to be lower in markets with higher average incomes.

Finally, we include cellular density terms, which measure the number of mobile connections per head of population (*DEN).* This is intended to allow for a possible relationship between market maturity and churn. We might expect an increase in switching propensity as customers become more familiar with mobile telephony and as cohorts with greater price sensitivity take up access. In more mature markets, falling demand growth may weaken the incentives for switching by changing the nature of competition (for example, via reductions in handset subsidies). Since density tends to approach a limit as each market matures rather than continuing to

rise linearly, we include higher order transformations of this variable in the regressions.[15]

Quarterly dummies (*Q1* and *Q3-Q4*) are also included. Other factors that could affect churn, but on which data are not available, include the rate of service innovation, the extent of pre-paid customer registration vs. anonymity, the frequency of customer repeat purchase or sampling, contract lengths, the level of other (non-number-related) switching costs and the extent of substitutability between services of different operators.

We allow for I.I.D. errors in the measurement of variables through a disturbance term (ε_{it}). It also seems likely that data limitations, particularly regarding local preferences and service characteristics, have led to omission of variables that might help explain the level of churn in each country, so we expect to observe significant individual effects at country level (u_i).

Hence, for country $i = 1...38$ and quarter $t = 1...20$:

$$\ln\left(\frac{CHURN_{it}}{1 - CHURN_{it}}\right) = \alpha + \beta_1 OPS + \beta_2 GDPPC_{it} +$$

$$\beta_3 DEN_{i(t-1)} + \beta_4 DEN_{i(t-1)}^2 + \beta_5 DEN_{i(t-1)}^3 \quad (3)$$
$$\beta_6 Q1_{it} + \beta_7 Q3_{it} + \beta_8 Q4_{it} + u_i + \varepsilon_{it} +$$
$$[\beta_9 MNP5D_{it} + \beta_{10} MNP6P_{it} \text{ or } \beta_9 MNPR_{it}]$$

Summary of prior expectations about coefficients:

$\beta_1, \beta_3, \beta_5, \beta_9, \beta_{10} > 0;$
$\beta_2, \beta_4 < 0$

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[15] For a recent survey of empirical work on mobile telephony density, see Banerjee and Ros (2004).

5.1   Econometric Results

Since diagnostic tests after fixed effects OLS estimation showed evidence of autocorrelation and heteroscedasticity,[16] we estimated the models using the Arellano-Bond "difference GMM" estimator with robust standard errors. T-statistics are reported rather than Z-statistics due to the relatively small sample. The results are shown in Table 4 below.

Table 4: Churn regression results using Arellano-Bond estimator, with MNP variables treated as endogenous

Variables and statistics	Using MNP delivery time threshold dummies (<=5 days, >6 days)		Using reciprocal of MNP delivery time for countries with MNP	
Dep. variable	*LGTCHURN$_{it}$*		*LGTCHURN$_{it}$*	
	Coef.	*Robust t-stat.*	Coef.	*Robust t-stat.*
LGTCHURN$_{i(t-1)}$	0.682	*12.12****	0.675	*12.03****
MNP5d$_{it}$	**0.166**	**2.09****		
MNP6p$_{it}$	**-0.171**	**-1.63**		
RMNP$_{it}$			**0.00752**	**1.49**
OPS$_{it}$	-0.00719	*-0.17*	-0.0117	*-0.27*
LRGDPPC$_{it}$	-0.171	*-1.75**	-0.164	*-1.57*
DEN$_{i(t-1)}$	1.22	*1.5*	1.31	*1.95**
DEN$_{i(t-1)}^{2}$	-2.13	*-1.65**	-2.14	*-2.27***
DEN$_{i(t-1)}^{3}$	1.43	*1.9*	1.39	*2.64****
Constant	-0.00646	*-0.98*	-0.00623	*-1.02*
Q1$_{it}$	0.0227	*1.49*	0.0230	*1.46*
Q3$_{it}$	0.0210	*1.17*	0.0229	*1.27*
Q4$_{it}$	0.0409	*2.49***	0.0420	*2.54***
Sample	38 countries		38 countries	
Observations	667		667	
Min. periods	7		7	
Avg. periods	17.6		17.6	
Max. periods	20		20	
F(12,654)	35.9			
F(11,655)			32.4	
Arellano-Bond residual serial correlation test, order 2	Z = 0.04 [0.972]		Z = 0.03 [0.978]	

Note: All variables are in first differences apart from the constant, and variables with an L prefix are in log terms. Figures in italics are t-statistics; *, ** and *** denote significant at the 10%, 5% and 1% level respectively. Numbers in brackets are p-values. Data sources: see Table 2 above.

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[16] Modified Wald test for groupwise heteroscedasticity: $\chi^2(38)$=10,300 [0.000]; Wooldridge test for autocorrelation in panel data: F(1,37) = 17.3 [0.0002]

The one-period lag of our transformed churn variable is highly significant, positive and less than one, showing substantial persistence in the churn process. We find no evidence of second order autocorrelation in the residuals.[17]

There is a significant difference between the churn dummies for countries with a five day or shorter maximum target porting time and those permitting a longer porting time.[18] Countries requiring faster porting times experienced significantly higher churn rates after MNP, whereas there was no significant effect for those with a slower standard. Our alternative MNP variable based on the reciprocal of the target maximum porting time seems to have little explanatory power.

It is difficult to directly interpret the levels of coefficients in a model where the dependent variable has undergone a logistic transformation. However, in Table 5 below, we provide simulation results for the average treatment effect of MNP on quarterly churn rates and the equivalent increase in the average level of churn for countries with porting times of 5 days or less.

Table 5: Estimated MNP average treatment effect on churn and equivalent change in quarterly churn rates for countries with <=5 day porting rate target		
Measure	**Short run**	**Long run**
Average treatment effect[19]	0.253%	0.714%
Implied percentage change compared to sample average churn rate (2.05% as per Table 2)	+13.6%	+34.7%

The short run predicted increase in churn seems consistent with predictions in *ex ante* studies. For example, in a CBA conducted for Hong Kong, scenarios were examined

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[17] Second order autocorrelation would have been indicative of inconsistency, as per Arellano and Bond (2001), pp.281-282.

[18] A Wald test rejected equality between the MNP coefficients: $F(1,654) = 6.88$ [0.0098]

[19] The treatment effects are evaluated with other variables set to their sample averages. For the long run effect, the current period and lagged churn rates converged at 1.77% with MNP, in comparison to 1.06% without MNP.

allowing for increases of 5-15% in the churn rate following introduction of MNP.[20]

Our estimate is slightly lower than the 15% increase in actual switching (as opposed to churn) after MNP reported for South Korea in Kim (2005),[21] but we should expect this given that not all churn involves an inter-operator switch. However, note that because of the strong persistence we find in churn rates, the model predicts that the long run impact of MNP on churn will be significantly higher.

Other results from the models are broadly as expected. Two of the quarterly dummies are not significant, but the Q4 dummy provides evidence of higher churn in the fourth quarter. This may reflect seasonal marketing activity or shifts in demand. The number of operators and the constant term were also found to be insignificant.

All other coefficients in the two models have the expected signs, although income is of only marginal significance. While the cellular density terms in the first model appear to be individually insignificant, this is probably due to multicollinearity; a joint test on them rejects a zero value.[22]

We also tested lags of the MNP variables from 1-4 quarters, but the highest significance level was achieved with no lag.

6.   Modelling the effect of MNP on prices

The cross-country data available for estimating the effect of MNP on retail prices limits us to a relatively simple modelling strategy. In particular, it is not possible to maintain the standard access/usage distinction and other more complex features of

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[20] NERA/Smith (1998), p.66.

[21] Non-switching status fell from 91% to 79.1% of those surveyed; Kim (2005), Table 2.

[22] $F(3,654) = 2.85 \ [0.0366]$

telephony demand models. Again we employ two models using different proxies for quality-adjusted MNP. These models are described below.

The price variable

The proxy for prices is quarterly real average revenue per minute (*RPM*). It is an aggregate measure encompassing all revenues associated with mobile voice services in each country (but excluding revenue from data services).

Use of an average revenue proxy for prices involves a departure from the approach used by most other analyses of regulatory impact on prices in the mobile sector.[23] Prices are more commonly measured for a specified service bundle (e.g. three minutes of calling time).

RPM has some advantages as a price proxy. For example, we have already noted that charges for service components such as handsets and call termination may be affected by MNP, and these might not be captured if we were to focus on some other measure, such as the average price of a three minute call or the price of a bundle of X minutes.

However, the benefits of aggregation come at a price. In particular, previous research into telephony demand has highlighted differences in the determinants of demand for network access and network usage (i.e. calls). *RPM* aggregates these differences away. Other potentially important features of telephony pricing are also obscured by averaging, including handset subsidies, time of day effects, innovation in tariff structures (e.g. bundling schemes and pre-payment offerings) and the mix of different call types (e.g. national vs. international).

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[23] However, the same approach was taken in Hazlett and Muñoz (2004) for their study of the impact of spectrum licensing policies.

Explanatory variables

We use the same two alternative sets of regulatory variables as in the analysis of churn described above. The first model includes dummy variables based on target maximum porting times: one where *MNPTM* was 5 days or less (*MNP5D*) and one where it was 6 days or longer (*MNP6P*). The second model uses the reciprocal of *MNPTM*.

Table 6 below shows how the average of *RPM*, our proxy for price of mobile services, varies in the sub-samples with and without MNP. These statistics paint a surprising picture, inasmuch as MNP appears to increase prices.

Table 6: Relationship between average prices (real revenue per minute) and mobile number portability

Case	Sample mean real revenue per minute (USD)
No MNP	0.192
MNP delivery time <= 5 days	0.206
MNP delivery time 6+ days	0.233

Source: see Table 2 above.

However, these descriptive statistics may be misleading. First, there is a declining trend in consumer prices across all countries during the period, and where MNP was implemented it tended to come later in the time series. This timing effect will tend to bias the MNP averages downward. A similar downward bias may arise because there is a positive association of MNP with quantity of call minutes sold and a negative relationship between quantity and price. In contrast, GDP is positively associated with both MNP and prices, which could lead to an upward bias in the average. To isolate the effects of MNP from other variables, we turn to regression analysis.

Unlike the MNP coefficients in the churn models, the coefficients on MNP variables in the price models are expected to be negative, reflecting stronger competition in markets with lower switching costs. We again treat them as potentially endogenous.

Since each price observation in our dataset is associated with a unique Q (unlike the studies cited above, which fixed Q), we must control for the total quantity of minutes actually delivered by operators in a given country at price RPM. This quantity variable is designated $TOTMIN$, and it too is taken to be endogenous to allow for the simultaneous determination of quantities and prices in mobile markets.

We have a choice of proxies for the effects of market concentration on pricing, including the Herfindahl-Hirschman index (HHI), the one-firm concentration ratio ($CR1$) and the number of network operators (OPS). These are tested alternately in the regression since we do not wish to prejudge the nature of competition in the market. If greater concentration implies weaker competition in mobile telephony markets, HHI and $CR1$ should have positive coefficients when each of them is included, and OPS should have a negative one.

Population density ($PDNST$), a proxy for local cost conditions, should have a negative coefficient reflecting economies of density. Real GDP per capita ($RGDPPC$), a proxy for income, might take a positive coefficient as per the reasoning in Shew (1994) that customers in high income areas will exhibit less price sensitivity, leading to higher prices in such areas.[24] Both of these variables might have a non-linear relationship to average prices, so higher order terms are included in the regression.

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[24] Shew (1994), p.55.

We also include a time trend (*TIME*) to allow for time-varying unobserved effects and quarterly dummies to capture seasonal variations in pricing policies and demand patterns.

Detailed information on service characteristics is not readily available on an internationally-comparable basis. However, since we have time-series cross-section data, characteristics that are jurisdiction-specific may be captured by the use of individual effects.

Other potentially relevant variables were unavailable for the relevant set of countries and periods, including details of marginal price schedules, prices of substitutes (e.g. fixed line services), differences in contract terms, quantities of spectrum allocated in each country, the extent of trans-national ownership or control of operators, availability and relative importance of pre-paid services, advertising expenditure, and regulatory variables other than MNP (e.g. requirements to offer wholesale roaming or access to service providers).

As in the churn model discussed earlier, we include a disturbance term (ε_{it}) and control for individual effects at country level (u_i). Logs are taken of continuous variables, including *RPM*.

To summarise, for country $i = 1...37$ and quarter $t = 1...20$:

$$
\begin{aligned}
\ln(RPM_{it}) = {} & \alpha + \beta_1 \ln(TOTMIN_{it}) + \beta_2 HHI_{it} + \\
& \beta_3 \ln(PDNST_{it}) + \beta_4 \ln(PDNST_{it})^2 + \beta_5 \ln(PDNST_{it})^3 + \\
& \beta_6 \ln(RGDPPC_{it}) + \beta_7 \ln(RGDPPC_{it})^2 + \\
& \beta_8 Q1_{it} + \beta_9 Q3_{it} + \beta_{10} Q4_{it} + \beta_{11} TIME_t + u_i + \varepsilon_{it} + \\
& [\beta_{12} MNP5D_{it} + \beta_{13} MNP6P_{it} \text{ or } \beta_{12} MNPR_{it}]
\end{aligned}
\tag{4}
$$

Summary of prior expectations about coefficients:

$$\beta_2, \beta_4, \beta_6, \beta_{12}, \beta_{13} > 0;$$
$$\beta_1, \beta_3, \beta_5, \beta_7, \beta_{11} < 0$$

6.1   Econometric Results

In this section, we estimate the model described in Section 6 above. Table 7 below

sets out the regression results. As we found when modelling churn, initial estimation

using OLS with fixed effects gave rise to heteroscedasticity and autocorrelation.[25]

Here too, we estimated the models shown below using the Arellano-Bond estimator

with robust standard errors, and diagnostic testing rejects the presence of second

order serial correlation in the residuals. Due to differencing of the data, the fixed

effects are eliminated and the differenced time trend yields a constant.

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[25] Modified Wald test for groupwise heteroscedasticity: $\chi^2(38)=12,800$ [0.000]; Wooldridge test of autocorrelation in panel data: $F(1,36) = 27.8$ [0.000]

Table 7: Price regression results using Arellano-Bond estimator, with *LTOTMIN* and MNP variables treated as endogenous

Variables and statistics	Using MNP delivery time threshold dummies (<=5 days, >6 days)		Using reciprocal of MNP delivery time for countries with MNP	
Dep. variable	*LRPM$_{it}$*		*LRPM$_{it}$*	
	Coef.	*Robust t-stat.*	Coef.	*Robust t-stat.*
LRPM$_{i(t-1)}$	0.453	*4.92****	0.458	*4.83****
MNP5d$_{i(t-1)}$	**-0.0658**	***-3.18******		
MNP6p$_{i(t-1)}$	**-0.0258**	***-0.94***		
RMNP$_{i(t-1)}$			**-0.00676**	***-3.41******
LTOTMIN$_{it}$	-0.343	*-5.46****	-0.332	*-5.35****
LPDNST$_{it}$	-1.30	*-0.25*	-1.32	*-0.27*
LPDNST$^2_{it}$	1.08	*0.77*	1.10	*0.82*
LPDNST$^3_{it}$	-0.133	*-1.15*	-0.132	*-1.17*
LRGDPPC$_{it}$	1.20	*2.37***	1.22	*2.52***
LRGDPPC$^2_{it}$	-0.0306	*-1.29*	-0.0330	*-1.45*
LHHI$_{it}$	-0.158	*-1.35*	-0.169	*-1.34*
Constant	0.00231	*0.9*	0.000556	*0.23*
Q1$_{it}$	-0.0250	*-3.28****	-0.0263	*-3.45****
Q3$_{it}$	0.0115	*1.5*	0.0113	*1.49*
Q4$_{it}$	0.00344	*0.53*	0.00305	*0.47*
Sample	37 countries		37 countries	
Observations	649		649	
Min. periods	4		4	
Avg. periods	17.5		17.5	
Max. periods	20		20	
F(14,634)	974			
F(13,635)			1,290	
Arellano-Bond residual serial correlation test, order 2	Z = -0.90 [0.366]		Z = -1.02 [0.307]	

Note: All variables are in first differences apart from the constant, and variables with an L prefix are in log terms. Figures in italics are t-statistics; *, ** and *** denote significant at the 10%, 5% and 1% level respectively. Numbers in brackets are p-values. Data sources: see Table 2 above.

We found evidence that MNP reduces retail prices, but only when its quality is high. For those countries with MNP delivery times of five days or less, the estimated short run effect of implementing MNP was a fall in real average prices of about 6.6%, after a one quarter lag.[26] The estimated long run reduction was significantly higher, at 12%. We also found a negative MNP coefficient for countries with longer MNP delivery times, but it was not significantly different from zero. However, these

results do not prove that a tighter MNP standard yielded a stronger price effect; we could not reject the hypothesis that the coefficients on the two MNP dummies were equal ($F(1,634) = 2.70$ [0.101]).

The alternative approach of including the reciprocal of each country's maximum time for MNP delivery also yielded a negative coefficient. This model implies a substantial price effect in countries with the tightest MNP delivery standards, but little effect elsewhere (see Table 8 below).

Table 8: Estimated effect of MNP on real average retail prices from 1/*MNPTime* model		
MNP standard	**Short run**	**Long run**
2 hours	-8.11%	-14.97%
2 days	-0.34%	-0.62%
5 days	-0.14%	-0.25%
10 days	-0.07%	-0.12%
20 days	-0.03%	-0.06%

As expected, we found a robust inverse relationship between the number of minutes of traffic and real average prices. Income and population density variables also had the expected signs. Although t-tests on each of the population density terms suggested a lack of statistical significance, a joint test on all the terms strongly rejected a zero value: $F(3,634) = 3.59$ [0.0135]

Neither HHI (shown above) nor alternative proxies for market concentration (*CR1* and *OPS*) proved to be significant. We did find evidence of lower average prices in the first quarter of each year, perhaps reflecting the effect of temporary discounts on packages sold in the fourth quarter, but the constant term is not significantly different from zero.

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[26] We tested lags of between 0 and 4 quarters on the MNP variables, and the signs were the same in all cases, although statistical significance varied. A one quarter lag yielded the highest t-statistic for *MNP5d* and is thus reported here.

7.   Conclusions

Our central finding is that prices fell and churn increased in countries with a five day or better MNP delivery standard, as summarised in Table 9 below.

Table 9: Estimated effect of MNP on real average retail prices and churn rates, for countries with a <= 5 day target maximum porting time		
MNP standard	**Short run**	**Long run**
Average prices (real revenue per minute)	-6.58%	-12.0%
Churn rate, quarterly average	+13.6%	+34.7%

The price result can be compared to the finding in Viard (forthcoming) that there was a 4.4% fall in prices after the introduction of toll-free number portability, which is a different but similar service.[27]

We found no significant effect of MNP on churn or average prices for countries that applied a less stringent target for maximum porting time. For jurisdictions requiring "high quality" MNP, our results are consistent with the presence of significant Type 1 and Type 2 benefits.

Areas for further research

The mobile market data currently available on a consistent basis over time and across countries has limitations when used for modelling the effects of MNP. First, our choice of a five day porting time threshold for examining MNP quality is essentially arbitrary, and additional data in the future should allow a finer distinction to be drawn between the effects of different porting time standards.

Second, we have not been able to control for the varying price of MNP across countries. In some jurisdictions, MNP is free to the subscriber. In others, it can involve significant fees. For example, the system adopted in Singapore in 1997

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[27] See Section 2.3 above.

permitted operators to levy monthly charges on users, but from August 2003 onwards only a one-time administrative fee was allowed.[28] There is also variation in the levels of one-off fees among those jurisdictions that permit them to be charged.[29] While charging could act as a deterrent to usage of MNP,[30] published information on such charges and on other aspects of MNP quality (for instance, whether or not it covers SMS messages) is scanty, and these dimensions are not explicitly addressed in our analysis.

Future research into the effect of MNP will also benefit from the existence of additional time series data from jurisdictions where MNP has been implemented; many countries in our sample had only recently introduced these services.

Also, publication of harmonised cross-country data by supranational bodies such as CEPT, which published most of the MNP implementation and porting time data we used in this paper (see Table 10 in the annex), should make it easier for future researchers to make inter-country comparisons.

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[28] Infocomm Development Authority, Singapore, 2003.

[29] Ovum (2005), Section 3.5.

[30] Buehler, Dewenter and Haucap (2006), p.395.

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Annex: Additional information on the dataset

DEN is each country's average number of mobile telephony users (including both post-paid and pre-paid customers) divided by the country's population. This variable may be subject to varying reporting practices in different jurisdictions. While it is easy to define and measure the number of post-paid subscribers, in most jurisdictions these represent a minority of mobile telephony users. The identities of the remainder, who use mobile telephony on a pre-paid basis, are often unknown to their network operators. As a result, network operators generally use a formula to estimate the number of active customers, typically treating a subscriber as active if his or her phone has been used within a set number of months. While we understand from Merrill Lynch (2004) that these formulae may vary across the sample, we have no details of the differences.

This caveat also affects the *CHURN* variable, which is a quarterly average of monthly actual and imputed[31] disconnections from networks as a proportion of the average number of users in each period.

Gross domestic product in real USD terms per capita (*RGDPPC*) was calculated for OECD countries based on local currency real GDP figures and GDP deflators from the OECD quarterly national accounts database. Exchange rates were taken from IMF International Financial Statistics. Figures for non-OECD countries are taken from the IMF World Economic Outlook database (September 2004), and are annual data, rather than quarterly. This treatment of GDP in non-OECD countries is not ideal, but as no quarterly national accounts data were available for these countries it was unavoidable. In any event, the coefficients on GDP are not the focus of our analysis.

Table 10: Sample coverage and MNP data					
Country	Churn observations	RPM observations	MNP implemented ("-" if not implemented by 2Q04)	Target maximum porting time (days)	Main source for date of MNP implementation
Argentina	10	20	-	-	
Australia	20	20	3Q01	0.0833	Regulator
Austria	20	20	-	-	ECC/CEPT 2005
Belgium	16	13	3Q02	2	ECC/CEPT 2005
Brazil	20	20	-	-	
Canada	20	20	-	-	Telegeography
Chile	20	19	-	-	
Colombia	16	12	-	-	
Denmark	8	20	3Q01	5	ECC/CEPT 2005
Egypt	7	8	-	-	
Finland	20	20	3Q03	5	ECC/CEPT 2005
France	20	20	3Q03	30	ECC/CEPT 2005
Germany	20	20	4Q02	6	ECC/CEPT 2005
Greece	20	20	3Q03	1	ECC/CEPT 2005
Hong Kong	20		1Q99	2	Regulator
Hungary	12	19	2Q04	14	ECC/CEPT 2005
India	20	20	-	-	Telegeography
Ireland	11	11	3Q03	0.0833	ECC/CEPT 2005
Israel	20	18	-	-	
Italy	20	20	2Q02	5	ECC/CEPT 2005
Japan	20	20	-	-	
Malaysia	20	20	-	-	
Mexico	20	20	-	-	
Netherlands	20	20	2Q99	10	ECC/CEPT 2005
New Zealand	20	20	-	-	
Norway	14	20	4Q01	7	ECC/CEPT 2005
Poland	20	9	-	-	ECC/CEPT 2005
Portugal	20	20	1Q02	20	ECC/CEPT 2005
Russia	17	17	-	-	
South Africa	20	20	-	-	
Spain	15	20	4Q00	4	ECC/CEPT 2005
Sweden	20	20	3Q01	5	ECC/CEPT 2005
Switzerland	20		1Q00	5	ECC/CEPT 2005
Taiwan	12	15	-	-	
Thailand	20	4	-	-	
Turkey	13	8	-	-	
UK	20	20	1Q99	9	ECC/CEPT 2005
US	16	17	4Q03	0.104	Regulator
Venezuela		19	-	-	

[31] Pre-paid users inactive for a specified period.