The Success of Internet Banking: An Econometric Investigation of its Pattern of Diffusion Within Western Europe

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Abstract

Though the Internet has been a fast-adoption medium, even banking which is particularly suited to online applications has witnessed slow customer adoption. This paper models transactional Internet banking diffusion for an unique sample of the largest 100 Western European banks for last 5 recent years (1997:2001 included). The model explicitly uses a logistic form to replicate the S-curve shape that is typical of adoption behavior of new technologies. Furthermore, the reduced form includes both supply and demand factors from the theoretical literature to analyze shifts in adoption and “push” behaviors. Among crucial results, we obtain that internet literacy (as measured by the penetration of internet usage in a country) is the major factor underlying on-line banking penetration in Western Europe. However, bank-specific variables also play a significant role in determining a bank success of on-line banking penetration. Among others, factors such as bank size, or its overall cost efficiency helps explain a large part of the differences currently observed in on-line banking diffusion in Europe. Finally, our model estimates that the maximum conversion of internet users to on-line banking seems to peak at 50%, I.e., the on-line reach of transactional banking is likely not be mass-market at the current state of the dial up internet technology.

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I. INTRODUCTION

The recent evolution of the Internet stocks clearly illustrates the current state of dismissal regarding on-line businesses. Currently, business to consumer (henceforth, B2C) e-tailing only accounts for less than 2% of total retail. Even businesses that seemed particularly well suited to online such as banks have found that customer adoption of their e-services remain low. A recent CyberAtlas (2001)'s debrief illustrates the point for the US with the headline that: "*On-line banking continues to disappoint [banks]..(…) with only 5 to 10 percent of their customers base using such services*". The same story holds true for Western European banks despite the experience of strong electronic banking such as pan-European automated teller machines networks. In fact, in Europe, bank customer on-line penetration has so far reached about 14 million household usage by end-year 2002.

Nevertheless, the costs of internet communications to deliver banking services are noticeably low (see e.g. OECD, 1998), and the possibility to develop aggressive new customer value propositions through new services (e.g., e-financial news) and through higher convenience (e.g., "24hours/7 days a week"), make it important to understand the diffusion mechanisms underpinning banking online (Schwaiger and Locarek-Junge, 1998).

As known, the economic analysis of on-line banking has been mostly theoretical, (see e.g. Bakos et alii (2000), Stamoulis (2000) or still Courchane et al. (2002)). Recently, the focus has shifted towards empirical analyses of various e-business models, including e-finance, e-credit, etc. (e.g. Chen and Hitt, 2000 or, Clemons and Hitt ,2000), but a systematic analysis of the mechanisms underpinning the diffusion of on-line financial services had yet to be conducted, at least for Western Europe [1]. This paper fills this gap by conducting a cohesive econometric investigation of the determinants for on-line transactional banking for a sample of 97 Western European banks, and for the period between 1997 and 2001 included. The rationale to use Western Europe is that on-line banking is quite evolved. We can also exploit the cross-country differences within Europe in terms of internet literacy for instance. The sample uses the largest banks because of data availability, but also because the concentration of assets in banking makes our sample a large part, more than 50%, of the total retail banking assets in Europe. Furthermore, the paper takes explicit caution that diffusion of on-line should follow a typical "S-curve" as experienced by many new innovations. It combines supply and demand factors that can "shift" the adoption curve. The major findings are noteworthy:

(1) First, Internet penetration of the bank homeland drives a large part of differences in customers' conversion to on-line among the banks sampled. Otherwise stated, the demand effect is large, as it is for other bank innovations such as Automated Teller Machines (see Saloner and Shepard, 1995);

(2) However, this is not the only aspect of on-line transactional banking diffusion. Other major "push" factors, especially the bank size, or its cost position, explain as well on-line diffusion differences among banks. The findings are consistent with the competitive real option analysis developed by Courchane et al. (2002). In their model, banks that are large or have already a strong incentive to move customers to new channels will exercise their investment options early to ensure first-mover advantages [2]. For instance, our model estimates that a bank with cost per asset one standard deviation above/below the sample average benefit today from an on-line banking penetration of about 15% higher than the average, a *material* difference in penetration indeed.

(3) Finally, the on-line transactional banking diffusion model illustrates two other crucial points. *First*, the proportion of on-line users using on-line transactional banking is likely to peak at just below 50%. Hence, the technology is not a mass-market phenomenon at least for *dial-up* Internet. This again is consistent with other anecdotal evidence that the reach of on-line transaction banking is relatively low still ,in the range of 25% in Europe by end of 2001 (Datamonitor, 2001). In passing, this also means that the internet technology adoption will lie between phone banking and automated teller machine usage among banking clients. *Second*, our estimation model can separate *timing* of, from *speed* of, variables affecting on-line banking conversion (see Bughin (2001a), Mahajan et al.(2000)). We define the *timing* of a variable, at a certain time *t*, as its effect on the *cumulated* on-line adoption at the same time *t*. In contrast, speed measures its effect on the *growth* of on-line adoption between *t-1* and *t* (Gruber and

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[1] Other empirical models of transactional banking include Furst,Lang and Nolle (2000), for the US only. Masciandaro (2000) does it for Italy only. Furthermore, none of those models uses a comprehensive logistic adoption model.

[2] In economic terms, this means that large banks prefer a "Stackelberg" leadership position.

Verboven 2001). In our model, it is also remarkable that the set of supply variables affecting timing and speed appears different.

The article reads as follows. The next section discusses the sample. Section 3 develops the econometric model, while section 4 discusses the results and the implications of the findings. A last section concludes and provides possible extensions.

II. SAMPLE KEY FEATURES

2.1. Definitions and Scope

The scope of our paper is about *retail* banking as opposed to corporate banking. It is well known that in retail banking lies the most transaction cost saving potential from the internet (Beck, 2001). Furthermore, retail banking is typically a local market with non-resident customers being virtually negligible, at the possible exception of small countries such as Luxembourg. This feature means that we can reasonably assign country variables to each bank based on their home country of origin. For a minority of banks with multi-national retail banking presence (say Fortis), we are able to separate the various locations (in this case, the Netherlands and Belgium), but we retain only the data linked to the major country as the reference (in this case, Belgium). We also define on-line banking as *any* form of transactional banking outside e-brokerage. A bank customer is considered adopting transactional banking if it does *at least one* transaction per quarter.

The sample is composed of 97 Western European banks ($i= 1,..97$) , and the data have been compiled from early 1997 to end of year 2001 on a quarterly basis ($t= 1,..20$). The sample covers thirteen European countries ($j= 1,..13$), namely Belgium, Switzerland, the Netherlands, Ireland, UK, Germany, France, Spain, Portugal, Italy, Sweden, Denmark, and Finland. Banks included in the sample are available from the authors upon request and include well-known brands such as *Fortis, BNP-Paribas, Deutsche Bank, ABN-Amro, Credit Suisse First Boston* or *Halifax*. The sample was inherently designed with the aim to include the high-end of the banking sector within each country, or about the top 10 banks. In practice, the data availability to us restricts the sample to 97 banks for the extensive period analysed. However, in total, we estimate that the banks in the sample stand for about 55% of the assets of the retail banking sector in Western Europe. The period of analysis is also noteworthy as it covers the most comprehensive period available for statistical analysis, from 1997-to 2001. In other words, the sample includes the early stage of internet banking conversion as well ad the recent period of the “downturn” observed after March 2000 .

2.2. Data constructs

We first discuss on-line transactional banking penetration. We leverage three sources: a McKinsey "E-performance" confidential database described and already used in Bughin and Hagel (2000). Second, we use Jupiter / MediaMetrix, a lead company that tracks visitors on-line. Finally, we leverage multiple investment banking reports by Merrill Lynch, Goldman Sachs, Chase / Morgan, and Societe Generale, which also report quarterly data on on-line banking penetration. As discussed in Bughin and Hagel (2000), all the data included in the E-performance have been collected through log files and checked for consistency. The sample includes 62 banks from that source. Concerning Jupiter / MediaMetrix, the data cannot directly measure transactions per se, but on-line sessions under secure environment. However, those types of sessions typically occur when an on-line banking user makes a transaction. The data includes an addition of 22 banks, outside of the overlap with E-performance. The final data comes from the investment banks reports, --the less accurate source of the three used in this paper.

From the total of *retail* customers converted to on-line, (henceforth, ON_{ijt}), we compute on-line conversion, ($CONV_{ijt}$). This stands for the percentage of on-line bank registered users divided by the bank's *retail* client base ($C_{L,jt}$), which we also have collected from the same sources as above.

Concerning now the explanatory variables used later in our analyses, we also leverage various, but reliable, sources already used in Bughin (2001a). First, we systematically tracked information on whether the transactional banking web site is an independent affiliate of the bank under survey. We define $ORG_{i,j}$= 1 if the on-line banking is standalone organizational unit , such as *If.com* from *Halifax,*or *Cortal* from *BNP-Paribas.* As in Courchane et al. (2002), the information was collected via the web on each bank as well as, if available, confirmed from investment bank reports.

Other bank statistics, such as the bank assets, its cost structure, and its rivalry intensity, was directly downloaded from one single, consistent and well-reputable source, i.e. the rating agency IBCA. Data include: balance sheet assets ($ASSETS_{ijt}$)[3]; cost-assets ratio($COAA_{i,jt}$); and employees ($EMPL_{ijt}$). Those statistics will be leveraged as "push" factors, that is, the extent to which some bank–specific factors can affect the supply of on-line banking.

Finally, we also wish to analyze how on-line banking is linked to internet literacy. To this end, we have collected quarterly on-line penetration and e-commerce usage ($INTERNET_{jt}$ and

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[3] The average assets measure reflects the average assets holdings during the period analyses, i.e., every quarter.

$ECOMM_{jt}$) for each of the thirteen countries of our sample. The data have been collected from IDC, an authoritative research company in the field of internet statistics[4].

As an illustration, descriptive median / average sample statistics are provided in **Table 1** for mid-year 2000 **Table 1** demonstrates that the sample indeed concentrates on large established banks—the bank franchise being between 1,5-1,7 million of retail customers. By mid of year 2000, online customer conversion was also still relatively low, averaging 7.6% of banks' customer base. Given internet usage at 35% for the 13 countries in the sample, this translates into about 22% of quarterly *Internet* users in top Western-European banks who were converted to banking on-line by mid of Year 2000, -a statistic that is comparable with typical monthly reach of various web sites [5] (see also Datamonitor 2001).

Table 1: Key Sample Features as- of Quarter 2:2000

**On-line **	**CONV**	7.6 / 6.6	6.3	0.1 / 29.1
Employees*	**EMPL**	14.2 / 8.3	24.8	3.0 / 71.1
Customers*	**CL**	1720 / 1540	1145	245 / 9,100
Assets***	**ASSETS**	0.92 / 0.58	87	0.6 / 4.52
Structure**	**ORG**	44 / --	--	--
Commerce**	**ECOMM**	14.6 / 11.0	8.7	1.5 / 32.0
Internet use**	**INTERNET**	34. 9 / 30.1	17.7	9.0 / 61.1

*: Thousands; **: percentage; ***: Hundreds of Billion Euro

Notice as well, the standard deviation of the key variables in our sample. For on-line transactional banking penetration, it is as large as its average. This means that *the top 5% of the*

[4] For evident consistency, the on-line penetration is based on quarterly users, while e-commerce users are defined as users who have bought at least once per quarter. Given high collinearity between both variables, only on-line penetration will be used in the analysis.

[5] Typically, monthly reach for portals are about 70% in Europe, for about 80% for front-end ISP. News sites reach about 35% of the internet users monthly, for 30% for retailing sites, and about 25% for brokerage and banking sites. For more information, those statistics are readily available at MediaMetrix on-line.

banks in the sample have managed *to cumulate more than triple* the proportion of on-line banking users in comparison to the sample mean for the period until mid 2000. At the top of on-line penetration, we obviously observe the Scandinavian banks.

Of course, Scandinavian countries have benefited early from a high degree of household “internet” literacy. E.g., Sweden had an Internet penetration more than double the one of France or Italy by 2000. But this does not explain everything. How could it be that banks from the same country, -say Portugal- have themselves witnessed tremendous difference in their conversion capability to e-banking? For instance, the Portuguese Banco di Spirito e Commerciale managed to convert a proportion four times higher than what a competitor such as BPI did by the turn of the century. The next section develops an empirical model to explain such disparity of on-line banking diffusion.

III. WHAT DRIVES ON-LINE BANKING DIFFUSION?

3.1. A demand model of diffusion

The best suited model for determining a “demand model” of diffusion is a logistic model of a “S-curve” (see Mahajan et al. ,2000, or Gruber and Verboven, 2001). Technically, the logistic model of on-line demand conversion, ON_{ijt}, is represented as:

$$(1) \qquad ON_{ijt} = ON^{*}_{ijt} / (1+\exp(-a-b.t))$$

where the parameters, *a* is a *timing* variable (i.e., *a*>0 shifts penetration rate at time t) and *b* measures diffusion *speed* (i.e., *b*>0 means that penetration rate is getting faster).

Note in equation (1) that ON^{*}_{ijt} is not measurable, but we reasonably assume that it is in part a function of bank total customer base, and the proportion of the bank homeland quarterly internet users:

$$(2) \quad ON^{*}_{ijt} = c.\,(INTERNET_{jt}.Cl_{ijt})^{\alpha}.\,W_{ijt}$$

W is a random term capturing the idea that some demand effects are not directly measurable in equation (2), such as the “wealth” of each bank customer franchise. The parameter α is an elasticity measure with α higher than 1 would mean that a client base in a country with more Internet literacy would facilitate more on-line banking conversion. This is empirically tested hereafter.

By definition, we should also have that: $c<=1$. As such c would measure the peak proportion of bank Internet users willing to convert to on-line banking. In practice, we expect the estimate of c to be relatively low. The reasons being that technology adoption in banking is no more than 65% even for automated teller machines after more than many years of introduction in Europe; also, many barriers remain for on-line transaction banking, e.g. web security. Finally, a recent Datamonitor (2001) survey clearly emphasizes that branches still rule banking in Western Europe. The Internet is the preferred method for dealing with banking products in only 4% of the European population, far below phone banking at 13%, and branches at 79%.

3.2. Adding supply-factors

Equations (1) and (2) assume that a and b are constant parameters, so that the only difference in customer conversion to on-line is linked to (country) internet literacy. However, one important theme of this article is that internet literacy cannot explain alone the large spread of internet banking diffusion among banks. Rather, *banks may have had some incentives in offering e- services and used proactive "push" strategies to "speed up" the conversion of their customers' franchise to on-line banking.*

In order to test this out, we then allow both the diffusion timing and speed parameters, a and b, to be correlated with bank –specific factors (symbols explained hereafter).

$$(3) \qquad a_{ijt} = a + d.\ COAA_{ijt} + e.\ COMP_{ijt} + f.\ ORG_{ij} + g.\ EMPL_{ijt} + U_{ijt}$$

and:

$$(4) \qquad b_{ijt} = b + h.\ COAA_{ijt} + j.\ COMP_{ijt} + k.\ ORG_{ij} + l.\ EMPL_{ijt} + V_{ijt}$$

where all the labels have been defined in **Table 1** while the variable COMP is discussed later.

The equations (3)-(4) are first-order, linear, approximations of how bank-specific factors can affect on-line transactional banking diffusion. We, of course, are agnostic as to whether the *complete* list of supply factors is included in our reduced-form model. We thus add the terms U and V as random terms to account for possibly omitted factors.

The hypotheses concerning the "first-order" effect of the bank-specific factors on timing and speed of on-line banking are as follows. Bughin (2001b) and Stamoulis (2000) provide empirical evidence that banks with lower cost structure have both converted existing, and acquired new, customers faster than other banks. This fits with the traditional model of spatial competition where low cost providers may reinforce their competitive advantage

(Bouckaert and Degryse, 1995). As a priori hypothesis, we thus posit that a low cost per asset base, COAA, should be correlated with on-line banking diffusion timing and speed [6][7].

COMP is a proxy measure for the effect of rivalry on the bank incentives to offer and push for on-line adoption. In the oligopoly model developed by Courchane et al. (2002), banks with significant market power (and thus lower rivalry) will push penetration faster. We measure COMP as a traditional rival concentration index, i.e., COMP is the sum of squares of the market share of the home country rival banks. According to Courchane et al. (2002), one expects *a* and *b* to be significantly negatively linked with COMP.

ORG is an organization variable supposed to positively affect the diffusion timing and speed of on-line banking. In fact, there is more than casual evidence that spun-off organizations such as e-brokerage, have been launched faster than on-line banking and been more flexible and aggressive than an integrated bank (Hagel and Singer, (2000), or Bakos et al., (2000)). One expects *a* and *b* to be positively correlated with ORG.

Finally, and again in consistency with the real option model of Courchane et al. (2002), we also test whether bank size may positively affect on-line banking diffusion. Empirically, we use an employment variable, EMPL, and not total assets. This avoids the issue of too high multi-collinearity with COAA, a variable which is scaled by the asset base of the banks.

IV. RESULTS

4.1. Estimation techniques and model variants.

The procedure amounts to fold equations (3-4) into (1-2), and then to estimate the model under the assumption of normally distributed disturbance terms. We use non-linear least squares with heteroscedastic-consistent estimates as in Putsis and Srinivasan, (2000). Furthermore, we include both *country and banks* fixed effects to account for unobserved variables in our empirical estimation.

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[6] In a previous version of this paper, we also used automated teller (ATM) density as a measure of alternative channels—however, as all referees pinpoint rightfully, this is possibly a too weak proxy for commitment to alternative channels. After all, ATMs are still physical networks (I.e., not shifting success factors for access to the customer dramatically) and can also be very different from country to country (e.g., a common shared ATM infrastructure exist or not). A better proxy could be phone banking; however this is not available in our sample.

[7] The asset denominator of COAA is computed as the average assets between two quarters.

Results are presented in **Table 2**, with only statistically significant variables laid out. Also, various versions of the model are also illustrated.

As a benchmark, column (**a**) presents a model where the timing and speed parameters are estimated as constant terms. This is the "pure demand" model where banks are assumed to play a *neutral* role in the on-line banking diffusion process. Column (d) presents the most complete model based on (1-4) above. In between, we use 2 hybrid models: column (**b**) presents a model where timing is defined by equation (3) above—but speed is assumed a constant; column (**c**) in contrast presents the model where timing is a constant, but speed effect is modeled by equation (4) [8].

In order to discriminate for the best model, we could already look at the significance of the bank-specific coefficients when moving from column (a) to (d). Their statistical significance implies that the complete model is by far the most informative. Furthermore, the pseudo-R^2 derived from the non-linear estimation technique is a global measure of goodness-of fit and confirms that the most general model (d) is the most relevant. Other more robust measures such as the Akaike information test confirm it; hence we discuss the estimates at the light of model (d) hereafter.

4.2. Demand effects

The coefficient *c* converges towards just less than 50% of *internet* users converting to on-line transactional banking. Versus current *average* penetration in our sample of about 31% of on-line converting to banking, this still represents a significant potential upside. In contrast, the top 5% banks in our sample that already achieved about 43% conversion by the end of the period. For those banks, the only growth factor to on-line banking is now the growth of Internet usage (and after that, conversion to on-line banking) in the future.

Note finally that the significance of *c* was found "borderline" (just below the 5%-level). This is typically the case in a logistic model of diffusion when the data presumably only tackles the early stage of the diffusion curve. Nevertheless, in absence of more relevant data, our estimates provide an interesting benchmark, and suggest that on-line banking is not necessarily to exceed the success of automated teller machines adoption for the years to come.

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8 A referee has suggested to also test the S-curve model against a simpler linear model. This would amount to test the model in column (a) against a more restrictive model whereby the term $(-a_{ijt} - b_{ijt}.t)$ in equation (1) is a constant. Those restrictions imply a pseudo R^2 of 0, 32, i.e., significantly less informative than the simplest model in column(a).

Now, referring to the coefficient α, the elasticity is estimated at 1,2. The elasticity is slightly different, and above one using a one-tail t-test for the most complete model in column (d). This may represent some evidence of scale or network effect of the Internet, or still the fact that, the more widespread the Internet usage, the wider the use of internet applications as demonstrated in the various GVU surveys (GVU, 1999).

Table 2 –On-line transactional banking diffusion estimates ("Demand Effects")

Model:	**(a)**	**(b)**	**(c)**	**(d)**
c (peak usage)	0,44	0,47	0,46	0,49
α (internet elasticity)	1,12	0,92	1,04	1,18

Notes:

1. Fixed Effects included for all models. All F-tests of the existence of fixed effects are statistically significant.
2. All reproduced estimates are statistically significant at risk level, α=5%.

4.3. Supply factors

As shown in **Table 2(continued)**, the *timing* and *speed* factors have intuitive value, with a quarterly diffusion speed of above 4% of the total client converting per quarter, for instance. We first look at the estimated impact that each supply variable separately has on diffusion. We then discuss the relative effect between "demand" and "supply" factors to affect on-line diffusion patterns.

. The analysis highlights as in Stamoulis (2000) that a bank commitment to a *low* cost per assets structure is a robust correlate to customers banking on-line, both with respect to timing and speed of diffusion. In order to gauge those effects further, and based on estimates in **Table 2**, , we have computed what would be the difference in on-line banking penetration for banks with two standard deviations lower cost structure than the observed mean [9]. In such case, on-line banking penetration will respectively be 36% higher for the second quarter of year 2000, a non-trivial difference indeed.

[9] Two standard deviations allow to take the top 5% banks in the distribution if the distribution is approaching a Normal distribution .

Table 3–On-line transactional banking diffusion estimates (“Supply Effects”)

Model:	**(a)**	**(b)**	**(c)**	**(d)**
“Timing”:				
Constant	0,618	0,796	0,622	0,746
COAA	-	-0,0132	-	-0,0148
COMP	-	-0,023	-	-0,019
EMPL	-	0,084	-	0,062
ORG	-	n.s.	-	n.s
”Speed”:				
Constant	0,0502	0,0425	0,0146	0,0152
COAA	-	-	-0,071	-0,062
COMP	-	-	-0,037.	-0,054
EMPL	-	-	n.s.	n.s.
ORG	-	-	0,026	0,023.
Implied diffusion coefficient averages:				
- “timing”	0,618	0,648	0,622	0,649
- “speed”	0,0502	0,0425	0,0461	0,0415
Pseudo-R^2:	**0, 51**	**0, 68**	**0,62**	**0, 76**

Notes:
1. Fixed Effects included for all models. All F-tests of the existence of fixed effects are statistically significant.
2. All reproduced estimates are statistically significant at risk level, α=5%. Otherwise, n.s.

As in Courchane et al. (2002), larger banks have higher on-line banking penetration, but according to our model estimates, size *only affects timing*, *not speed*, of conversion. Furthermore, the marginal impact of two standard deviations larger size is in the range of 10 percentage more penetration, - a relatively lower impact on diffusion than is the impact of a lower cost structure.

In contrast, the dummy ORG affects *only speed* of on-line banking penetration. Since most banking units with separate organizations are also offering e-brokerage services, this may reflect e-brokerage being an important driver to push people using internet transactional banking as well (Diniz,(1998)) .

Finally, the COMP variable appears as well as quite significant both for timing and speed of diffusion. The negative sign corroborates the US findings by Courchane et al. (2002) that the concentration of a bank's rivals in its home market has a negative impact on the incentive of the bank to engage in proactive on-line banking conversion of its customer base.

. All in all, thus, the results here-before are clearly consistent with the theory that banks can adopt "push" strategies to convert their base to on-line. Empirically, we can also estimate the *relative* importance of the supply and demand factors for the total diffusion of on-line banking in Western Europe.

In fact, the estimated variance around the estimated mean is explained by 54% by the variance in internet penetration, 31%, by difference in bank-specific effects on timing, and 15%, by difference in bank-specific effects on speed of diffusion. In other words, bank-specific factors explain roughly as much as Internet literacy to explain differences in on-line banking diffusion.

5. CONCLUSIONS

This paper has developed a empirical model of on-line transactional banking diffusion in Western Europe, for a large sample of close to 100 top banks, using quarterly data from 1997 to end of 2001. This period includes the early stage of on-line development, as well as the significant downturn since March 2000. The diffusion model is estimated through robust non-linear least squares and includes fixed country-, and bank-, effects to account for unobservable effects.

The estimates demonstrate that on-line banking diffusion is clearly linked to internet literacy, but as much as to bank-specific factors affecting both the timing and speed of on-line banking diffusion. In particular, cost-effectiveness and size are two critical bank factors, as are the degree of rivalry intensity together with the organizational spin-off of on-line offering.

Various avenues for research include the extension of the model to other on-line applications, or other geographies. Also, the results of the model still hinge on the early years of the diffusion curve, and may not be fully robust in the future. However, experiments for out-of-sample predictions (not reported in this paper) demonstrate the likely robustness of the above results. Hence, we are willing to claim that clear “push” effects are likely to be a strong play to boosting on-line adoption. After all, the fact that some companies have spent tens of millions of Euro building high-level look and feel web sites is a competitive commitment that should come out of any clear-cut model of new technology diffusion.

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