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Regulation and Performance: Evidence from the  
Telecommunications Industry

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September 2009

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# Declaration

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# Abstract

This thesis proposes an empirical analysis of performance in the telecommunications industry, in relation to the regulatory framework in place in the sector. We first provide an introduction and overview of the related literature on performance measurement and regulatory institutions. Chapters 1 and 2 focus on firm-level measures of productivity components and, in particular, of technical change and efficiency. This analysis is motivated by the form of incentive regulation in force in the industry, which links future price increases allowed by the regulator to certain measures of the operators' performance. Chapter 1 investigates embodied technical change in the U.S. industry relying on firm-level panel data. It builds on the definition of capital as a sum of vintages of different qualities and incorporates this definition into a cost function. Estimates of embodied technical change in this sample vary depending on the specification, but do not appear large enough to affect productivity. Chapter 2 analyses the variation of efficiency over time in the same sample of U.S. operators. This is done by applying estimators that allow for (freely) time-varying efficiency to an input stochastic distance function. Estimates confirm that standard panel estimators, which are commonly used by regulators to assess relative efficiency, do not adequately capture the time-varying component of efficiency. Finally, Chapter 3 studies how cross-country differences in sector performance, measured in terms of access to mobile networks, can be explained by regulatory and country governance. In particular, using a system approach, it considers whether the impact of regulatory governance on penetration can be an indirect result of country institutions. In addition, feedback effects between access to infrastructure and income are incorporated in the analysis. The analysis is carried out on a panel of low and middle-income countries. The empirical results show that the establishment of a separate telecommunications regulator is associated to higher penetration levels, and that this is more important for low-income countries. The effect is partly related to the quality of wider country governance rather than sector-specific institutions.

# Introduction

In the last decades, infrastructure sectors have been affected by significant and widespread reforms. This trend has gradually led to the commercialization and the privatization of the national incumbent operator, and in some cases to radically restructuring the sector.

The telecommunications industry is one of the sectors where reform has developed more rapidly, in line with the increasingly central role that telecommunications services play for both individuals and firms. The importance of the sector goes beyond its share of the Gross Domestic Product of a country, as telecommunications services transform pervasively markets and social interactions, and are therefore considered enabling services for other economic sectors. Moreover, the industry is affected by major technological and market changes, which are more rapid than in other infrastructure sectors, and are arguably among the factors that have allowed a faster pace of reform.

While some countries have been precursors of the reform process, the trend has acquired a worldwide dimension in the nineties, when European Union member states liberalized their telecommunications markets and reform spread also to less developed countries, mainly through support from international organizations. Extensive reforms have mainly consisted of the privatization of the incumbent, the introduction of competition, particularly in the mobile market, and the establish-

ment of separate (non-Ministry) regulatory agencies.

In parallel, economic theory has extended its focus from efficient pricing under natural monopoly (Braeutigam, 1989) and optimal incentive schemes for regulated monopolies (Laffont and Tirole, 1986) to the study of issues arising in an industry characterized by competition between a dominant vertically integrated firm and new entrants (e.g. survey by Armstrong et al., 1994). The question itself of the conditions under which competition is preferable to monopoly has also been addressed (Armstrong and Sappington, 2006).

In addition, the literature has increasingly emphasized the limitations of the assumption that the regulator is “omniscient, benevolent, and able to fulfill any promises he makes,” (Armstrong and Sappington, 2006) and therefore the need to account for the complexity of the environment has emerged. For instance, with respect to sector reform, the choice between a regulated monopoly and a competitive market is affected by a variety of institutional factors, such as the degree of regulatory independence and accountability.

As the practice of regulation has progressed, some areas of analysis have attracted growing attention by professional and academic economists. For instance, the move to incentive regulation has renewed interest in the measurement of productivity and efficiency. These are indicators that are used by regulators as a basis for setting future price rules and, at the same time, they are considered among the measures of the effectiveness of regulation.

The aim of this thesis is two-fold. Firstly, relying on firm-level panel data, it investigates the question of embodied technical change and of relative efficiency in the telecommunications sector. As mentioned above, these are relevant measures in

the regulatory process of setting the rules that constrain future prices.

Secondly, following the literature on the effects of institutions on regulatory policy, it studies the interplay between regulatory governance (e.g. the establishment of a separate regulator) and country institutions in promoting telecommunications penetration. This analysis is carried out on a cross-country panel dataset.

The remainder of this introduction is organized as follows. The next section outlines the main characteristics of the telecommunications sector. Sections 3 and 4 provide some background on the strains of the literature this thesis builds on. Finally, Section 5 summarizes the contents of the three remaining chapters.

## **1. Economic Characteristics of the Sector**

Telecommunications networks are organized in a hierarchical way. In simple terms, subscribers are connected, through the local loop, to the network's local exchanges which, in turn, are connected to higher levels of the network architecture. This structure allows customers connected to different local exchanges to communicate with each other.

In the traditional fixed telephony market, the sector's structure was characterized by a vertically integrated operator. Mostly due to the substantial fixed costs incurred for network deployment, some portions of the infrastructure were provided by a single firm. The government's presence in the sector has thus been generally justified on natural monopoly arguments, because market forces alone could not achieve an efficient outcome.<sup>2</sup>

In some countries, notably in Europe, in this early phase of market development,

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<sup>2</sup>The first-best competitive market outcome of uniform marginal cost pricing cannot be achieved if the firm maximizes profit in the absence of lump-sum transfers, given the substantial fixed costs involved in network investment. Instead, under these assumptions, the optimal outcome with uniform prices is given by the Ramsey pricing rule. The firm sets a mark-up on marginal costs which is inversely proportional to the price elasticity of demand, under the assumption of zero cross-price elasticities and income elasticities (Braeutigam, 1989).

the government was initially the owner of the network operator and subsequently its role evolved into that of regulator of the privatized incumbent. However, this is not a universal feature of the sector as, for instance, in the U.S. and in Canada, private ownership in the telecommunications sector dates back from the early days of telephony (Wallsten, 2006).

Following the sector's liberalization, competitors have entered one or more market segments. In particular, when microwave links were deployed in the long-distance network, it became apparent that this part of the infrastructure could be replicated also by competitors. Entry has therefore taken place, at least initially, in markets where alternative infrastructure could be rolled-out at reasonable cost and margins were attractive (e.g. business customers).

Where new entrants have not built their own infrastructure, they have relied on the incumbent's network to provide services to end-users. The distinction between retail and wholesale services has therefore become relevant in the new market structure. Retail telecommunications services are services provided to end-users. Wholesale telecommunications services, on the other hand, are network component services that the incumbent provides to other industry players so that they can compete and provide an end-to-end service to their retail customers.

For new entrants, purchasing wholesale services represents, at least to a certain degree, an alternative to rolling out their own infrastructure. Examples of wholesale services required to provide retail long-distance calls are origination, transport and termination services. The availability of wholesale services usually rests on the imposition of regulatory obligations on the vertically integrated incumbent.

However, given that the incumbent is also a competitor in the retail market, there

are concerns it may act anti-competitively in the provision of inputs to alternative operators. One of the key issues so far, not surprisingly, has been competitors' access to the incumbent's infrastructure, in particular to the copper network between the local exchange and the customer's premises. As a result, the theory of access pricing in the telecommunications industry has attracted considerable interest in the literature (e.g. see review by Armstrong, 2002).

The problem of access is also related to the nature of competition in the market, i.e. whether new entrants are service providers or infrastructure-based operators. Service providers have minimal infrastructure and therefore rely on the incumbent's network for the provision of services to their retail customers. New entrants may also choose to bypass certain parts of the incumbent's infrastructure by investing in network deployment. However, this is especially expensive in the case of the local loop, given the substantial costs of trenches and ducts.<sup>3</sup>

Regulation also needs to be adapted to a changing environment, affected by technology and market transformations. In particular, many operators are currently investing in next generation networks (NGNs). These allow both voice and data communications to be transformed into 'packets' of information that can be transmitted without the need to keep a dedicated circuit for a specific phone call, as required instead in traditional networks. As a consequence, this move should improve the potential for optimizing network capacity as well as cost efficiency, but it also involves substantial investment by incumbent operators. In addition, as demand

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<sup>3</sup>However, alternative networks have been deployed in some high-margin urban areas. In general this is a simplified description of competitors. In fact, there are a variety of intermediate solutions between pure resale and a fully-fledged infrastructure operator. For instance, in the broadband market, incumbents in EU member states are required to offer a range of wholesale access services which allow competitors to choose the preferred level of investment. As a result, a different mix of wholesale access services, as well as of alternative self-provided infrastructure, has emerged in European countries.



for high bandwidth services (e.g. IPTV and video on demand) increases, operators are upgrading their networks investing in fiber deployment, which allows for higher transmission speed compared with traditional copper-based access networks.

In light of this, it is difficult to predict in which direction regulatory practice will move. The move to next generation networks, in conjunction with current trends towards separation of the network (or the access network) from service provision, may in fact slow down the roll-back of regulation rather than increase its pace. As some commentators conclude, “the telecoms sector is undergoing significant and concurrent changes, with the only common theme being an evolution from the relative simplicity of the past towards complexity in the future.” (Cave and Corkery, 2006)

## **2. Empirical investigation of telecommunications technology**

As mentioned above, this thesis is related to two strands of the literature on the telecommunications industry: at a microeconomic level, the structure of the production technology and, at an aggregate level, the role of regulatory and country institutions.

In the seventies and the eighties, the methodology to estimate cost and production functions was applied to telecommunications data to explore the nature of returns to scale in the sector. This was mainly related to the policy question of whether a competitive market structure would be efficient or, as argued by incumbents, whether the provision of telecommunications services was a natural monopoly (e.g. see review in Correa, 2003).

In the early seventies, the seminal papers estimated production functions and were broadly consistent with a finding of constant returns to scale. Subsequent studies, in line with a more general trend in microeconomics, focused on the dual

representation of technology through cost functions. For instance, Fuss and Waverman (1981) applied the translog cost function proposed by Christensen et al. (1973) to time-series data from the Canadian telecommunications sector.

The interest in the structure of the sector, and especially the presence of economies of scale, was further stimulated by the decision by the U.S. Department of Justice to break-up the dominant and vertically integrated operator AT&T in 1984. Several academics focused on the estimation of cost functions and proposed various tests to investigate the question of whether the Bell System was a natural monopoly and, related to this, assess the economic merits of its structural separation.<sup>4</sup> While Fuss and Waverman (1981) and Evans and Heckman (1984) concluded that competition in long-distance services was possible, other studies (e.g. Röller, 1990) found that the data was consistent with a natural monopoly.

In the above-mentioned studies, time series data was used to estimate cost functions. A major change in the established methodologies to study the technology of telecommunications operators was the use of panel data by Shin and Ying (1992), following the development and diffusion of panel data in applied econometrics. This allowed overcoming the small sample problem that had affected earlier studies. In addition, Shin and Ying (1992) used output quantities rather than deriving them from revenue and price data, thus addressing the measurement error problem that could be caused by this procedure.

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<sup>4</sup>A resurgence of the topic in recent years, even though in the less extreme version of “functional” separation, has been spurred by concerns about abusive practices by vertically integrated incumbents. In this form of separation, some telecommunications incumbents have created separate access divisions that provide, on an equal basis, inputs to the incumbents’ retail arms and to competitors.

However, the debate has not focused on the subadditivity of the cost function. The move to next generation networks (i.e. parts of the network are replaced by fibre and voice and data are transmitted over the same infrastructure) could well reignite the debate on the incumbent’s scale advantages due to the substantial sunk investments and fixed costs involved in the migration to the new environment.

However, following market liberalization in many jurisdictions, the question of whether the telecommunications industry (or at least some market segments) is a natural monopoly has gradually lost interest. The estimation of the operators' costs, on the other hand, is still very relevant for public policy purposes.

Productivity measures are relied upon by some regulators when reviewing the price cap regime that will apply to the operators they oversee. When calculating productivity, a simple growth accounting formula would not allow disentangling productivity's various components, including economies of scale and scope, allocative efficiency and technical change (Fuss and Waverman, 2002). For this reason, knowledge of the production technology is still essential.<sup>5</sup>

In general, interest in efficiency and productivity measurement is also motivated by the expectation that economic regulation should encourage firms to improve their efficiency. In consequence, one of the main questions investigated in empirical papers has been whether the shift from rate of return regulation to incentive regulation has resulted in increased efficiency. This issue has mostly been analyzed with reference to the U.S. telecommunications sector, given data availability and the gradual move of U.S. states to incentive regulation.<sup>6</sup> The estimation of productivity and efficiency in the U.S. telecommunications industry has been the subject of several papers, which adopt Data Envelopment Analysis (DEA) (Majumdar, 1997; Resende, 2000; Uri, 2001) and stochastic frontier techniques (Resende, 1999). DEA is a linear programming technique which compares each operator with the production or cost level of a peer group, while Stochastic Frontier Analysis (SFA) compares a given firm

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<sup>5</sup>The investigation of the firm's costs is further complicated by the multi-product nature of telecommunications operators, which use the same infrastructure to provide different services, such as access to the network and different types of calls.

<sup>6</sup>The sector is regulated both at federal level, by the Federal Communications Commission, and at state level. States have moved to various forms of incentive regulation at different points in time.

to a frontier, defined either in terms of output or cost levels, derived from sample data using a statistical methodology.

In other recent applications, comparisons among international operators have been drawn to investigate other policy questions, for instance whether liberalization and privatization have an effect on efficiency.<sup>7</sup>

In this thesis, the estimation of cost and production functions has two objectives. In Chapter 1, it is the approach chosen to estimate embodied technical change, i.e. the improvement in assets' quality in new vintages of assets purchased by a firm. Most papers have modeled technical change as a time trend, therefore assuming implicitly that all technical change is disembodied. This latter assumption implies an organizational type of progress, which would affect the capital stock already in place by making it more efficient without requiring any new investment in physical assets.

In Chapter 2, a stochastic distance function is estimated to compare the results from different estimators of relative efficiency. A distance function is an extension of the concept of production function to a multi-product technology. The analysis is motivated by the question of whether estimators commonly employed in empirical studies of relative efficiency provide potentially misleading results as to the time profile of efficiency. In particular, the study in Chapter 2 applies estimators allowing for (freely) time-varying, rather than constraining its variation over time to pre-specified functions.

### **3. The role of institutions in the telecommunications sector**

The question of the wider institutional framework within which a regulator operates is the focus of a strand of the literature on infrastructure sectors which has

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<sup>7</sup>Picazo-Tadeo and Quiros-Romero (2004).

acquired prominence since the reform wave of the nineties, in particular after the publication of the seminal work by Levy and Spiller (1994). In essence, the authors emphasized the link between a given country's institutions and regulatory governance in infrastructure industries.

As argued by Brown et al. (2006), one could think of regulatory governance as the “how” of regulation, as opposed to the “what” which instead refers, for instance, to the regime governing interconnection and price controls. Both detailed regulations and regulatory governance need to take into account a country's specific conditions, especially in less developed countries, where the divergence between written rules and their actual implementation may be more significant. An empirical study on the link between country institutions and regulatory governance has been carried out by Gual and Trillas (2006).

The interest in regulatory governance is spurred not only by its interrelation with country institutions, but also by the impact it may have on the performance of an infrastructure sector, commonly defined in terms of access to infrastructure. For the telecommunications network, this is referred to as penetration, measured as the number of subscribers (fixed or mobile) per 100 population.

In fact, understanding the factors leading to improved access to infrastructure in some countries appears difficult – we observe countries with very different socioeconomic characteristics having similar proportions of telecom subscribers. One explanation that has been proposed in the literature is the wave of telecom sector reform which, in the mid-90s, led many countries to commercialize and (in many cases) privatize their national incumbent telecom company; to liberalize telecom markets; to introduce competition, particularly in the mobile sector; and to establish separate

(non-Ministry) regulatory agencies.

Much attention has been devoted by international institutions, such as the World Bank, to the question of how to reform a sector so as to improve access to infrastructure, be it communications networks or energy and other utilities. Previous studies that have analyzed the impact of reform packages on fixed telecommunications penetration include Ros (1999), Wallsten (2001) and Estache et al. (2006).

Within this reform program, particularly important is the establishment of a separate regulator, i.e. one that is autonomous (or “independent”) both of the government and of the incumbent telecom operator. Gutierrez (2003) has further developed the approach of the aforementioned studies by introducing a better description of a given country’s regulatory characteristics.<sup>8</sup>

However, this literature does not, with some exceptions (e.g. Gasmi et al. (2007), pay much explicit attention to the institutional setting within which the new regulatory agencies operate such as political structure, the rule of law or the degree of economic openness.

The last chapter of the present thesis takes a fresh look at the relationship between regulation and performance in the telecommunications sector and, in particular, to the role of the institutional setting. Unlike previous studies, it treats regulatory governance as an endogenous variable which is related, among other factors, to general country governance.

Another key difference between Chapter 3 and previous papers is given by the explicit consideration of the potential impact of telecommunications penetration on aggregate income. In fact, the importance of the telecommunications sector in

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<sup>8</sup>Similar questions have also been analysed in other infrastructure industries. In the electricity sector, the first study that has addressed the question of whether regulatory governance has an effect on generation capacity using panel data is Cubbin and Stern (2006).

improving a country's income level is a major complication when analyzing the factors that influence telecommunications penetration rates. While it is conventional to assume that income is among the variables affecting the demand for infrastructure capacity and services, the economic feedback impact of telecommunications infrastructure capacity also needs to be modelled if we are not to have a misleading picture.

#### **4. Summary of the thesis**

##### **4.1 Technical change and efficiency: firm-level estimates**

The aim of the present thesis is to analyze and contribute to two strands of the economic literature which have found widespread application in infrastructure industries, as summarized above. Firstly, it focuses on the firm-level measurement of the productivity of telecommunications operators, specifically of embodied technical change and productive efficiency. Secondly, it analyses the role of regulatory institutions in promoting access to infrastructure, as well as the interplay between governance, penetration and aggregate income.

Chapter 1 presents the results from the estimation of technical change in a sample of telecommunications operators, taking explicitly into account innovations embodied in new vintages of capital. The relevance of embodied technical change for telecommunications operators is due to the fact that regulators often take past performance as an indication of future productivity gains that an operator can achieve. As a result, regulators may set tight price rules for a regulated firm, in the expectation that it will achieve further productivity gains. Therefore, an investment in more efficient equipment might result in a tighter price rule in future periods. However, given that this investment would increase capital costs for the regulated firm

– unlike disembodied technical change – it may potentially result in a problem of lower incentives to invest in more efficient vintages of capital.

Despite the large interest in the embodiment hypothesis in numerous fields of economics, there have been very few studies on the role of embodiment and its implications for regulated sectors. Chapter 1 empirically investigates this topic in the context of the U.S. telecommunications industry using a panel of 28 incumbent Local Exchange Carriers over 14 years (1990 – 2003).

Unlike previous studies on telecommunications that interpret average age as a proxy for embodied technical change, without an explicit derivation, Chapter 1 builds on a framework developed by Sakellaris and Wilson (2004). In their study, building on the theoretical definition of embodiment, they assume a constant yearly rate of technical change throughout the sample period and they define capital stock in efficiency units with respect to a base year.

In addition, the analysis relies on the most disaggregated data available for investment in different categories of telecommunications plant by the large U.S. telecommunications operators in order to calculate surviving capital as accurately as possible. The assumption that all assets share the same rate of embodied technical change is partially relaxed: capital assets with long lives (e.g. buildings, poles, conduits) are kept separate from telecom equipment, which typically needs faster replacement and is characterized by higher technical change.

In general, an inherent difficulty in estimating embodied technical change is the impossibility to isolate the other inputs and the outputs which are directly associated to a specific vintage of capital, as would be required by the theoretical formulation of the problem in order to identify the effects of different vintages on output. In the



estimation, this problem is somewhat alleviated by the cross-section variation in the sample.

According to our results, depending on the specification, the estimated rate of embodied technical change ranges between 0% and 5.7%. At the lower end of the interval, the variation in investment across firms and across time is not sufficient to identify differential capital quality (i.e. embodied technical change). At the upper end of the interval, the estimate is about one percentage point higher than the results obtained by Sung (2002) following a different methodology.

While some variation in results is expected, evidence for embodied technical change does not appear conclusive in this sample. In addition to methodological and data issues already highlighted, this may also reflect slower network modernization in the period under analysis compared to a longer timeframe (e.g. as analyzed by Sung, 2002). While our sample period includes the 1999-2000 investment “bubble” in telecommunications companies and infrastructure, average gross investment for the incumbent Local Exchange Carriers (LECs) does not appear very dynamic. In particular, average gross investment was either flat or moderately growing until 1999, underwent a step increase in 2000 and declined very markedly until 2003.<sup>9</sup>

This is consistent with two stylized facts about the telecommunications “bubble”. As reported in the 2003 OECD Communications Outlook, “overinvestment in facilities” took place mainly in backbone markets and intercontinental links. Therefore LECs were less affected given that they did not focus on such markets. In addition, the “irrational exuberance” of the late 1990s concerned more the avail-

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<sup>9</sup>The sample period contrasts markedly with the significantly longer period of analysis covered in Sung (1998, 2002) which spans from 1951 to 1991, during which major technical innovations transformed the industry, e.g. the evolution from mechanical to electronic switches and the digitalization of the access network.

ability of financial capital to invest in new services and new geographic markets rather than heavy investment in physical capital (OECD, 2003). Again, this phenomenon in the U.S. affected mostly Competitive Local Exchange Carriers (CLECs) and long-distance operator rather than the operators in our sample.

Using the same sample of operators, Chapter 2 estimates an input distance function to analyze the firms' technical efficiency. This is motivated by the increasing use by utilities regulators of the measurement of relative performance when implementing incentive-based regulation in a given infrastructure sector. However, not all studies emphasize adequately that different assumptions and therefore different estimation techniques may produce contrasting results. In particular, the assumption, made in some approaches, that inefficiency is constant over a period of time appears difficult to justify.

Compared with previous studies on the telecommunications industry, the new element in the analysis is the application of panel estimators that account for (freely) time-varying inefficiency, proposed by Greene (2004) and Greene (2005a). Rather than constraining the pattern of inefficiency over time on the basis of a given function, these estimators allow the inefficiency component to vary randomly from one period to another.

In addition, the framework of the analysis is an input distance function. The advantages of a distance function are two-fold: firstly, it provides a natural extension of the production function for a multi-output firm; secondly, compared to a cost function, it does not require input prices data and the behavioral assumption of cost minimization. However, related to the latter point, the distance function provides a partial picture as it only allows for the estimation of technical efficiency, rather

than overall economic efficiency. Another limitation of the analysis is that the model assumes all inefficiency to be time-varying, therefore leaving any element which does not change over time (e.g. inefficient management practices) in the firm-specific effect.

Results show that the sample of U.S. operators included in the study shows considerable variation in productive efficiency over time, which suggests that an estimator that abstracted from such variation may be misleading. In addition, it is interesting to consider the time pattern of average efficiency in conjunction with market and regulatory changes that took place over the period of analysis. As explained above, one such event was the telecommunications “bubble”. Average estimated efficiency in our sample follows broadly the same trend as the investment in the telecommunications market: it increases during the period of market growth, reaches its highest value in 2000 in coincidence with the peak of the market, after which it falls.

Average efficiency in the sample can also be related to another major change that took place in the U.S. telecommunications market in 1990 – 1991, namely the move from rate of return regulation to incentive regulation. In the first years of the sample, until 1994, estimated efficiency declines compared to its 1990 level. This may appear surprising, given that the introduction of incentive regulation was aimed at promoting efficiency. However, the effect may have been lagged or other factors may have counterbalanced regulation’s impact on efficiency. Finally, the finding confirms the results in Uri (2001) who concludes that average technical efficiency has not increased between the 1988-1990 time period and 1991-1999, i.e. before and after the introduction of incentive regulation across the U.S.

## 4.2 Regulatory institutions and telecommunications penetration

The last chapter of this thesis studies the relationship between regulation and performance in the mobile telecommunications sector, taking account of the economic impact of telecommunications infrastructure on aggregate income and of the role of country institutions in promoting economic growth. We address these questions by estimating a system of equations for a panel of 93 low and middle-income countries over the 1995 - 2004 period.

The focus on the mobile market is motivated by two types of considerations. Firstly, mobile markets have often been characterised by a relatively competitive market structure almost from service launch, while the liberalization of fixed markets has somewhat lagged behind. This raises the question of whether results on regulatory governance from previous studies, which concerned the fixed market, would hold also in a more competitive market, where regulation is usually more limited.

Secondly, mobile communications have enjoyed impressive rates of growth across low and middle-income countries in recent years and therefore seem to represent better telecommunications infrastructure in our sample. The average number of phones per 100 population in our sample has increased from less than 1 in 1990 to around 40 in 2004. Compared to the availability of fixed lines (around 17 lines per 100 population as of 2004), the success of mobile communications is even more staggering. High connection charges and long waiting lists, as well as the substantial investment required to develop extensive fixed networks, have held back traditional communications networks and favoured the expansion of mobile services.

In terms of methodology, the main advantage of the approach followed in the

present thesis is that we estimate a system of equations rather than a single reduced-form equation. This should allow investigating more thoroughly the interactions among the above mentioned variables, which perforce are either ignored or only implicitly modeled in the single equation reduced form model. In this respect, this study's contribution is the explicit inclusion of regulatory governance and country institutions in the framework of analysis, as well as the treatment of regulation as endogenous.

In addition, the present dataset includes a reasonably large set of developing countries only (93 countries). Hence, we have a more homogenous group of countries than in most previous studies. The latter have often combined both developed and developing countries and therefore implicitly assumed that a common model held for very different countries.

The limitations of the analysis are mainly related to the measurement of the governance variables. Firstly, in common with other studies (but see a new dataset in Montoya and Trillas, 2007) regulatory governance is measured on the basis of formal characteristics of the legal framework, such as the existence of the regulator and the way it is funded. However, this may not coincide with the actual governance of the regulatory authority i.e. how the regulator operates and is allowed by the government to operate in practice.

Secondly, related to the previous point, the only available measure of regulatory governance for all our countries is a dichotomous variable which takes value one when a certain characteristic is present (e.g. regulator separate from Ministry, autonomous funding) and zero otherwise. This type of variable does not allow us to quantify differences between countries' regulators in any detail. However, compared to an

index-type variable, it is more suitable for system estimation.

Thirdly, country institutions are among the explanatory variables in the system. There is an open question of the potential endogeneity of country institutions. This is a key and hotly debated theme in the empirical literature on institutions and growth (Durlauf et al., 2005). In the present study, the issue is addressed by treating the proxies for country institutions as predetermined for the year in question. This approach is motivated by institutions' strong persistence over time, especially in relation to the limited timeframe of the present sample.

The evidence we present confirms the positive effect of regulatory institutions on telecommunications penetration. We find evidence that the existence of a separate industry regulator is associated to higher penetration rates of mobile telecommunications in developing countries, with estimates varying depending on the specification. In particular, we find a different pattern between low-income and middle-income countries. The marginal effect of a separate regulator in lower-income countries is higher compared to middle-income countries. On this basis, the establishment of a separate body in charge of regulating the industry appears especially crucial in lower-income countries. This can be explained by the fact that, in middle-income countries, market forces may be more important for encouraging the sector's development, compared to the regulatory agency.

According to the estimates obtained in the system, there is a positive relationship between country institutions, proxied by the development of the financial sector, and sector-specific governance. Therefore, part of the positive effect of regulatory institutions on the telecommunications sector may be related to overall country governance. However, it is likely the importance is underestimated in the present

study. The importance of better proxies for country institutions is highlighted by the higher coefficients and significance obtained when the Kaufmann et al. (2006) indexes for the rule of law and quality of governance are considered.

Finally, the impact of mobile telecoms infrastructure on per capita GDP is not found to be significant except for a subset of countries over a longer 15-year period. This may be related to unobserved characteristics of the sub-sample, which was selected on the basis of the availability of data for a longer period, but may also indicate that the impact of mobile infrastructure on GDP cannot be detected over a short time span. Moreover, the analysis is carried out on aggregate data, which are not well suited to uncover the mechanism through which mobile phones can affect income and growth (for a micro-level approach see the seminal contribution by Jensen, 2007).

## **Part I**

# **Technical Change and Efficiency: Firm-Level Estimates**



## Chapter 1

# Embodied Technical Change in the U.S. Telecommunications Industry

### 1.1 Introduction

The measurement of productivity and the identification of its components have attracted considerable attention in a variety of industries. Resende (1999) and Uri (2000, 2001), among others, have examined productivity growth in the U.S. telecommunications sector. One of the findings of these studies is that technical change represents a high share of productivity gains (Uri, 2001). However, the source of technical progress has not been investigated so far. Most papers have modelled technical change as a time trend, therefore assuming implicitly that all technical change is 'disembodied'. This assumption implies an organizational type of progress, which would affect the capital stock already in place by making it more efficient *without*

requiring any new investment in physical assets.

However, technical progress in the telecommunications sector may be due to investment in new and more efficient capital, that is it may be 'embodied' in new capital rather than disembodied. Embodiment could be defined as the 'extent to which in the long run productivity growth is due to improvements in the quality of machinery and equipment' (Oulton, 2007). Therefore embodied technical change is reflected in improvements which affect the efficiency only of new capital goods rather than all the capital stock. The deployment of advanced communications networks by telecommunications operators<sup>1</sup> indicates the importance of innovation through new infrastructure in this industry.

This question becomes even more relevant for the sector because of its potential interaction with the type of regulation incumbents are subject to. In the 1990s, traditional cost of service regulation was replaced by price-based incentive regulation of telecommunications services in most U.S. states. Under price cap regulation, baskets of certain services provided by incumbent Local Exchange Carriers (LECs) are subject to a cap, which limits the expected price increase in the regulated firm's average prices. The level of the cap for regulated firms is positively related to their productivity gains relative to the economy.<sup>2</sup> Therefore, given the form of regulation, the more productive the telecommunications industry compared to the rest of the economy the tighter, *coeteris paribus*, is the price cap.

If embodied technical change were found to be empirically relevant for regulated operators and if its level had a large impact on productivity growth, past investment

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<sup>1</sup>For instance, the Network Modernization Plan agreed by Verizon with the Pennsylvania regulator in the 1990s.

<sup>2</sup>The level of the cap is "last year's cap, increased by inflation in the overall economy, but adjusted for productivity differences between the average firms in the overall economy and the industry being regulated [this difference is the 'X factor' itself]" (Tardiff and Taylor, 2003).

decisions of regulated firms would affect their future allowed price levels. If this was the case, the regulator would have to consider whether the current form of regulation reduced incentives to invest.<sup>3</sup>

Despite the large interest in the embodiment hypothesis in numerous fields of economics,<sup>4</sup> there have been very few studies on the role of embodiment and its implications for regulated sectors.<sup>5</sup> This chapter empirically investigates this topic in the context of the U.S. telecommunications industry.

A key point in the analysis is adopting a definition of the capital stock which allows incorporating the potential embodied (or investment-specific) technical change.

The capital stock for a given asset is usually constructed by summing together different 'vintages' of capital which may not be homogenous, given that older equipment may not be as productive as the latest equipment available. For this reason, the question of aggregating investments made in different periods into a single measure of capital stock is fraught with difficulties and this problem has been recognised in the literature since the debate on vintage effects in the 1950s. In order to address this problem, statistical agencies are working towards the construction of capital measures that are adjusted for quality changes in successive vintages of capital. They develop quality-adjusted price indexes, for instance based on hedonic methods, to derive more accurate estimates of the capital stock.

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<sup>3</sup>If prices are not quality-adjusted, investment in more productive vintages of capital should be accompanied by higher Total Factor Productivity (TFP) levels.

<sup>4</sup>The importance of embodied technical progress has been investigated widely in the theoretical and empirical growth literature, since the debate of vintage effects in the 1950s and 1960s (e.g. Solow, 1957; Jorgenson, 1966). In addition, more recent papers on embodied technical progress have also focused, for instance, on optimal investment with energy saving technical progress (Boucekkine and Pommeret, 2004) and on the estimation of embodied technical progress in the manufacturing sector, using plant data (Sakellaris and Wilson, 2004) or stock market data (Laitner and Stolyarov, 2004).

<sup>5</sup>With the exception of Mandy (2004) on the use of cost models for regulating telecoms interconnection charges and Sung (1998, 2002), who estimated the rate of embodied technical change in a sample of US telecommunications operators.

In the present study, we follow a different approach, which traces back to the early embodiment literature. Following Sakellaris and Wilson (2004), we define capital as the sum of investment flows in assets, where each year's investment is adjusted by a factor that reflects the lower quality of older vintages compared to most recent ones. This factor represents investment-specific technical change. As a result, this approach reduces the value of capital expenditure in the early years of the sample.

However, this adjustment may not necessarily reflect reality. In the presence of investment-specific technical change, one would assume that the price of equipment would tend to decline faster or rise more slowly than would be otherwise and that, in consequence, a firm may use more capital. However, this effect may not be reflected in the firm's capital expenditure: lower prices (i.e. better quality) and higher quantities may compensate each other.<sup>6</sup> The implication is that the present approach, by artificially reducing the value of capital expenditure, may distort the estimated coefficients.

Using a panel of 28 U.S. telecommunications operators over 14 years, the present study obtains estimates for the rate of embodied technical change ranging from 0% to 5.7% per year, depending on the methodology. These estimates are not very different from those obtained in previous papers using a narrower sample of U.S. telecommunications operators and a different methodology (Sung 1998, 2002).

The finding of a positive rate of embodiment as opposed to a finding of no embodiment has implications for the value of the capital stock. When the latter is constructed assuming a positive rate of embodiment, the capital stock (in efficiency

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<sup>6</sup>Given that we only observe capital expenditure, we cannot separate what we could call the *quantity* effect from the *quality* effect (i.e. price decline).

units) would grow faster than under the assumption of no embodiment.<sup>7</sup> Moreover, an increase in capital quality (i.e. embodiment) means that capital services will grow faster than the capital stock itself (Jorgenson and Stiroh, 2000). This is because capital services are proportional to the capital stock, and the constant of proportionality is given by capital quality.

If prices are *not adjusted to reflect quality changes*, investing in new (and more productive) equipment should be accompanied by an increase in Total Factor Productivity (TFP). This was found, among others, by Jorgenson and Stiroh (2000) who, using official figures for the stock of capital, showed that a large contributor to the acceleration of TFP growth in the 1990s was investment in Information and Communications Technologies (ICT). By adjusting capital for embodiment through quality-adjusted prices, Cummins and Violante (2002) found that part of TFP growth was explained by investment-specific technical change which was not accounted for in official statistics. Therefore, incorporating quality-changes in the measurement of capital resulted in lower TFP compared with the figures calculated using official statistics.

In order to check whether the embodiment effect has any impact on TFP measurement in the present dataset, we calculate TFP growth rates obtained under the alternative assumptions of zero and positive embodied technical change, where the value of the parameter is the highest estimate obtained in the present study, i.e. 5.7% per year.

TFP growth under the assumption of embodiment is 0.2% higher than under the assumption of no embodiment. The result that TFP is higher when capital

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<sup>7</sup>This assumes that the first year in the sample is used as a numeraire, otherwise the growth rate of the capital stock is decreasing in the parameter representing embodiment (Sakellaris and Wilson, 2004), as in the present application.

is adjusted for embodiment (rather than when it is not adjusted) depends on the choice of the numeraire year which means that, in the present approach, capital stock increases faster when the rate of embodied technical change is lower, rather than vice versa (Sakellaris and Wilson, 2004).

Finally, the difference between the TFP rates indicates that part of the productivity growth that feeds into the price cap formula for the regulated operators is related to their past investment choices. As a result, accurate measurement of capital taking account of quality improvements would be of benefit to regulators and the industry.

The structure of the remainder of the chapter is as follows. The next section briefly reviews the most closely related studies, while Section 1.3 outlines the approach to the estimation of embodied technical change followed in this study. Section 1.4 describes the data. In Section 1.5, the main findings and estimation issues are discussed. Section 1.6 concludes.

## 1.2 Related Literature

The different approaches to the study and estimation of embodied technical change can be broadly divided into methodologies based on comparisons of price indexes and methodologies based on estimating cost and production functions.<sup>8</sup> This section summarizes very briefly and comments on the studies which are most closely related to the present work.

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<sup>8</sup>A more limited strand of research was followed so far by Hobijn (2001) who developed a model of investment, including embodied technical change, and estimated it using panel data for manufacturing industries.

### 1.2.1 Production and Cost Functions

The early debate on embodiment (or vintage effects) dates back to the late 1950s and 1960s in the field of growth models. Some studies on the U.S. economy found that the increase in capital per worker (capital deepening) was not among the major reasons leading to the increase in output per worker over a given period of time (Solow, 1957). On the contrary, the idea of embodiment suggested that capital modernization was at least as important as capital deepening as a source of productivity.

Following the seminal contribution by Solow (1957), a production function incorporating vintage effects was formulated which was functional to a growth model. The Cobb-Douglas production functions formulated by Solow (1957) and Phelps (1962) included explicitly the rate of embodied technical change, in the form of larger weights assigned to more efficient vintages of capital in producing a given output. The empirical testing of such formulation aimed at establishing the relationship between growth and investment, and drawing implications for economic policy.

One of the main findings of the early theoretical literature was that asymptotic growth rates were the same regardless whether technical progress was embodied or disembodied (Phelps, 1962).<sup>9</sup> Other studies found instead that disembodied technical change was important to explain growth.<sup>10</sup>

Nelson (1964) transformed the Cobb-Douglas production function in order to decompose output growth to depend linearly on the age of the capital input. This formulation allowed obtaining the rate of embodied technical change as the ratio

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<sup>9</sup>In the long-run, the growth rate is independent of the investment-output ratio, which is a property of Cobb-Douglas models (Phelps, 1962, page 556). This result is not affected by the presence of vintage effects.

<sup>10</sup>E.g. see account given by Hulten (1992).

between the coefficient on the average age of capital and the coefficient on the capital input. Thus, in Nelson's study, the links between the variables were used to calculate the rate of embodied technical change for different values of the shares of capital and labour on output value.

Numerous studies have attempted to estimate the rate of embodied technical change, both at an aggregate level, and at a firm level. Some of the early studies, such as Solow's (1962), encountered the problem of estimating the parameters of interest on the basis of assumptions on other important parameters (i.e. share of capital on output and depreciation rate) and assuming that there was no disembodied technical change. This difficulty highlights the problem of disentangling physical decay from embodied technical change (i.e. obsolescence), which is common to all empirical studies in this area.

Expanding the same framework in Solow (1962),<sup>11</sup> Wickens (1970) found that embodied technical progress was not significantly different from zero in the period 1900 - 1960 for the U.S. economy.<sup>12</sup> You (1976) also found little evidence of embodiment in the period 1929 - 1968 in the U.S. on the basis of time-series data. Using Nelson's (1964) approximation of effective capital (adjusted to incorporate embodied technical progress) as a function of average age, You concluded that the age distribution of capital did not enter the determination of growth of output per man-hour in the sample he considered and that there was no evidence of embodied technical change.

A more recent article, by Bahk and Gort (1993), examined technical change

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<sup>11</sup>By transforming the Cobb-Douglas production function in order to obtain unconditional estimates.

<sup>12</sup>Wickens found that the sum of the rate of depreciation and of embodied progress was not significantly different from zero.



using plant-level data. In a Cobb-Douglas model that incorporated learning by doing effects, they measured embodied technical change by average vintage (i.e. the year in which investment in capital was made) and estimated that a one-year fall in average vintage was related to an increase in the plant's gross output ranging from 2.5% to 3%. Assuming, for instance, a one-third weight of capital in the production function, this would mean that the rate of embodied technical change in capital would be between 7.5% and 10.5%. Their estimates contrast with the results by Wickens (1970) and You (1976), and seem to be related to the greater level of disaggregation in their study, which used panel data from U.S. manufacturing plants over the period 1973 - 1986. However, Bahk and Gort's result would still hinge on assumptions on the weight of capital on output.

Finally, the contribution most closely related to the present study is Sakellaris and Wilson (2004). They assume that technical change proceeds exponentially as in Solow (1957, 1962) and Phelps (1962), but in discrete time. Moreover, they allowed for variable capital utilization, using energy consumption as an indicator of utilization. In their study, a constant yearly rate of technical change was assumed throughout the sample period and capital stock was defined in efficiency units with respect to a base year. Sakellaris and Wilson (2004) found a 12% rate of embodied technical change, using panel data on U.S. manufacturing plants to estimate a Cobb-Douglas production function.<sup>13</sup>

In the context of the telecommunications industry, Sung (1998, 2002) used average vintage as an index of quality for physical capital, based on the observation that embodied technical change "simply refers to changes in the quality of capital goods"

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<sup>13</sup>Their data refer to manufacturing plants and are contained in the Longitudinal Research Database of the U.S. Bureau of Census.

and that technical efficiency of different vintages varies. Moreover, the labour input was also defined in terms of an index of labour quality.<sup>14</sup>

The author estimated different specifications for a translog cost function. The period of the analysis extended from 1951 to 1991 and the sample consisted of eight U.S. telecommunications local exchange carriers.<sup>15</sup> Both a long-run and a short-run cost function were estimated, giving comparable results. In the study of the long-run cost function, Sung (1998) obtained estimates of the contribution of capital quality to shifts in the cost function between 1.6% and 3.2% while, in the short-run (Sung, 2002), such contribution ranged from 2.6% to 4.2%.

A major difference with other studies was the use of cost functions rather than production functions in order to alleviate problems of simultaneity in the choice of inputs and outputs and to account for the multiproduct nature of the firm. However, in both studies output was defined as deflated revenues, which can give rise to measurement problems (Shin and Ying, 1992). Moreover, even though the author refers to Bahk and Gort (1993), the analysis lacks an explicit derivation of the interpretation of average age as a proxy for embodied technical change.

### **1.2.2 Price-based measures**

The seminal paper on price indexes incorporating quality changes was a study by Gordon (1990) who constructed quality-adjusted indexes for a range of assets. Adjusting prices for quality change is crucial especially when the pace of technological improvement in a given sector is very rapid. In these cases, a traditional price index

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<sup>14</sup>Sung defined a weighted sum of working hours of the different occupational groups within a firm. The index of labour quality was then defined as the ratio between the weighted sum and the simple sum of working hours.

<sup>15</sup>Six operators were Bell operating companies (i.e. incumbent companies) and the remaining two were independent operators.

may not adequately capture asset price reductions. This is because a new asset may sell at the same price as a previous model at the time of launch, but incorporate more advanced characteristics. Gordon (1990) used, among others, hedonic techniques to construct price indexes on the basis of detailed information on prices and characteristics of 22 categories of durable equipment.<sup>16</sup>

Building on this work, Hulten (1992) derived estimates of embodied technical change by comparing the rate of growth of the quality-adjusted price index for different durable goods with a consumption price index. Intuitively, this approach is based on the fact that "the opportunity cost of innovating [...] is foregone consumption" (Cummins and Violante, 2002).<sup>17</sup> He found an estimate of 3.4% for the period 1949 - 1983 in the U.S. manufacturing sector. The most recent study using a similar methodology, by Cummins and Violante (2002), found a 4% rate of annual change over the period 1947 - 2000.

While price-based measures of embodied technical change have proved relatively popular in the literature, it should be noted that the underlying methodology may not be entirely robust. Sakellaris and Wilson (2004), quoting Gordon, highlight that differences between adjusted and unadjusted quality indexes may reflect not only

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<sup>16</sup>In hedonic models, regressions are used to explain a time series of prices as a function of a set of quality characteristics and time dummies. There are different ways in which these models can be implemented. For instance, one approach could be to use the estimated coefficients on the time dummies as an indication of price movements, keeping quality characteristics constant.

<sup>17</sup>The intuition for this approach is formalised in the literature by means of a simple two-sector model of investment and final goods (e.g. Cummins and Violante, 2002; Sakellaris and Wilson, 2001). It is assumed that both goods are produced competitively.

The production of final goods ( $c_t$ ) follows a constant returns to scale technology which uses labour and capital. Final goods can be used either for consumption or to produce investment goods. The production function for investment goods in efficiency units is  $i_t^* = q_t c_t$ , where  $q_t$  is the technological level specific to the investment sector.

The price of investment goods (in efficiency units) is  $p_t^{i^*}$  and the price of consumption goods is  $p_t^c$ . Equating the marginal product and the relative price of the input gives  $p_t^{i^*} i_t^* = p_t^c c_t$  which, combined with the production function of the investment good, results in  $\frac{p_t^{i^*}}{p_t^c} = \frac{1}{q_t}$ . From this expression, the change in the investment-specific technological level can be inferred from the change of the price of investment goods relative to consumption goods.

quality change, but also other factors.

### **1.3 Empirical Strategy**

An inherent difficulty in estimating embodied technical change is the impossibility to isolate the other inputs and the output which are directly associated to a specific vintage of capital, as would be required by the theoretical formulation of the problem in order to identify the effects of different vintages on output. In the estimation, this problem is addressed by the cross-section variation in the sample: if two firms show a different time pattern of investment up to a certain year, then their outputs (or costs) in that year may differ, after controlling for other factors. By comparing outputs (or costs), the variation in the distribution of investment between the two firms can therefore be used to measure embodied technical change. However, this only alleviates the aggregation problem described above, as well as the aggregation between asset types which may potentially incorporate different levels of technical change, as some equipment (e.g. switches) has been affected by faster progress than other types.

Taking into account the above considerations, the present study uses panel data to estimate embodied technical change on the basis of a cost and a production function. Unlike previous studies on telecommunications, it relies on the most disaggregated data available for investment in different categories of telecommunications plant by the large U.S. telecommunications operators in order to calculate surviving capital as accurately as possible.

In addition, compared to Sung (1998, 2002), the sample includes a larger number of firms (even though on a shorter time interval, due to data availability) and the

assumption that all assets share the same rate of embodied technical change is partially relaxed: capital assets with long lives (e.g. buildings, poles, conduits) are kept separate from telecom equipment, which typically needs faster replacement and is characterized by higher technical change.

The sensitivity of the results to different assumptions concerning depreciation is also made possible by a flexible formulation, which does not use depreciation rates obtained from company data, but calculates the percentages of surviving investment across time for each asset type.

Finally, output variables in the cost function are measured as physical quantities in order to overcome measurement problems.

### 1.3.1 Definition of Capital

The previous discussion briefly reviewed the alternative approaches to analyzing embodied technical change, providing possible alternatives to the studies carried out on the telecommunications industry by Sung (1998, 2002). The definition of capital adopted by Sakellaris and Wilson (2004) seems better suited to address the issues which were highlighted in the literature review. In particular, it has the appeal of resembling closely the theoretical formulations adopted in the literature on embodied technical change and therefore has a straightforward economic interpretation.

The seminal papers by Solow (1957) and Phelps (1962), under the assumption of neutral technical change,<sup>18</sup> specify the following production function in continuous time:  $y_t = B(t)F(K_t, L_t)$  where  $K_t$  and  $L_t$  are capital and labour respectively.

The measure of embodied technical progress grows exponentially at the rate  $\lambda$  and

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<sup>18</sup>Hicks-neutral technical change can be represented by a multiplicative term  $B$  which multiplies the entire function, i.e.  $y = BF(K, L)$ , and therefore results in a shift of the isoquant, keeping the input ratios constant.

only affects the efficiency of new capital goods:  $B(t) = B_0 e^{\lambda t}$ . It is also assumed that the production function is Cobb-Douglas with constant returns to scale. The output corresponding to a given vintage of capital, and related labour, is given by  $y_t = B_0 e^{\lambda v} K_{vt}^c L_{vt}^{1-c}$ . The level of capital input in the production function above is only the capital of vintage  $v$  surviving at time  $t$  and the corresponding level of labour is that employed on capital of vintage  $v$ . Given that technical progress is neutral,  $c$  (elasticity of output with respect to  $K$ ) is constant for each vintage  $v$ . It can be shown that summing the homogeneous outputs of all vintages, one obtains the aggregate production function at time  $t$ :  $y_t = B_0 J_t^c L_t^{1-c}$ , where  $J_t = \int_{-\infty}^t e^{\lambda v} K_{vt} dv$ . In this expression,  $J_t$  is the sum of all surviving capital goods where older capital carries a smaller weight.<sup>19</sup>

This definition can be made operational by constructing capital recursively as the sum of past investment flows. Therefore, gross investment has to be adjusted to take account of depreciation over time, so that only surviving capital is included in the construction of the stock active at a given time period.

The depreciation measure that is employed in this procedure is crucial for the correct measurement of capital. Economic depreciation, which is usually the appropriate measure, covers two effects: physical decay, due to wear and tear, and obsolescence. The latter factor is precisely the investment-specific technical change that studies on embodiment aim to measure. If this component was known, for instance relying on the price-based measures described above, one could construct the capital stock in efficiency units on the basis of physical decay and obsolescence,

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<sup>19</sup>In this setting, disembodied technical change would be expressed by a similar multiplicative factor, which however would not depend on the specific vintage  $v$ , and would be applied to the entire capital stock still productive at time  $t$  (Phelps, 1962).

as in Cummins and Violante (2002).<sup>20</sup> However, in the present study, the obsolescence component is the parameter that needs to be estimated. For this reason, it is important to ensure that gross investment is only depreciated for physical decay and that no other adjustments are made that would account for obsolescence, such as deflation by quality-adjusted indexes.

Following Sakellaris and Wilson (2004), capital adjusted for embodied technical change for a given asset type can be defined in efficiency units, relative to a given base year  $t_0$ , as

$$J_t = \sum_{s=0}^T I_{t-s} D_{t,t-s} (1 + \gamma)^{t-s-t_0} \quad (1.1)$$

where  $I_{t-s}$  is gross investment of vintage  $(t-s)$ ,  $D_{t,t-s}$  is the fraction of vintage  $(t-s)$  still productive at time  $t$  (accounting only for physical decay),  $(1 + \gamma)^{t-s-t_0}$  is a factor that represents the productivity of vintage  $(t-s)$  with respect to the numeraire year,  $t_0$ , and  $T$  is service life for the specific asset category. Where service life exceeds the available sample size, initial capital ( $J_0$ ) is added to obtain the capital stock.<sup>21</sup> The rate of embodied technical change is parameter  $\gamma$  and is not known.<sup>22</sup>

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<sup>20</sup>Cummins and Violante (2002) estimate embodied technical change from quality-adjusted price indexes and then use this value to construct capital, by combining their estimate with parameters of physical decay.

<sup>21</sup>The possibility of estimating past investment flows was also explored, but proved unsuccessful. In particular, given the very low correlation of firms' investments with macro data for the sector it was not possible to use aggregate data to estimate firm-level data. Moreover, the high variability in investment in different asset types over time, for a given firm, would make unreliable any attempt to estimate past firm investment.

<sup>22</sup>This definition has the advantage of measuring embodied technical change without changing its numeraire year,  $t_0$ , and for this reason it is very suitable to data with a time-series component.

This can be shown by an example in which the numeraire is  $t_0 = 2002$  and it is assumed we are calculating  $J_t$  at  $t = 2001$ . According to the definition, investment in  $(t-1)$  - i.e. 2000 - is divided by  $(1 + \gamma)^2$ , the term in  $(t-2)$  - i.e. 1999 - is divided by  $(1 + \gamma)^3$  and so on. Given that  $(1 + \gamma) > 1$ , this series increases with its exponent. Past investments are therefore divided by increasing numbers and the 'efficiency' of older investment is lower relative to today's. These 'discount factors' depend not only on the numeraire but also on  $t$ , the year for which the capital stock is calculated.

If we change  $t$ , we can verify that the discount factors do not change. Assume we were constructing

Capital stock is given by the sum of investment flows, after removing physical decay and adjusting to take into account the different embodied technical progress compared to the base year. This approach assumes constant technical change throughout the sample period with respect to the base year (i.e. the last in the sample), in which the level of embodied technology is normalized to 1.

As noted by Sakellaris and Wilson (2004),  $(1 + \gamma)^{t-s-t_0}$  represents the relative productive efficiency of vintage  $(t - s)$  whose opportunity cost is one unit of the consumption good. This interpretation is based on the same framework underpinning the price-based measures of embodiment, described in the literature review section. An implication of this framework is that, in the above expression, investment is measured in 'foregone consumption units' and, for this reason, it needs to be deflated by a consumption price deflator that is not adjusted for quality changes in capital goods.<sup>23</sup>

Moreover, the above definition of capital adjusted for embodied technical change assumes perfect substitutability among different vintages of capital. It also implies that firms always invest in the most advanced vintage, while production or resale of older vintages may also take place in practice.

The adjustment for physical decay is carried out by multiplying investment in a given type of asset in a given year by a hyperbolic (or beta) decay function. This function depends on the asset life ( $S$ ) for the specific type of asset and on the actual

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$J_t$  at  $t = 2000$ , while keeping 2002 as the numeraire. In this case, investment in  $(t - 1)$  – i.e. 1999 – is divided by  $(1 + \gamma)^3$ . This is consistent with what obtained above for the case  $t = 2001$ , which implies that the above definition keeps the numeraire constant.

<sup>23</sup>An implication of the definition of capital adopted in this study is that, for capital to be expressed in efficiency units, the value of  $\gamma$  inserted in the definition should be the 'correct' value of embodied technical change.

Therefore, for other values of the parameter  $\gamma$ , capital may not be correctly measured and, as a result, the estimated parameters of the cost and production functions may not accurately reflect the substitution between capital and other inputs.



age of the asset ( $s$ ).

$$D_{t,t-s} = \frac{S - s}{S - bs} \quad (1.2)$$

The value of  $b$  reflects different curvatures of the function. For equipment, a value of 0.5 is used by the U.S. Bureau of Labor Statistics (Mohr and Gilbert, 1996) and the same was also adopted here as the base case. Following the Bureau of Labor Statistics, a value of 0.75 was used for structures. As will be explained below, the sensitivity of the results to this assumption was tested by using a range of values for  $b$ .<sup>24</sup> The decay function does not remove obsolescence effects (i.e. the reduction in capital efficiency due to embodied technical change) as this would be in contrast with the objective of the study.<sup>25</sup>

The factors used to adjust past investments are, as mentioned above, functions of the numeraire year (in this case, 2003) and the vintage ( $t - s$ , i.e. investment made  $s$  years prior to period  $t$ ). For this reason, the exponents ( $t - s - t_0$ ) are constant for each vintage, given a numeraire. Finally, it is assumed that investment is immediately productive which seems a reasonable hypothesis with annual data (i.e. the sum includes investment in year  $t$ ).

Definition 1.1 is also used to construct the stock of capital assumed to incorporate lower embodied technical change, for instance land and buildings ( $S_{it}$  in Equation 3). However,  $\gamma$  is assumed equal to zero for these assets.

As will be explained below, the above definitions are used to construct the capital stock and to estimate both a production and a cost function so as to provide

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<sup>24</sup>As the parameter  $b$  increases towards 1, the function  $D_{t,t-s}$  approaches a horizontal line, dropping to 0 at the end of the asset life. See Figure in Appendix.

<sup>25</sup>In addition, it does not incorporate stochastic retirements of stock for simplicity.

robustness checks on the results on embodied technical change. The remainder of this section describes the approach to the estimation. In particular, we firstly describe the approach towards the estimation of the cost function. Secondly, we detail alternative approaches to test the reliability of the results.

### **1.3.2 Estimation of a Cost Function**

Most of the recent literature on the telecommunications industry focuses on the estimation of cost functions, with some exceptions (e.g. Uri, 2001a). This is due to a variety of reasons, including their natural extension to multi-output technologies and their underlying assumptions. In particular, output in regulated industries is often not storable and demand-driven, so it could be regarded as exogenous. In addition, when the analysis is carried out at a micro level as in the present study, it seems reasonable that firms regard prices as exogenous.

In consequence, cost functions do not suffer from the problem of endogeneity which has long been recognized in the estimation of production functions. In the context of a production function, the regressors are inputs which are chosen optimally by the firms and, for this reason, the assumption of exogeneity may fail thus making the estimates inconsistent (Griliches and Mairesse, 1995). In a cost function, the underlying assumptions would alleviate this problem. However, the information requirements are more demanding for a cost function than for a production function, as input prices for each firm in the sample are needed. Such information is available for the sample used in the present study. Data on output quantities, rather than simply revenues, is also available and this can be exploited in the context of a multi-product cost function. However, some form of aggregation across outputs is still needed to limit the number of parameters and, given that not all output quantities

are available, this approach may be suffer from an omitted variables problem.<sup>26</sup>

In the present study, a short-run cost function is estimated to take into account that adjustment to long-run equilibrium may not be instantaneous for all inputs and this would be even more applicable to regulated firms which may face short-run constraints in minimizing costs.

The estimated cost function is a translog cost function, introduced by Christensen et al. (1973). This function is very commonly used in applications, as it provides a second order approximation to an arbitrary cost function. In particular, it does not suffer from the curvature problems of Cobb-Douglas cost functions in the presence of multiple outputs (Kumbahakar and Lovell, 2003). In addition, it does not constrain the elasticities of substitution and the economies of scale over the interval of production.

The translog variable cost function can be written as follows, for firm  $i = 1, 2, \dots, N$  and period  $t = 1, 2, \dots, T$ :

$$\begin{aligned}
\ln VC_{it}(w, y, z) = & \alpha_0 + \sum_{m=1}^M \alpha_m \ln y_{m,it} + \frac{1}{2} \sum_{m=1}^M \sum_{n=1}^M \alpha_{mn} \ln y_{m,it} \ln y_{n,it} \quad (1.3) \\
& + \sum_{k=1}^K \beta_k \ln w_{k,it} + \frac{1}{2} \sum_{k=1}^K \sum_{l=1}^K \beta_{kl} \ln w_{k,it} \ln w_{l,it} \\
& + \sum_{p=1}^P \psi_p \ln z_{l,it} + \frac{1}{2} \sum_{p=1}^P \sum_{h=1}^H \psi_{ph} \ln z_{p,it} \ln z_{h,it} + \\
& + \sum_{k=1}^K \sum_{m=1}^M \rho_{km} \ln w_{k,it} \ln y_{m,it} + \sum_{p=1}^P \sum_{m=1}^M \theta_{pm} \ln z_{p,it} \ln y_{m,it} \\
& + \sum_{k=1}^K \sum_{p=1}^P \sigma_{kp} \ln w_{k,it} \ln z_{p,it} + \varepsilon_{it}
\end{aligned}$$

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<sup>26</sup>See the Appendix for the construction of the output variables.

where  $y_m$  are outputs,  $w_k$  input prices and  $z_p$  fixed inputs.  $\varepsilon_{it}$  is an error term.

All variables are deviations from the sample mean after taking logs. Therefore, first-order coefficients can be interpreted as elasticities at the sample mean, given that second-order and cross-product terms drop out when evaluating the derivative of  $\ln VC_{it}(w, y, z)$  at the sample mean.

The estimated function needs to satisfy the theoretical properties of a cost function (Coelli et al., 2005): (i) Nonnegative; (ii) Nondecreasing in input prices  $w$ ; (iii) Nondecreasing in output  $y$ ; (iv) Homogeneous of degree one in input prices  $w$ , i.e. multiplication of all input prices by a nonnegative amount  $c$  will result in a  $c$ -fold increase in costs; and (v) Concave in  $w$ , which implies that input demand functions cannot be increasing in input prices  $w$ . In addition, a short-run cost function is non-increasing in the fixed inputs.

Homogeneity of degree one with respect to input prices would require certain restrictions on the coefficients, as follows:

$$\begin{aligned} \sum_{k=1}^K \beta_k &= 1 \\ \sum_{l=1}^K \beta_{kl} &= 0, \forall k \\ \sum_{k=1}^K \rho_{km} &= 0, \forall m \\ \sum_{k=1}^K \sigma_{kp} &= 0, \forall p \end{aligned}$$

In addition, for a twice differentiable cost function the second order derivatives are equal, which implies a symmetry restriction on the cross-product coefficients:

$$\alpha_{mn} = \alpha_{nm}, m, n = 1, 2, \dots M$$

$$\beta_{kl} = \beta_{lk}, k, l = 1, 2, \dots K$$

$$\psi_{ph} = \psi_{hp}, h, p = 1, 2, \dots P$$

Applying Shephard's lemma, one obtains the cost share equations of the variable inputs, i.e. labour and materials:

$$\begin{aligned} \frac{\partial \ln VC_{it}}{\partial \ln w_{k,it}} &= \frac{w_{k,it}}{VC_{it}} \frac{\partial VC_{it}}{\partial w_{k,it}} = \frac{w_{k,it} x_{k,it}}{VC_{it}} = S_k \\ &= \beta_k + \sum_{k=1}^K \beta_{kl} \ln w_{k,it} + \sum_{m=1}^M \rho_{km} \ln y_{m,it} + \sum_{p=1}^P \sigma_{kp} \ln z_{p,it} + \varepsilon_{k,it} \end{aligned} \quad (1.4)$$

where  $\varepsilon_k$  is the error term for the share equation.

Due to the definition of equipment capital  $J_{it}$  (1.1), when  $\ln J_{it}$  is inserted in the cost function this results in an expression in which the parameter of embodied technical change ( $\gamma$ ) enters non-linearly and from which it is not possible to identify separately the impact of  $\gamma$  and of the coefficient on capital  $\psi_p$ .

One way to address this issue, rather than directly attempting to estimate all the parameters in the model simultaneously, is to assign different values to the parameter  $\gamma$  and estimate a cost function for each of the assumed values. These alternative models can then be compared to investigate whether different values of the parameter have any impact on the quality of the estimation. We proceed in two steps. Firstly, the function is estimated for  $\gamma = 0$  in order to focus on model specification. In addition, for each value of  $\gamma$ , a different shape of the physical decay

function is assumed by varying the value of  $b$  from the base case of 0.5 to 0.75 (the latter value is often used for structures rather than equipment). Secondly, the function is estimated for different values of the parameter  $\gamma$  and the corresponding log-likelihood is calculated for each value. The search is then refined by focusing on the interval of values which results in higher log-likelihood.

### **1.3.3 Alternative Approaches**

In order to provide a robustness check to the results, alternative approaches are attempted. Firstly, following the same approach as for the cost function above, a production function is estimated. Secondly, we attempt to address the identification of the rate of embodied technical change directly by linearizing the definition of capital and inserting this definition in a production function.<sup>27</sup>

#### **Nonlinear production function**

As explained above, an iterative procedure is followed in order to identify the value of the rate of embodied technical change  $\gamma$  resulting in the best fit. As a first step, a translog and a Cobb-Douglas production function are estimated in order to test which functional form is the more appropriate. Building on this result, the estimation is repeated for a range of assumed values of  $\gamma$  to identify the best fit.

#### **Linearized production function**

Another approach to the identification problem is to transform algebraically definition 1.1 into the product between a function of the rate of embodied technical change and a function of the flow of investment. For these purposes, and taking into

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<sup>27</sup>The latter approach was also implemented in a Cobb-Douglas cost function but the estimation did not lead to significant coefficients on the relevant variables.

account that for most assets the time dimension of the sample is shorter than asset life, the definition of capital  $J_t$  can be rewritten as

$$\begin{aligned}
J_t &= J_0 + \sum_{s=0}^{t-1} I_{t,t-s}^* (1 + \gamma)^{t-s-t_0} \\
&\approx J_0 + \frac{1}{(1 + \gamma)^{t_0}} \times \sum_{s=0}^{t-1} I_{t,t-s}^* + \frac{\gamma}{(1 + \gamma)^{t_0}} \times \sum_{s=0}^{t-1} I_{t,t-s}^* (t - s)
\end{aligned} \tag{1.5}$$

where  $I_{t,t-s}^*$  is surviving gross investment of vintage  $t - s$  in period  $t$ . The sum equals initial capital and the flow of investment up to and including time  $t$ . By approximating the sum using a first-order Taylor expansion around  $\gamma = 0$ ,<sup>28</sup> one obtains the second expression above which can be inserted in the production function.

Given that this complicates the formulation of the production function, for simplicity we estimate a Cobb-Douglas production function, as follows:

$$\ln y_{it} = \alpha_i + \beta \ln L_{it} + \delta \ln M_{it} + \theta \ln J_{it} + \xi \ln S_{it} + \varepsilon_{it} \tag{1.6}$$

In the above function, for firm  $i$  in period  $t$ ,  $y_{it}$  is output,  $\alpha_i$  is a firm-specific effect,  $L_{it}$  indicates labour hours,  $M_{it}$  represents materials,  $J_{it}$  is the level of capital stock and  $S_{it}$  is the level of stock of capital that is assumed to have a low rate of embodied technical progress (e.g. land and support). The level of capital ( $J_{it}$ ) is obtained as the sum of the  $J_t$  from Equation 1.1 for individual asset types. Finally,  $\varepsilon_{it}$  is the error term.

Inserting the definition of capital gives a linearized version of the Cobb-Douglas

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<sup>28</sup>By the Taylor approximation,  $(1 + x)^n \approx 1 + nx$  for  $x$  small. In this case,  $\frac{(1+\gamma)^{t-s}}{(1+\gamma)^{t_0}} = \frac{1+\gamma(t-s)}{(1+\gamma)^{t_0}}$

function. Subtracting  $\ln y_0$  from  $\ln y_t$  it is possible to obtain Equation 1.7, which does not depend on  $J_0$ .<sup>29</sup>

$$\begin{aligned}
\ln y_{it} - \ln y_{i0} &= \beta(\ln L_{it} - \ln L_{i0}) + \delta(\ln M_{it} - \ln M_{i0}) + \theta(\ln J_{it} - \ln J_{i0}) + \quad (1.7) \\
&\quad + \xi(\ln S_{it} - \ln S_{i0}) + \eta_{it} \\
&= \beta(\ln L_{it} - \ln L_{i0}) + \delta(\ln M_{it} - \ln M_{i0}) + \frac{\theta \cdot a_{it}}{(1+\gamma)^{t_0}} + \frac{\theta \cdot \gamma \cdot b_{it}}{(1+\gamma)^{t_0}} + \\
&\quad + \xi(\ln S_{it} - \ln S_{i0}) + \eta_{it}
\end{aligned}$$

$$\text{where } a_{it} = \frac{\sum I_{it,t-s}^*}{J_{i0}} \text{ and } b_{it} = \frac{\sum I_{it,t-s}^*(t-s)}{(1+\gamma)^{t_0}}.$$

This new equation for  $(\ln y_t - \ln y_0)$  can be estimated and the rate of embodied technical change  $\gamma$  can be recovered as the ratio between the coefficients on the two terms in the footnote, similarly to the derivation in Nelson (1964). However, the approximation that allows simplifying  $(\ln J_{it} - \ln J_{i0})$  is valid only for some time periods in the sample. Consequently, while Equation 1.7 is estimated in this study, it should be emphasized that the required assumptions are satisfied only in part and therefore results cannot be used to estimate embodied technical change.

$$\begin{aligned}
^{29} \ln J_t - \ln J_0 &= \ln \left( J_0 + \frac{\sum I_{t,t-s}^*}{(1+\gamma)^{t_0}} + \frac{\gamma \cdot \sum I_{t,t-s}^*(t-s)}{(1+\gamma)^{t_0}} \right) - \ln J_0 = \\
&= \ln \left[ J_0 \times \left( 1 + \frac{\sum I_{t,t-s}^*}{J_0 \cdot (1+\gamma)^{t_0}} + \frac{\gamma \cdot \sum I_{t,t-s}^*(t-s)}{J_0 \cdot (1+\gamma)^{t_0}} \right) \right] - \ln J_0 = \\
&= \ln \left( 1 + \frac{\sum I_{t,t-s}^*}{J_0 \cdot (1+\gamma)^{t_0}} + \frac{\gamma \cdot \sum I_{t,t-s}^*(t-s)}{J_0 \cdot (1+\gamma)^{t_0}} \right) \approx \frac{\sum I_{t,t-s}^*}{J_0 \cdot (1+\gamma)^{t_0}} + \frac{\gamma \cdot \sum I_{t,t-s}^*(t-s)}{J_0 \cdot (1+\gamma)^{t_0}} = \\
&= \frac{a_t}{(1+\gamma)^{t_0}} + \frac{\gamma \cdot b_t}{(1+\gamma)^{t_0}}
\end{aligned}$$

The approximation  $\ln(1+x) \approx x$  is possible if  $x$  is small. In this case, however, this quantity is small only for some time periods in the sample and therefore the approximation is not valid.



## 1.4 Data

The sample is a panel of yearly data on 28 large Incumbent Local Exchange Carriers (ILECs) over the period 1990 to 2003. The main source for the data is the Automated Reporting Management Information System (ARMIS)<sup>30</sup>, published by the Federal Communications Commission (FCC), the U.S. telecommunications regulator, and available from the regulator's website. Variables are briefly described in this section and more details are provided in Appendix.

For the cost function, variable costs ( $VC$ ) are calculated as deflated operating expenses minus depreciation. The measures of output considered in this chapter are access lines, local calls and toll calls. The price of labour is obtained as total compensation divided by the number of full-time employees, while the price of materials is given by materials divided by the number of access lines. In line with previous papers, materials are constructed as deflated operating costs minus total compensation to employees and depreciation costs. The shares of inputs over cost ( $S_k$ ) are calculated as the relevant expenses divided by variable costs.

The stock of capital was constructed separately for assets that could be expected to exhibit low levels of embodied technical change (e.g. land, buildings) and those that could be expected to show higher embodied technical change (i.e. telecommunications equipment).

For each firm in the sample, gross capital stock was constructed on the basis of investment flows. The objective was to reflect the heterogeneity of capital, formed by a series of vintages, and to abstract as much as possible from different firms' accounting policies. For instance, while capital stock in the LECs' balance sheet is

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<sup>30</sup>The tables used for the analysis are Reports 43-01 (Annual Summary Report), 43-02 (USOA Rpoert - Balance Sheet Accounts), and 43-08 (Operating Data Report).

affected by the firms' decisions to withdraw or write down assets, capital constructed using the perpetual inventory method incorporate only gross additions. Instead of relying on the companies' depreciation figures, a common depreciation profile was assumed for all the LECs.<sup>31</sup> As described above, this was a hyperbolic decay function. The detailed steps followed in the construction of capital are explained in Appendix.

For the production function, output is proxied by deflated revenues, which show very high positive correlation with the actual quantities of output, i.e. numbers of access lines and of calls. This assumption is often adopted in econometric studies on the telecommunications industry (e.g. Resende 1999; Sung, 1998 and 2002), although it may lead to a distorted representation of output because revenues are affected by the firms' pricing strategies (in the specific case, they are also affected by price cap constraints, which are exogenous to the firms). However, it is not straightforward to construct a basket of the LECs' output quantities due to data availability, as explained in Appendix.

Additional variables are considered both in the production and cost function, including a measure of customer density (sheath length per access line), a measure of congestion (number of calls per switch) and an indicator of modernization (the share of fiber on total network Kilometers).

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<sup>31</sup>These are adjusted to reflect the asset lives of the different assets.

As explained above, only physical depreciation is taken into account, rather than economic depreciation. This is because economic depreciation also incorporates obsolescence, which is precisely what the present study aims to estimate. As a result, investment is not deflated by a price index which is quality-adjusted to reflect investment-specific technical change.

Table 1.1: Summary Statistics

Variable	Mean	Standard Deviation	Min	Max	Obs.
Variable costs (\$ 000)	1,547,159	1,650,790	91,094	6,766,698	392
Access Lines (units)	6,445,288	7,576,483	232,465	41,400,000	392
Local Calls (m)	16,100	19,800	529.5	99,300	392
Toll Calls (m)	2,919	3,126	84.3	16,800	392
Equipment (\$ 000)	4,579,015	5,193,016	228,051	24,800,000	392
Sheath / line (Km/line)	0.034	0.020	0.002	0.117	359
Utilization (000)	0.396	0.219	0.038	0.976	388
Modernization (units)	0.090	0.040	0.006	0.220	387

### 1.4.1 Description of the Sample

The 28 operators included in the sample, although all classified as "large" in the FCC tables, vary greatly in size, as can be seen from the summary statistics in the table below.<sup>32</sup>

When looking at the behavior of the variables over time, variable costs do not show a clear pattern: for many companies in the sample, they fluctuate around a constant level, while for others there is a positive or a negative trend. For the firms showing a positive trend, variable costs have a peak corresponding to the years between 1998 and 2000, corresponding to a period of expansion in the telecommunications market. The behavior of variable costs over time is plotted in the appendix to this chapter. In terms of correlations between variable costs and measures of output (i.e. access lines and calls), while variable costs are positively related to the number of access lines, as expected, there is no clear relationship with the number of calls.

While the measures for output and capital<sup>33</sup> show an upward trend over time, labour tends to fall, even if this trend cannot be clearly identified for all LECs. For

<sup>32</sup>The variation is partly due to different reporting conventions. Verizon reports results at the operating company level, while BellSouth and Qwest results are aggregated.

The choice of sample was dictated by data availability, in particular with reference to gross investment flows, which only the operators in this sample (i.e. incumbent LECs) are required to report.

<sup>33</sup>For this purpose, capital was constructed as the simple sum of surviving investment, not accounting for embodied technical change.

some of the operators, output's positive trend peaks around the years between 1998 and 2000, and declines afterwards. Moreover, as expected, output exhibits positive correlation with all inputs.

Lastly, it is interesting to note that the share of labour on output declines over the sample period, while the share of capital increases. The share of labour on variable costs declines over time for most of the sample, thus confirming the trend for the corresponding share on output. Labour productivity increases, as could be anticipated given the trend towards more competition in the sector.

Given the observations above on the behavior of variables over the sample period, unit roots tests were carried out. The Im, Pesaran and Shin (2003) and Maddala and Wu (1999) tests for panel data failed to reject the null hypothesis of  $I(1)$  for the log of output, once the optimal lag for removing serial correlation for each cross-section was selected on the basis of the Akaike Information criterion. On the other hand, when the same tests were applied to the logarithm of variable cost,<sup>34</sup> the null hypothesis of non stationarity was rejected by the Maddala and Wu (1999) test, while it could not be rejected by the Im, Pesaran and Shin (2003) test. However, the test results may not be entirely reliable, as most panel unit roots tests are designed for the case in which  $T \rightarrow \infty$ , followed by  $N \rightarrow \infty$ , while in the present sample  $T = 14$  and  $N = 28$ . Moreover, these tests are sensitive to correlation between the different groups (i.e. firms) and, as in the case of univariate test, to problems such as non-linearities or structural breaks.<sup>35</sup>

In order to address the potential dynamic behavior of the variables, while also taking into account the limited time dimension of the sample, we include a time

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<sup>34</sup>The test was applied to the left-hand variable in the translog cost function, which was constructed as the ratio between variable cost and its average.

<sup>35</sup>For a discussion, see Smith and Fuertes (2005).

trend or time dummies in the model, in line with common practice, e.g. Bloom et al. (2007) for estimation of the production function estimation, Coelli and Perelman (1996) and Resende (1999) for applications to utilities.

Finally, when investment flows, measured by gross additions, are plotted over time, the firms in the sample show very different patterns. As noted above, the use of panel data allows exploiting differences in investment profiles, and therefore vintage structures, across firms in order to circumvent the identification problem which arises in studies relying on time-series aggregate data.

## 1.5 Main Results

### 1.5.1 Estimation Results of the Cost Function

As explained above, in order to overcome the identification problem, a grid search is performed for a range of assumed values of the parameter. Capital with higher technical change ( $J_{it}$ ) is calculated for different values of the parameter  $\gamma$ . The cost function (Equation 1.6) is then estimated for each assumed  $\gamma$  to find the value of the parameter resulting in the highest log-likelihood.

As a first step, in order to focus on the specification, the cost function is estimated under the assumption of no embodied technical change (i.e.  $\gamma = 0$ ), while the parameter of physical decay  $b$  is set to its base value of 0.5. The results are presented in the table below.

The translog cost function is jointly estimated in a system with the labor cost share, using seemingly unrelated regression (SUR) techniques. Given that the sum of cost shares equals 1 for each observation, when there are  $n$  factor share equations, only  $n - 1$  of them are linearly independent. The estimation is therefore carried out

removing the cost share equation for materials. The system of equations is estimated by maximum likelihood and therefore parameter estimates, log-likelihood values and estimated standard errors are invariant to the choice of the share equation (Berndt, 1991). Following common practice, conditions for linear homogeneity and symmetry are imposed on the parameters of the translog cost function before estimation.

The results for the translog variable cost function are reported in Table 1.2. The parameters of the share equation are not reported as they are equal to the corresponding terms in the translog cost function, once the derivative over the price of labour is taken (see Equation 1.4).

The system is estimated including a time trend, in order to capture disembodied technical change, and firm-level fixed effects.<sup>36</sup> The estimated coefficients are reported in Table 1.2. However, before proceeding to comment on the coefficients, it is important to check whether the properties of the cost function are satisfied.<sup>37</sup> As already mentioned, symmetry and homogeneity are already imposed before estimation. In addition, estimated variable costs and marginal costs are non-negative at all observations, and the function is concave in input prices on about 90% of observations, as indicated by the elasticities of substitution calculated on fitted shares.<sup>38</sup> Moreover, based on the estimated coefficients, the cost function exhibits mildly increasing returns to scale.<sup>39</sup>

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<sup>36</sup>The same model was also estimated both including and excluding fixed effects in the share equation. In order to check whether it would be possible to disregard fixed effects, a Wald test was conducted of the hypothesis that all the firm dummy variables were jointly zero in the share equation. As it was not possible to reject this hypothesis, this would seem to suggest some form of unobserved heterogeneity in the coefficients of the cost function. The same test gave consistent results under the specification that included time dummies instead of a time trend.

<sup>37</sup>In addition, we check that residuals are well-behaved and that there are no outliers. On the basis of this analysis, two LECs are removed from the sample, i.e. Verizon Hawaii and Verizon DC.

<sup>38</sup>For the translog cost function, the elasticities of substitutions are given by:  $\theta_{kl} = \frac{\beta_{kl}S_kS_l}{S_kS_l}$  and  $\theta_{kk} = \frac{\beta_{kk}S_k(1-S_k)}{S_k^2}$ , where  $S_k$  indicates the fitted cost share of input  $k$  and  $\beta_{kk}$  and  $\beta_{kl}$  are estimated parameters from the cost function (Greene, 2003).

<sup>39</sup>In a variable cost function, the definition of returns to scale (RTS) needs to be modified to take

Table 1.2: Translog Cost Function - Results for gamma = 0

Dependent Variable: $\ln VC_{it}$			
$\ln Lines_{it}$	0.753 (0.031)***	$\ln PriceRatio_{it} * t$	-0.013 (0.001)***
$\ln LCalls_{it}$	0.031 (0.018)*	$\ln Lines_{it} * \ln LCalls_{it}$	-0.130 (0.066)**
$\ln TCalls_{it}$	-0.033 (0.011)***	$\ln Lines_{it} * \ln TCalls_{it}$	0.021 (0.047)
$\ln PriceRatio_{it}$	0.397 (0.008)***	$\ln Lines_{it} * \ln J_{it}$	-0.196 (0.079)**
$\ln J_{it}$	0.153 (0.036)***	$\ln Lines_{it} * \ln S_{it}$	-0.026 (0.078)
$\ln S_{it}$	0.050 (0.032)	$\ln Lines_{it} * t$	-0.006 (0.005)
$t$	-0.030 (0.002)***	$\ln LCalls_{it} * \ln TCalls_{it}$	0.063 (0.026)**
$\frac{1}{2} \ln Lines_{it}^2$	0.324 (0.121)***	$\ln LCalls_{it} * \ln J_{it}$	0.163 (0.045)***
$\frac{1}{2} \ln LCalls_{it}^2$	-0.003 (0.047)	$\ln LCalls_{it} * \ln S_{it}$	-0.082 (0.048)*
$\frac{1}{2} \ln TCalls_{it}^2$	-0.032 (0.029)	$\ln LCalls_{it} * t$	0.001 (0.003)
$\frac{1}{2} \ln PriceRatio_{it}^2$	0.220 (0.004)***	$\ln TCalls_{it} * \ln J_{it}$	-0.056 (0.034)
$\frac{1}{2} \ln J_{it}^2$	0.033 (0.113)	$\ln TCalls_{it} * \ln S_{it}$	0.008 (0.039)
$\frac{1}{2} \ln S_{it}^2$	0.008 (0.114)	$\ln TCalls_{it} * t$	-0.002 (0.002)
$\frac{1}{2} t^2$	-0.000 (0.000)	$\ln S_{it} * t$	-0.004 (0.004)
$\ln PriceRatio_{it} * \ln Lines_{it}$	-0.196 (0.015)***	$\ln J_{it} * t$	0.014 (0.004)***
$\ln PriceRatio_{it} * \ln LCalls_{it}$	0.043 (0.007)***	$\ln J_{it} * \ln S_{it}$	0.041 (0.102)
$\ln PriceRatio_{it} * \ln TCalls_{it}$	-0.013 (0.005)**	Constant	0.153 (0.103)
$\ln PriceRatio_{it} * \ln J_{it}$	0.084 (0.015)***	Fixed Effects	Yes
$\ln PriceRatio_{it} * \ln S_{it}$	0.064 (0.014)***	Observations	364
		Log likelihood	1812.40

Note: Standard errors in parentheses. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. VC: variable costs normalized by the price of materials; Lines: access lines; LCalls: local calls; TCalls: toll calls; Pratio: price of labor normalized by the price of materials; Equip: equipment capital; Struct: structures; t: time trend.

The majority of the estimated coefficients is significant. In particular, as expected, the sign of the first-order capital terms is positive and significant<sup>40</sup>, as are the first-order output terms on access lines and local calls. However, the coefficient on toll calls is negative in this specification, while it is not significant in others which were tested.

The output interaction term between access lines and local calls ( $\ln Lines_{it} \cdot \ln LCalls_{it}$ ) is negative and significant, therefore indicating the presence of economies of scope. This implies that an increase in one of the outputs leads to cost reductions for a firm offering both outputs and is expected, given that the provision of local calls essentially relies on the access network. Regarding access lines and long-distance calls ( $\ln Lines_{it} \cdot \ln TCalls_{it}$ ), it is not expected that these services would exhibit economies of scope and, in the present sample, the coefficient on their interaction term is not significant. Finally, this study confirms the finding in Resende (1999) that local and toll calls ( $\ln LCalls_{it} \cdot \ln TCalls_{it}$ ) are not characterized by economies of scope.

Turning to price coefficients, the ratio of labor to materials input price ( $\ln PriceRatio_{it}$ ) is positive and significant. Estimated factor shares are positive at all observations and are in line with actual shares. In addition, the interaction term with time reveals a decrease of the labor share over the relevant time period. This may be probably related to an effort to cut costs, induced both by the move to incentive regulation and by increased competition over the sample period.

The measure of disembodied technical change ( $t$ ) is significant and the negative

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account of the fixed inputs. Following Caves et al. (1981),  $RTS = \frac{1 - \sum (\partial \ln VC / \partial \ln K_i)}{\sum (\partial \ln VC / \partial \ln y_i)}$ , where  $y$  indicates output and  $K$  fixed factors.

<sup>40</sup>Given the definition of the variables, the derivative of the cost function with respect to capital equals the partial derivative of  $\ln VC$  with respect to  $\ln J$ , because the second-order term and the cross-products equal zero at the sample mean.



coefficient indicates improvements over the period considered in the present study. This is consistent with general observations about positive technical change in the telecommunications sector and, for instance, with the findings in Resende (1999) on the 1988 - 1994 period for a sample of U.S. Local Exchange Carriers and in Correa (2003) on U.K. operators from 1990 to 1997.

Finally, the model in Table 1.2 is used to compare the log-likelihood from a range of assumed values of the parameter  $\gamma$  (i.e. the rate of embodied technical change). The best fit is found for  $\gamma = 4.9\%$  and this is robust to changes to the shape of the decay function for capital equipment.<sup>41</sup>

## 1.5.2 Estimation Results of the Production Function

### Nonlinear Production Function

As for the cost function, different values of the parameter of embodied technical change ( $\gamma$ ) are assumed. On this basis, the corresponding  $J_{it}$  is calculated and the production function is estimated for each of the values of  $\gamma$ . However, as a first step, the production function is estimated under the assumption of no embodied technical change (i.e.  $\gamma = 0$ ), with the parameter of physical decay  $b$  set to its base value of 0.5, in order to concentrate on the specification of the function.

The translog production function does not give significant estimates for the coefficient on capital  $J_{it}$ , for alternative specifications.<sup>42</sup> For this reason, this part of the analysis relies on the results from a Cobb-Douglas production function, which are presented in the table below (Table 1.3).

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<sup>41</sup>For other specifications, namely when the time trend was replaced with time dummies (with and without firm fixed effects), the estimation did not converge.

<sup>42</sup>In addition to the full translog function, alternative specifications were tested, e.g. the time trend was omitted or the cross-products terms for all variables were excluded from the specification.

Table 1.3: Cobb-Douglas Production Function - Results for  $\gamma = 0$

	OLS (1)	OLS (2)	IV (3)	IV (4)
Dependent Variable: $\ln y_{i,t}$				
$\ln L_{it}$	0.167*** (0.057)	0.168*** (0.053)	0.141** (0.054)	0.166*** (0.056)
$\ln M_{it}$	0.189*** (0.032)	0.188*** (0.032)	0.194*** (0.028)	0.193*** (0.027)
$\ln J_{it}$	0.494** (0.235)	0.510*** (0.174)	0.626*** (0.049)	0.608*** (0.049)
$t$	0.016* (0.009)	0.016** (0.008)	0.012** (0.004)	0.014*** (0.004)
$\ln S_{it}$	0.046 (0.171)	-	-	-
$O_{it}$	-0.222 (0.367)	-	-	-
Constant	1.329*** (3.268)	1.763 (3.243)	-	-
Observations	392	392	364	364
Fixed effects	Yes	Yes	Yes	Yes
F test <sup>44</sup>	106.33***	131.99***	-	-
Log likelihood	487.99	487.62	-	-

Note: Standard errors in brackets under coefficients in all columns are robust to heteroscedasticity and serial correlation. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. Variables: y: output; L: labour; M: materials; J: capital adjusted for embodied technical change (see Eq. 1); S: structures other capital), O: sheath per access line; t: time trend.

As mentioned above, there are endogeneity issues arising in the estimation of a production function due to the correlation between inputs and factors, such as managerial ability, which are potentially observable by firms when they make input decisions, but which are treated as unobservable in the estimation. The endogeneity problem has been traditionally addressed by fixed effects estimation and by instrumental variables (Akerberg et al., 2005). In light of this observation, the Cobb-Douglas production function is initially estimated by a fixed effects estimator. The reported standard errors are robust to heteroscedasticity and serial correlation.<sup>43</sup>

In Column 1, the results from the fixed effects estimator are reported. All coefficients are significant and of the expected sign. The time trend captures the impact of disembodied technical change, as commonly modeled in cost and production functions. Both  $J_{it}$  and  $S_{it}$ , i.e. capital with 'higher technical change' and 'lower technical change' respectively, are included. However,  $S_{it}$  does not seem to

<sup>43</sup>This adjustment is implemented in Stata following the result in Wooldridge (2002). It allows for heteroscedasticity and for serial correlation across observations for the same firm.

have a significant effect on output, once the other inputs are included in the model. Moreover, total sheath divided by the number of total access lines was included as an indicator of the different operational environments faced by the operators ( $O_{it}$ ): higher sheath per line indicates lower economies of density and therefore a worse environment for the operator. The coefficient on this variable is not significant, which could be explained by the inclusion of fixed effects and a time trend. When the model is estimated by OLS, without time dummies or a trend, then total sheath per line has a negative impact on output, as expected.

Column 2 shows the estimates after removing the variables whose coefficients were not significant.

Moreover, in order to take account of the endogeneity of the capital and labor inputs, the model is estimated by instrumental variables. In Column 3, capital ( $J_{it}$ ) is treated as endogenous and instrumented using lagged capital and the price of capital. In Column 4, both capital and labour are allowed to be endogenous and, in addition to the instruments used in Column 3, lagged labor and the price of labor are also included. In both models, the coefficient on capital increases compared with the case in which endogeneity is not taken into account. While the test of overidentifying restrictions does not reject the null hypothesis of instrument validity for Columns 3 and 4,<sup>45</sup> there are questions as to whether lagged inputs can be used as instruments, given that residuals are serially correlated. Moreover, the price of capital has a weak correlation with both labor and capital, and its validity as an instrument is therefore questionable.

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<sup>45</sup>This test is applied to a regression estimated via instrumental variables, when the number of instruments is greater than the number of included endogenous variables. The null hypothesis is that excluded instruments are valid, i.e. uncorrelated with the error term.

For Columns 3 and 4 in the table, the test of overidentifying restrictions does not reject the null hypothesis of instrument validity.

Based on these considerations, the specifications reported above in Table 1.3, Column 2, is chosen for the purpose of comparing the results obtained for a range of values of the rate of embodied technical change  $\gamma$  and the parameter of physical decay  $b$ .

The model is estimated for different values of  $\gamma$ . In addition, for each  $\gamma$ , different values of the rate of physical decay  $b$  (see Equation 1.2) are assumed in order to check the sensitivity of the results to this parameter, ranging from 0.5 (base case) to 0.75 (which implied less physical decay throughout an asset's life and a very fast decay at the end of it). Therefore, a different  $J_{it}$  is constructed for each combination of the rate of embodied technical change  $\gamma$  and the rate of physical decay  $b$ . The production function is then estimated for each of those combinations and the sum of squared residuals is used to compare the goodness of fit for the different combinations  $(\gamma, b)$ . On the basis of the sum of squared residuals, the value of  $\gamma = 0$  gives the best estimates.

Finally, the models reported in the table above are estimated without properly accounting for the behaviour of the variables over time. For instance, when time dummies are included, rather than a simple trend, the coefficient on capital  $J_{it}$  becomes insignificant. In order to alleviate the problems raised by trended variables, the production function is also estimated in terms of deviations from the sample mean for each year, following the approach in Bloom et al. (2006).<sup>46</sup> However, the estimation of the simple function reported in Table 1.3 above results in values of the parameter of embodied technical change  $\gamma$  in the range of 15% - 20%, which are

<sup>46</sup>Variables are defined as  $x_{it} \equiv \ln X_{it} - \ln \bar{X}_t$ , where  $\bar{X}_t$  is the sample mean of the variable at time  $t$ .

As can be seen in the Data Appendix to this chapter, for  $\ln y_{it}$ , this procedure dramatically changes the pattern of the variable.

Table 1.4: Cobb-Douglas Linearized Production Function

	OLS (1)	OLS (2)	IV (3)	IV (4)
Dependent Variable: $\ln y_{i,t-1} - \ln y_{i0}$				
$\ln L_{it} - \ln L_{i0}$	0.163** (0.072)	0.133* (0.066)	0.135* (0.068)	0.817* (0.402)
$\ln M_{it} - \ln M_{i0}$	0.210** (0.087)	0.207** (0.096)	0.206** (0.090)	0.140 (0.100)
$a_{it}$	0.176 (0.144)	0.355*** (0.078)	0.362*** (0.068)	0.780** (0.270)
$b_{it}$	0.025*** (0.008)	0.020*** (0.007)	0.019** (0.007)	0.021** (0.009)
$\ln S_{it} - \ln S_{i0}$	0.306 (0.194)	-	-	-
$\gamma$	-	5.73%	5.25%	2.69%
Observations	392	392	364	364
F test	47.63***	55.46***	54.85***	29.60***

Note: Standard errors in brackets under coefficients in all columns are robust to heteroscedasticity and serial correlation. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. Variables: y: output; L: labour; M: materials; a: sum of investment flows; b: product between investment flows and time trend; S: structures ( other capital).

outside the range found following other approaches and by Sung (1998, 2002).<sup>47</sup>

In line with the literature on the estimation of production functions, a dynamic specification was also estimated using the Arellano and Bond (1991) estimator. Given the small number of cross-sections in the present sample, this estimator is not particularly suitable for this study. Results from the dynamic specification are reported in Appendix.

### Linearized Production Function

As explained above, estimating Equation 1.7 allows recovering the rate of embodied technical change  $\gamma$  and the output elasticity of capital  $\theta$  separately. Due to the way variables are defined in Equation 1.7, the first year in the sample is excluded. The model is estimated by OLS and standard errors are robust to heteroscedasticity and serial correlation.

The table below presents the results from estimating Equation 1.7, including values for  $\gamma$  obtained as the ratio of the coefficients on  $b_{it}$  and  $a_{it}$ .

<sup>47</sup>This approach was only followed for physical depreciation corresponding to  $b = 0.5$  for capital  $J_{it}$ .

When  $\ln S_{it} - \ln S_{i0}$  is included in the production function (Column 1), its coefficient and the coefficient on  $a_{it}$  are not significant. Similar results are obtained when other variables (i.e. time trend, time dummies or variables to reflect the different operating conditions of the operators) or fixed effects are included. The impact of including further (nonsignificant) variables on the coefficients on  $a_{it}$  and  $b_{it}$  seems related to the high collinearity between  $a_{it}$  and  $b_{it}$ , which also implies that in some specifications it is not possible to disentangle the effect of each variable separately.<sup>48</sup>

In order to deal with the potential endogeneity of inputs, the model is estimated by two-stage least squares. The structure of input demand functions suggests input prices as natural instruments in this context. In Column 3, capital is instrumented by the corresponding input price (i.e. the price of capital, constructed as described in Appendix). When  $\ln L_{it} - \ln L_{i0}$  is also treated as endogenous (Column 4), and compensation per employee is used as a further instrument, the coefficient on the variable increased to 0.8, while the rate of embodied technical change falls to 2.69%.<sup>49</sup> Overall, the results from instrumental variables estimation seem to confirm broadly those from OLS estimation.<sup>50</sup>

However, as already mentioned in Section 1.3.3, the assumptions required for the estimation of Equation 1.7 are valid only for some time periods in the sample, which makes the results above not suitable to draw conclusions on embodied technical change.

Finally, as a check to the results above, the production function is estimated in levels following a methodology similar to Sung (1998, 2002). Rather than relying

<sup>48</sup> While it is possible to reject the joint hypothesis that the coefficients on  $a_{it}$  and  $b_{it}$  are zero, the two single hypotheses that the coefficients are zero cannot be rejected simultaneously.

<sup>49</sup> Given that materials were constructed as residual expenses,  $\ln M_{it} - \ln M_{i0}$  was not treated as endogenous.

<sup>50</sup> A GMM estimator is also applied, but the small number of cross-sections relative to the number of instruments makes the GMM procedure not suitable to the present study.

on the expression derived by algebraic manipulation of the non-linear production function, variables  $a_{it}$  and  $b_{it}$  are replaced by their ratio (i.e. average vintage as a proxy for capital quality) and  $\ln J_{it}$  is also included, under the assumption of  $\gamma = 0$ . In this specification, the coefficient on the ratio was 0.026, corresponding to  $\gamma = 2.6\%$ ,<sup>51</sup> which is within the range estimated in the above mentioned studies.

## 1.6 Concluding Remarks

Embodied technical change was assessed in the present paper through the estimation of a production function and of a cost function for a panel of 28 Local Exchange Carriers (LECs) over a 14-year period. Following Sakellaris and Wilson (2004), the proposed methodology defined capital as the sum of investment flows, adjusted to take into account different quality across vintages. This approach resulted in a non-linear function in the parameter of interest (i.e. the rate of embodied technical change) and the specific formulation posed an identification problem, as the rate of embodied technical change could not be estimated separately from the coefficient on capital.

For this reason, the problem was addressed in two alternative ways. Firstly, as explained in the paper, a range of values for the parameter of embodied technical change  $\gamma$  were assumed and the corresponding value of capital was calculated for each of them. A translog cost function was then estimated for each of those values and the best fit was obtained for  $\gamma = 4.9\%$ . As a robustness check, a production function was also estimated which, however, did not allow distinguishing between embodied and disembodied technical change. Secondly, a linearized Cobb-Douglas

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<sup>51</sup>The equation in levels is estimated including fixed effects and with clustered standard errors.

production function was estimated, resulting in rates of embodied technical change in the interval between 2.6% and 5.7% per year approximately.

In order to draw implications for setting the level of the price cap, Total Factor Productivity (TFP) for the sample of Local Exchange Carriers was calculated under the assumptions of a rate of embodied technical change equal to zero and equal to 5.7% per year (i.e. the lowest and highest estimates from the present study). For illustration, a simplified setting was used and TFP was calculated without taking account of other factors, such as economies of scale and inefficiency.<sup>52</sup> For the positive rate of embodied technical change (i.e. 5.73%), average TFP was higher compared to the case of no technical change. The result that TFP is higher when capital is adjusted for embodiment (rather than when it is not adjusted) depends on the choice of the numeraire year which means that, in the present approach, capital stock increases faster when the rate of embodied technical change is lower, rather than vice versa (Sakellaris and Wilson, 2004).

The difference between the TFP rates indicates that part of the productivity growth that feeds into the price cap is due to the past investment profile of the regulated operators. In the U.S. regulatory framework, the level of the cap which limits increases in average prices is "last year's cap, increased by inflation in the overall economy, but adjusted for productivity differences between the average firms

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<sup>52</sup>TFP growth between period  $t - 1$  and  $t$  was measured by the Törnquist index (see Fuss and Waverman, 2002):  $\Delta \ln TFP = \Delta \ln Y - \Delta \ln X$ , where  $\Delta \ln Y = \sum (1/2)(R_{jt} + R_{j,t-1})[\ln Y_{jt} - \ln Y_{j,t-1}]$  and  $\Delta \ln X = \sum (1/2)(S_{it} + S_{i,t-1})[\ln X_{it} - \ln X_{i,t-1}]$ .

$Y_{jt}$  is the amount of output  $j$  produced at time  $t$

$X_{it}$  is the amount of input  $i$  utilized at time  $t$

$R_{jt}$  is the revenue share of output  $j$  at time  $t$

$S_{it}$  is the cost share of input  $i$  at time  $t$

For simplicity, output was measured by the basket of services used for the production function estimation and therefore the above formula simplifies.

This representation of TFP, in particular the form of the output index, relies on the assumption of proportionality between price and marginal cost (Fuss and Waverman, 2002). Given that the firms in the sample are regulated and their prices are cost-based, this assumption can be expected to hold.



in the overall economy and the industry being regulated." A higher TFP for the telecommunications operators would therefore result in an increase in the overall "X factor", *coeteris paribus*. Therefore, the correct measurement of capital quality would be an important input in the regulatory process.

However, it should be noted that factors other than technical change may affect TFP and, in consequence, the "X factor". In particular, on the input side, investment in new vintages can be expected to go in parallel with substitution between labour and capital, which may be another reason why TFP is higher. On the output side, an increase in demand over time could lead to the modernization of the capital stock, as is currently the case for broadband connections. Therefore, the increase in TFP and in the X factor may depend on the rate of growth of demand since this will determine the speed of investment and hence the rate of embodiment.

In addition, while we tried to address the well-known problems in aggregating different vintages of capital, the definition of capital stock used in the analysis may still provide a distorted measure. This was derived from available data on capital expenditure, adjusted to reflect the better quality of recent vintages of capital compared to old ones. However, the implication is that the present approach, by artificially reducing the value of capital expenditure in earlier years, may distort the estimated coefficients.<sup>53</sup> If the price of equipment was available, a more promising alternative would be to estimate the input demand functions to capture the effect of lower prices (i.e. better quality) of equipment in a direct way.

Finally, the validity of the production function estimation was limited by issues of unobserved productivity differences and simultaneity in the production function

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<sup>53</sup>For instance, if the lower price of equipment and an increase in capital compensated each other, capital expenditure would remain constant. The firm's increased use of capital would not be represented in our measure.

approach, which were addressed by standard techniques relying on instrumental variables and GMM procedures.<sup>54</sup> As a possible improvement on the present setting, the production function and the cost function may be incorporated in a wider framework to fully take into account the endogeneity of inputs. In particular, considering the investment decisions of the regulated firms in some detail would seem a promising avenue for future research, for instance analyzing technical change on the basis of investment behavior (e.g. following Hobjin, 2001) or in the context of a structural model of supply and demand (e.g. Nadiri and Nandi, 1999).

## **1.7 Appendix 1 - Construction of the Variables and Data Quality**

### **1.7.1 Construction of the Variables: Production Function**

For the production function, output was proxied by deflated revenues due to the difficulties of obtaining a basket of the LECs' output measured in quantities. Specifically, output for the LECs should include lines and call minutes. There are three types of lines provided by the LECs: switched access lines, special access lines (i.e. "dedicated lines from the customer to the interexchange carrier point of presence"<sup>55</sup>) and leased lines. Volume data for leased lines were not available; a possibility could then be to ignore lines. It is not clear whether this would be a significant omission or not, given that it was not possible to identify how much of the LECs' revenues was represented by leased lines. More importantly, ignoring leased lines may introduce

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<sup>54</sup>For the production function, the Levinsohn and Petrin (2003) methodology could not be utilized due to limited variation in the intermediate input (i.e. materials) which did not allow identification of the parameters.

<sup>55</sup>Europe Economics, Operating Costs for the Access Network in Ireland: an Econometric Approach, 27 February 2004.

distortions between different LECs, since different operators may provide different shares of leased lines on total output.

For these reasons, output was constructed as an index of revenues. In particular, it is the weighted sum of local services revenues, long-distance revenues, and access and miscellaneous revenues. Each of these were deflated using appropriate price indexes for U.S. cities from the Bureau of Labor Statistics. A chained index was constructed, with the weights given by the revenue shares of the previous year. This was preferred to a Laspeyres index or a Paasche index. It was observed that there was some variability in the shares of the different services over the sample period and, therefore, the choice of either the initial period or the end period as the reference weights would introduce some distortions in the results.

The deviation of output from the sample mean in each period was also constructed in order to alleviate problems raised by the nonstationarity of the variable. The resulting variable is shown in the Figure 1.1.

In line with previous papers, materials were constructed as operating costs minus total compensation to employees and depreciation costs deflated by the producer price index for commodities (source Bureau of Labor Statistics).

Labour hours in a year were calculated as the number of full-time employees, multiplied by the average hours of work per week in the communications sector times 52 weeks. The source for the number of weekly working hours was the Bureau of Labor Statistics and the data were for the communications workers.

## **Capital**

The stock of capital was constructed separately for assets that could be expected to exhibit low levels of embodied technical change (e.g. land, buildings) and those

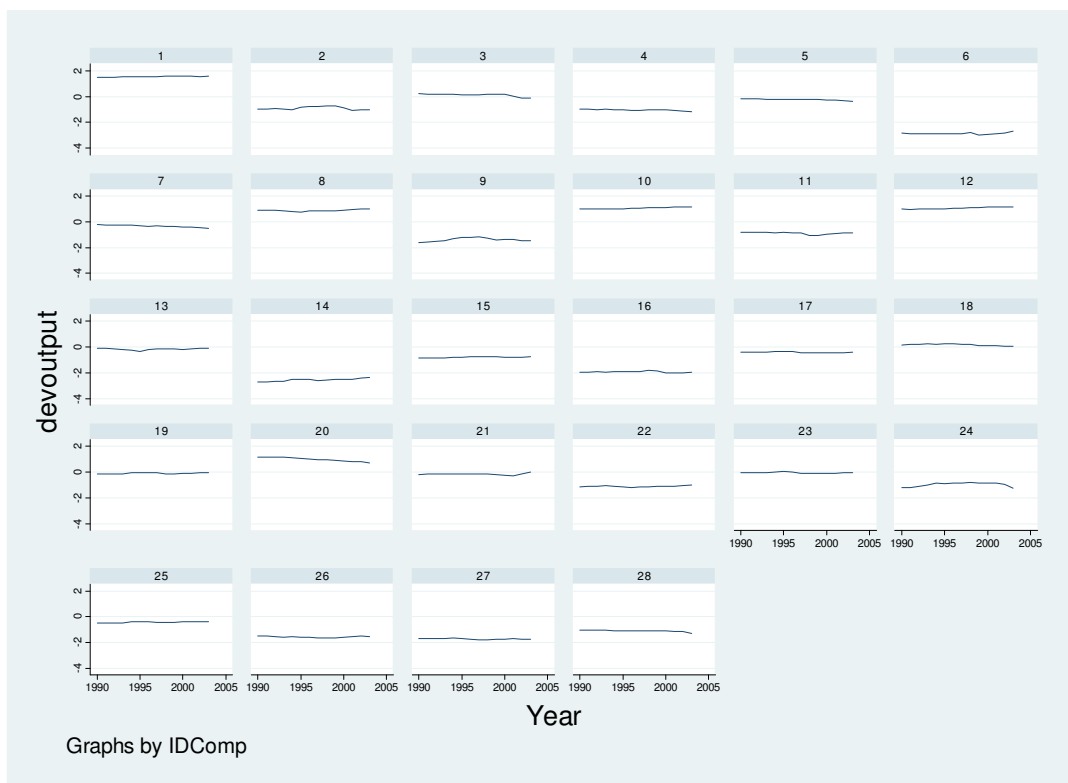


Figure 1.1: Deviation of Output from the Sample Mean ( $\ln y_{it} - \ln \bar{y}_t$ )

that could be expected to show higher embodied technical change (i.e. telecommunications equipment).

The procedure for deriving each LECs' capital input for "assets with higher embodied technical change" in a given year could be summarized as follows:

1. Gross investment per year per category of assets was derived from each company's accounts. The data source was the ARMIS 43-02 report, available from the FCC's website. Asset types such as land and support (e.g. furniture, buildings, vehicles) but also conduits and poles, were excluded. The procedure for deriving capital for asset types with expected low levels of embodied technical change is described below.
2. Investment was deflated to obtain real gross investment using an appropriate deflator at constant prices of the initial year in the sample (i.e. 1990). Following Sakellaris and Wilson (2004), the Personal Consumption Expenditure deflator from the U.S. Bureau of Labor Statistics was used. This was to ensure that the deflator was not a price index adjusted for quality changes in capital goods. As explained in the text, adjusting for quality changes in capital goods would remove the obsolescence effect which the present study aims to measure.
3. Real gross investment per year per asset category was then multiplied by the percentage of that vintage that was still physically productive in each of the subsequent years (see description of the hyperbolic decay function above). It should be noted that the investment surviving in each of the subsequent years had to be calculated, e.g. additions made in 1995 survive in different percentages in the years 1996 to 2001.

4. Each term thus obtained was multiplied by  $(1 + \gamma)^{t-s-t_0}$ , where  $t_0$  is the numeraire year, in this case year 2003,  $t$  is the year for which capital is constructed and  $s$  is the number of years prior to year  $t$ . Note that  $\gamma$  is unknown and is the parameter to be estimated.
5. Following the steps above, one obtains a flow of past investments relative to each year in the sample for a given asset category. Given that asset categories will have different asset lives, the number of terms containing  $\gamma$  will vary depending on the category.
6. The terms obtained above for the different asset categories were then summed to obtain the flow of surviving investment for a company.

The result of this procedure for a company at year  $t$  was a number of terms representing surviving investment flows from  $t$  to 2003, each a function of parameter  $\gamma$ . By repeating this procedure for every year in the sample, it was possible to reconstruct the stock of capital surviving at each point in time. Moreover, the initial capital stock in the first year in the sample (i.e. 1990) had also to be included to account for investments prior to that year.

Other assets were constructed as described above, except for the fact that the investment flows were not multiplied by the adjustment term to take into account embodied technical change.

### **Shape of the Decay Function**

In the analysis, the adjustment for physical decay was carried out using a hyperbolic (or beta) decay function (see Section 1.3). This function depends on the asset life ( $S$ ) for the specific type of asset and on the actual age of the asset ( $s$ ).

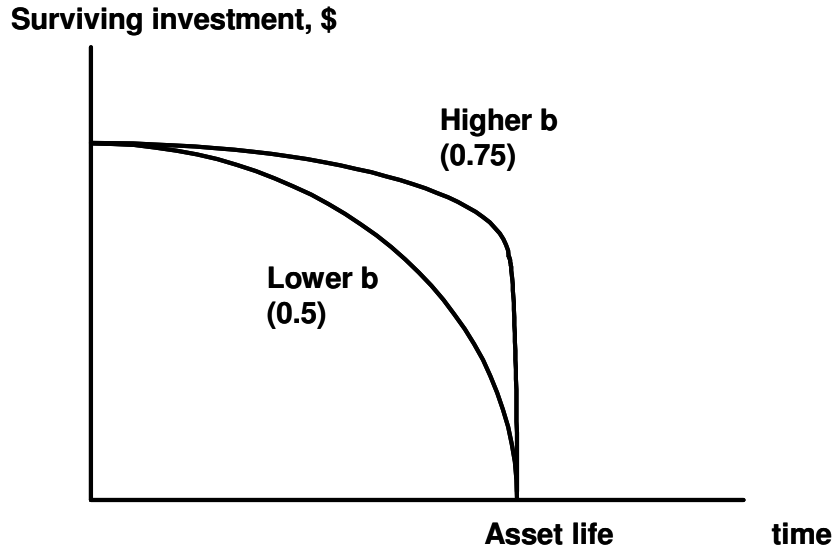


Figure 1.2: Hyperbolic Decay Function

$$D_{t,t-s} = \frac{S - s}{S - bs} \quad (1.8)$$

The parameter  $b$  reflects different curvatures of the function, as illustrated in the Figure below.

### 1.7.2 Construction of the Variables: Cost Function

While the variables relating to input prices were constructed by relying on the variables already obtained for the estimation of the production function, the output measures used in the cost function differ substantially from the basket of weighted revenues described above. In particular, exploiting the possibility to allow for multiple outputs in a translog cost function, both the number of lines and the number of calls were introduced in the cost function. The sources for all output data was ARMIS and, for some missing values, the Statistics of Common Communications Carriers, also provided by the FCC on its website.

Access lines were measured as the sum of switched access lines and special lines, while calls were given by local calls and toll calls. In turn, the latter were obtained as the sum of interLATA and intraLATA calls. See 'Production Function' above about the omission of leased lines due to non availability of the relevant data.

The behavior of variable costs, before transforming the variable, is plotted in the graphs below.

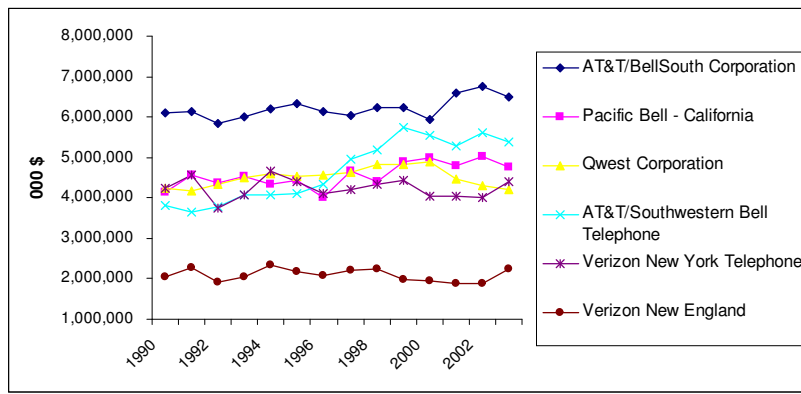


Figure 1.3: Deflated variable costs for the largest ILECs in the sample

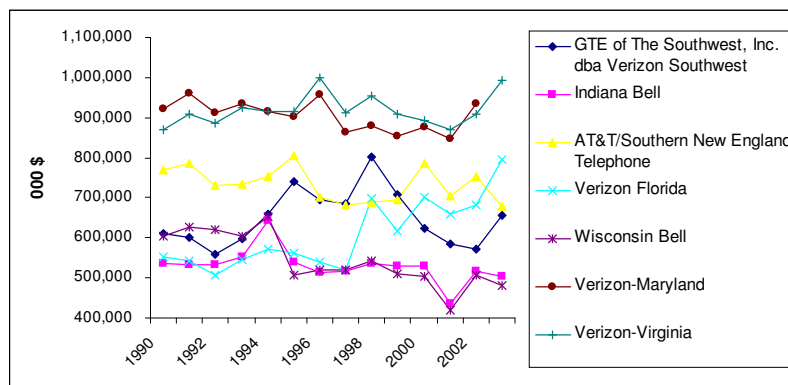


Figure 1.4: Deflated variable costs for large ILECs in the sample



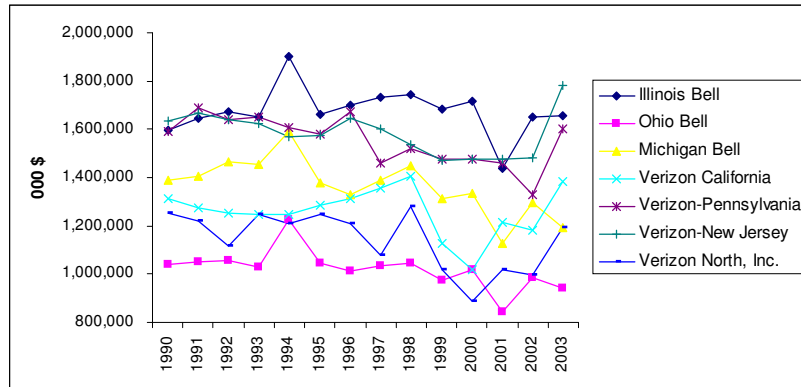


Figure 1.5: Deflated variable costs for medium-sized ILECs in the sample

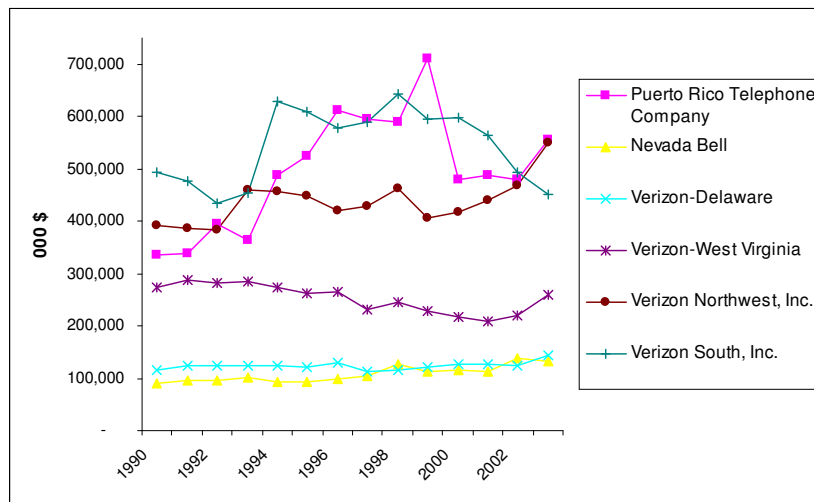


Figure 1.6: Deflated variable costs for small ILECs in the sample

### 1.7.3 Data Quality

The raw data used to construct the variables, as described above, were checked for consistency across time in order to identify possible errors. For instance, errors in reporting gross investment were detected in the FCC figures. In particular, prior to 2000, the instructions to the balance sheet accounts did not require data for 'additions' (i.e. gross investment) to be positive. In some cases, 'additions' were used to correct mistakes in previous years, even though there was a specific column (transfers/adjustments) for the purpose. In the cases where additions were negative,

it was assumed that there were no additions for that year.

Two operators were removed from the initial sample of large Incumbent LECs. In particular, the perimeter of operation for GTE Midwest changed in 2000 and observations were not available for the years 1990 to 1992. Contel was also removed from the sample because of lack of data for some years.

Regarding Puerto Rico, in the years 1990 to 1993, there used to be two operating companies, Puerto Rico Telephone Company and Puerto Rico Communications Corporation, and only the former was required to file its financial statements with the FCC. In 1994, the two companies merged. However, the difference between data in 1993 and in 1994 does not seem very significant (asset values increase by less than one fifth between 1993 and 1994). Moreover, the scale of the operator is very limited. For these reasons, the observations for Puerto Rico Telephone Company prior to 1994 were kept in the dataset.

Another comment concerns the price indexes used to deflate revenue data. These indexes are national averages rather than state-level indexes, which would be available for some of the services. This may introduce some distortions in the data. The reason why national averages were used is that some of the firms operate in more than one state and it is a long process to allocate revenues to different states on the basis of published figures. Similarly, labour hours are the average of the industry multiplied by the number of full time employees. This may not reflect different utilization across firms.

Finally, as described above, the construction of both variables  $J_{it}$  (expected to incorporate high embodied technical progress) and  $S_{it}$  (structures, expected to incorporate lower embodied technical progress) involves deflation by the Personal

Consumer Expenditure (PCE) index. While deflating structures by the same deflator used for telecoms assets may be desirable for consistency, it may not be correct. This is because structures themselves may incorporate a certain degree of technical change and would require a specific deflator reflecting improvements in structures quality. If, in fact, the quality-adjusted deflator for structures grew at a lower rate than the PCE, failing to take quality changes into account would result in underestimating the stock of structures in efficiency units. However, the price index for structures did not grow appreciably faster than the price index for consumption (Cummins and Violante, 2002) and therefore this effect may not be very important.

## 1.8 Appendix 2 - Dynamic Specification of the Production Function

In line with the literature (Akerberg et al., 2005), the production function is also estimated in a dynamic specification to reflect the fact that current levels of output might be affected by previous periods' outputs. In other words, the impact of past values of the regressors is assumed to be persistent and, in a dynamic specification, is captured by lags of the dependent variable. For this reason, we specify the following dynamic model:

$$\ln y_{it} = \alpha_i + \varphi \ln y_{i,t-1} + \beta \ln L_{it} + \delta \ln M_{it} + \theta \ln J_{it} + \xi \ln S_{it} + \eta_t + \varepsilon_{it}$$

where, for firm  $i$  in period  $t$ ,  $y_{it}$  is output,  $\alpha_i$  is a firm-specific effect,  $L_{it}$  indicates labour hours,  $M_{it}$  represents materials,  $J_{it}$  is the level of the capital stock and  $S_{it}$  is the capital that is assumed to incorporate a low rate of embodied technical progress.

The model above cannot be estimated by simple fixed effects because the estimator would be biased in this context, due to the correlation between the lagged dependent variable and the error term for a small number of periods (Nickell, 1981). The Arellano and Bond (1991) estimator is commonly employed in similar settings. The Arellano-Bond estimator transforms the model by differencing and relies on the Generalized Methods of Moments (GMM), where instruments include suitable lags of the variables. In addition, it allows for regressors which are not strictly exogenous and therefore can be fruitfully applied to the estimation of production functions.<sup>56</sup>

Results are reported below.<sup>57</sup>

Both the fixed effects and the GMM estimator indicate that the coefficient on lagged output is significant and of considerable size, while most other coefficients are not significant.

However, due to methodological problems, the above results are not reliable. As already mentioned, the fixed effects estimator is biased in short panels. Roodman (2006) refers to studies that find a 20% bias even when  $T = 30$  in simulations.

Regarding the Arellano-Bond estimator, the applicability of this type of estimator to the present sample is limited by the small number of firms relative to the number of periods. Firstly, the consistency of the estimator has been demonstrated

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<sup>56</sup>In the Blundell and Bond (1998) estimator, the original equation in levels is estimated together with the transformed equation, hence the definition of the estimator as "System GMM". Compared to the Arellano-Bond estimator, it is more efficient.

However, in the current setting, the limited size of the sample made its application difficult. When capital and labour were treated as predetermined, the number of instruments was very high, even when restricting the number of lags to be used as instruments.

The system GMM estimator was developed in: Roodman, D. (2005), `xtabond2`, Stata module to extend `xtabond` dynamic panel data estimator, Center for Global Development, Washington, <http://econpapers.repec.org/software/bocbocode/s435901.htm>

<sup>57</sup>The tests of the null hypothesis of serial correlation check whether one of the assumptions needed to use the estimator are satisfied. In particular, for lags to be used as instruments the idiosyncratic disturbance (i.e. excluding the firm effect) has to be serially uncorrelated.

The test is run on differenced residuals. If the assumption of no serial correlation in the levels is correct, the first difference is an MA(1) process and it has zero second-order autocorrelation. For this reason, it is expected that the test rejects the null hypothesis of no first-order autocorrelation and accepts the hypothesis of second-order autocorrelation.

Table 1.5: Cobb-Douglas Production Function, GMM Estimators - Results for  $\gamma = 0$

	Fixed Effects (1)	Arellano-Bond (2)	Arellano-Bond (3)
Dependent Variable: $\ln y_{i,t}$			
$\ln y_{i,t-1}$	0.737*** (0.061)	0.686*** (0.069)	0.759*** (0.076)
$\ln L_{it}$	0.030 (0.044)	0.044 (0.060)	0.065 (0.059)
$\ln M_{it}$	0.079*** (0.025)	0.065** (0.024)	0.069** (0.031)
$\ln J_{it}$	0.113 (0.189)	0.340 (0.255)	0.263 (0.218)
$\ln S_{it}$	0.029 (0.075)	-0.145 (0.128)	-0.108 (0.133)
Constant	0.058 (2.592)	-0.0004 (0.104)	0.155 (0.013)
Observations	364	336	336
Firm effects	Yes	Yes	Yes
Time dummies	Yes	Yes	Yes
F test	316.73***	47.86***	76.27***
First order serial correlation ( <i>p</i> -value)	-	-3.11 (0.002)	-3.49 (0.000)
Second order serial correlation ( <i>p</i> -value)	-	-0.74 (0.457)	-0.74 (0.460)

Note: Standard errors in brackets under coefficients in all columns are robust to heteroscedasticity and serial correlation. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. Variables: *y*: output; *L*: labour; *M*: materials; *J*: capital adjusted for embodied technical change (see Eq. 1); *S*: structures (other capital). In Column 2, lagged output is treated as endogenous. In Column 3, lagged output is treated as endogenous, while capital and labour are treated as pre-determined.

for samples in which  $N$  goes to infinity, under the assumption of a small number of time periods.<sup>58</sup> Secondly, the size of the instruments matrix increases quickly as more variables are treated as endogenous, leading to poor identification.<sup>59</sup> This is a common problem in GMM estimation. As the number of instruments increases, the number of elements in the covariance matrix also increases. In consequence, a large matrix needs to be estimated and a small sample may not provide enough information for the estimation of the covariance matrix. As a rule of thumb, when the number of instruments exceeds the number of cross-sections the estimates are not reliable (Roodman, 2006).

For these reasons, the dynamic formulation of the production function cannot be used as a basis for the estimation of embodied technical change.

<sup>58</sup> Arellano and Bond (1991) and Blundell and Bond (1998) apply the estimator to an unbalanced panel of 140 companies and between 7 and 9 yearly observations.

<sup>59</sup> For the GMM estimators reported in the table, the Sargan-Hansen test of the joint validity of the instruments found that the instruments were valid. However, this test is known to be weak when there are many instruments (Roodman, 2006).

## Chapter 2

# Estimating Time-Varying

# Technical Inefficiency for a

# Panel of Telecommunications

# Operators

## 2.1 Introduction

The measurement of relative performance is commonly found in a variety of different areas, including non-profit organizations (e.g. hospitals, nursing homes), industries over time or across geographical regions, and units within a firm (e.g. bank branches). Utilities regulators increasingly rely on the measurement of relative performance when implementing incentive-based regulation in a given infrastructure sector.

Econometric studies on efficiency usually rely on panel data and adopt sophis-

ticated estimation techniques. However, not all studies emphasize adequately that different assumptions and therefore different estimation techniques may produce contrasting results. In particular, the assumption, made in some approaches, that inefficiency is constant over a period of time appears difficult to justify. In addition, most panel estimators are not very good at disentangling heterogeneity from inefficiency (Greene, 2005b). In applications, these two factors distort estimates of inefficiency (Greene, 2005). The present analysis aims at comparing the efficiency scores provided by different estimation methods, in order to assess the impact of these concerns. Unlike previous studies, we apply panel estimators that distinguish between heterogeneity and (freely) time-varying inefficiency, in order to compare the results with those from traditional panel data methods.

The framework of the analysis is an input distance function. The advantages of a distance function are two-fold: firstly, it provides a natural extension of the production function for a multi-output firm; secondly, compared to a cost function, it does not require the input prices data and the behavioral assumption of cost minimization. However, related to the latter point, the distance function provides a partial picture as it only allows the estimation of technical efficiency, rather than overall economic efficiency.

Distance functions may be defined with an output orientation or with an input orientation. The latter appears to be more suitable to a regulated industry, as it seems more likely that such firms would have more control on inputs than on outputs. This argument is similar to those encountered in the literature on the empirical estimation of cost functions for multi-product regulated firms (Coelli and Perelman, 1996). For this reason, the analysis will rely on *input* distance functions

and focus on technical efficiency.

The present study implements this approach for a panel of 27 U.S. telecommunications operators over the 1990 - 2003 period. Our results indicate that panel models for which the inefficiency term is constant over time give very low efficiency scores on average. The underestimation of efficiency is expected, given that in these models the inefficiency term also captures time-invariant factors. When we estimate the model using fixed and random effects panel estimators with time-varying inefficiency, we obtain more reasonable values of relative efficiency. In addition, compared to standard panel estimators, they reduce the standard errors of estimated efficiency, which suggests that they address better the heterogeneity among firms. In contrast with more flexible estimators, the Battese-Coelli (1992) specification, which assumes that inefficiency follows a time trend common to all firms, generates efficiency estimates and a ranking of firms that almost coincide with those obtained when the inefficiency is time-invariant.

The structure of the chapter is as follows. Section 2.2 reviews papers which are relevant for the present study. Section 2.3 presents the model specification and the methodology. Section 2.4 describes the sample and the empirical results. Section 2.5 concludes.

## **2.2 Related Literature**

The estimation of productivity and efficiency in the U.S. telecommunications industry has been the subject of several articles, which adopt a variety of methodologies including Data Envelopment Analysis (DEA) (Majumdar, 1997; Resende, 2000; Uri, 2001) and stochastic frontier techniques (Resende, 1999), in particular cost fron-



tiers. Empirical contributions in the field of distance functions have mostly focused on other utilities. Therefore, before discussing the existing applications of distance functions to telecommunications data, we review the most relevant empirical contributions in other sectors.

Coelli and Perelman (1996) and Coelli and Perelman (2000) are among the first applications of distance functions to utilities. They study a panel of 17 European railways over the period from 1979 to 1983. The papers estimate a translog distance function, both input and output oriented, and compare the results with those obtained for a production function where output is measured as total revenues or as an index.

In a more recent application, Sickles et al. (2002) estimate an output distance function for a panel of 16 airlines from Eastern and Western Europe over the period 1977 - 1990. The functional form of the stochastic frontier is a Cobb-Douglas augmented by a second-order term for the output measure. Unlike Coelli and Perelman, Sickles et al. (2002) adopt a semi-parametric estimator and they compare its results with the Data Envelopment Analysis (DEA) method and the efficiency component of the Malmquist index. With all methods, they find that Eastern European carriers are less efficient compared to their Western European competitors. A comparison between stochastic frontier methods and the DEA approach is also provided by Berg and Lin (2008), who study a panel of 44 Peruvian water utilities in 1996 to 1998. They estimate an input distance function, mostly relying on cross-section data for the year 1998. The paper finds that the stochastic frontier models produce efficiency ranking that have a correlation around 0.5 and above with results from the DEA approach.

The contributions that are most closely related to the present study are Uri (2002) and Resende (2008). Relying on a sample of 19 Local Exchange Carriers (LEC) over the 1988 - 1999 period, Uri (2002) estimates a translog *output* distance function and finds no change in technical efficiency between 1988 - 1990 and 1991 - 1999, which he takes as an indication that incentive regulation did not promote technical efficiency. The estimation methodology implemented in the paper is the corrected OLS (COLS) approach. However, it is not clear whether the results are reliable, given that he finds that the first-order output coefficients are positive rather than negative as dictated by theory.<sup>1</sup>

Resende (2008) estimates a translog *output* distance function, using a sample of 30 LECs, over the period 1988 - 2000.<sup>2</sup> The paper compares the efficiency scores produced by Data Envelopment Analysis (DEA), deterministic frontier models (OLS and random effects) and the following stochastic frontier methods:<sup>3</sup> (i) Random effects time-invariant inefficiency; (ii) Battese and Coelli (1992) time-varying inefficiency<sup>4</sup>; and Battese and Coelli (1995) time-varying inefficiency.<sup>5</sup> Resende (2008) finds that efficiency scores do not show a monotonic variation across time. However, given that the equation shown in the paper does not include any time trend or time interactions it is not clear whether period specific events are adequately captured in the model and are therefore reflected in the error and in the inefficiency term.

In addition, he finds that the scores estimated with the Battese-Coelli (1992)

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<sup>1</sup>In an output distance function, keeping inputs constant, there is an inverse relationship between the amount of a given output the firm can produce and the other outputs.

<sup>2</sup>It is not possible to compare the results of the present chapter with those in Resende (2008), given that this working paper does not provide the estimated coefficients on the basis that the focus of the study is on efficiency scores and rankings.

<sup>3</sup>In all cases, it is assumed that the inefficiency follows a half-normal distribution.

<sup>4</sup>Inefficiency is assumed to be an exponential function of time, with a common parameter across firms.

<sup>5</sup>In this version of the model, the inefficiency depends on various explanatory variables, however only a time trend is used in Resende (2008).

model has a very high rank-order correlation with the stochastic frontier estimated under the random effects model. This appears consistent with our result that this estimator provides efficiency scores which are very close to those obtained using the fixed effects and the random effects models, with time-invariant inefficiency. Other methodologies show much lower correlations, as expected. With respect to the variation of efficiency patterns over time, Resende (2008) finds substantial persistence in the ranking provided by most estimators, which indicates only moderate changes in efficiency over time.

The differences between the sample used in Resende (2008) and in the present study arise in two respects. Firstly, our dataset derives the capital input using the perpetual inventory method, which requires data on gross investment. Such information is only available for the sample of large incumbent Local Exchange Carriers (LECs) we used. Secondly, the dataset in the present study extends to 2003 while, for the sample in Resende (2008), not all variables are available for recent years.

Moreover, unlike Uri (2002) and Resende (2008) the present study estimates an *input* distance function rather than an output distance function, given that it seems more likely that telecommunications operators set inputs rather than outputs. In addition, it explicitly models the interactions between a time trend, and inputs and outputs. This aspect appears especially important if the analysis aims at investigating changes in technical efficiency over time. In terms of estimation methodologies, the main differences with the above mentioned studies are: a) the implementation of estimators that allow for inefficiency to vary freely across periods and across firms; and b) the explicit modelling of heterogeneity.

## 2.3 The Estimation of Distance Functions

This section briefly reviews the main concepts from the literature on stochastic frontiers, specifically on production functions. In addition, it describes the role of distance functions in the estimation of technical efficiency for multi-output firms and specifies the functional form of the model estimated in the present analysis.

### 2.3.1 Frontier Production Function

The focus of this chapter is on technical efficiency, that is the relationship between the observed output a firm produces and the potential quantity of output it could produce as specified by a production function, given certain amounts of inputs. Assuming that the technology of interest concerns the production of a single output, the measure of technical efficiency can be embedded in a production function. It is common to empirically specify the function as follows:

$$\begin{aligned} y_i &= f(\mathbf{X}_i, \boldsymbol{\beta}) \cdot TE_i & (2.1) \\ \ln y_i &= \alpha + \boldsymbol{\beta}^T \cdot \mathbf{x}_i + \ln TE_i = \alpha + \boldsymbol{\beta}^T \cdot \mathbf{x}_i - u_i \end{aligned}$$

where  $i = 1, 2, \dots, N$  indexes the firms in the sample,  $\boldsymbol{\beta}$  is a vector of parameters and  $\mathbf{X}_i$  a vector of inputs. Technical efficiency, as defined above, is the ratio between observed output and potential output and satisfies  $0 < TE \leq 1$ . In the second line, as in most applications, it is assumed that  $f(\mathbf{X}_i, \boldsymbol{\beta})$  is linear in  $\mathbf{x}_i$ , the logs of the inputs (or functions of them). Moreover,  $u_i \geq 0$  is a measure of technical inefficiency, where  $u_i = -\ln TE_i$ . Therefore,  $TE_i = \exp(-u_i)$ . Under the assump-

tion of uncorrelation between the error and the regressors, OLS provides consistent estimates of the slope parameters but not of the intercept, given that  $-E(u_i) < 0$ . This limitation can be overcome by shifting the least squares line upward so that the largest residual is zero. This procedure is known as corrected OLS (COLS).

However, there is a more fundamental problem in the above formulation. The deviation of a given observation from the maximum output that could be achieved is all attributed to technical inefficiency, while random factors outside the control of the firm or measurement error play no role. As an extension of the previous model, the stochastic frontier production function was introduced by Aigner et al. (1977), as opposed to the deterministic approach above. For a sample of firms producing a single output, the stochastic model takes the form

$$\ln y_i = \alpha + \boldsymbol{\beta}^T \cdot \mathbf{x}_i + v_i - u_i \quad (2.2)$$

where  $v_i$  is a symmetric random error accounting for statistical noise and  $u_i \geq 0$  represents technical inefficiency. The composed error  $\varepsilon_i = v_i - u_i$  is therefore asymmetric. In the basic formulation of the model, it is assumed that  $v_i$  is normally distributed with zero mean and constant variance  $\sigma_v^2$ , while  $u_i$  follows a half-normal distribution, i.e. the nonnegative part of a normal with zero mean and constant variance  $\sigma_u^2$ .<sup>6</sup> It is commonly assumed that the errors are distributed independently of each other and independence is also assumed across firms  $i$ .

Finally, both errors are assumed to be uncorrelated with the explanatory variables  $\mathbf{x}_i$ . As above, OLS provides consistent estimates of the slope parameters but

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<sup>6</sup>Other distributions are the truncated normal ( $N(\mu, \sigma^2)$ ), the exponential and the gamma distribution.

Most of the empirical applications that compare the results from different distributions find reasonably robust estimates of inefficiency (but not necessarily of the parameters). For instance, see Greene (2008) and Kumbhakar and Lovell (2000).

not of the intercept, given that  $E(\varepsilon_i) = -E(u_i) < 0$ . Therefore, this estimation method cannot be used for the purposes of efficiency analysis. Maximum likelihood methods are usually adopted to estimate the parameters in the above model (Kumbhakar and Lovell, 2000). After estimation, the result by Jondrow, Lovell, Materov and Schmidt (1982) (JMLS) provides an estimate of  $E(u_i|\varepsilon_i)$  so as to disentangle the inefficiency from the random error.<sup>7</sup>

When firms are observed repeatedly over time, panel estimators can be applied to Equation 2.2 to address some shortcomings of cross-sectional models. For instance, firm heterogeneity can be identified by applying panel data estimators to the frontier model. Pitt and Lee (1981) adapted the random effects model to this context by assuming that the inefficiency component followed a half-normal distribution.<sup>8</sup> The resulting specification is given by  $y_{it} = \alpha + \boldsymbol{\beta}^T \cdot \mathbf{x}_{it} + v_{it} - u_i$ , where the inefficiency component is fixed over time. Following the same line of research, Schmidt and Sickles (1984) developed a fixed effects estimator for the frontier model, which provides consistent estimates even if inefficiency is correlated with input levels. For the fixed effects estimator (Schmidt and Sickles, 1984), the model is  $y_{it} = \alpha_i + \boldsymbol{\beta}^T \cdot \mathbf{x}_{it} + v_{it}$ , where  $\alpha_i = \alpha - u_i$  and no distributional assumptions on  $u_i$  are required. The inefficiency component is fixed over time and the estimated inefficiency is calculated as  $\hat{u}_i = \max(\hat{\alpha}_i) - \hat{\alpha}_i$ . According to the definition, at least one firm is assumed to be technically efficient and the efficiency of the others is measured in relative terms. In both of these early panel data models, the heterogeneity among firms is interpreted as inefficiency and therefore the advantage of repeated observations per firm is lost.

Moreover, the standard estimation of a stochastic frontier model with panel data

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<sup>7</sup>The result can also be applied to the truncated normal model, even though the parameters in the formula are different.

<sup>8</sup>As for the cross-sectional case, alternative assumptions on the distribution are possible.

Table 2.1: Summary of Econometric Specification: Standard Panel Models

	<b>Fixed Effects</b>	<b>Random Effects</b>	<b>Battese-Coelli</b>
Firm-specific component	Fixed ( $\alpha_i$ )	$u_{it} = u_i$	$u_{it} = \exp[\eta(t - T)] \times u_i$
Random Error	$\varepsilon_{it} = v_{it}$	$\varepsilon_{it} = v_{it} - u_i$	$\varepsilon_{it} = v_{it} - u_{it}$
	-	$u_i \sim N^+(0, \sigma_u^2)$	$u_i \sim N^+(0, \sigma_u^2)$
	$i.i.d.(0, \sigma_v^2)$	$v_{it} \sim N(0, \sigma_v^2)$	$v_{it} \sim N(0, \sigma_v^2)$
Inefficiency	$u_i = \max(\alpha_i) - \alpha_i$	$E[u_i   v_{it} - u_i]$	$E[u_{it}   v_{it} - u_{it}]$

also has drawbacks, in particular it relies on the assumption that the inefficiency term is fixed over time. While the estimation of the firm-specific component  $u_i$  improves as the number of available time periods increases, the same assumption of time invariance is less likely to hold.

In an attempt to alleviate this concern, a number of alternative specifications for the inefficiency term have been proposed. The Battese and Coelli (1992) formulation, which is commonly adopted in applications, including Resende (2008), characterizes the inefficiency as  $u_{it} = \exp[\eta(t - T)] \times u_i$ , where  $t$  is the period,  $T$  is the last period and the stochastic component  $u_i$  is time invariant. In the above model inefficiency is constrained to vary with an exponential pattern which is common to all firms in the sample. Battese and Coelli (1995) extend the framework by assuming that  $u_{it}$  is distributed as a truncated normal, and that the mean of the distribution of  $u_{it}$  is a function of explanatory variables which may for instance include time dummies.<sup>9</sup>

The main assumption on the panel models estimated in the chapter are summarized in Table 2.1.

The models proposed in Greene (2004) and Greene (2005a) differ from the papers by Battese and Coelli in that they allow  $u_{it}$  to vary freely from period to period for the different firms and they accommodate firm heterogeneity, instead of considering it as inefficiency. The assumption implicit in this formulation is that all inefficiency

<sup>9</sup>In Resende (2008), a time trend is the only variable considered. Given that the time trend is not firm-specific, this formulation generates a common pattern for all the firms in the sample as for Battese and Coelli (1992).

Table 2.2: Summary of Econometric Specification

	<b>Pooled</b>	<b>True FE</b>	<b>True RE</b>
Firm-specific component	None	Fixed ( $\alpha_i$ )	$w_i$
Random Error	$\varepsilon_{it} = v_{it} - u_{it}$ $u_{it} \sim N^+(0, \sigma_u^2)$ $v_{it} \sim N(0, \sigma_v^2)$	$\varepsilon_{it} = v_{it} - u_{it}$ $u_{it} \sim N^+(0, \sigma_u^2)$ $v_{it} \sim N(0, \sigma_v^2)$	$\varepsilon_{it} = v_{it} - u_{it}$ $u_{it} \sim N^+(0, \sigma_u^2)$ $v_{it} \sim N(0, \sigma_v^2)$
Inefficiency	$E[u_{it} v_{it} - u_{it}]$	$E[u_{it} v_{it} - u_{it}]$	$E[u_{it} w_i - u_{it} + v_{it}]$

is assumed to be time varying. This is the counterpart to the underlying assumption of the Schmidt and Sickles (1984) and the Pitt and Lee (1981) models, in which all inefficiency is assumed to be time invariant. In the so-called 'true' fixed effects model (Greene, 2004), the estimated specification is  $y_{it} = \alpha_i + \boldsymbol{\beta}^T \cdot \mathbf{x}_{it} + v_{it} - u_{it}$ , where a set of dummy variables is added to the cross-sectional stochastic frontier model. For the 'true' random effects model (Greene, 2005a), the model is  $y_{it} = (\alpha + w_i) + \boldsymbol{\beta}^T \cdot \mathbf{x}_{it} + v_{it} - u_{it}$ , where  $w_i$  is a random variable with zero mean and finite variance. The firm-specific effects capture heterogeneity and are assumed uncorrelated with the regressors.

In Section 2.4.2 below, the two models incorporating time-varying inefficiency are estimated and compared to their earlier counterparts.

### 2.3.2 Multioutput Production and Distance Functions

The production technology of telecommunications operators is, as for other utilities, multi-input and multi-output. A convenient way to describe this technology for the purposes of efficiency measurement is the distance function, which is intuitively the distance between the production frontier and the specific point in the technology set where a given firm is producing.

Distance functions are becoming increasingly common in the empirical literature for various reasons. Firstly, the distance function provides a natural extension of the production function for a multi-output firm. Secondly, compared to a cost function,



it does not require the behavioral assumption of cost minimization.

Distance functions may be defined with an output orientation or with an input orientation. Intuitively, the output distance function considers how the output may be proportionally expanded if the input vector is held fixed, while the input distance function measures "the amount by which the input set of each firm may be proportionally contracted with the output set held fixed" (Coelli and Perelman, 2000). The latter appears to be more suitable to a regulated industry, as it seems more likely that such firms would have more control on inputs than on outputs. This argument is similar to those encountered in the literature on the empirical estimation of cost functions for multi-product regulated firms. For this reason, the remainder of the analysis will focus on the discussion of input distance functions.

Formally, the production technology of the firm can be described from an input perspective by  $L(\mathbf{y})$ , the set of all input vectors  $\mathbf{x} \in \mathbb{R}_+^k$  that can produce the output vector  $\mathbf{y} \in \mathbb{R}_+^M$ , i.e.  $L(\mathbf{y}) = \{\mathbf{x} \in \mathbb{R}_+^k : \mathbf{x} \text{ can produce } \mathbf{y}\}$ . The input distance function may be defined on the input set as:

$$D_I(\mathbf{x}, \mathbf{y}) = \max\{\rho : (\mathbf{x}/\rho) \in L(\mathbf{y})\} \quad (2.3)$$

The input distance function,  $D_I(\mathbf{x}, \mathbf{y})$ , is non-increasing in  $\mathbf{y}$  and is non-decreasing, linearly homogeneous and concave in  $\mathbf{x}$ . Moreover, if the input vector  $\mathbf{x}$  is an element of the feasible input set  $L(\mathbf{y})$ , the input distance function will take values greater than or equal to unity. Using the notation above,  $D_I(\mathbf{x}, \mathbf{y}) \geq 1$  if  $\mathbf{x} \in L(\mathbf{y})$ . If the vector  $\mathbf{x}$  is located on the inner boundary of the input set, the distance function will be equal to one.

The distance function can be used to calculate the input-oriented measure of

technical efficiency, given by  $TE = \frac{1}{D_I}$ . If a firm is efficient, it will be on the frontier and both  $TE$  and  $D_I$  will be equal to 1. Throughout the majority of the literature, the Debreu-Farrell measure of technical efficiency implied by the definition above is adopted, that is technical efficiency is measured in terms of *equiproportionate* contraction of all inputs (input-orientation). In other words an input vector, for instance, is technically efficient for a given output if no *equiproportionate* contraction of all inputs is feasible. The advantage of radial measures is that they are invariant to the units of measurement.

Distance functions can be estimated by econometric methods, as in the present study, or mathematical programming techniques. Data Envelopment Analysis (DEA) is a linear programming method which constructs a piece-wise frontier and compares the each firm in the sample with a peer group. Based on the frontier, distance functions and efficiency measures can be defined with an input or an output orientation.

### **2.3.3 Econometric Estimation**

In this study, we rely on the translog functional form, introduced by Christensen et al. (1973), which is used very frequently in efficiency estimation because it is a flexible form, i.e. does not restrict the elasticities of substitution between inputs and allows for returns to scale to vary over the output range. In particular, it has been used in the distance function context by Coelli and Perelman (1999) and other studies. The translog distance function can be specified as:

$$\begin{aligned}
\ln D_{I,it}(\mathbf{x}, \mathbf{y}) = & \alpha_0 + \sum_{m=1}^M \alpha_m \ln y_{m,it} + \frac{1}{2} \sum_{m=1}^M \sum_{n=1}^M \alpha_{mn} \ln y_{m,it} \ln y_{n,it} \quad (2.4) \\
& + \sum_{k=1}^K \beta_k \ln x_{k,it} + \frac{1}{2} \sum_{k=1}^K \sum_{l=1}^K \beta_{kl} \ln x_{k,it} \ln x_{l,it} \\
& + \sum_{k=1}^K \sum_{m=1}^M \gamma_{km} \ln x_{k,it} \ln y_{m,it} + \sum_{k=1}^K \delta_k x_{k,it} t + \sum_{m=1}^M \phi_m y_{m,it} t \\
& + \lambda_1 t + \frac{1}{2} \lambda_{11} t^2
\end{aligned}$$

where  $i = 1, 2, \dots, N$  denotes firm  $i$  in the sample and  $t = 1, 2, \dots, T$  indicates the time period. Given the relevance of the time component for the present study, the standard translog function is augmented by a time trend and its interactions with the other variables.

From the continuity of the distance function, the symmetry between cross-derivatives follow:<sup>10</sup>

$$\alpha_{mn} = \alpha_{nm}, \quad m, n = 1, 2, \dots, M$$

and

$$\beta_{kl} = \beta_{lk}, \quad k, l = 1, 2, \dots, K$$

As the distance is defined in terms of an equiproportionate contraction of the inputs, the distance function is homogeneous of degree 1 in inputs. A very common and convenient way to impose the homogeneity constraints directly on the distance function derives from the observation that

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<sup>10</sup>These are directly imposed in the formulation of the translog function. In the empirical implementation of the translog, the two cross-products between the same variables are summed together.

$$D_I(\omega x, y) = \omega D_I(x, y), \forall \omega > 0$$

Therefore, one of the inputs can be chosen arbitrarily so that  $\omega = 1/x_K$ , which is equivalent to dividing all inputs by  $x_K$ . The estimated form of the distance function can then be rewritten as:

$$\begin{aligned} -\ln x_{K,it} &= \alpha_0 + \sum_{m=1}^M \alpha_m \ln y_{m,it} + \frac{1}{2} \sum_{m=1}^M \sum_{n=1}^M \alpha_{mn} \ln y_{m,iy} \ln y_{n,it} & (2.5) \\ &+ \sum_{k=1}^{K-1} \beta_k \ln x_{k,it}^* + \frac{1}{2} \sum_{k=1}^{K-1} \sum_{l=1}^{K-1} \beta_{kl} \ln x_{k,it}^* \ln x_{l,it}^* \\ &+ \sum_{k=1}^{K-1} \sum_{m=1}^M \gamma_{km} \ln x_{k,it}^* \ln y_{m,it} + \sum_{k=1}^{K-1} \delta_k x_{k,it}^* t + \sum_{m=1}^M \phi_m y_{m,it} t \\ &+ \lambda_1 t + \frac{1}{2} \lambda_{11} t^2 - \ln D_I \end{aligned}$$

where  $x_{ki}^* = x_{k,it}/x_{K,it}$ . In Equation 2.5, in line with common practice (Coelli et al., 2003)  $\ln D_I$  is no longer interpreted as a function, but as the value taken by the function itself.

$$\begin{aligned} -\ln x_{K,it} &= \alpha_0 + \sum_{m=1}^M \alpha_m \ln y_{m,it} + \frac{1}{2} \sum_{m=1}^M \sum_{n=1}^M \alpha_{mn} \ln y_{m,iy} \ln y_{n,it} & (2.6) \\ &+ \sum_{k=1}^{K-1} \beta_k \ln x_{k,it}^* + \frac{1}{2} \sum_{k=1}^{K-1} \sum_{l=1}^{K-1} \beta_{kl} \ln x_{k,it}^* \ln x_{l,it}^* \\ &+ \sum_{k=1}^{K-1} \sum_{m=1}^M \gamma_{km} \ln x_{k,it}^* \ln y_{m,it} + \sum_{k=1}^{K-1} \delta_k x_{k,it}^* t + \sum_{m=1}^M \phi_m y_{m,it} t \\ &+ \lambda_1 t + \frac{1}{2} \lambda_{11} t^2 + v_{it} - u_{it} \end{aligned}$$

Given that the distance can be interpreted as the difference between the observed

data and the predictions given by the transformation function, the distance term  $(-\ln D_I)$  is replaced with a composite error term,  $v_{it} - u_{it}$ , where  $v_{it}$  is an error term and  $u_{it}$  is the inefficiency component. As a result, the function can be estimated using the same methods developed for stochastic production frontiers (Coelli et al., 2003) and the efficiency score can be estimated by  $\exp(-u_{it})$ .

## 2.4 Data and Empirical Evidence

### 2.4.1 Descriptive Statistics

The data set used in this chapter relies on the statistics published by the U.S. regulator, the Federal Communications Commission.<sup>11</sup> It includes 27 incumbent Local Exchange Carriers (LECs) over the period from 1990 to 2003.

Local exchange carriers provide a range of different access lines, as well as local, intrastate and interstate calls. On the basis of data availability and for simplicity, the outputs considered in the analysis are access lines, local calls and toll calls. In particular, these measures exclude leased lines, which are not available for the entire sample period.

The inputs are aggregated into labor, capital and materials. Labor is measured by the number of full-time employees. In line with previous papers, materials are calculated as operating costs minus total compensation to employees and depreciation costs, deflated by the producer price index for commodities. Finally, capital is constructed using the perpetual inventory method and is measured in constant 1990

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<sup>11</sup>The statistical reports are contained in the Automated Reporting Management Information System (ARMIS), available from the regulator's website. The tables used for the analysis are Reports 43-01 (Annual Summary Report), 43-02 (USOA Report - Balance Sheet Accounts), 43-08 (Infrastructure Data Report) and 43-08 (Operating Data Report).

Table 2.3: Summary Statistics

Variable	Mean	Standard Deviation	Min	Max	Obs.	% Between Variation
Access Lines (units)	6,204,843	7,589,101	232,465	41,400,000	378	0.884
Local Calls (m)	15,800	20,100	529.459	99,300	378	0.978
Toll Calls (m)	2,826.463	3,129.041	84.337	16,800	378	0.937
Labor (units)	14,057.51	16,370.05	692	76,585	378	0.967
Materials (\$ 000)	860,616.5	969,545	54,248	4,455,259	378	0.942
Capital (\$ 000)	4,356,075	5,152,482	228,051.1	24,800,000	378	0.979
Sheath / line (Km/line)	0.034	0.020	0.002	0.117	346	0.651
Utilization (000)	0.393	0.221	0.038	0.976	374	0.940
Modernization (units)	0.090	0.040	0.006	0.220	373	0.304

Note: Fraction of variance due to between variation is defined as  $Var(u_{it})/(Var(u_{it}) + Var(\epsilon_{it}))$ , where  $u_{it}$  and  $\epsilon_{it}$  are the residuals of a GLS regression of the corresponding variable on a constant. (Farsi et al., 2005)

US\$.<sup>12</sup>

In order to account for the environment in which the firms operate and for the characteristics of their networks, we construct three variables. Firstly, a measure of density (sheath length per access line) is included to account for different geographic conditions across the operators. Secondly, as a proxy of the utilization of the network, the number of calls per switch is calculated. Finally, we also incorporate in the analysis an indicator of network modernization, given by the share of fiber on the total kilometers of cable.

The summary statistics are given in Table 2.3. As can be seen, there is substantial variability in the sample. Given the panel nature of the sample, it is useful to understand the share of "between" variation over the total, in order to gain an insight into the importance of heterogeneity across firms. The last column reports the share of between variation out of total variation. This measure highlights striking differences within the firms and indicates that the treatment of heterogeneity may pose concerns, if not fully captured by the variables included in the model.

While most of the variation in the sample derives from differences across countries, rather than changes over time, almost all variables show a trended pattern.

<sup>12</sup>Details are provided in Chapter 1.

For instance, all output measures increase over the sample period and, for some of the firms, there is a peak between 1999 and 2001, corresponding to the boom in the telecommunications industry at the end of the '90s. Capital also has an upward trend, while the number of employees decreases over time for all firms, which is consistent with labor saving technical change as found, for instance, by Resende (1999). In the estimation, we include a time trend and its interactions with inputs and outputs in order to account for common movements of these variables across firms.

#### 2.4.2 Estimation Results

In applications, it is common to transform the data by dividing each observation by its geometric mean so that the arithmetic sample averages of the variables in logarithms are equal to 0. This transformation allows direct interpretation of the first-order parameters  $\alpha_m$ ,  $\beta_k$  and  $\lambda_1$  as the elasticities evaluated at the sample mean. Therefore, in Equation 2.5, variables were transformed accordingly. Moreover, in the estimation capital was used as the "numeraire", in line with Coelli and Perelman (2000).

The results presented in this section were obtained under the assumption of half-normal distribution for the inefficiency term. The translog functional form, augmented with a time trend and interaction terms between time and inputs and outputs, was the starting point of the analysis. The likelihood ratio test was used to compare different restrictions to this general form and the following specification was chosen:<sup>13</sup>

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<sup>13</sup>The specifications were: 1. Full translog including time trend and interactions with time; 2. Translog with separability between inputs and outputs, i.e.  $\gamma_{km} = 0$ ; 3. Terms including time set equal to zero; 4. Restriction 2 and other cross-terms set equal to zero ( $\alpha_{mn} = 0, m \neq n$  and  $\beta_{kl} = 0, k \neq l$ ); 5. Cobb-Douglas.

$$-\ln x_{K,it} = \alpha_0 + \sum_{m=1}^M \alpha_m \ln y_{m,it} + \sum_{k=1}^{K-1} \beta_k \ln x_{k,it}^* + \frac{1}{2} \sum_{m=1}^M \alpha_{mm} \ln^2 y_{m,it} + \frac{1}{2} \sum_{k=1}^{K-1} \beta_{kk} \ln^2 x_{k,it}^* \quad (2.7)$$

$$+ \lambda_1 t + \frac{1}{2} \lambda_{11} t^2 + \sum_{k=1}^{K-1} \delta_k x_{k,it}^* t + \sum_{m=1}^M \phi_m y_{m,it} t + v_{it} - u_{it}$$

In Table 2.4, we report results from the following panel estimators: 1) Fixed effects (time-invariant inefficiency); 2) Random effects (time-invariant inefficiency); and Battese and Coelli (1992).

Results from the following estimators, which allow the inefficiency component  $u_{it}$  to vary freely from period to period, are shown in Table 2.5: 1) Pooled, i.e. the sample is treated as a cross-section; 2) Fixed effects (time-varying inefficiency); and 3) Random effects (time-varying inefficiency).

The latter two estimators are also known as 'true' fixed effects and random effects estimators given that, unlike their standard panel counterparts, they separate inefficiency from heterogeneity. In addition, as can be seen from the equations above, it is assumed that all inefficiency is time-varying, as opposed to the opposite assumption that all inefficiency is time-invariant, as in the standard panel models. More details on the differences between these estimators are provided in the previous section on methodology.

The estimated coefficients on the first-order input terms are positive in all models, as expected in an input distance function. However, results for the panel estimators

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For specifications 1, 3 and 5, the OLS residuals were skewed in the wrong direction and therefore the models were excluded.

We also checked whether all the firms in the initial sample could be pooled together. This was done by comparing the log-likelihood of the estimated model, removing from the sample one firm at a time. One firm which produced a negative contribution to the log-likelihood was removed from the initial sample.



Table 2.4: Standard Panel Models

Dependent Variable:	Fixed Effects (1)	Random Effects (2)	Battese-Coelli (3)
$-\ln K_{it}$			
$\ln L_{it}^*$	0.119*** (0.028)	0.124* (0.057)	0.109*** (0.028)
$\ln M_{it}^*$	0.127*** (0.015)	0.129*** (0.021)	0.139*** (0.016)
$\ln Lines_{it}$	-0.323*** (0.036)	-0.313*** (0.086)	-0.306*** (0.035)
$\ln LCalls_{it}$	-0.002 (0.023)	-0.014 (0.035)	-0.017 (0.022)
$\ln TCalls_{it}$	0.026* (0.014)	0.022 (0.023)	0.027 (0.014)
$(\ln Lines_{it})^2$	-0.260 (0.364)	-0.270 (0.406)	-0.198 (0.355)
$(\ln LCalls_{it})^2$	0.020 (0.017)	0.023 (0.032)	0.017 (0.016)
$(\ln TCalls_{it})^2$	0.008 (0.009)	0.009 (0.016)	0.007 (0.009)
$(\ln L_{it}^*)^2$	0.199*** (0.064)	0.201 (0.206)	0.204*** (0.064)
$(\ln M_{it}^*)^2$	0.037 (0.057)	0.045 (0.103)	0.048 (0.056)
$t$	0.004 (0.003)	0.004 (0.007)	0.004 (0.003)
$t^2$	-0.004** (0.002)	-0.004 (0.002)	0.001 (0.002)
$\ln Lines_{it} \cdot t$	0.052* (0.024)	0.052 (0.029)	0.044 (0.024)
$\ln LCalls_{it} \cdot t$	0.000 (0.002)	0.000 (0.002)	0.004 (0.003)
$\ln TCalls_{it} \cdot t$	-0.001 (0.002)	-0.001 (0.003)	0.001 (0.002)
$\ln L_{it}^* \cdot t$	0.014*** (0.004)	0.014 (0.012)	0.012*** (0.004)
$\ln M_{it}^* \cdot t$	0.005 (0.003)	0.004 (0.005)	0.002 (0.003)
Constant	-	2.174*** (0.407)	2.206*** (0.086)
$\frac{\sigma_u}{\sigma_v}$	n/a	52.533	26.509
$\mu$	-	-	2.174
$\eta$	-	-	-0.006
log Likelihood	644.098	505.874	511.645

Note: Standard errors in parentheses. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. L\*: labor normalized by capital; M\*: materials normalized by capital; Lines: access lines; LCalls: local calls; TCalls: toll calls; t: time trend.

Table 2.5: Models with Time-Varying Inefficiency

Dependent Variable:	Pooled (1)	True FE (2)	True RE (3)
$-\ln K_{it}^*$	0.282*** (0.055)	0.295*** (0.036)	0.198*** (0.010)
$\ln M_{it}^*$	0.369*** (0.047)	0.386*** (0.037)	0.137*** (0.007)
$\ln Lines_{it}$	0.168 (0.123)	0.133 (0.089)	-0.245*** (0.020)
$\ln LCalls_{it}$	-0.681*** (0.026)	-0.676*** (0.033)	-0.140*** (0.005)
$\ln TCalls_{it}$	-0.263*** (0.025)	-0.274*** (0.018)	0.001 (0.005)
$(\ln Lines_{it})^2$	-3.933*** (1.252)	-5.114*** (0.953)	0.366 (0.225)
$(\ln LCalls_{it})^2$	-0.003 (0.022)	-0.001 (0.012)	-0.004* (0.002)
$(\ln TCalls_{it})^2$	-0.008 (0.022)	-0.017 (0.012)	0.005 (0.003)
$(\ln L_{it}^*)^2$	0.311** (0.108)	0.330*** (0.075)	0.294*** (0.020)
$(\ln M_{it}^*)^2$	0.169 (0.265)	0.003 (0.122)	0.063 (0.034)
$t$	0.009 (0.009)	0.012 (0.006)	0.008*** (0.002)
$t^2$	-0.016** (0.006)	-0.021*** (0.004)	-0.001 (0.001)
$\ln Lines_{it} \cdot t$	0.220** (0.091)	0.293*** (0.063)	0.001 (0.015)
$\ln LCalls_{it} \cdot t$	0.001 (0.007)	0.002 (0.005)	0.000 (0.001)
$\ln TCalls_{it} \cdot t$	-0.013* (0.006)	-0.014** (0.005)	-0.003*** (0.001)
$\ln L_{it}^* \cdot t$	0.038*** (0.010)	0.040*** (0.008)	0.020*** (0.002)
$\ln M_{it}^* \cdot t$	-0.047*** (0.010)	-0.051*** (0.009)	0.005** (0.002)
Constant	0.214*** (0.041)	-	-
$\frac{\sigma_u}{\sigma_v}$	1.189	1.325	4.844
log Likelihood	142.133	184.603	464.782

Note: Standard errors in parentheses. \* significant at 10 %, \*\* significant at 5 %, \*\*\* significant at 1 %. L\*: labor normalized by capital; M\*: materials normalized by capital; Lines: access lines; LCalls: local calls; TCalls: toll calls; t: time trend.

that assumed time-invariant inefficiency, although very similar to each other, do not appear reasonable. For instance, while the coefficients on labor and materials are positive as expected, they are unrealistically low and imply a coefficient on capital which is above 0.7 in both cases.<sup>14</sup>

Results in Columns 1 and 2 (pooled and true fixed effects, respectively) are substantially different from those from other models and appear more in line with expectations. In the true random effects model, first-order input terms are significant

<sup>14</sup>The coefficient on capital can be recovered by exploiting the constraint that the sum of the first-order coefficients on inputs is equal to one.

but their magnitude is unreasonable, similarly to the results in Table 2.4.

Evidence on the output coefficients is equally contradictory and, depending on the specification, first-order terms in either lines or calls are significant. Again, results from the pooled and the true fixed effects models appear more reasonable. For instance, the sum of the significant first-order output coefficients is slightly less than one in Columns 3 and 4, indicating the presence of increasing returns to scale, as found for instance by Uri (2002) in a similar sample of U.S. operators. Regarding the significance of the first-order coefficient on lines, it should be noted that the pattern of lines over time matches closely the time trend. The parallel movement of lines and the time trend could explain why their interaction is significant, while it is not possible to identify the contribution of each variable separately.

The presence of inefficiency in the sample is found across estimators, as shown by the ratio of the standard error of the inefficiency term ( $\sigma_u$ ) and the standard error of the symmetric error component ( $\sigma_v$ ). However, this value is very high for the random effects model with time-invariant inefficiency, which may be due to the strong assumption that inefficiency is uncorrelated with the regressors and is also consistent with the observation that there is little "within" firm variation for some variables, i.e. firm heterogeneity is reflected in high values of  $\sigma_u$ .

For this reason, we attempt to model observed heterogeneity directly by incorporating in the estimation some characteristics of the operators' networks.<sup>15</sup> There are different alternatives as to how to address observable heterogeneity (Coelli et al., 1999), including: (a) in the model specification (i.e. the underlying assumption is that environmental variables affect production levels rather than inefficiency);

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<sup>15</sup>The variables are: sheath per line (dispersion of the network), share of fibre cable to total kilometres (modernization) and the ratio of calls to switches (capacity usage).

and (b) in the mean of the inefficiency distribution (i.e. the inefficiency term is assumed to follow a truncated normal distribution and its mean is a function of the environmental variables). In these models, environmental variables shift either the input function or the inefficiency. In terms of other assumptions, the first option does not present any differences compared to the standard pooled model, while the shift of the underlying mean of the inefficiency component (case b) implies that inefficiency follows a truncated normal distribution:  $u_i = |U_i|$ , where  $U_i \sim N(\mu_i, \sigma_u^2)$ ,  $\mu_i = \mu_0 + \mu_1' \mathbf{z}_i$ .

Table 2.6 provides results from these alternative models.

When the environmental variables are directly included in the distance function (Column 1), sheath per line is the only significant variable, but with a very small coefficient ( $\approx 0.0001$ ). Given that the mean of sheath per line is about 34 meters per line, the average effect of the variable is about 0.003, which is significant but small. An alternative approach is a two-step model, which differs from the one reported in the text in that conventional inefficiency estimates are obtained, omitting the influence of environmental variables, and then regressed on such environmental variables. Results from this approach confirm the significance of sheath per line, as shown in the text. However, incorporating the environmental variables directly in the model appears preferable, as argued by Greene (2007) and Wang and Schmidt (2002).

In Column 2, the mean of the inefficiency component  $u_{it}$  is assumed to be a linear function of sheath per line. As expected, higher sheath length (i.e. a less dense network) results in higher inefficiency.

Regarding the other estimates, the main difference with Table 2.4 is that the

Table 2.6: Models with Heterogeneity

Dependent Variable:	Environmental Variables (1)	Heterogeneity Mean (2)	Truncated Normal (3)
	$-\ln K_{it}$		
$\ln L_{it}^*$	0.285*** (0.055)	0.233*** (0.053)	0.280*** (0.057)
$\ln M_{it}^*$	0.374*** (0.048)	0.252*** (0.047)	0.364*** (0.049)
$\ln Lines_{it}$	0.127 (0.125)	0.274* (0.121)	0.175 (0.126)
$\ln LCalls_{it}$	-0.681*** (0.026)	-0.672*** (0.023)	-0.679*** (0.026)
$\ln TCalls_{it}$	-0.266*** (0.025)	-0.280*** (0.022)	-0.267*** (0.025)
$(\ln Lines_{it})^2$	-4.080*** (1.253)	-2.712** (1.170)	-4.034*** (1.295)
$(\ln LCalls_{it})^2$	-0.003 (0.022)	0.011 (0.022)	-0.005 (0.023)
$(\ln TCalls_{it})^2$	-0.011 (0.021)	-0.028 (0.021)	-0.008 (0.022)
$(\ln L_{it}^*)^2$	0.321** (0.109)	0.360*** (0.103)	0.308*** (0.111)
$(\ln M_{it}^*)^2$	0.138 (0.272)	0.291 (0.229)	0.168 (0.274)
$t$	0.015 (0.009)	-0.002 (0.009)	0.008 (0.009)
$t^2$	-0.020** (0.007)	-0.009 (0.006)	-0.017** (0.007)
$\ln Lines_{it} \cdot t$	0.235** (0.093)	0.125 (0.090)	0.227** (0.095)
$\ln LCalls_{it} \cdot t$	0.002 (0.007)	0.004 (0.006)	0.003 (0.007)
$\ln TCalls_{it} \cdot t$	-0.014* (0.006)	-0.015** (0.006)	-0.015** (0.006)
$\ln L_{it}^* \cdot t$	0.037*** (0.010)	0.048*** (0.010)	0.038*** (0.010)
$\ln M_{it}^* \cdot t$	-0.048*** (0.010)	-0.050*** (0.009)	-0.045*** (0.010)
<i>Sheath per line<sub>it</sub></i>	0.00009** (0.000)		
<i>Utilization<sub>it</sub></i>	0.000 (0.006)	-	
<i>Modernization<sub>it</sub></i>	0.000 (0.006)	-	
Constant	0.246*** (0.031)	0.356** (0.133)	0.297 (0.145)
Mean of inefficiency:			0.176
<i>Sheath per line<sub>i</sub></i>	-	4.128*** (0.737)	
$\frac{\sigma_u}{\sigma_v}$	1.582	1.397	1.338
log Likelihood	145.516	168.552	142.532

Note: Standard errors in parentheses. \* significant at 10%, \*\* significant at 5%, \*\*\* significant at 1%. In Column 2, the mean of the inefficiency distribution is assumed to take the form  $\mu_i = constant + \theta \times sheath_i$ . L\*: labor normalized by capital; M\*: materials normalized by capital; Lines: access lines; LCalls: local calls; TCalls: toll calls; t: time trend.

first-order coefficient on lines is significant and has the wrong sign in Column 2. In general, there are inconsistencies across the different estimators as to the sign and significance of coefficients on first-order, quadratic and interaction terms involving lines. There does not seem to be a consistent pattern across firms in the relationship between the dependent variable ( $-\ln Capital$ ) and the lines variable ( $\ln Lines$ ), and this may make the identification of the coefficients problematic. While lines are an important part of an operator's assets, the composition of capital, and therefore the share of transmission over total capital, may vary among different firms depending, for instance, on the distribution of customers over the service area. For instance, in the last year of the sample the proportion of transmission on total assets in operation varied from around 15% to about 30%.

In Column 3, we report results from the truncated normal model. The sample is treated as a cross-section, i.e. all the observations are pooled together without accounting for unobserved heterogeneity. The main difference with the pooled model in Table 2.5 is the assumption on the distribution of the inefficiency term, which in this case is  $u_i \sim N^+(\mu, \sigma_u^2)$ . The symmetric error remains unchanged compared to the standard model:  $v_{it} \sim N(0, \sigma_v^2)$ .

### **Efficiency Estimates**

Table 2.7 summarizes the efficiency measures obtained with the different estimators. It also reports results from the truncated normal model, for which coefficient estimates are provided in Appendix.

Results in terms of efficiency estimates are in line with the conclusions reached for model specification. For the panel estimators with time-invariant inefficiency, the estimates do not exhibit the expected shape and average technical efficiency is

Table 2.7: Efficiency Measures

	Mean	Std. Dev.	Min	Max	Obs.
Time-Invariant Efficiency					
Fixed Effects	0.188	0.221	0.012	1.000	378
Random Effects	0.193	0.206	0.013	0.866	378
Time-Varying Efficiency					
Battese-Coelli	0.183	0.217	0.011	0.970	378
True Fixed Effects	0.857	0.038	0.694	0.948	378
True Random Effects	0.933	0.047	0.784	0.995	378
Cross-Section Models					
Pooled	0.881	0.050	0.719	0.966	378
Environmental Variables	0.863	0.066	0.655	0.952	346
Truncated Normal	0.814	0.073	0.622	0.963	378
Heterogenous Mean	0.803	0.081	0.633	0.970	378

Notes: In the model with heterogeneous mean, network density (proxied by sheath per line) is the environmental variable used in the estimation. The number of observations in the model with environmental variables is lower because of missing observations for all firms (e.g. 1990 observations are missing for all).

very low. As found for instance in Greene (2004), the low estimates of efficiency are expected, given that these models do not allow separating heterogeneity from inefficiency, therefore overestimating the latter. In addition, in line with Greene (2004) and Resende (2006), we find little difference between estimates from time-invariant models and the Battese-Coelli (1992) formulation. Rankings are also very highly similar across these three models.

In all other models, efficiency measures show the expected shape and their averages are significantly higher. In line with our findings on coefficient estimates, these results confirm that for the present sample the estimators that perform better are those that allow the inefficiency term to vary over time, either by treating the sample as a cross-section or by separating time-varying inefficiency from time-invariant firm-specific effects. In addition, among these estimators, the true random effects model provides the highest estimates of efficiency, as reported for instance by Greene (2004) for the WHO dataset. However, we consider these estimates unreliable, given they are obtained from a model whose coefficients are out of range (Table 2.4, Column 2). In addition, the true random effects estimator did not converge for some of the model specifications tested and therefore its results for the present sample

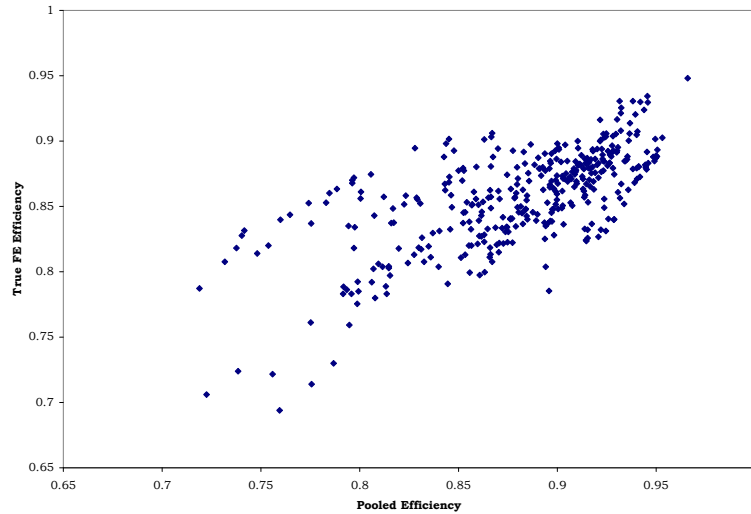


Figure 2.1: Efficiency Estimates - Pooled and True FE Models

should be taken with caution.

For these reasons, we focus on the results from the true fixed effects model and we compare its results with the cross-section estimators. All the latter give very similar estimates of efficiency, and this also holds for the models in which the inefficiency term is assumed to have a mean different from zero, i.e. the truncated normal and the model with heterogeneous mean. In the figure below, we plot efficiency estimates from the pooled model against the true fixed effects model.

When comparing the results of the fixed effects model with time-varying inefficiency with those of standard pooled model, a strong correlation emerges between the efficiency estimates. However, the overall variation of the efficiency estimates is lower in the true fixed effects model. Moreover, when average efficiency by firm is computed, the average estimates of the pooled model are more dispersed compared with the true fixed effects model, which produces results that are more similar across firms. Both these observations can be explained by the fact that the firm-specific dummy captures the unobserved heterogeneity across operators, thus leading to



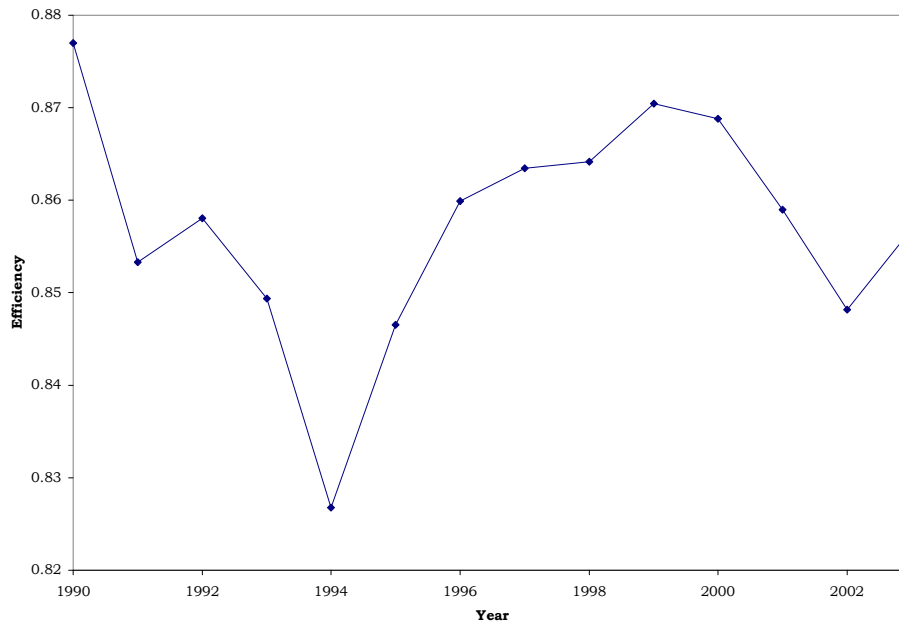


Figure 2.2: Average Technical Efficiency over Time

fewer differences among the efficiency scores of the operators.

In the figure below, we plot average efficiency over time, noting that the model specification used in the study already included a time trend and its interactions with inputs and outputs. As shown in the figure, average efficiency shows substantial variability over the sample period and therefore estimators which do not account for time variation do not seem suitable. Moreover, average efficiency does not follow a monotonic pattern, as also found by Resende (2006) in his sample of U.S. local exchange carriers over the 1988 – 2000 period.

The variation in technical efficiency over the 14 years covered by the study is expected, given its relatively long time-span and also the fast pace of change in the telecommunications market. The present sample extends from 1990 to 2003, a period which coincides with major changes in the U.S. industry. For instance, a "bubble" in investment in telecommunications companies and infrastructure took

place between 1997 and 2000 in the U.S. and abroad. In the following years, despite the fact that capital investment was drastically reduced,<sup>16</sup> operators experienced low rates of capacity utilization. Consistently with this observation, average efficiency in the present sample, after a peak in 1999, declined until 2002.

In addition, significant market changes were initiated in the U.S. by the Telecommunications Act of 1996. The Act opened local markets to competition, imposing at the same time conditions to make competition effective, e.g. on interconnection of networks, non-discrimination and cost-based pricing of network elements leased by new entrants.<sup>17</sup> Moreover, the Act required that certain conditions be met before the incumbent local exchange carriers (LECs) could enter the long-distance market (Economides, 2005).<sup>18</sup> Given that long-distance calls originate and terminate on local networks, local telephone companies would control access to a bottleneck input required by long-distance companies. If local operators were allowed to compete in the long-distance market with no conditions attached, they would have an incentive to act anticompetitively towards their rivals (e.g. raise the price of an essential input) (Armstrong, 2002).

In terms of the specific form of regulation adopted in the U.S., an increasing number of U.S. states moved from rate-of-return regulation to some form of incentive regulation from 1990 (Ai and Sappington, 2002). In light of this, the decline

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<sup>16</sup>As reported by Lenain and Paltridge (2003), the Telecommunications Industry Association announced that capital spending by U.S. telecommunications service providers in 2002 was back to the same level as 1997.

<sup>17</sup>In addition to building their own facilities, new entrants had two ways of entering the local market: 1) reselling retail services provided by the incumbent; 2) lease unbundled network components (Economides, 2005).

<sup>18</sup>After the breakup of AT&T in 1984, local exchange carriers that were formerly part of the Bell System (Regional Bell Operating Companies, RBOCs) were prevented from offering services outside certain areas defined by regulation (e.g. Nynex - now Verizon - operated in the states of New York and New England). Within these areas, they provided both local calls and toll calls.

The term 'long-distance call' in the U.S. market refers to a call outside the service area in which a given RBOC was allowed to operate, e.g. a call between California (Pacific Bell - now AT&T) and New England.

in technical efficiency from 1990 to 1994 may appear surprising. Therefore, the expectation that incentive regulation encourages firms to pursue higher efficiency would not be confirmed by Figure 2.2. This is among the conclusions of a study by Uri (2002), who finds that average technical efficiency has not increased between the 1988-1990 time period and 1991-1999, i.e. before and after the introduction of incentive regulation across the U.S.<sup>19</sup> However, the effects of incentive regulation cannot be assessed on the basis of simple measures, such as those reported in the Figure above or in Uri (2002). Firstly, no general conclusions can be drawn as to the correlation between the introduction of incentive regulation and technical efficiency without a proper statistical study. This type of analysis has been carried out in other papers, such as Majumdar (1997), who finds a marginally positive effect of incentive regulation on technical efficiency. Secondly, the impact of incentive regulation may be lagged, as found again by Majumdar (1997), and therefore not be reflected in efficiency measures in the first years of the sample.

## 2.5 Concluding Remarks

In this chapter, we have estimated an input distance function for a panel of U.S. telecommunications operators over 14 years to analyze their technical efficiency. Compared with previous studies on the telecommunications industry, the new element in the analysis is the application of panel estimators that account for (freely) time-varying inefficiency.

The application of alternative estimators in our sample suggests that results are

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<sup>19</sup>Uri (2002) argues that the lack of change in technical efficiency can be explained by observing that, over the 1988-1999 period, total factor productivity increased, but this was due to innovation related to significant investments in switching and transmission equipment rather than increases in efficiency.

sensitive to the specific estimator used. We find large differences in efficiency scores and rankings, depending on whether the inefficiency component is assumed to be time-invariant or varying over the sample period. Average estimated efficiency levels are very high for the fixed effects and random effects models with time-varying inefficiency, while they are much lower for the time-invariant panel estimators, as expected. These two approaches rely on very different assumptions on the nature of inefficiency. In one case, all inefficiency is assumed to be time invariant and any other component that varies over time is absorbed by the symmetric error. In the other, the opposite view is taken and the model assumes all inefficiency is time-varying, therefore leaving any element which does not change over time (e.g. inefficient management practices) in the firm-specific effect. Variation in efficiency estimates observed over the sample period confirms the importance of panel estimators which account for such variation over time, especially over long periods. However, the rather large discrepancies between results provided by different estimators also indicate that further analysis is needed as to the underlying sources of inefficiency. Ideally, a richer dataset could help uncover more observed heterogeneity between firms and therefore focus better on the inefficiency component.

Finally, the analysis relies on an input distance function and this approach suffers from some general limitations. Firstly, normalized inputs appearing as regressors in an input distance function may not be exogenous. This is a debated issue (Kumbhakar and Lovell, 2000) which has been addressed, for instance, by instrumenting potentially endogenous variables using GMM techniques (Atkinson and Primont, 2002). Moreover, the study focuses on technical efficiency, which only provides a partial view of the firms' behavior as technical efficiency does not imply that, given

input prices, a firm chooses the input mix that minimizes cost. In order to reach conclusions on cost efficiency, the study would need to be complemented by a stochastic cost function approach.

## **Part II**

# **Institutions and Infrastructure**

## **Development: Evidence from**

### **Cross-Country Data**

## Chapter 3

# Regulatory Institutions and Mobile Penetration in Low and Middle-Income Countries

### 3.1 Introduction<sup>1</sup>

In developing countries, liberalization, restructuring, privatization and the introduction of independent regulatory agencies for infrastructure industries appear to have generally been successful in improving sector performance in terms of higher investment and service availability, particularly in telecommunications. However, the specific policies and factors behind both reform successes and failure are the subject of intense debate.

The relationship between the existence of an independent regulator and the development of infrastructure industry investment and productivity levels has been a

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<sup>1</sup>Parts of this Chapter have been published in Maiorano and Stern (2007) on a different sample of countries. This concerns in particular, Sections 3.1, 3.2 and parts of 3.3. The sample used in the published paper is the same as in Section 3.7.3, but with a different model specification.

particular focus of debate and this specific issue has been investigated in numerous papers, with reference both to the telecommunications and electricity sectors (Gutierrez, 2003; Cubbin and Stern, 2006). These studies focus on the characteristics of regulatory institutions that tend to be associated with higher levels of certain performance indicators in public utilities, such as the electricity produced or the number of telephone lines per inhabitant. The independence of the regulator is generally a major explanatory variable, following the literature on central bank independence (e.g. review in Stern and Trillas, 2003). However, this literature does not, with some exceptions, pay much explicit attention to the institutional setting within which the new regulatory agencies operate.

This chapter takes a fresh look at the relationship between regulation and performance in the telecommunications sector, by investigating the issue of independence in the *mobile* communications sector, rather than the fixed network which is the focus of previous papers on regulatory institutions in this industry. In particular, it can be reasonably assumed that regulatory institutions have a different impact in markets in which there are competing firms, rather than a single state-owned operator. Given that mobile telephony is usually characterized by a certain degree of competition almost from its commercial launch, the role of regulation may be different from that exercised in the fixed market, where the development of fixed telephony often takes place for a long time in the absence of competition and the infrastructure is deployed entirely by a state-owned monopolist.

In addition, we draw lessons from strands of the economic literature that are sometimes neglected in previous studies and, in particular, attending to the role of the institutional setting. In consequence, we also take account of further potential



interrelated effects, in particular (a) the economic impact of telecommunications infrastructure on aggregate income; and (b) the role of country institutions in promoting economic growth and the quality of regulatory institutions.

Concerning the relationship between telecommunications penetration and national income, income is considered one of the most important determinants of demand for telecommunications services. But, in addition, investment in telecommunications infrastructure can contribute to economic growth directly by an increase in production and, indirectly, by facilitating communications between firms, thus increasing their production possibilities (Röller and Waverman, 2001).

We explicitly include this feedback effect in the present analysis to provide a fuller picture of the interrelationship between income and telecommunications infrastructure capacity. We do this by focusing on the case of mobile telephony where recent research has suggested an impact of the rapid expansion of mobile telephone subscribers<sup>2</sup> on GDP levels and growth rates in middle and low income countries (Waverman et al., 2005).

Another related issue that is considered in this chapter is the role of country institutions. When investigating the impact of regulation on telecommunications development, it is crucial to ensure that this effect does not capture other factors which are not explicitly included in the analysis. More specifically, our study tries to separate the impact of regulation from the potential indirect effects due to country institutions.

The present study attempts to bring together these questions into a unified framework of analysis. We do this by estimating a system of equations for a panel

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<sup>2</sup>Following the International Telecommunications Union (ITU), the definition of subscribers throughout the chapter includes both pre-paid and post-paid users of telephone mobile services.

of 93 low and middle-income countries over the 1995 - 2004 period.

In summary, the evidence we present confirms the positive effect of regulatory institutions on telecommunications penetration. In particular, the presence of a separate regulator is associated to higher mobile penetration in low-income countries and the estimated effects is higher than for medium-income countries. Country institutions are also found to have a positive impact on GDP per capita and regulatory quality, even if those results hold only for some of the proxies used for country institutions. However, we do not find evidence of the positive effect of mobile penetration on GDP per capita, once the latter is treated as endogenous in a fixed effects model.<sup>3</sup>

The structure of the chapter is as follows. Section 3.2 reviews the most relevant results from the related literature. Section 3.3 sets out the approach adopted in the paper; Section 3.4 provides a description of the data; Section 3.5 sets out the empirical strategy; Section 3.6 discusses the main results; Section 3.7 presents results from alternative specifications of the model; and Section 3.8 provides some short concluding comments and a summary table of the key results.

## 3.2 Related Literature

The standard perspective on utility industries is that the existence of very long-lived and sunk assets gives rise to a time inconsistency problem, which is similar to that which affects monetary policy as described by Barro and Gordon (1983) (Levine et al., 2005). In the telecommunications industry, if the public authority cannot commit to future price levels credibly, that is to refrain from lowering prices beyond

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<sup>3</sup>Waverman et al. (2005) rely on a similar setting, but do not include fixed effects in their equation for mobile penetration.

the originally declared targets, the operator will anticipate the authority's incentive to appropriate its return on sunk investment. As a result, the operator may choose a lower than optimal level of investment. The establishment of an independent regulator is seen as a way of addressing this commitment problem and of safeguarding consumers at the same time, mainly because it should be better insulated from political pressure and therefore less inclined to pursue policy objectives through arbitrary intervention in the regulated sector.<sup>4</sup>

In consequence, there is a growing body of empirical evidence looking at the relationship between the presence of an independent regulator and investment in the telecommunications network, mirroring the extensive literature testing the impact of independent central banks on inflation and growth. The most relevant papers for the present work are discussed in this section.

Gutierrez (2003) is the recent contribution which is the most closely related to our study. Using a panel of 22 Latin American countries over the period 1980 -1997, he finds that good regulatory governance has a positive impact on fixed lines' deployment and efficiency (measured as employees per main lines). In Gutierrez (2003), the main explanatory variables are privatization, competition and regulatory development, where the latter is represented by an index covering, for instance, whether there is separation of telecom operations and regulatory activities and whether the creation of the regulator is backed by law or by a minor legal norm.

The main advantage of Gutierrez's index is the attempt to characterize regulatory governance in a more comprehensive way than allowed by a simple dummy variable for the presence of the regulator, thus recognizing that the mere existence

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<sup>4</sup>For instance, see Levine and Rickman (2002) for the theoretical underpinnings in a model of price regulation under asymmetric information.

of an independent regulator is not by itself informative of the quality of institutions. Moreover, he addresses the potential endogeneity of the regulatory variable in his dynamic model. However, his analysis is based on a reduced-form equation that neglects the potential interactions among network deployment, income level and regulatory governance. In addition, the results from his dynamic model should be treated with caution, as explained in Section 3.7.2 below.

Opposite results, obtained with a different dataset and model specifications, are presented by Estache et al. (2006). They study a panel of 204 developed and developing countries during the period 1990 - 2003, in a model that also incorporates measures of country governance such as investment risk and corruption. For the purposes of this chapter, the key results of the paper are that the presence of a separate regulator does not affect fixed penetration, while the majority privatization of the incumbent does have a positive effect.

Regarding the endogeneity of reforms, in a recent paper Gual and Trillas (2006) investigate the determinants of reforms concerning regulators' independence and entry barriers in the telecommunications sector. In particular, they define independence by an index covering the regulator's functions, its funding, the years since establishment and the percentage of private ownership, among other factors. Regulatory independence is regressed on explanatory variables which include proxies of country institutions used in the growth literature, such as the legal origin of the country, the general quality of government and the rule of law.<sup>5</sup>

Gual and Trillas find that the rule of law variable has a significant negative im-

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<sup>5</sup>Those variables relating to the wider institutional environment have been used as the explanatory variable of interest by Henisz and Zelner (2001), who focus on institutions at the macro-political level. They create an index of political constraints for 147 countries over the period 1960-1994 and they find that it has a positive impact on telecommunications infrastructure development.

fact on independence. They explain this result arguing that "independence is a substitute for other ways to achieve commitment not to expropriate".<sup>6</sup> This interpretation is consistent with the view of the independent regulator as an answer to the commitment problem, as summarized above. However, other researchers (Cubbin and Stern, 2006) find that the rule of law is a complement to better quality regulation rather than a substitute.

An alternative approach is to look at specific policy outcomes rather than actual network development. In Edwards and Waverman (2006), for instance, interconnection rates charged by incumbent operators are the dependent variable which is explained by an index of regulatory governance and other controls in a panel data context. This approach has the appeal of narrowing the focus to an outcome which is more related to institutional quality than measures of performance, and of simplifying the empirical methodology.

However, the Edwards-Waverman approach does not seem suited for the research question of the present study. Firstly, there are no available time-series data on policy outcomes for a sufficiently large set of low and middle-income countries. In addition, as a more general point, this type of approach would implicitly assume that a given policy outcome automatically leads to a higher degree of development of telecommunications networks. The more relevant question for this study is precisely to explain the development of telecommunications, rather than assuming that it would follow from the "right" type of policies. For instance, in the case of interconnection rates, the positive outcome would traditionally be considered to be a low level of charges in order to promote service competition, as in Edwards and Waver-

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<sup>6</sup>Interestingly, other measures of institutions, such as the Polcon index developed by Henisz and Zelner (2001), a measure of procedural complexity (i.e. the number of steps a new firm has to take to operate) and a proxy for government effectiveness were not significant.

man (2006). However, interconnection rates which were too low (i.e. below cost) would give distorted price signals, and therefore may encourage entry by inefficient competitors while acting as a disincentive for investment by the incumbent.

Most papers on regulatory institutions usually analyze the direct impact of regulatory governance on outcomes, while other types of institutions are not considered, with a few notable exceptions.<sup>7</sup> Cubbin and Stern (2006) try to estimate the impact of country institutions on outcomes in the electricity sector and find that there is no significant statistical evidence of the impact of country governance in models that incorporate country fixed effects, once regulatory governance is controlled for. For the telecommunications industry, Estache et al. (2006) and Gasmi et al. (2006) include proxies for country institutions in their models. Estache et al. (2006) surprisingly find that measures both of corruption and investment risk are associated to higher fixed penetration, but the interactions between these proxies and regulatory reform policies (presence of a regulator and privatization of the incumbent) have a negative sign. The authors interpret this result by arguing that "even though corruption may lead to some performance improvements in the presence of red tape and resistance to change, reform policies can lead to stronger and better performance outputs in a much more ethical way." Gasmi et al. (2006) study a sample of 29 developing countries over the period 1985 - 1999, and they find in a dynamic model that corruption has a negative effect on mobile penetration while checks and balances in the political system have a positive coefficient. The methodology proposed in this chapter will address the relationship between country governance and mobile penetration in the context of a system of simultaneous equations, assuming that the channels

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<sup>7</sup>The paper by Henisz and Zelner (2001) focuses on the effect of country institutions on telecommunications investment. Esfahani and Ramirez (2003) study the impact of executive constraints on growth and on the growth of telephones per capita.

through which country governance can influence penetration are the level of income and regulatory governance.

Finally, all these studies rely on formal measures of institutional quality which, in developing countries, may not be indicative of the effective degree of regulatory governance (Pande and Udry, 2005). Attempts to provide de facto measures of independence, i.e. taking account of how regulators and governments actually operate in practice, have been introduced in the literature on central banks independence by Cukierman (1994) and Haan and Koi (2000) and are currently being developed for utilities (Montoya and Trillas, 2007).

### **3.3 Main Issues and Methodology**

The focus of this study is the relationship between measures of telecommunications development and regulatory governance, while taking explicit account within a system framework of (a) the role of income and (b) country governance. In this section we outline a framework to address these issues. We firstly describe the scope of the study; and, secondly, briefly explain the main theoretical references and the approach followed.

The present work studies the penetration of telecommunications infrastructure, as measured by the number of mobile telephone subscribers per head. Among the factors that may affect penetration, we consider the effect that income may have on the uptake of mobile telephone subscriptions, as could be expected based on standard demand models. However, in the case of developing countries this is also important to investigate in order to understand whether penetration is demand-constrained or supply-constrained.

Another consideration relates to the possible feedback effects between penetration and income. Network infrastructure services, including telecommunications, play a crucial role for the economy (Jensen, 2007).<sup>8</sup> There is considerable evidence that higher telecommunications penetration - fixed and mobile - can have significant effects on income.<sup>9</sup> This potential feedback effect is represented by the two-sided arrow in the top row of Figure 3.1. In fact, the methodology of this study is adapted from the literature on the economic impact of infrastructure, as described below.

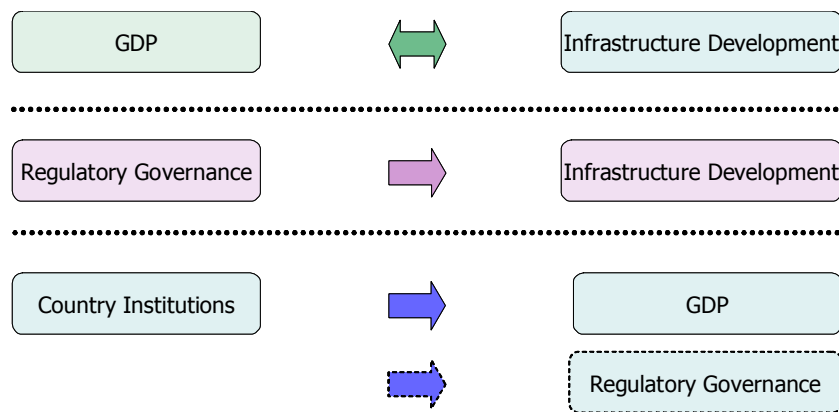


Figure 3.1: Factors potentially affecting infrastructure development

Considering the second row of Figure 3.1, the economic importance of the telecommunications industry has been among the factors contributing to the active role of governments in this sector. The reform process that has taken place in developed countries and in many low and middle-income countries aims at achieving public interest targets by complex policy changes, in which the establishment of a regulatory framework is accompanied by sector restructuring, the liberalization of the market and the privatization of the incumbent. In short, introducing private finance

<sup>8</sup>Canning (1999) and Canning and Bennathan (2000).

<sup>9</sup>Correa (2006) assesses the contribution of investment in telecommunications to productivity using input-output analysis. Röller and Waverman (2001) and Waverman et al. (2005) estimate system of equations to address the endogeneity of infrastructure. Esfahani and Ramirez (2003) develop a framework for deriving reduced form equations to deal with the endogeneity of infrastructure.



and privatizing expanding telecommunications industries has been the main force behind the development of new regulatory organizations as well as, arguably, encouraging general improvements in country governance in the areas of commercial law enforcement.

The present study concentrates on the regulatory framework for telecommunications per se and, in particular, on key aspects of regulatory governance (e.g. the establishment of a separate regulator). As described in the literature review, regulatory reform has the stated objective of promoting better infrastructure development, among other targets, by attracting investment and lowering the cost of capital. This effect is symbolized by the arrow in the middle panel of Figure 3.1.<sup>10</sup>

Effective regulatory frameworks need to be adapted to the specific circumstances of each country and, in particular, to their governance characteristics. This is shown in the third part of Figure 3.1. One interpretation of this relationship views the establishment of an independent regulator as a substitute for strong country institutions (e.g. strong property rights and competition authorities, supported by strong and independent commercial courts). However, an alternative and perhaps more plausible view is that countries with strong institutions may be more likely to engage in substantive reform, which will include genuinely independent and high quality sector regulatory agencies.

Finally, as highlighted by widespread evidence of the positive impact of high-

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<sup>10</sup>It may be argued that countries with more widespread telecommunications penetration are more likely to set up regulators and that therefore there may be some feedback effects from infrastructure development to regulatory governance. In fact, this apparent feedback may instead be related to other factors, such as liberalization or privatization, affecting both regulatory governance and infrastructure development. Moreover, in the type of countries considered in the present study, the reform of the telecommunications sector often takes place under the influence of international lending institutions.

For this reason, no causal relationship is assumed from mobile penetration to the establishment of a separate regulator.

quality country governance and institutions on GDP growth rates (Acemoglu et al., 2005), we need to include the potential direct effect of country institutions on income into the analysis. The last panel of Figure 3.1 summarizes that, in this setting, country governance is assumed to affect mobile penetration indirectly through the channels of GDP per capita and regulatory governance.

### **3.3.1 Analytical Framework**

In this section, we present the main methodological reference for the study. In general, papers on regulatory institutions do not rely on a formal theoretical analysis, but are based on informal considerations on the factors that may affect the development of infrastructure. A more formal approach could be derived from growth models and, given the emphasis of the present study on the endogeneity of per capita income and regulatory institutions, specifically the literature on the impact of infrastructure on growth. A similar approach is provided by the paper by Esfahani and Ramirez (2003), who base their estimation on a growth accounting framework in which they include different types of capital.

The question of endogeneity is addressed by the analytical framework in Esfahani and Ramirez (2003), which incorporates two infrastructure sectors, i.e. fixed telecommunications and electricity generation. The paper aims at addressing the issue of simultaneity between infrastructure and aggregate output by developing a system approach and incorporates, in the empirical specification of the model, country characteristics and policies which are assumed to affect the interrelation between infrastructure and output. The main focus in Esfahani and Ramirez (2003) is to specify both the adjustment to the equilibrium path and the steady-state as functions of country characteristics. However, unlike the present study, the model is

specified so that the equation for the infrastructure sectors can be estimated in a reduced form.

The basis of the Esfahani and Ramirez (2003) paper is an augmented Cobb-Douglas production function, in which infrastructure capital is included among the inputs and constant returns to scale have been imposed.

$$\ln y = \alpha \ln k + \beta \ln n + (1 - \alpha - \beta) \ln Q \quad (3.1)$$

where  $y$  is aggregate output ( $Y$ ) divided by labor ( $L$ ),  $k$  represents non-infrastructure capital ( $K$ ) over labor ( $L$ ),  $n$  infrastructure capital ( $N$ ) divided by labor ( $L$ ) and  $Q$  indicates all other factors that affect productivity.  $Q$  and labor  $L$  are treated as exogenous in this model. Expressing the production function in growth form, one obtains

$$\gamma_y = (1 - \alpha - \beta)q + \alpha\gamma_k + \beta\gamma_n \quad (3.2)$$

where  $\gamma_i$  indicates the growth rate of per-capita variable  $i$  and  $q$  the growth rate of  $Q$ . The accumulation equations for capital and infrastructure are

$$\gamma_k = s_k y/k - \delta - l \quad (3.3)$$

$$\gamma_n = s_n y/n - \delta - l$$

where  $s_i$  is the share of output devoted to the accumulation of input  $i = k, n$ ,  $\delta$  is a constant depreciation rate and  $l$  is the growth rate of labor. The authors note that such rates of accumulation "generally depend on institutional and policy

factors as well as preference and production opportunities in the economy.<sup>11</sup> In the steady state, the endogenous per-capita variables grow at the same rate, which is equal to the long-run productivity growth ( $q^*$ ).<sup>11</sup> From this, the steady state ratios between infrastructure and output ( $n^*/y^*$ ) and between capital and output ( $k^*/y^*$ ) are given by  $s_i^*/(q^* + \delta + l)$ , with  $i = k, n$ . In this expression,  $s_i^*$  represents the steady-state rate of accumulation of asset  $i = k, n$ .

The difference between actual accumulation in sector  $i$  ( $\gamma_i$ ) and the steady-state rate of growth ( $\gamma_i^*$ ) is derived from Equation 3.3 as

$$\gamma_i - q^* = s_i y/i - (q^* + \delta + l) \quad (3.4)$$

This difference can be expressed as a function of the gap between the initial and the steady-state ratio of asset  $i$  to output. Firstly,  $(q^* + \delta + l)$  can be substituted with  $s_i^* y^*/i^*$  in Equation 3.4. Then, the right-hand side can be rewritten as  $(q^* + \delta + l) \left[ \frac{s_i y/i}{s_i^* y^*/i^*} - 1 \right]$ . Collecting  $\frac{s_i}{s_i^*}$  the equation can be transformed as follows:

$$\begin{aligned} \gamma_i - q^* &= (q^* + \delta + l) \left( \frac{s_i}{s_i^*} \right) \left[ \frac{y/i}{y^*/i^*} - \frac{s_i^*}{s_i} \right] \\ &= (q^* + \delta + l) \left( \frac{s_i}{s_i^*} \right) \left[ \frac{i^*/y^*}{i/y} - 1 + \frac{s_i}{s_i} - \frac{s_i^*}{s_i} \right] \\ &= (q^* + \delta + l) \left( \frac{s_i}{s_i^*} \right) \left[ G_i + \frac{s_i - s_i^*}{s_i} \right] \end{aligned} \quad (3.5)$$

where  $G_i$  is the gap between the initial and the steady state ratio between asset  $i$  and output. Esfahani and Ramirez (2003) focus on the case of fixed investment rates ( $s_i = s_i^*$ ). They note that, even if  $s_i/s_i^*$  can be approximated as one,  $s_i$  may

<sup>11</sup>This is obtained from Equation (3.2) when all growth rates are set equal. In addition, it is assumed that the long-run productivity growth is constant across countries.

still deviate from  $s_i^*$ .

While neoclassical models consider this a second-order effect in steady-state, Esfahani and Ramirez (2003) argue that for an infrastructure sector, deviations of rates of accumulation in the short-run are important. They argue that asset imbalances leading, for instance, to service suspension provide strong incentives to increase the rate of accumulation in the short-run. On this basis, they conclude that  $\frac{s_i - s_i^*}{s_i}$  may be large compared with  $G_i$  and therefore cannot be considered negligible. This consideration implies that, in Equation 3.5,  $\frac{s_i}{s_i^*} \approx 1$  while the term in the square brackets remains. Moreover, it is assumed that  $\frac{s_i - s_i^*}{s_i}$  can be approximated as  $\frac{s_i - s_i^*}{s_i} = g_i(X)G_i$  for  $i = k, n$ . The function  $g_i(X)$  incorporates the effect of a vector of variables  $X$ , which can also include measures of the effectiveness of institutions and other country characteristics. When substituting in Equation 3.5 one obtains (for  $s_i \approx s_i^*$ )

$$\gamma_i = q^* + (q^* + \delta + l)[1 + g_i(X)]G_i \quad (3.6)$$

Therefore, in the neighbourhood of the steady-state, the adjustment rate to asset imbalances ( $G_i$ ) for sector  $i = k, n$  is given by  $(q^* + \delta + l)[1 + g_i(X)]$  and is a function of country characteristics  $X$ . From the definition of  $G_i$  and the approximation  $G_i \approx \log(1 + G_i)$  for small  $G_i$ ,

$$G_i = \log\left(1 + \frac{i^*/y^*}{i/y} - 1\right) = \log\left(\frac{i^*}{y^*}\right) - \log\left(\frac{i}{y}\right) = \log\left(\frac{s_i^*}{q^* + \delta + l}\right) - \log\left(\frac{i}{y}\right) \quad (3.7)$$

The asset gap can be substituted in Equation 3.6 to give

$$\gamma_i = q^* + (q^* + \delta + l)[1 + g_i(X)] \left[ \log(s_i^*) - \log(q^* + \delta + l) - \log\left(\frac{i}{y}\right) \right] \quad (3.8)$$

In Equation 3.8 above, the growth rate in asset  $i = k, n$  is expressed as the sum between (a) the long-run rate of growth ( $q^*$ , productivity growth) and (b) the product between adjustment rate (which depends on various country characteristics) and the infrastructure gap  $G_i$ . In turn, the infrastructure gap depends on the initial ratio between the asset and output and the factors that determine the long-run investment rate in the asset ( $s_i^*$ ). In the empirical formulation of the model, the function  $g_i(X)$  is replaced by a linear function of variables  $X$ . As can be seen from Equation 3.8, Esfahani and Ramirez (2003) derive the growth rates for asset sectors so that they can be estimated in a reduced form.

The growth rates thus obtained for the asset sectors  $i = k, n$  can be substituted into Equation 3.2 to give the growth rate of per capita output. The equation in Esfahani and Ramirez (2003) is simplified so as to incorporate only one infrastructure sector  $n$ .

$$\gamma_y = \beta\gamma_n + (1 - \beta)q^* + (1 - \alpha)(q - q^*) + (q^* + \delta + l)[1 + g_k(X)]\alpha G_k \quad (3.9)$$

In Expression 3.2 above,  $G_k$  is the analogous of Equation 3.7 and is a function of  $s_k^*$ , which is assumed to depend on country characteristics as was the case for  $s_i^*$  in Equation 3.8. Moreover, the function  $g_k(X)$  is assumed linear in the empirical part of the paper.

The Esfahani and Ramirez (2003) estimation relies on Equations 3.8 and 3.2. In

particular, the former is a reduced-form equation and its fitted values are inserted in the latter in place of  $\gamma_n$ . In order to make the equations operational, the authors proxy the unobservable variables  $s_i^*$  and  $s_k^*$  with the overall investment GDP ratio and some country-level variables which can affect long-run levels. Regarding the other steady-state variable,  $q^*$ , this is set equal to the average growth rate of productivity across countries and time.

Building on Esfahani and Ramirez (2003), we adapt their model to the framework commonly used in studies on the effectiveness of sector regulators. Firstly, following Röller and Waverman (2001) and Waverman et al. (2005), the analysis focuses on the equilibrium path of the variables rather than on their growth rates. Secondly, the estimation approach differs in that we estimate the equations jointly, while Esfahani and Ramirez (2003) do not account for the potential feedback effect from income growth to infrastructure development, as explained above. Thirdly, the panel dataset used by Esfahani and Ramirez (2003), in line with many growth studies, have averaged data over five-year periods, while studies on infrastructure sectors usually rely on annual data. Finally, given that allowing for endogenous regulatory quality is among the objectives of the analysis, we also consider an additional equation in which the dependent variable is the presence of a separate regulator (or other institutional characteristics of the telecommunications sector) as explained below.

### **3.3.2 Summary of the Approach**

In order to deal with the interactions described above and represented in Figure 3.1, a system of simultaneous equations is estimated in which the dependent variables are infrastructure development, per capita income and regulatory governance. This approach assumes that these variables are endogenous.

As will be explained in more detail in Section 3.5 below, the basic econometric specification consists of three equations, which have been derived from theory or from previous empirical studies. In the first equation, the penetration of telecommunications infrastructure is explained by income, regulatory governance, investment in telecommunications and other variables. The second equation relates income levels to the penetration of telecommunications, a measure of country institutions and other variables. Finally, in the third equation, regulatory governance is explained by income, country institutions and other variables.

Jointly estimating the system of equations presents the advantage of improving the efficiency of the estimates, compared to the results obtained by instrumental variables estimators on each equation. However, with systems estimation, if the structure of the model is misspecified any modelling error in any one equation will be propagated through the system. In consequence, we first estimate the equations individually in order to focus on their specification more carefully.

The limitations of the analysis are mainly related to the measurement of the governance variables. Firstly, in common with other studies (but see a new dataset in Montoya and Trillas, 2007), regulatory governance is measured on the basis of formal characteristics of the legal framework, such as the existence of the regulator and the way it is funded. However, this may not coincide with the actual governance of the regulatory authority i.e. how the regulator operates and is allowed by the government to operate - in practice.

Secondly, related to the previous point, in this chapter the only available measure of regulatory governance for all our countries is a dichotomous variable which takes value one when a certain characteristic is present (e.g. regulator separate from



Ministry, autonomous funding) and zero otherwise. This type of variable does not allow us to quantify differences between countries' regulators in any detail. However, compared to an index-type variable,<sup>12</sup> it is more suitable for system estimation.

Thirdly, country institutions are among the explanatory variables in the system. There is an open question of the potential endogeneity of country institutions. This is a key and hotly debated theme in the empirical literature on institutions and growth (Durlauf et al., 2005). In the present study, the issue is addressed by treating the proxies for country institutions as predetermined for the year in question.<sup>13</sup> This approach is motivated by institutions' strong persistence over time, especially in relation to the limited timeframe of the present sample.

Unlike previous studies on regulatory governance in telecommunications, this chapter focuses on mobile communications, in order better to tailor the analysis for low and middle-income countries.<sup>14</sup> Given the substantial sunk investments to deploy the fixed network and the chronic waiting lists, mobile phones have proved formidable substitutes for fixed lines in developing countries.

In terms of methodology, the main advantage of this approach is that we estimate a system of equations rather than a single reduced form equation that is informed by the underlying economic relationships. This should allow to investigate much more thoroughly the interactions described above, which perforce are either ignored or only implicitly modelled in the single equation reduced form model. For instance,

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<sup>12</sup>For example, as explained in Section 3.2, Gutierrez (2003) measures regulatory quality by an index which includes six different components, such as whether the regulator is separate from the Ministry and whether it is independently funded. A similar approach is also followed by Cubbin and Stern (2006). In both studies, the index of regulatory quality is only an explanatory variable, rather than a dependent variable as in the present paper. Such indexes typically can take only a discrete number of values and, therefore, cannot be treated as continuous variables in the estimation. For this reason, for estimation in a system it is simpler to have a dummy variable rather than the ordered data that would result from an index.

<sup>13</sup>For instance, see Rajan and Zingales (1998).

<sup>14</sup>However, a recent paper by Gasmi et al. (2006) analyses a sample of 29 developing countries over the period 1985 - 1999 using a single-equation model.

our approach can shed light on the economic factors that determine the quality of regulatory governance and, potentially, on mobile telecoms penetration. In particular, it allows testing whether and how far country general institutions are a driver of mobile penetration through their indirect effects on infrastructure regulation and on income levels.

Secondly, a key difference compared with previous papers is that they do not consider the effect of telecommunications infrastructure on income. The failure to treat income as endogenous can lead to inconsistency in a reduced-form equation. The approach proposed in the paper should provide more reliable results by explicitly allowing for income to be endogenous.

Thirdly, the paper also relates to studies measuring the impact of telecommunications penetration on income. In this respect, this paper's contribution is the explicit inclusion of regulatory governance and country institutions in the framework of analysis, as well as the treatment of regulation as endogenous.<sup>15</sup>

Finally, the present dataset includes a reasonably large set of developing countries only (93 countries). Hence, we have a more homogenous group of countries than in most previous studies. The latter have generally combined both developed and developing countries and therefore implicitly assume that a common model holds for very different countries (e.g. Wallsten (2003) and Waverman et al. (2005) for the cross-section results).

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<sup>15</sup>Waverman et al. (2005) include a rule of law measure, while Esfahani and Ramirez (2003) include a dummy for private ownership.

### 3.4 Description of the Sample

Our dataset consists of an unbalanced panel of yearly data on 93 low-income and middle-income countries over the period 1995 to 2004. The main sources for the data are the International Telecommunications Union (ITU) World Telecommunications Indicators and the World Development Indicators from the World Bank. This section describes the main variables, while details on the other variables included in the analysis are provided in Appendix.

Telecommunications penetration is measured by the number of mobile subscribers per 100 inhabitants, as explained in Appendix 1. In line with other studies (Gutierrez, 2003; Röller and Waverman, 2001), GDP per capita is measured in constant U.S. dollars.

For regulatory governance, a limited number of indicators have been chosen. These include: whether (a) the country has passed a framework law for the telecommunications sector; (b) the country has established a regulator as a separate entity from the policy maker;<sup>16</sup> and (c) the regulator is not funded by the Government's budget. In addition, the years since the creation of the regulator are also considered in order to capture the time necessary to build up staff numbers and competencies and reputation, as in Cubbin and Stern (2006). The indicator for a separate regulator we construct is conceptually different from the definition of autonomy in the International Telecommunications Union (ITU)'s database, which is based on responses from country regulators and therefore on a subjective assessment that the regulatory agency is independent of the executive and of the industry. We believe

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<sup>16</sup>The year in which the law establishing the regulator was passed may differ from the year when the regulator was actually set up. In most countries in our sample, they coincide. For Belize, it was not possible to identify the law setting up the regulator.

that the establishment of a regulatory body is a more objective measure and can be verified independently, as was done to construct our dataset. Data sources for these regulatory variables include the International Telecommunications Union (ITU) online database on country and regulators profiles,<sup>17</sup> Henisz, Zelner and Guillén (2004), Wallsten et al. (2004)<sup>18</sup> and the regulators' websites.

The variable for privatization is an indicator which takes value one when the fixed incumbent has been privatized and zero otherwise. Privatization is defined here as the sale of more than 50% of the incumbent's shares by the government. Similarly, the liberalization dummy takes value one if competition for long-distance services in the fixed market is permitted. Regarding privatization, the data collected by Henisz, Zelner and Guillén (2004) for the period up to 1999 were updated using the World Bank Privatization Database and other publications.<sup>19</sup> The liberalization variable was also drawn from Henisz, Zelner and Guillén (2004) and was updated using case studies from a variety of sources.

The countries considered in the analysis are very diverse, as shown in the summary statistics in Table 3.1. Even though all the countries in the sample are characterized by the World Bank as low and middle-income, the level of GDP per capita in constant dollars varies from 156.30 USD to more than 9,000 USD. Similarly, if we only restrict our attention to the last year in the sample, the number of mobile subscribers per 100 inhabitants (mobile penetration) ranges from 1 in Niger to 105 in the Czech Republic in 2004.

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<sup>17</sup><http://www.itu.int/ITU-D/treg/profiles/guide.asp?lang=en>

<sup>18</sup><http://www.aei-brookings.org/publications/abstract.php?pid=724> Scott Wallsten, George Clarke, Luke Haggarty, Rosario Kaneshiro, Roger Noll, Mary Shirley, and Lixin Colin Xu. "New Tools for Studying Network Industry Reforms in Developing Countries: The Telecommunications and Electricity Regulation Database." AEI-Brookings Joint Center Related Publication 04-05. March 2004.

<sup>19</sup><http://rru.worldbank.org/Privatization/>

Table 3.1: Summary Statistics

<b>Variable</b>	<b>Obs.</b>	<b>Mean</b>	<b>Std. Dev.</b>	<b>Min</b>	<b>Max</b>
Mobile penetration (subs. per 100 pop)	921	9.03	15.33	0	105.64
Fixed penetration (subs. per 100 pop)	914	10.79	10.92	0.07	50.48
GDP per capita (US\$ 2000)	926	1953.47	1933.19	156.30	9650.72
Share of private credit (%GDP)	925	29.65	27.89	1.38	165.72
Openness (%GDP)	924	82.12	39.20	12.80	228.87

Source: ITU, World Bank

In Table 3.1, and in all the estimated equations, monetary variables are included on an exchange rate basis, in constant 2000 US dollars, rather than in PPP terms. This approach follows, for instance, Röller and Waverman (2001), Cubbin and Stern (2006), Gutierrez (2003), Estache et al. (2006) and the panel data analysis in Waverman et al. (2005).

When looking at the behavior of the variables across time, it appears, not surprisingly, that mobile penetration and GDP per capita, as well as capital and labor per capita, show an upward trend. Conversely, other variables, such as the index of political constraints (Polcon), show little variation over time. As explained further in the next section, we explicitly take into account in the estimation the various considerations on the dynamic behavior of the variables, especially mobile penetration and GDP per capita.

The variables regarding sector reform are summarized in Figure 3.2, which shows higher mobile penetration in countries that have implemented different types of reforms (telecommunications law, separate regulator, liberalization of long-distance services and majority privatization of the incumbent) compared to the others. At first glance, this would suggest the reform may have had a positive impact on penetration. However, the differences seem more significant for liberalization and privatization than for the establishment of a separate regulator.

It is interesting to note that the countries that have reformed the telecommuni-

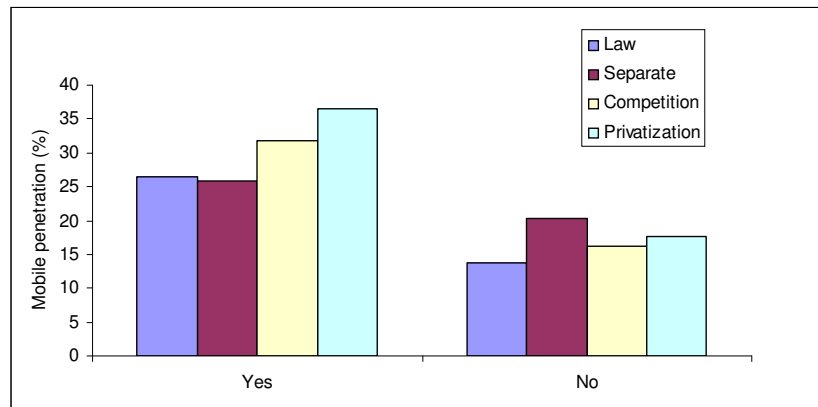


Figure 3.2: Mobile penetration, in countries that have and have not implemented reform, 2004. Source: ITU, regulators' websites.

cations sector by 2004 almost coincide with the entire sample. Out of 93 countries, 84 have enacted a law reforming the telecommunications sector and most of them had done so by year 2000. In theory this could reduce the degree of inter-country variation needed to identify the effect of regulatory reform; in practice, the time dimension from the very different dates at which countries introduced their reforms could provide the necessary variation to discriminate between countries. However, this source of identification is limited in this sample by the presence of countries that have carried out the reform process at the very beginning or at the end of our sample period. For instance, 54 countries in total have either set up a regulator in 1995-6 or in 2004.

In order to analyze differences across countries in the timing of reform, the age of the regulator is a good proxy. It also provides an indication of the authority's staff expertise and reputation, which may be important in addressing the commitment problem described in the literature section. Figure 3.3 shows a histogram of the age of the regulator as of 2004, the last year in the sample. The variable indicates the number of years since the establishment of the regulator, starting with the year after

the law introducing a regulator was passed.<sup>20</sup> The average age is almost 5.5 years and the median is 5, that is around half the regulators have more than 5 years of operation.

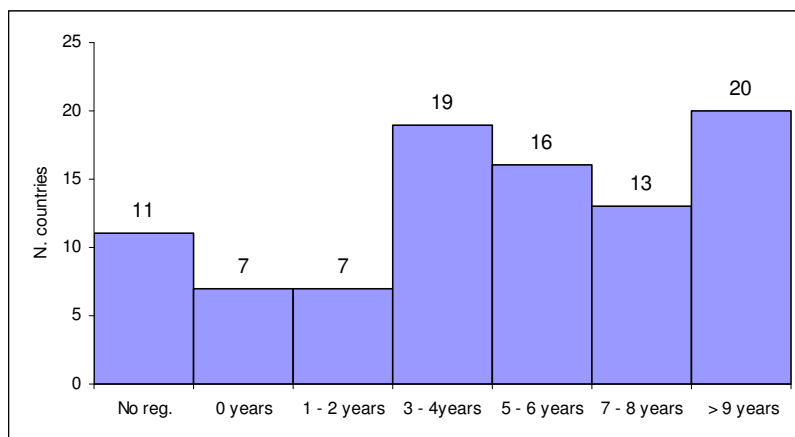


Figure 3.3: Distribution of countries by age of regulator, 2004. Source: ITU, regulators' websites.

The correlation between the different elements of regulatory reform is shown in Table 3.2. However, the correlation between enacting a modern telecom law and having a separate regulator is high, but weaker than expected. This is due to a number of countries that do not have a separate regulator even though they have passed a sector law and vice versa (around one quarter of the sample as of 2004). In addition, around one quarter of the countries that have privatized the fixed incumbent operator have also introduced some liberalization measures in the long-distance fixed telecommunications market, even though the two events have not necessarily been contemporaneous. In consequence, the degree of correlation between the regulatory variables that we find for telecommunications is quite high but lower than Cubbin and Stern (2006) found in their sample for electricity.

<sup>20</sup>For instance, the first group includes countries that do not have a separate regulator or that passed the law in 2004. The second group includes countries that passed the law establishing the regulator in 2001 - 2002, i.e. age of the regulator is between 1 and 2 years, and so on.

Table 3.2: Correlation matrix

	Law	Separate Reg.	Funding	Competition	Privatization
Telecom Law	1				
Separate Regulator	0.5823	1			
Independent Funding	0.4631	0.6966	1		
Competition	0.3028	0.2160	0.0716	1	
Privatization	0.2232	0.2409	0.0203	0.3550	1

The relationship between GDP per capita and the rate of mobile penetration is crucial to this study and is plotted in Figure 4. The two variables exhibit a positive correlation, in line with expectations. In addition, the wide variation which was highlighted in the summary statistics is also clearly visible in the graph. Finally, the positive relationship is consistent both with (a) GDP driving mobile subscription, a demand-side effect; and (b) mobile availability and usage increasing GDP, a supply-side effect.

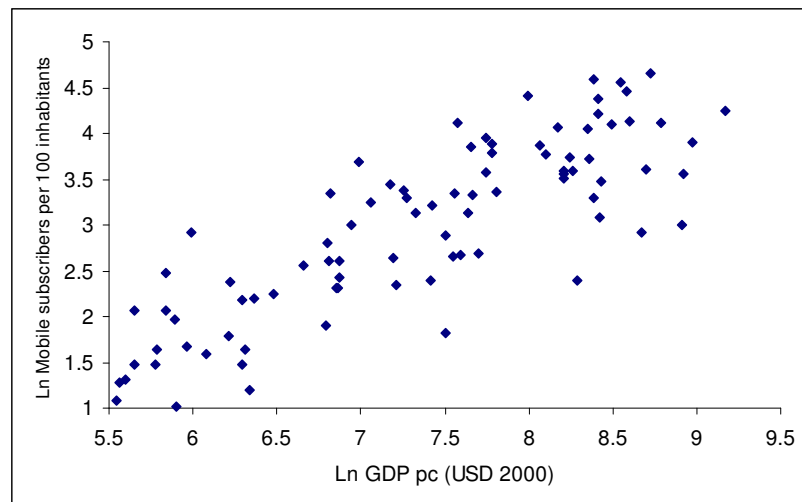


Figure 3.4: Correlation between mobile penetration and per capita GDP, 2004. Source: ITU, World Bank.

### 3.5 Econometric Methodology

On the basis of the discussion above, the variables that are considered endogenous in the present analysis are (i) the penetration rate of mobile telecommunications, (ii)



GDP per capita and (iii) a measure of regulatory governance.<sup>21</sup> Hence, we formulate a three equation model, which consists of an equation describing the behavior of telecommunications penetration and two further equations for the other potentially endogenous variables.

The model that we estimate is set out below. In the basic formulation of the system, it is assumed that the three endogenous variables have an impact only on the contemporaneous values of other endogenous variables i.e. we assume no lagged effects. Other restrictions, derived from the previous discussion, are also imposed in the following equations.

The estimation relies on a panel of countries over time from 1995 to 2004, and we assume that the parameters of the model are constant both across countries and over time. This assumption is relaxed in Section 3.7.1.

The penetration of mobile subscribers ( $PEN_{it}$ ) in country  $i$  at time  $t$  is assumed to be a function of the other potentially endogenous variables in the system, per capita income ( $GDPpc_{it}$ ) and regulatory governance ( $RG_{it}$ ) and of some exogenous variables. This gives us the following equation for mobile penetration rates:

$$\begin{aligned} \ln PEN_{it} = & \alpha_{0,i} + \alpha_1 \ln GDPpc_{it} + \alpha_2 \cdot RG_{it} + \alpha_3 Privat_{it} + \alpha_4 Liberal_{it} & (I) \\ & + \alpha_5 (RG_{it} \cdot Privat_{it}) + \alpha_6 (RG_{it} \cdot Liberal_{it}) + \alpha_7 (Privat_{it} \cdot Liberal_{it}) \\ & + \alpha_8 X_{it} + \mu_t + \varepsilon_{1,it} \end{aligned}$$

where  $X_{it}$  is a vector of controls,<sup>22</sup>  $Privat_{it}$  is a dummy denoting majority priva-

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<sup>21</sup>The measures of regulatory governance used in the present study are described in the Data Section above.

<sup>22</sup>These included various price measures, both for fixed and mobile services, the penetration of fixed services, the percentage of rural population, population density and the size of the waiting list for fixed lines.

tization, and  $Liberal_{it}$  is a dummy indicating competition in the long-distance fixed market. In addition, the interaction terms between the variables that represent policy reforms allow investigating the effects of one component of reform conditional on the others.  $\alpha_{0,i}$  are country-specific effects,  $\mu_t$  are time dummies and  $\varepsilon_{1,it}$  is an error term for this first equation.

In Equation (I), for the reasons discussed in the previous section, the estimated coefficients on  $\ln GDPpc_{it}$  and  $RG_{it}$  are expected to be positive. The effects of fixed line privatization and liberalization are indeterminate. They may have a negative effect on mobile penetration, as it may be sensible to assume that those reforms would improve availability and quality of fixed services. However, there are also reasons why the impact could be positive e.g. if one of the mobile operators is part of a newly commercialized incumbent fixed line operator.

The equations describing the other two potentially endogenous variables are:

$$\begin{aligned} \ln GDPpc_{it} = & \beta_{0,i} + \beta_1 Inst_{i,t-1} + \beta_2 \ln PEN_{it} + \beta_3 \ln K_{it} + & \text{(II)} \\ & \beta_4 \ln HK_{it} + \beta_5 Z_{it} + \eta_t + \varepsilon_{2,it} \end{aligned}$$

$$\begin{aligned} RG_{it} = & \gamma_{0,i} + \gamma_1 Inst_{i,t-1} + \gamma_2 \ln GDPpc_{it} + \gamma_3 \ln IntlEffect_{it} & \text{(III)} \\ & + \gamma_4 Privat_{it} + \gamma_5 Liberal_{it} + \nu_t + \varepsilon_{3,it} \end{aligned}$$

The equation for GDP per head takes the form of an aggregate production function, in which mobile telecommunications are included as an input, following Röllner and Waverman (2001). In this equation,  $Inst_{i,t-1}$  is a proxy for country institutions,  $HK_{it}$  is a proxy for human capital measured per capita,  $K_{it}$  is a measure of

per capita physical capital,  $Z_{it}$  is a vector of controls,  $\beta_{0,i}$  are country-specific effects,  $\eta_t$  are time dummies and  $\varepsilon_{2,it}$  is an error term. In the model that we estimate, it is assumed that country institutions are pre-determined. For this reason, we include in the system  $Inst_{i,t-1}$ , which pre-date the period of analysis, rather than  $Inst_{i,t}$  on the grounds that institutions in previous years cannot be affected by income levels in subsequent periods (see Rajan and Zingales, 1998; Esfahani and Ramirez, 2003 for previous use of this approach). All variables in Equation (II) are expected to have a positive effect on income levels.

In Equation (III) regulation is modelled as a dummy variable, which takes the value of 1 if a given characteristic (e.g. sector law or separate regulator) is present and zero otherwise. As explained in the previous section, different measures of 'good regulation' will be employed in the estimation. We treat regulation as an endogenous variable in this model, and it is assumed that the choice of whether to have a regulatory framework in place depends on country institutions and some other factors. The other factors that we consider include the country income level ( $GDPpc_{it}$ ) and pressure by international organizations (e.g. conditionality conditions imposed by international financial institutions), as proxied by multilateral lending ( $IntlEffect_{it}$ ).  $\varepsilon_{3,it}$  is the error term for Equation (III).

The functional form of the third equation is a linear probability model. While this model does not constrain predicted values to lie in the interval between 0 and 1 (Greene, 2003), it is particularly suitable for estimation in a multi-equation system. In order to alleviate the potential problem of out-of-range estimates, we check predicted probabilities after estimation to verify that they belong within the correct interval.<sup>23</sup>

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<sup>23</sup>However, we test whether a logit model would produce similar results. These are reported in

As we have a panel data set, we initially include country-specific fixed effects in all three equations. In most of the analysis, the fixed effects model was considered the basis for the estimates, as is very often the case with models like Equation (I) and (II). The fixed effects model is less restrictive in its underlying assumptions than the random effects model in that it allows for correlation between the regressors and the country-specific unobserved effect. Country-specific heterogeneity, including omitted variables, is captured by the fixed effects. However, this removal of heterogeneity across countries may also be a limitation, as highlighted by Pritchett (2000) and Durlauf et al. (2005). In the context of the analysis of income differences, some important variables (e.g. country governance) affecting income show very little variation over time within a given country and, given that the fixed effects estimator only uses "within variation", the effects of these variables cannot be identified.

Related to this general point, in our sample the policy variables are constant throughout the sample for the majority of countries. For this reason, their coefficients cannot be identified separately from the country-specific fixed effect in Equation (I). Therefore we cannot restrict the estimation to a fixed effects model and we report results from the pooled OLS estimator, the random effects model and the fixed effects model, where the latter is estimated on the subsample of countries for which "separate regulator", the main policy variable of interest, shows some variation.

For the other equations, the within estimator is the preferred methodology. We estimate the equations jointly by 3SLS in order to allow for the possibility that errors are correlated, for instance due to unobserved shocks affecting both income and mobile penetration.

## 3.6 Results

This section is divided into two main parts. In the first part, we report results for each equation estimated separately and, in the second part, we show the system estimates.

The results obtained estimating each equation separately are reported in order to provide a better insight on the data and also to explore the specification of the equations before the joint estimation of the system.

### 3.6.1 Equation-by-Equation Estimates: Equation I

The results for the penetration of mobile services (Equation I) are shown in Table 3.3. The specifications shown provide estimates from pooled OLS (Column 1), random effects (Column 2) and fixed effects models (Column 3). All columns include time dummies to take account of any common cross-country time period effects. The reported standard errors are robust to heteroscedasticity and serial correlation.<sup>24</sup>

In addition, the pooled OLS model includes continent dummies. The fixed effects model is estimated for a subset of countries for which  $RG_{it}$  exhibits some within-country variation.<sup>25</sup>

Evidence regarding the two main explanatory variables,  $\ln GDPpc$  and  $RG$ , is mixed. In the pooled OLS and the random effects models, GDP per capita has a positive significant coefficient, while it is not significant in the fixed effects model.

This seems to indicate that the heterogeneity across countries, as represented by different income levels, is absorbed by the fixed effects, while within country variation

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<sup>24</sup>This adjustment follows the result in Wooldridge (2002). It allows both for heteroscedasticity and for correlation across observations for the same country.

<sup>25</sup>The subsample excludes the countries for which the dummy "separate regulator" does not change across time or only one period differs from the others (i.e. the regulator was established in the second year of the sample or in the last one).

Table 3.3: Mobile Penetration - Basic Specification

	POLS (1)	RE (2)	FE (3)
Dependent variable: Independent Variables	$\ln PEN$		
$\ln GDPpc_{it}$	1.121*** (0.105)	1.274*** (0.093)	0.042 (1.774)
$RG_{it}$	0.175 (0.191)	0.055 (0.161)	-0.415* (0.208)
$Privat_{it}$	0.228 (0.298)	0.415* (0.223)	0.172 (0.286)
$Liberal_{it}$	1.047** (0.456)	0.718 (0.480)	-0.033 (0.703)
$RG_{it} \cdot Privat_{it}$	0.359 (0.349)	-0.160 (0.268)	-0.519* (0.285)
$RG_{it} \cdot Liberal_{it}$	-1.052** (0.449)	-0.604 (0.445)	0.056 (0.589)
$Privat_{it} \cdot Liberal_{it}$	-0.451 (0.375)	-0.495* (0.285)	-0.629 (0.391)
$\ln PriceMobile_{it}$	0.069 (0.069)	0.098* (0.056)	0.044 (0.095)
$\ln Faults_{it}$	-0.063** (0.032)	-0.022 (0.025)	-0.015 (0.047)
Constant	-10.298*** (0.851)	-10.846*** (0.825)	-2.923 (11.809)
Time Dummies	Yes	Yes	Yes
Corrected std. errors	Yes	Yes	Yes
R-squared	0.856	0.840	0.893
Countries	69	69	38
Observations	315	315	168

Note: Standard errors in brackets under coefficients. Corrected std. errors are robust to heteroscedasticity and serial correlation. The model in Column 1 also includes continent dummies. In Column 3, the sample includes only countries for which RG varies across time.

provided by such different income levels plays a more limited role, as indicated by an insignificant coefficient in the fixed effects model.

Most of the coefficients on the policy variables are not significant and the same holds for the interactions between them. Moreover, when interactions are significant they have a negative coefficient. For instance, in Column 1 the negative coefficient on  $RG_{it} \cdot Liberal_{it}$  indicates that, conditional on the liberalization of the fixed market, the establishment of a separate regulator is associated on average with lower mobile penetration. The insignificant or negative coefficients on the interaction terms question the commonly held assumption that there are additional benefits in a reform process encompassing the establishment of a regulator, privatization and liberalization.

Other controls in Equation (I) were generally not significant. In addition to the

regressors shown in the table, the significance of various price indicators for fixed and mobile services was investigated, but these variables were not found to be significant except for the price of a 3-minute peak mobile call in the random effects model (Column 2).<sup>26</sup> This result contrasts with what was found by Röller and Waverman (2001) and Waverman et al. (2005). However, in these papers prices are defined as average revenues (i.e. revenues divided by the number of subscribers) rather than average market prices faced by consumers. A negative relationship between average revenue per user and market penetration is expected, as mobile operators target less lucrative customer groups as the market matures. This phenomenon is common to mobile markets in developed countries, e.g. in Europe, where operators try to counterbalance ARPU (average revenue per user) declines from traditional voice services with high-value services such as mobile broadband. Given this, the logarithms of average revenues were used instead of prices in the estimation of Equation I. Both fixed and mobile revenue per subscribers were found significant, as in the above cited papers. Despite their significance, these variables were not included in the chosen specification due to their limited availability, which would have reduced sample size significantly, and because of endogeneity concerns.

The waiting list for a connection to the fixed network and the number of reported faults were also considered among the regressors, as they may both encourage potential users to subscribe to the mobile network rather than the fixed network. However, they were not significant except for the number of faults per 100 fixed lines in the pooled OLS model (Column 1).

Similarly, population density and the percentage of rural population were in-

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<sup>26</sup>Logarithm of the average price of a 3-minute mobile call (peak time) in US\$ 2000. Details on the other variables are provided in Appendix.

cluded to incorporate country characteristics that may affect the cost of coverage for mobile operators. For instance, the impact of rural population may incorporate two opposite considerations: on the demand side, a high percentage of rural population may increase penetration due to the appeal of mobile telephony in areas not easily reached by the fixed network; on the supply side, rural areas are less densely populated than cities and therefore more expensive to cover for mobile operators. Both variables were not significant.

As explained in Section 3.5, the fixed effects model does not allow exploiting the "between" variation. Given that, in the present sample the policy variables do not vary over time for most of the countries, we cannot rely on the within estimator. For this reason, in what follows we focus on estimates from a pooled OLS model which incorporates dummies for the different continents (Table 3.4, Columns 1-3) and a random effects model (Table 3.4, Columns 4-6).

Table 3.4 shows the coefficients in Equation (I) re-estimated omitting insignificant variables (Columns 1 and 4).<sup>27</sup> Given that robust standard errors are not available for the system estimation, we also check whether this adjustment affects the significance of the estimates (Columns 2 and 5). Moreover, in order to investigate the effect of treating  $\ln GDPpc$  as an endogenous variable in a single-equation context, Equation (I) was re-estimated using two-stage least squares, where the instruments were given by the right-hand side variables in Equation (II) (in addition to all the exogenous variables in Equation I). These instrumental variable estimates provide an intermediate step between the results in Columns 2 and 5, and the full system estimates. In Table 3.4, the instruments for the endogenous variables are

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<sup>27</sup>Policy variables are still included even when they are not significant, while faults in the fixed network and the price of a mobile call are removed because their inclusion reduces sample size by more than 550 observations, i.e. sample size increases by almost 65% when they are excluded.



Table 3.4: Mobile Penetration - Significant Variables

Independent Variables	POLS (1)	POLS (2)	POLS - IV (3)	RE (4)	RE (5)	RE - IV (6)
$\ln GDPpc_{it}$	1.100*** (0.085)	1.100*** (0.036)	1.160*** (0.047)	1.257*** (0.079)	1.257*** (0.067)	1.264*** (0.082)
$RG_{it}$	0.312** (0.148)	0.312*** (0.080)	0.323*** (0.080)	0.110 (0.099)	0.110 (0.080)	0.111 (0.080)
$Privat_{it}$	0.418** (0.159)	0.418*** (0.098)	0.410*** (0.099)	0.395** (0.148)	0.395*** (0.115)	0.396*** (0.116)
$Liberal_{it}$	0.830** (0.289)	0.830*** (0.161)	0.819*** (0.161)	0.431 (0.252)	0.431* (0.194)	0.439* (0.194)
$RG_{it} \cdot Liberal_{it}$	-0.888** (0.247)	-0.888*** (0.173)	-0.896*** (0.174)	-0.498* (0.252)	-0.498* (0.196)	-0.506* (0.197)
$Privat_{it} \cdot Liberal_{it}$	-0.210 (0.284)	-0.210 (0.160)	-0.209 (0.162)	-0.621** (0.192)	-0.621*** (0.160)	-0.617*** (0.160)
Constant	-10.427*** (0.565)	-10.427*** (0.256)	-11.798*** (0.322)	-11.143*** (0.577)	-11.143*** (0.481)	-11.200*** (0.580)
Time Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Corrected std. errors	Yes	No	No	Yes	No	No
R-squared	0.853	0.853	0.851	0.834	0.834	0.833
Countries	93	93	93	93	93	93
Observations	881	881	876	881	881	876

Note: Standard errors in brackets under coefficients. Corrected std. errors are robust to heteroscedasticity and serial correlation. In Columns 1 to 3, continent dummies are included in the model. In Columns 3 and 6,  $\ln GDPpc$  is treated as endogenous variable and the instruments are the regressors in Equation 2, in addition to all the exogenous variables in Equation 1.

only the exogenous variables while, in the system (Table 3.8), the regressors include also the endogenous variables in order to take account of feedback effects.

Columns 1 - 3 provide results for the estimation of Equation (I) using pooled OLS, after the removal of the insignificant policy interactions, of the price of mobile peak calls and of the share of faults out of fixed lines. While the latter variable was significant using the pooled OLS estimator (Table 3.3, Column 1), it was available only for a limited number of countries and therefore reduced the sample size by more than 20 countries and 550 observations. For this reason, and to improve comparability with the random effects estimates, it was removed from the equation.

The main difference between Column 1 and Table 3.3 (Column 1) is given by the significant coefficient on the regulatory governance and of the majority privatization dummy. Moreover, Column 2 shows that the adjustment to the standard errors of the coefficients does not affect much their significance. Finally, the coefficient on  $\ln GDPpc_{it}$  increases when it is treated as endogenous in Column 3. The same

Table 3.5: Marginal Effects on Ln Mobile Penetration

	POLS			RE		
	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$
Coefficient	0.131	0.260	0.365	0.002	0.007	0.264
Standard error	0.072	0.092	0.086	0.078	0.109	0.104
p-value	0.071	0.005	0.000	0.974	0.948	0.011

Note: Marginal effects calculated from the IV estimates (Columns 3 and 6).

specification was also estimated using the random effects model (Columns 4 - 6). Unlike the OLS model, the  $RG_{it}$  dummy remains not significant even after removing the adjustment to the standard errors (Column 5). As in Column 3, when the equation is estimated by 2SLS the coefficient on  $\ln GDPpc_{it}$  in the random effects model increases (Column 6).

In order to understand better the impact of the policy variables, marginal effects are computed.<sup>28</sup> Marginal effects are computed from Columns 3 (pooled OLS with continent dummies) and 6 (RE) to provide a better comparison with the system estimates. Marginal effects are found to be significantly different from zero in the pooled OLS estimator, while only privatization is significant when the model is estimated by random effects (Table 3.5).

As explained in Section 3.4, other proxies for regulatory governance were considered. The marginal effects of the telecommunications law, of the age of the regulator and of the independent funding of the regulator (i.e. autonomous from the government budget<sup>29</sup>) were insignificant in all the specifications shown in Table 3.4.

Finally, as a check, additional specifications of Equation (I) were estimated: (a) a flexible formulation in which the coefficients were allowed to vary for low-income and middle-income countries; (b) a dynamic equation; and (c) a longer time period

<sup>28</sup>The marginal effect was calculated by taking the derivative of Equation (I) with respect to the variable of interest and evaluating the derivative at the sample mean. We then tested the null hypothesis that the derivative was zero.

<sup>29</sup>This is a commonly used measure of independence. However, it should be noted that independent funding implies that the regulator relies on the industry for its funding. In small markets, this may not be synonymous with independence.

Table 3.6: Income Equation

Independent Variables	Polcon			Share of credit		
	POLS (1)	RE (2)	FE (3)	POLS (4)	RE (5)	FE (6)
$Inst_{i,t-1}$	-0.121 (0.185)	-0.002 (0.033)	-0.013 (0.033)	0.189 (0.052)***	0.057 (0.020)***	0.053 (0.020)***
$\ln PEN_{it}$	0.336 (0.102)***	0.028 (0.008)***	0.022 (0.008)***	0.261 (0.088)***	0.028 (0.007)***	0.021 (0.007)***
$\ln K_{it}$	0.418 (0.158)***	0.400 (0.058)***	0.293 (0.038)***	0.442 (0.152)***	0.420 (0.059)***	0.272 (0.060)***
$\ln HK_{it}$	-0.821 (0.256)***	-0.226 (0.195)	-0.198 (0.203)	-0.700 (0.253)***	-0.260 (0.183)	-0.267 (0.191)
$\ln Openness_{it}$	-0.158 (0.076)**	0.104 (0.030)***	0.099 (0.030)***	-0.163 (0.078)**	0.091 (0.028)***	0.091 (0.029)***
Constant	4.164 (1.354)***	3.139 (0.446)***	4.120 (0.355)***	3.353 (1.195)***	2.859 (0.434)***	4.108 (0.492)***
Time Dummies	Yes	Yes	Yes	Yes	Yes	Yes
Corrected std. errors	Yes	Yes	Yes	Yes	Yes	Yes
R-squared	0.80	0.70	0.57	0.82	0.74	0.57
Countries	92	92	92	93	93	93
Observations	797	797	797	802	802	802

Note: Standard errors in brackets under coefficients. Corrected std. errors are robust to heteroscedasticity and serial correlation.

(15 years) for which data was available for a subset of countries. The results are reported in Section 3.7 below.

### 3.6.2 Equation-by-Equation Estimates: Equation II

The results for the GDP per capita production function are shown in Table 3.6. We estimate this equation taking as proxy measures of the quality of country governance both (a) the Polcon index described in the data section and (b) the logarithm of the share of credit to the private sector to GDP. All explanatory variables are considered exogenous and year dummies are included. Results from pooled OLS, random effects and fixed effects estimators are provided, and standard errors are robust to heteroscedasticity and serial correlation.

The coefficient of the impact of mobile penetration on GDP per capita is positive and significant in all specifications, as expected and as found in Waverman et al. (2005). Its size decreases significantly when panel data estimators, rather than OLS, are used. Compared to Waverman et al. (2005), who obtain an estimate of 0.075 in their panel data model, the magnitude of our estimate (0.021 - 0.028) in the panel

models is significantly smaller.<sup>30</sup>

The variables used as proxies for country governance produce mixed results. When Polcon is used as a proxy for country institutions it is never significant, while the development of financial markets is significant at the 1% level in all specifications. The significance of the credit variable is in line with the literature on financial markets and growth (see Levine (2005) for a review). Regarding the impact of Polcon, it should be noted that is a measure of constraints on the executive power in a given country. In the study, it was assumed that this was also an indirect measure of the constraints to the expropriation of private property by the executive power. However, this assumption is probably not very robust and Polcon does not show any correlation with income levels.

As in the previous section, the model was estimated without adjusting the standard errors in order to provide comparability with the results from the system approach. This was carried out using the fixed effects estimator which, for Equation (II), does not suffer from the lack of variability in a regressor ( $RG_{it}$ ) described above for the penetration equation. When the standard errors are not adjusted, all variables are significant at the 1% level.

Moreover, as a check on the importance of country institutions, the Kaufmann et al. (2006) indexes of rule of law and quality of governance were also considered. These variables are only available for even years from 1996 to 2004 and, for this reason, Equation (II) was estimated for those 5 years rather than for the whole 10-year period. For both indexes, the coefficient on institutions in a fixed effects model

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<sup>30</sup>The results are not directly comparable, as Waverman et al. (2005) do not include country-specific effects. In addition, they transform penetration as  $PEN / (35 - PEN)$ . Finally, Waverman et al. (2005) prefer the estimates from the cross-section model, which implies that if, in a given country, there were one additional mobile phone for 100 people the country would experience a per capita GDP growth higher by 0.059 percent.

was significant at the 1% level and was comparable to the coefficient on institutions in Columns 5-6, Table 3.6. While this is a very crude check, it confirms the significant effect of country institutions on income levels.

Finally, a methodological issue arises from the choice of the time horizon in the GDP equation. The theoretical models that inform empirical studies generally address the question of steady-state growth. As a result, empirical growth studies have traditionally used cross-sections of countries, where data for each country were averaged over long periods of time. However, the estimation of growth models relies increasingly on panels where data have been averaged over five or ten years. This raises the question of whether it is appropriate to use long-run models to interpret short intervals of data (Levine, 2005; Durlauf, Johnson and Temple, 2005) and this potential pitfall seems even more acute if annual data are used. This is an open question in the empirical growth literature and needs to be weighed against the advantages of using panels, possibly in dynamic specifications that help distinguish short-run from long-run effects, rather than country cross-sections.

### **3.6.3 Equation-by-Equation Estimates: Equation III**

The estimates of Equation (III) for the presence or absence of a telecommunications regulator are reported in Table 3.7. They provide no evidence of any systematic link between country institutions and regulatory governance. This is counter to the results in Gual and Trillas (2006). However, we confirm their finding that Polcon (which we interpret as an index of the checks and balances in the political system of a given country) is not significant in a model explaining a measure of regulatory quality.

The specifications shown in Table 3.7 (Equation III) compare estimates from

Table 3.7: Regulatory Governance Equation

Independent Variables	Polcon		Share of credit	
	POLS (1)	RE (2)	POLS (3)	RE (4)
$\ln GDP_{pc_{it}}$	-0.052 (0.040)	-0.014 (0.035)	-0.068 (0.043)	-0.028 (0.040)
$Inst_{i,t-1}$	0.253 (0.179)	0.081 (0.139)	0.035 (0.046)	0.036 (0.047)
$Privat_{it}$	0.159* (0.090)	0.188** (0.091)	0.157* (0.088)	0.192** (0.091)
$Liberal_{it}$	0.180* (0.104)	0.267*** (0.080)	0.169 (0.104)	0.259** (0.082)
$Privat_{it} \cdot Liberal_{it}$	-0.242 (0.152)	-0.247** (0.107)	-0.206 (0.159)	-0.235** (0.108)
$\ln International_{it}$	-0.008 (0.026)	0.002 (0.025)	-0.011 (0.027)	0.001 (0.026)
Constant	0.567 (0.304)	0.304 (0.249)	0.654* (0.302)	0.326 (0.252)
Time Dummies	Yes	Yes	Yes	Yes
Corrected std. errors	Yes	Yes	Yes	Yes
R-squared	0.27	0.19	0.26	0.19
Countries	91	91	91	91
Observations	808	808	807	807

Note: Standard errors in brackets under coefficients. Corrected std. errors are robust to heteroscedasticity and serial correlation. Columns (1) and (3) also include dummies for continents.

a pooled OLS model, where dummies for continents are also included, and a random effects model. As explained above, because of the lack of variability in the dummy indicating regulatory governance, we did not estimate a fixed effects model. Moreover, all columns include time dummies and standard errors are robust to heteroscedasticity and serial correlation. After estimation, fitted values were computed and it was verified that they belonged to the interval between 0 and 1.<sup>31</sup>

As with Polcon (Columns 1 - 2), the share of private credit over GDP (Columns 3 - 4) does not have any statistically significant impact on the establishment of a separate regulator.<sup>32</sup>

The equation is relatively unsuccessful in explaining the presence of a telecommu-

<sup>31</sup>For the pooled OLS model in Column 3, this occurs for 95% of the predicted values. In addition, fitted values are compared with the actual values of 'separate'. When there is a separate regulator (i.e. separate = 1), the model correctly predicts the existence of a separate regulator in about 71% of observations (i.e. the fitted values are between a threshold value, which we set at 0.5, and 1). When there is not a separate regulator (i.e. separate = 0), the fitted values are between 0 and the threshold value of 0.5 for 68% of the observations.

<sup>32</sup>As for Equation II, the Kaufman et al. indexes of rule of law and quality of governance were also considered as proxies for country institutions. Their significance was checked in a sample consisting of even years only, due to limitations to the variables' availability. Even in a fixed effects model, both had positive and significant coefficients at the 10% level, which contradicts the results obtained with the other two indicators.

nications regulator.<sup>33</sup> The only significant variables, once time dummies are included in the model, are the dummies indicating the liberalization of long-distance fixed services and the majority privatization of the incumbent. The positive coefficients on other policy reforms are expected and are consistent with the fact that telecom regulators are frequently established as part of a package with liberalization and privatization.

As a further check on the significance of the regressors, a panel logit model including the same variables as Table 3.7 was estimated. It confirmed that the only significant variables were the policy variables and their interaction. The estimates from the logit model are reported in Appendix.

Finally, given that standard errors robust to heteroscedasticity and serial correlation are not available for the system, Equation III was re-estimated without adjusting the standard errors. This affected greatly the significance of the regressors and all coefficients became significant, except for the proxy for multilateral lending ( $\ln International_{it}$ ). Given these contradictory results on the specification of Equation III, all regressors were included in the model for joint estimation in the system, as reported below.

### 3.6.4 System Estimates

The system of simultaneous equations given by Equations (I) - (III) was estimated by three-stage least squares, where all right-hand variables were considered exogenous,

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<sup>33</sup>The model in Column 1 was also estimated for the other variables used as indicators of regulatory governance. When the dependent variable was the dummy indicating that a telecommunications law was passed in the country, the coefficient on liberalization was significant at the 1% level. The liberalization indicator was also the only significant regressor, at the 10% level, when the dummy for the funding of the regulator was the dependent variable, and it was not significant at all when the model in Column 3 was estimated.

Table 3.8: System Estimates

Dependent Variables:	$\ln PEN_{it}$		$\ln GDPpc_{it}$		$RG_{it}$
	(1)		(2)		(3)
$\ln GDPpc_{it}$	1.076*** (0.036)	$\ln PEN_{it}$	0.006 (0.016)	$\ln GDPpc_{it}$	-0.080*** (0.021)
$RG_{it}$	0.782*** (0.104)	$Inst_{i,t-1}$	0.057*** (0.008)	$Inst_{i,t-1}$	0.054** (0.022)
$Privat_{it}$	0.352*** (0.083)	$\ln HK_{it}$	-0.338** (0.108)	$Privat_{it}$	0.167*** (0.048)
$Liberal_{it}$	0.977*** (0.173)	$\ln K_{it}$	0.258*** (0.035)	$Liberal_{it}$	0.141** (0.054)
$RG_{it} \cdot Liberal_{it}$	-1.266*** (0.204)	$\ln Openness_{it}$	0.086*** (0.016)	$Privat_{it} \cdot Liberal_{it}$	-0.186* (0.080)
				$\ln International_{it}$	-0.009 (0.016)
Constant	-5.939*** (0.273)		-		1.212*** (0.160)
Fixed Effects	Continent		Country		Continent
Time Dummies	Yes		Yes		Yes
R-squared	0.845		0.997		0.260
Countries	92		92		92
Observations	780		780		780

Note: Standard errors in brackets under coefficients. The proxy for institutions is financial markets development, i.e. the share of credit to the private sector on GDP. Columns (1) and (3) include dummies for continents.

except for the interactions between the regulatory dummy and liberalization.<sup>34</sup> In Equation (I), the insignificant interactions between policy variables were left out of the model.

The results are presented in Table 3.8.<sup>35</sup> Time dummies are included in all equations, while fixed effects are only included in Equation II and the other two equations incorporate continent dummies, as in previous sections, to allow the identification of the effect of regulatory governance  $RG_{it}$ .

In the first column, the signs and significance of the coefficients confirm broadly the results obtained for the penetration equation when all variables are considered exogenous (Table 3.4). Moreover, all policy variables have significant marginal effects (Table 3.9) as in the single-equation counterpart of the pooled OLS model (Table

<sup>34</sup>It was verified that each equation satisfied the order condition and that the system passed the rank condition.

<sup>35</sup>In the system results reported in Table 8, around 95% of the predicted values from the regulation equation are inside the 0 - 1 interval. In addition, as for the single equation case, it was checked how many times the model correctly predicted the outcome. When the dummy for a separate regulator was 1, the model predicted the correct values 65% of the times. When separate regulator was equal to 0, it was correctly predicted only for 45% of the observations.



Table 3.9: Marginal Effects

	<i>RG<sub>it</sub></i>	<i>Liberal<sub>it</sub></i>	<i>Privat<sub>it</sub></i>
Coefficient	0.511	0.257	0.351
Standard error	0.090	0.090	0.083
p-value	0.000	0.004	0.000

Note: Marginal effects are calculated from the penetration equation.

(3.5).

Similarly, in the GDP equation in Column 2, most coefficients are in line with the single-equation results (Table 3.6). A major difference is given by the coefficient on mobile penetration in Table 3.8, which is insignificant once the feedback effect between income and mobile penetration is taken into account. This result also contradicts the findings on the positive effect of mobile phones in Waverman et al. (2005) whose model, however, differs from the present specification of the GDP equation in that it does not include country-specific fixed effects. Interestingly, when a longer time period is considered (for a subset of countries) while still relying on a system of three equations, this finding is reversed and the coefficient on mobile penetration becomes significant (Section 3.7). While this discrepancy may be an indication that a longer time period matters to detect such an effect or that the sample in Section 3.7 includes countries in which the mobile market is more developed, it also highlights the difficulties in disentangling causality within the present framework.

Finally, in the last equation to explain the presence or absence of an independent regulator (Column 3), most variables are significant. In particular, the proxy for country governance has a positive effect on the establishment of a separate regulator. The liberalization of fixed services and the privatization of the incumbent are also associated with higher probability that the country establishes a separate regulator. Modelling the development of economic institutions in general is difficult

(e.g. see examples in Acemoglu et al., 2005). In particular, regarding telecommunications regulatory agencies, Gual and Trillas (2006) study this issue using a cross-section of countries, as explained above, focusing more on between country heterogeneity compared to the present setting. This aspect of the analysis needs further consideration.

### **3.7 Alternative Specifications**

In this section, some alternative specifications for Equation I are presented. Firstly, coefficients are allowed to vary for low-income and middle-income countries, exploiting the large number of countries available in the sample. Secondly, a dynamic model for Equation I is estimated. Finally, for a subsample of countries a longer time interval is considered.

#### **3.7.1 Flexible Formulation of Equation 1 by Income Level**

The sample used in the present study comprises a large number of countries which, although they are all classified by the World Bank as low and middle-income countries, show a large variation in income, as described in Section 3.4. For this reason, the sample was split into two subsamples, consisting of low-income and of middle-income countries, and interactions between all variables and a dummy for countries with low GDP per capita were included in Equation I.<sup>36</sup> As before, the equation was first estimated individually in order to better focus on its specification. Only significant variables in the single equation were kept for estimation of the system, while the other two equations were unchanged compared to previous sections. The

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<sup>36</sup>An alternative formulation would have been to interact the variables directly with per capita GDP.

Table 3.10: System Estimates: Interactions with Income

Dependent Variables:	$\ln PEN_{it}$		$\ln GDPpc_{it}$		$RG_{it}$
	(1)		(2)		(3)
$\ln GDPpc_{it}$	1.078*** (0.039)	$\ln PEN_{it}$	-0.018 (0.011)	$\ln GDPpc_{it}$	-0.079*** (0.021)
$RG_{it}$	0.695*** (0.106)	$Inst_{i,t-1}$	0.061*** (0.008)	$Inst_{i,t-1}$	0.051* (0.022)
$Privat_{it}$	0.265** (0.128)	$\ln HK_{it}$	-0.420*** (0.099)	$Privat_{it}$	0.166*** (0.048)
$Liberal_{it}$	1.262*** (0.201)	$\ln K_{it}$	0.238*** (0.033)	$Liberal_{it}$	0.143** (0.055)
$RG_{it} \cdot Liberal_{it}$	-1.396*** (0.207)	$\ln Openness_{it}$	0.083*** (0.017)	$Privat_{it} \cdot Liberal_{it}$	-0.185* (0.081)
$Privat_{it} \cdot Liberal_{it}$	-0.305 (0.215)			$\ln International_{it}$	-0.010 (0.016)
$lowPrivat_{it}$	0.143 (0.177)				
$lowLiberal_{it}$	-0.283 (0.193)				
$lowPrivat_{it} \cdot Liberal_{it}$	1.254*** (0.356)				
Constant	-		-		-
Fixed Effects	Continent		Country		Continent
Time Dummies	Yes		Yes		Yes
R-squared	0.851		0.997		0.260
Countries	92		92		92
Observations	780		780		780

Note: Standard errors in brackets under coefficients. The proxy for institutions is financial markets development, i.e. the share of credit to the private sector on GDP.

results for the system are presented in Table 3.10.

Most coefficients in Columns 2 and 3 are similar to those obtained in Table 3.8. In the penetration equation, the coefficients on the policy variables vary compared to the basic formulation due to the inclusion of interactions with the low income dummy. Regarding the other variables, the results from this alternative specification confirm the positive effect of GDP per capita on mobile penetration. In addition, as in Subsection 6.4, we find no evidence that higher mobile penetration is associated to higher GDP per capita, once the feedback effect of GDP on penetration is taken into account.

The marginal effects of the policy variables are reported in Table 3.11. Concerning regulation, this more flexible specification highlights differences between low-income and middle-income countries. The coefficient on the separate regulator is higher for low-income countries which suggests that the role of a separate regu-

Table 3.11: Marginal Effects by Income Level

	All Sample			Low Income			Middle Income		
	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$
Coefficient	0.396	0.396	0.200	0.487	0.299	0.549	0.311	0.387	0.181
Standard error	0.090	0.123	0.109	0.093	0.125	0.119	0.089	0.117	0.107
p-value	0.000	0.001	0.067	0.000	0.017	0.000	0.001	0.001	0.089

lator is more important for less-developed markets, where a credible framework for interconnection, coverage and quality of service is crucial to ensure service availability. In more developed markets, other factors may be relevant in explaining the penetration of mobile services. In particular, reforms in the fixed telecommunications sector suggest a more competitive and developed fixed market, which may act as a competitive constraint on mobile operators. Moreover, due to data availability, this model does not incorporate competition in the mobile market itself, which may be a major factor in explaining penetration.

### 3.7.2 Dynamics in Mobile Penetration

The specification proposed so far is static while, similarly to the models estimated in the growth literature, it may be argued that the penetration of mobile phones in the current period is affected by penetration in previous periods. For instance, Gutierrez (2003) and Cubbin and Stern (2006) find evidence that the lags of the dependent variable have significant and large coefficients. This formulation assumes that the impact of past values of the regressors is persistent and is captured by the coefficient on the lagged dependent variable.

For this reason, we specify the following dynamic model, which includes the variables that were found to be significant in the analysis above in addition to the lagged dependent variable.

$$\begin{aligned} \ln PEN_{it} = & \alpha_{0,i} + \delta_1 \ln PEN_{i,t-1} + \alpha_1 \ln GDPpc_{it} + \alpha_2 \cdot RG_{it} + \alpha_3 Privat_{it} \quad (3.10) \\ & + \alpha_4 Liberal_{it} + \alpha_5 (RG_{it} \cdot Liberal_{it}) + \alpha_6 (Privat_{it} \cdot Liberal_{it}) + \mu_t + \varepsilon_{1,it} \end{aligned}$$

When the number of time periods is small ( $T = 10$  in our sample), the standard fixed effects estimator of a dynamic model is biased.<sup>37</sup> This is due to the correlation between the lags of the dependent variable and the error term in the model. For this reason, as in Gutierrez (2003), the Arellano and Bond (1991) estimator is used in order to overcome the bias created by the introduction of lagged values of the dependent variable in Equation 3.10.

We report the results from the Arellano-Bond estimator in Table 3.12, once insignificant coefficients have been removed. As this model cannot be readily incorporated within a linear system, we only report single equation estimates. In the first column, all variables are treated as exogenous except for the lagged dependent variable. In addition,  $\ln GDPpc_{it}$  is considered endogenous in the second column, while in the third  $RG_{it}$  is also endogenous.

Compared with previous estimates obtained from the pooled OLS and random effects models, the coefficients on the policy variables are not significant and the coefficient on GDP per capita, when significant, is smaller. Therefore, the lagged dependent variable appears to incorporate the impact of past values of the right-hand variables, apart from GDP per capita in Column 1. This result raises the issue of a better dynamic specification of the penetration equation. However, once the lagged dependent variable is included, lags of the policy variables are insignificant.

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<sup>37</sup>Nickell (1981) finds that, in an autoregressive model of order 1, if the coefficient on the lag of the dependent variable is 0.5 the bias is -0.167 (assuming 10 time periods).

Table 3.12: Dynamic Specification

Independent Variables	All Exogenous (1)	<i>GDPpc</i> Endog. (2)	<i>GDPpc</i> and <i>RG</i> Endog. (3)
$\ln PEN_{i,t-1}$	0.849*** (0.067)	0.824*** (0.069)	0.815*** (0.067)
$\ln GDPpc_{it}$	0.981* (0.473)	1.034 (0.888)	0.834 (0.776)
<i>RG</i> <sub>it</sub>	-0.065 (0.091)	-0.102 (0.083)	-0.162 (0.119)
<i>Liberal</i> <sub>it</sub>	0.076 (0.134)	0.053 (0.128)	0.092 (0.137)
<i>Privat</i> <sub>it</sub>	0.027 (0.108)	0.032 (0.103)	0.046 (0.107)
<i>RG</i> <sub>it</sub> · <i>Liberal</i> <sub>it</sub>	-0.166 (0.104)	-0.175 (0.098)	-0.141 (0.129)
<i>Privat</i> <sub>it</sub> · <i>Liberal</i> <sub>it</sub>	-0.066 (0.117)	-0.094 (0.121)	-0.146 (0.128)
Constant	0.021 (0.037)	0.098 (0.072)	0.114 (0.071)
Time Dummies	Yes	Yes	Yes
Wald test	328.44	338.08	392.08
2nd order serial correlation	-0.89	-0.85	-0.77
p-value	0.376	0.396	0.439
Countries	93	93	93
Observations	693	693	693

Note: Robust standard errors in brackets under coefficients. The lagged dependent variable is treated as endogenous in all columns. The Arellano-Bond test for 2nd order serial correlation fails to reject the null hypothesis of no serial correlation in the residuals.

As a check, the model was estimated by fixed effects to verify whether the coefficient on lagged penetration in Table 3.12 was in the correct range. The fixed effects estimate (0.612) provides a lower bound for the coefficient and therefore our Arellano-Bond estimates are consistent with this result. Moreover, compared with Gutierrez (2003), who estimates the model for fixed line penetration, the coefficient on the lagged dependent variable is pretty similar (Gutierrez finds 0.7748 and 0.7287 depending on the specification) but the coefficient on GDP per capita is insignificant in this model, once its endogeneity is taken into account.<sup>38</sup> Finally, it should be noted that the Arellano-Bond estimator addresses the endogeneity of the lagged dependent variable under the assumption of a small number of time periods and a number of cross-sections going to infinity.<sup>39</sup> Therefore, the results obtained by

<sup>38</sup> Gutierrez (2003) treats liberalization, privatization and regulation as predetermined variables.

<sup>39</sup> In their article, Arellano and Bond (1991) apply the estimator to an unbalanced panel of 140 companies over at least 7 periods.

Recognizing the importance for applied work of panels where the cross-section and time dimensions are of similar size, Alvarez and Arellano (2003) investigate the properties of the estimator when they both go to infinity. However, their focus is the exploration of panels with larger T rather than with smaller N. In addition, concerns on the small sample properties of the Arellano-Bond

Gutierrez (2003) should be treated with caution, given that the Arellano-Bond estimator has not been proven to be consistent for datasets with a small number of cross-sections, as in Gutierrez (2003).<sup>40</sup>

While the results from the dynamic specification of Equation (I) indicate the importance of taking dynamics into account, there is a question of whether this setting is appropriate to deal with the interrelationships between variables that this study aims to analyse. Restricting the estimation to a single equation would still allow to treat some variables as endogenous, but not to incorporate any feedback effects. Moreover, in the Arellano-Bond estimator the size of the instruments matrix can increase very quickly with the number of endogenous or predetermined variables, leading to poor identification. This risk would need to be carefully weighed when choosing the lag structure for the model and for the instruments.

### **3.7.3 Longer Time Period**

As a further check, the same model in Section 3.5 was re-estimated over a longer period. For this purpose, a subsample of 30 countries was selected out of the initial 93 countries. This sample was constructed by selecting the countries based on the availability of data over a 15-year period (1990 - 2004), rather than the shorter timeframe adopted in the rest of the chapter (1995 - 2004).

Moreover, for these countries, data on investment in telecommunications infrastructure (both fixed and mobile) was available for a sufficient number of years and was included in the model specification. Investment levels, relative to country GDP, give an indication of whether the infrastructure is being developed to keep

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estimator, due to the large size of the instrument matrix, are discussed in Kiviet (1995).

<sup>40</sup>That dataset includes 22 countries over 18 years.

Table 3.13: System Estimates: 15-Year Sample

Dependent Variables:	$\ln PEN_{it}$		$\ln GDPpc_{it}$		$RG_{it}$
	(1)		(2)		(3)
$\ln GDPpc_{it}$	0.991*	$\ln PEN_{it}$	0.027**	$\ln GDPpc_{it}$	0.533*
	(0.500)		(0.009)		(0.244)
$RG_{it}$	1.194***	$Inst_{i,t-1}$	0.038**	$Inst_{i,t-1}$	-0.010
	(0.337)		(0.012)		(0.065)
$Privat_{it}$	0.916***	$\ln HK_{it}$	-0.321**	$Privat_{it}$	0.419***
	(0.217)		(0.101)		(0.064)
$Liberal_{it}$	-0.383**	$\ln K_{it}$	0.576***	$Liberal_{it}$	0.364***
	(0.119)		(0.030)		(0.083)
$RG_{it} \cdot Privat_{it}$	-0.733**	$\ln Openness_{it}$	-0.005	$Privat_{it} \cdot Liberal_{it}$	-0.401***
	(0.317)		(0.023)		(0.094)
$\ln InvRatio_{i,t-1}$	0.317***			$\ln International_{it}$	-0.047
	(0.058)				(0.030)
Constant	-10.358*		2.808***		5.956**
	(4.439)		(0.292)		(2.121)
Fixed Effects	Yes		Yes		Yes
Time Dummies	Yes		Yes		Yes
R-squared	0.944		0.997		0.730
Countries	30		30		30
Observations	343		343		343

Note: Standard errors in brackets under coefficients. The proxy for country institutions is financial markets development, i.e. the share of credit to the private sector on GDP.

up with demand for telecommunications services. As mentioned above, there is a question of whether penetration is demand-constrained or supply-constrained in developing countries. As in Esfahani and Ramirez (2003), we attempt to mitigate the simultaneity between the dependent variable and the investment ratio by replacing the latter with its lagged value.

The penetration equation was firstly estimated individually and only significant variables were then kept in the specification estimated jointly in the system and shown in Table 3.13.

While the significance of most coefficients is in line with the results in the main section of the chapter, there are some striking differences. Firstly, in Column 2 the coefficient on penetration is now significant compared with Table 3.8 and this difference may be related to the longer time required for the impact of mobile phones on GDP to show in the data.<sup>41</sup> Secondly, the proxy for country governance has a lower and less significant coefficient in Column 2 compared to results from the overall

<sup>41</sup>The within estimator is used for Column 2 both for the smaller sample and for the results in the main Section of the chapter. Given that this method relies on within-country variation, i.e. over time, a larger number of periods leads to better defined coefficients.



sample, while it is not significant at all in Column 3. Therefore, in this sample we find evidence of an indirect effect of country conditions on mobile penetration through GDP per capita but not through regulatory governance. Thirdly, the estimates suggest that the likelihood of establishing a separate regulator is higher in countries with higher GDP per capita, in contrast with the overall sample.

The sample used in this subsection differs in many respects from the larger sample on which most of the analysis in this chapter has relied. The countries in the subsample form a selected, and possibly better, set of countries compared to the overall sample of 93 countries for two main reasons. Firstly, these are countries for which a longer time series for data on mobile penetration is available, which means that they are likely to have more developed markets. For instance, competition in the mobile market is not controlled for in this study due to data availability and may be the source of some unobserved heterogeneity among the countries. Secondly, the countries were also selected on the basis of data quality, e.g. availability of information on investment in telecommunications. Again, this seems to indicate a selected group of countries compared to the larger sample used in the rest of the study.

Finally, the marginal effects of the policy variables tend to be larger compared to the estimates in Tables 3.9 and 3.11. A major difference that points again in the direction of a different stage of development for the countries in the subsample is the fact that the liberalization of long-distance services has now a negative coefficient, which seems to indicate some substitutability between fixed and mobile services.

Table 3.14: Marginal Effects

	$RG_{it}$	$Liberal_{it}$	$Privat_{it}$
Coefficient	1.022	-0.383	0.499
Standard error	0.283	0.119	0.159
p-value	0.000	0.001	0.002

Table 3.15: Summary of Main Results

Equation:	Penetration		GDP		Regulatory Governance
	$RG_{it}$	$\ln GDPpc_{it}$	$Inst_{i,t-1}$	$\ln PEN_{it}$	$Inst_{i,t-1}$
1. Basic Formulation					
Coefficient	0.511	1.076	0.057	0.006	0.054
Standard error	0.090	0.036	0.008	0.016	0.022
p-value	0.000	0.000	0.000	0.716	0.016
2. Income Interactions					
Coefficient	0.396	1.078	0.061	-0.117	0.051
Standard error	0.090	0.039	0.008	0.111	0.022
p-value	0.000	0.000	0.000	0.113	0.023
3. Longer Period					
Coefficient	1.022	0.991	0.038	0.027	-0.010
Standard error	0.283	0.550	0.012	0.009	0.065
p-value	0.000	0.047	0.002	0.004	0.878

Basic formulation refers to the results in Section 3.6.4, Income interactions to Section 3.7.1 and Longer Period to Section 3.7.3.

### 3.8 Concluding Remarks

This study analyses the link between GDP, mobile telecommunications penetration and regulatory governance (country and sector-specific) in a sample of low and middle-income countries over a 10-year period. The main new element in the analysis is the estimation of a system of simultaneous equations, in which mobile penetration, income level and a proxy for regulatory governance are all treated as endogenous. As far as we are aware, this is an approach that has not been adopted elsewhere for infrastructure industry models.

The results from the main model formulation are checked for robustness by estimating two alternative specifications. Firstly, the sample is split into low-income and middle-income countries and coefficients are allowed to vary by income level. Secondly, the basic formulation is estimated on a subsample of 30 countries for which data are available over a 15-year period, rather than a 10-year period as for the rest of the analysis. The key estimated coefficients are summarised in Table 3.15.

We find evidence that the existence of a separate industry regulator is associated to higher penetration rates for mobile telecommunications in developing countries, with estimates varying widely depending on the specification. In particular, we find a different pattern between low-income and middle-income countries. The marginal effect of a separate regulator in lower-income countries is higher compared to middle-income countries. On this basis, the establishment of a separate body in charge of regulating the industry appears crucial in lower-income countries, while in 'richer' countries the effect of the regulator on penetration is not as substantial. This can be explained by the fact that, in middle-income countries, market forces may be more important for encouraging the sector's development.

Other indicators of regulatory governance, such as the existence of a sector law or the funding of the regulator through licence fees, were not found to have any significant impact on mobile penetration rates, both in the single equation and in the system estimates. Therefore, our results are less robust than Gutierrez (2003) or Cubbin and Stern (2006); it is not clear whether this is due to our relatively simplistic regulatory variable or to the possibility that the role of regulators is not as crucial for mobile operators as it is in other sectors.

According to the estimates obtained in the system (basic formulation and specification with income interactions), there is a positive relationship between country institutions, proxied by the development of the financial sector, and sector-specific governance. Therefore, part of the positive effect of regulatory institutions on the telecommunications sector may be related to overall country governance. However, it is likely the importance of country institutions is underestimated in the present study. The importance of better proxies for country institutions is highlighted by

the higher coefficients and significance obtained when the Kaufmann et al. (2006) indexes for the rule of law and quality of governance are considered.

Finally, the impact of mobile telecoms infrastructure on per capita GDP is not found to be significant except for a subset of countries over a longer 15-year timespan. This may be related to unobserved characteristics of the subsample, which was selected on the basis of the availability of data for a longer period, but may also indicate that the impact of mobile infrastructure on GDP cannot be detected over a short timespan. Moreover, the analysis is carried out on aggregate data, which are not well suited to uncover the mechanism through which mobile phones can affect income and growth (for a new approach, see Jensen, 2007). In addition, the particular specification chosen for the system of equations may affect the results significantly and leaves open an avenue for further research.

## **3.9 Appendix**

### **3.9.1 Telecommunications Penetration**

In this study telecommunications penetration is measured in terms of mobile phones. In less developed countries, mobile phones have shown high growth rates since their introduction and have proved formidable substitutes for fixed lines. For this reason, mobile telephony cannot be ignored in the analysis and focusing on fixed lines only would not capture the reality of developing countries.

Measuring penetration solely on the basis of mobile lines seems the preferable option, compared to using the total number of lines (i.e. fixed and mobile). In particular, it can be reasonably assumed that regulatory institutions have a different impact in markets in which there are competing firms, rather than a single state-

owned operator. Given that mobile telephony is usually characterized by a certain degree of competition almost from its commercial launch, the role of regulation appears to be significantly different from that exercised in the fixed market, where the development of fixed telephony often takes place for a long time in the absence of competition and the infrastructure is deployed entirely by a state-owned monopolist.

In fact it could be argued that, in the mobile market, the role of regulators is not as crucial as it is for the fixed market, where the very asymmetric structure often requires regulatory intervention to grant new entrants access to the portions of the incumbent's infrastructure that cannot be economically replicated. By combining the fixed and mobile markets together, one would in fact constrain the impact of regulatory institutions to be the same in both, even if the two markets were in different stages of development. On the basis of the above considerations, the variable of interest in the analysis is mobile penetration.

### **3.9.2 Explanatory Variables**

There are different variables that are used in the literature as proxy for country institutions. The institutional characteristics considered in this study are: (a) protection against expropriation risk (e.g. POLCON index on executive constraints); (b) financial market development (e.g. share of credit to the private sector on GDP).<sup>42</sup>

More specifically, it is assumed that the most relevant institutional characteris-

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<sup>42</sup>Another important aspect of a country's endowment is given by the functioning of the legal system, measured for instance by an index of procedural complexity. Djankov et al. (2003) find that higher procedural formalism is a strong predictor of longer duration of dispute resolution [and] higher corruption. Formalism is defined by Djankov et al. (2003) as the extent to which regulation causes dispute resolution to deviate from the neighbor model. They refer to the neighbor model as a situation in which a controversy is resolved by a third on fairness grounds without resorting to courts. Given that regulators' decisions are normally subject to appeal to courts, the threat of litigation may represent a constraint for the regulator to provide 'good' decisions. This threat is credible only if the legal system is capable of dealing with it efficiently and justly. Unfortunately, this dimension cannot be captured in a panel setting due to data limitations.

tics that impact the main variables mentioned above are protection against expropriation risk (e.g. POLCON index on executive constraints) and financial market development.<sup>43</sup> In particular, it has been shown that protection against expropriation risk affects GDP per capita and investment positively in a cross-section of countries (Acemoglu and Johnson, 2005).

In addition, a large body of literature (e.g. review by Levine, 2005) estimate that financial sector development has a positive impact on per capita GDP, by facilitating those sectors which are typically more dependent on access to external funding. This factor may also be important for investment in the telecommunications industry, especially due to the large investments required.<sup>44</sup>

The POLCON index is POLCON III from Henisz, W. J. (2002) "The Institutional Environment for Infrastructure Investment," *Industrial and Corporate Change* 11(2). The source for the share of credit to private sector is the World Bank Development Indicators.

In addition to the variables described in the text, other variables that may affect mobile telecommunications penetration are the average price of mobile services and the average price of fixed services. These are also measured in constant U.S. dollars, consistently with GDP per capita. Mobile and fixed prices are measured as: (1) price of a 3-minute peak call; (2) price of a 3-minute off-peak call; (3) monthly subscription; (4) connection charge.<sup>45</sup> The source for these variables is the International

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<sup>43</sup>These variables differ from those used in Gual and Trillas (2006), which are not available for the whole time frame of the present study.

<sup>44</sup>For developing countries, other factors may be important, such as loans from international institutions and foreign direct investment.

<sup>45</sup>It could be argued that prices should be expressed in relative terms, compared with the general level of prices in a given country. However, GDP per capita already provides an indication of the cross-country differences which are also likely to be reflected in consumers' purchasing power. Therefore, in line with Röller and Waverman (2001) and Waverman et al. (2005), prices are included in the penetration equation.

Telecommunications Union.

In the GDP per capita equation, in addition to the variables already described above, the following variables are included. Human capital is proxied by labor force and physical capital stock is from Miketa (2004), and is calculated using the perpetual inventory method in US\$ at constant 2000 prices.<sup>46</sup> Openness is defined as the sum of exports and imports of goods and services as a share of GDP.

The variable measuring the effect of international pressure is multilateral debt service, which is defined as the sum of interest and principal due to the World Bank, regional development banks, and other multilateral agencies, as a percentage of public and publicly guaranteed debt service.

All variables for which we do not mention the source are from the World Bank Development Indicators.

### **3.9.3 Additional Results: Logit Model**

Equation 3 was estimated using a logit model in order to check the results in the main text. As in Table 7, in a logit random effects model the policy variables and their interaction are the only significant regressors. Finally, a fixed effects model did not converge.

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<sup>46</sup>Miketa, A., 2004: Technical description on the growth study datasets. Environmentally Compatible Energy Strategies Program, International Institute for Applied Systems Analysis (IIASA), Laxenburg, Austria, October 2004. [http://www.iiasa.ac.at/ECS/data\\_am/index.html](http://www.iiasa.ac.at/ECS/data_am/index.html)

Table 3.16: Dependent Variable: Separate Regulator (Yes/No)

	Share of Credit RE (1)	Polcon RE (2)
<hr/>		
Independent variables		
$\ln GDP_{pcit}$	-0.238 (0.402)	-0.107 (0.350)
$Inst_{i,t-1}$	0.433 (0.444)	1.594 (1.454)
$Privat_{it}$	2.368** (0.800)	2.232** (0.805)
$Liberal_{it}$	1.793* (0.891)	2.085** (0.854)
$Privat_{it} \cdot Liberal_{it}$	-2.125* (1.269)	-2.413* (1.255)
$\ln International_{it}$	0.263 (0.279)	0.234 (0.287)
Constant	-4.356 (2.826)	3.666 (2.932)
Time Dummies	Yes	Yes
Corrected std. errors	No	No
Wald Chi2	183.88	181.03
Countries	91	91
Observations	807	808

Note: Standard errors in brackets under coefficients.



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