BANCA D'ITALIA

EUROSISTEMA

Temi di discussione

(Working papers)

Delayed privatization

by Bernardo Bortolotti and Paolo Pinotti

Number 663

April 2008

DELAYED PRIVATIZATION

by Bernardo Bortolotti[*] and Paolo Pinotti[**]

Abstract

This paper studies the timing of privatization in 21 major developed economies in the period 1977-2002. Duration analysis shows that political fragmentation plays a significant role in explaining a government's decision to privatize: privatization is delayed longer in democracies characterized by a larger number of parties and operating under proportional electoral rules, as predicted by war of attrition models of economic reform. Results are robust to various assumptions on the underlying statistical model and to controlling for other economic and political factors.

JEL classification: D72, D78, L33.

Keywords: privatization, political economy, war of attrition.

Contents

[*] University of Turin and FEEM

[**] Bank of Italy and UPF

1 Introduction[1]

In the last two decades, a big privatization wave has redrawn the borders of the economic activity of the state in developed economies. Thanks to privatizations worth approximately $ 1 trillion, OECD countries have shrunk their state-owned enterprises (SOE) sector on average from more than 12 to 6 percent of GDP. In most cases, privatization represented a qualifying element of a package of measures including liberalization, deregulation, and corporate governance reforms. Thus, the implementation of privatization policy is certainly one of the most important experiments of structural reform ever attempted in market economies.

As Figure 1 shows, the process followed a similar cyclical pattern across countries. In all OECD economies (with the notable exception of the United Kingdom and the United States) privatization started in the late 80s or early 90s, peaked in the late 90s and dramatically declined after the turn of the century. Yet the timing of sales varied greatly across countries. Some governments have promptly entered the advanced stage of the process, and raised a significant fraction of their revenues earlier, while others have lagged behind.

Why are privatizations delayed? Why did it take just a few years for the United Kingdom to launch the largest scale privatization program in history, while the process started in Switzerland only in the late 90s?

We claim that political fragmentation, which is related to the number of agents with veto power in a given political system, hampers the implementation of policies with significant distributional consequences, such as privatization. A lower political fragmentation favours executive stability and allows incumbent governments to privatize a sizeable fraction of their SOE sector sooner, as the constituency of the "losers" from the policy change is less likely to enjoy bargaining power. On the contrary, highly fragmented political systems tend to disperse decision-making power among different actors, so that executives are weaker and characterized by higher turnover. In this context, the different political actors will hardly reach an agreement about how to distribute the burden of the policy change, and privatization will be delayed by a "war of attrition" as in Alesina and Drazen

[1] We wish to thank Utpal Bhattacharya, Antonio Ciccone, Luca Farinola, Arendt Lijphart, Bill Megginson, Marco Pagano, David Parker, Enrico Perotti, Xavier Sala-i-Martin, Alessandro Sembenelli, Guido Tabellini, and seminar participants at the World Bank, 2003 EEA meeting, WZB, Berlin, and Fondazione Eni Enrico Mattei for useful comments. Ettore Panetti provided excellent research assistance. The opinions expressed do not necessarily reflect those of the Bank of Italy.

(1991) and Spolaore (2004).

In this paper, we test this prediction by estimating a duration statistical model on data for 21 OECD economies during the 1977-2002 period. The results are broadly consistent with the empirical implications of the war of attrition theoretical model. Political systems with a smaller number of parties and operating under majoritarian electoral rules privatize sooner, while large-scale privatization is delayed in more fragmented democracies.

A tale of two countries, the United Kingdom and Switzerland, illustrates the role of political fragmentation in the timing of privatization. After winning the 1979 election, Mrs. Thatcher kicked off her program immediately with a first batch of sales in (reasonably) competitive industries. The process then gained momentum after the 1983 re-election and continued apace in the late 80s in the newly liberalized electricity market and in the water industry. Throughout, privatization was fiercely opposed by the trade unions and by the Labour party. Nevertheless, the majority enjoyed by the Conservatives in Parliament combined with the power granted to the cabinet by the British political system allowed the government to push back the opposition and to accomplish the announced program (Vickers and Yarrow (1988)).

Conversely, Switzerland was the last developed country to privatize, since it took decades for the four parties forming the Federal Council to find a consensus on reform.[2] After a long negotiation, the 1998 Telecommunications Act was eventually enacted yielding a timid liberalization of the sector, and the flotation of a minority stake of Swisscom. By the end of 2005, the Swiss state still held a 66.1% stake in the company. At the beginning of 2006, the executive set forth a plan to further the privatization of the company but the policy was immediately blocked by the opposition of the centre-left Social Democrats party, one of the permanent members of the Council. Besides, any further attempt would likely have to pass a popular referendum, definitely a distinguishing feature of the Swiss political system.

This study relates to empirical studies of privatization, surveyed by Megginson and Netter (2001). In particular, recent work has explored specifically the role of politics. Clarke and Cull (2002) examine the political and economic incentives for provincial governments in Argentina to privatize banks. They find that the likelihood of privatization is higher for poorly performing banks, while the overstaffed and larger banks tend to remain

[2]The Federal Council is the executive body of the Swiss political system. It is formed by seven members that represent all and only the four major parties, which span the whole ideological spectrum. For detailed information, see chapter 2 of Lijphart (1999).

under state ownership. Boehmer, Nash and Netter (2005) extend the analysis to a larger group of countries, finding that in non-OECD countries bank privatization is more likely the lower is the quality of the nation's banking sector, the more the government leans to the right and the greater the government's accountability to the people. Financial distress is instead the main determinant of bank privatizations in OECD countries. The authors also study the timing of bank privatization and conclude that countries with banks that have less equity-capital and extend more loans to the government, and with higher government accountability privatize state-owned banks faster. Political factors instead do not seem to affect the timing of bank privatization. In a case study on India, Dinc and Gupta (2005) analyze the decision to privatize at the central government level and find that the likelihood of privatization is higher in states where the party of the incumbent central government faces less local political competition.

While previous literature focused on ideological orientation as a political determinant, our paper is the first to study empirically the role of political fragmentation on the timing of privatization in developed economies.

Our work is also related to the empirical literature on the political economy of reform in the context of fiscal stabilization. In particular, in the last few years, several papers have tested the war of attrition model using fiscal data. For example, Padovano and Venturi (2001) provide a detailed case study of the effect of political fragmentation on public finance in Italy during the post-war period; Huber et al. (2003) and Woo (2003) extend the analysis to OECD and to almost 60 countries, respectively. All these papers find evidence of a positive relationship between political fragmentation and budget deficits or public debt.

The empirical strategy common to all these studies has been to fit OLS regressions of some measure of fiscal imbalance on political fragmentation, along with other political and economic explanatory variables. An estimated positive coefficient on measures of political fragmentation, like for instance the number of parties, is interpreted as evidence in favour of the war of attrition model. However, such a methodology does not allow one to disentangle the empirical implication of the war of attrition model from those of alternative, more general models encompassing a "public good" type of market failure: the higher the number of veto players involved in the decision making, the larger should be the total draw from the common pool of government's budget. Put differently, in the war of attrition model higher political fragmentation results in deeper fiscal imbalances only indirectly, as a consequence of longer delays to reform, while the specific prediction of the model concerns the length of the delay itself.[3]

We improve in this respect by identifying a formal link between the war of attrition theoretical model and the duration econometric framework pioneered by Cox (1972). Such a link arises naturally from the central role played by time both in the theoretical and the statistical model.

We also contribute to this strand of literature by providing a new set of continuous and time-varying political indices computed from electoral data. Our dataset survives an extensive cross-checking with independent data, proving itself a reliable tool for empirical work in political economy.

The remainder of the paper is structured as follows. Section 2 reviews the theoretical war of attrition model. Section 3 derives the estimating equation and shows that duration analysis provides a suitable statistical framework to perform this test. Section 4 introduces measures of the delay to privatize and of political fragmentation, it describes the data and compares them to existing datasets. Section 5 presents the empirical results. Section 6 concludes.

2 Theoretical framework

The political economy of policy adjustment (particularly, fiscal stabilization) has been studied by Alesina and Drazen (1991). In their model, the benefit of stabilization accrues to all citizens and stems from abandoning a highly distortionary method of financing public expenditure. However, the costs of stabilization (i.e. higher taxation) are apportioned differently between interest groups, with one group bearing a disproportionate fraction of the tax burden. Under these assumptions, the process leading to stabilization becomes a "war of attrition" between groups, characterized by political stalemate until one group concedes. Concession occurs at equilibrium when the group-specific costs and benefits of waiting balance each other. Importantly, Alesina and Drazen note that large coalition cabinets made of diverse parties may hardly reach an agreement on how to allocate the

[3] For instance, Velasco (1997) presents a dynamic model in which higher political fragmentation leads to higher public deficits without resorting to any war of attrition between political agents. A simpler, static example is presented in chapter 7 of Persson and Tabellini (2000). The more general relationship between political fragmentation and fiscal distress has been as well extensively tested since the seminal work by Roubini and Sachs (1989); more recent contributions are Alesina et al. (1998) and Perotti and Kontopoulos (2002).

tax increase among the different constituencies. Therefore delayed stabilization should be associated with higher political fragmentation.

The empirical implications of this model appear a bit far fetched to allow for a proper empirical test. Spolaore (2004) makes an important step in this direction, by developing a model that allows comparing patterns of adjustment policies in different systems of government. The primary focus is on the way control over decision making is allocated across political agents with different preferences. Two benchmark systems are considered: the "cabinet" system, giving full control over policies to one decision maker, and the pure "consensus" system, in which each political agent retains veto power over adjustment policies. The two systems differ therefore in terms of *political fragmentation*, which is defined simply as the number of political agents with veto power.

The cabinet system is shown to provide prompt adjustment, even if it may adjust too often as the policy-maker fails to internalize the adjustment costs of other political agents. On the contrary, the consensus system may fail to adjust even when adjustment is optimal. Interestingly, in the presence of large adjustments, like privatization, the model shows that the only equilibrium in the consensus system is a war of attrition *à la* Alesina and Drazen, and that the expected delay to reform depends on political fragmentation.

In particular, let T be the delay of reform, with $f(T)$ and $F(T)$ being, respectively, its density and cumulative distribution. The *concession hazard rate* $\lambda(T) = \frac{f(T)}{1-F(T)}$ is the probability that adjustment occurs after T periods given that the economy did not adjust before. Then, the prediction of the model is that

$$\lambda(T) = \left(\frac{n}{n-1}\right)\theta, \tag{1}$$

where n is the number of agents with veto power and θ is an exogenous parameter that depends on the size of adjustment at stake (or, in another way, on the initial conditions of the economy). Thus, the implied concession hazard rate is decreasing in political fragmentation.

Privatization is a major adjustment policy, defined as any efficient policy change with significant distributional consequences. First, privatization curbs political interference, improves managers' incentives, and tends on average to increase the efficiency of firms (Megginson and Netter (2001)). Second, privatization has important distributional effects as it typically involves a transfer of wealth from insiders of state-owned enterprises (such as employees) to outsiders, especially shareholders. Indeed, state sell-offs have been often associated with restructuring and layoffs, with efficiency gains accruing to shareholders of newly privatized firms (Megginson, Nash and van Randenborgh (1994), Haltinwanger and Singh (1999)). If one country's political system is highly fragmented, the interest group of "losers" from privatization has voice in the political arena and engages in a war of attrition which delays the efficient policy change.

In this context, it is thus straightforward to interpret T as the time elapsed until privatization occurs. The next sections will describe how we take equation (1) to the data.

3 Empirical strategy

Equation (1) relates the concession hazard rate at T, i.e. the probability of observing the adjustment after T periods, to some explanatory variables. Duration analysis provides the exact translation of this relationship into a statistical model. The dependent variable of duration models is the conditional hazard rate

$$\lambda(T \mid x) = \frac{f(T \mid x)}{1 - F(T \mid x)}, \tag{2}$$

where T , $f(.)$ and $F(.)$ are defined as in (1) and x is a vector of covariates including proxies for n and θ, along with other political and economic controls.

Following the literature on survival analysis (Cox (1972), Kiefer (1988) and Van Den Berg (2001)), we first assume a proportional hazard rate, which implies separability of $\lambda(.)$ in T and x:

$$\lambda(T \mid x) = \Gamma(x)\Lambda(T) \tag{3}$$

The proportionality assumption (3) allows the difference in hazard rates between countries i and l observed in period r to depend on the difference $[x_i(r) - x_l(r)]$ but *not* on the particular period r (at least not directly) and is key to the interpretation of many results. The additional term $\Lambda(T)$ is introduced to allow for flexible time dependence of the hazard rate and encompasses time independence (that is, a constant $\Lambda(T)$ like in (1)) as a particular case.

We will fit two different versions of the proportional hazard rate model. First, we follow the original Cox (1972) semi-parametric approach, which leaves the baseline hazard

$\Lambda(.)$ unspecified. In spite of its simplicity, the Cox model is already sufficient to identify the effect of changes in x on the hazard rate (this is a direct consequence of proportionality). We will then check the robustness of results by estimating a fully parametric model which specifies a functional form for the baseline hazard. In particular, we refer to the conventional Weibull (1951) specification

$$\Lambda(T) = \alpha T^{\alpha-1}, \tag{4}$$

where α is an ancillary nonnegative parameter which allows for duration dependence.[4]

Both models assume a non-negative exponential form for $\Gamma(.)$:

$$\Gamma(x) = \exp(x\prime\beta), \tag{5}$$

where β is the vector of coefficients of interest, which is estimated by maximum likelihood (partial for Cox, full for Weibull). The direction of the effect of the k-th regressor on the hazard rate relates directly to the sign of the k-th element of β: an increase in x^k increases (decreases) the hazard rate as long as $\beta > 0$ ($\beta < 0$). In particular, we will be mainly interested in the coefficient of some proxy for the theoretical number of veto players n in (1) .

The proportionality assumption imposed by the Cox and the Weibull models is convenient for several reasons. First, it allows to model very simply the effect of the explanatory variables on the hazard rate, which is often the relationship of primary interest. Second, in our particular case, proportionality characterizes as well the hazard rate in equation (1) , which we want to test. Nevertheless, it remains a restrictive assumption and we may want to check how the results change as we relax it.

Consider the parametric Weibull model and notice that it can be restated as

$$\alpha \ln T = -x\prime\beta + \nu, \tag{6}$$

where ν has a type I extreme value distribution, which is implied directly by the proportionality assumption. We will relax proportionality by letting $\ln T$ follow a normal distribution, conditional on the vector of covariates x. In this case, maximizing the like-

[4] In particular, for $\alpha > 1$ the process shows positive duration dependence, i.e. the probability of failure increases through time; the opposite holds true as $\alpha < 1$;finally, for $\alpha = 1$ the hazard rate is independent of time (in this last case the Weibull model collapses to the simpler exponential form).

lihood for the lognormal distribution of T will provide efficient estimates of the vector of parameters $\phi = -\frac{\beta}{\alpha}$

4 Data

This section presents our dataset. Our sample includes 21 sound democracies with established political institutions enabling an orderly succession of powers: most of Western Continental European countries (Austria, Belgium, Denmark, Finland, France, Germany, Greece, Italy, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland), Anglo-Saxon countries (Australia, Canada, Ireland, New Zealand, United Kingdom, United States) and Japan.

Given our focus on the timing of privatization policy and related reforms, the sample period is certainly a key dimension of the dataset. We set 1977 as the initial year, reporting what is conventionally considered the first privatization in recent times, the IPO of British Petroleum.[5] The final year of the sample period is 2002, when the privatization wave ends in most countries. Indeed, privatization activity in OECD countries peaked in 1999 and abruptly slowed down after the turn of the century, with revenues back to the levels reported in the early 80s, at times when only the United Kingdom was seriously engaged in privatization (Bortolotti and Siniscalco (2004)). Our sample period thus captures in its entirety the big privatization cycle of the last two decades and is thus suitable for the empirical analysis of the timing of reform.

Next, we introduce our privatization, political and economic variables. The most important ones are the empirical counterparts for T and n; they are also those involving the most critical measurement issues.

4.1 Delay of privatization

A reasonable starting point to measure the delay period is t_0=1977, when privatization definitely entered the world economic and political agenda. About the end year t_i, which is needed to set the length of the delay period in each i-th country, we may want to choose a date that takes into account genuinely the advancement of the privatization process in

[5] Some important historical antecedents were the sales of Volkswagen and VEBA implemented in the Federal Republic of Germany by Adenauer in the early 60s. However, these companies quickly returned in public hands and were bailed-out under the pressure from disappointed investors.

that country. Thus, we have first collected revenues data for all privatizations (public offers and private sales) reported in Securities Data Corporation, certainly one of the most comprehensive sources of information at the transaction level. We have aggregated them to construct *REVGDP*, equal to total privatization revenues as a fraction of GDP in each country-year. Then, the end year of the delay period is defined as

$$t_i = \min\left\{s : REVGDP_{is} \geq median\left[REVGDP_{ir}\right],\ r = 1977, ..., 2002\right\}, \tag{7}$$

that is, we consider the first year in which total privatization revenues raised in country i equaled or exceeded its median yearly revenues. Median revenues are adopted rather than the first transaction because initial privatizations are typically sporadic and small-scaled, so that they do not prefigure a real start of the reform. For analogous reasons we discarded using the year in which maximum privatization revenues were raised. Finally, median revenues were preferred to average revenues because of the invariance of the former to extreme (and possibly anomalous) values of the observed distribution. The delay of privatization in country i is thus defined as

$$T_i = t_i - t_0. \tag{8}$$

4.2 Political fragmentation

Conceptually, political fragmentation relates to the number of veto players n in expression (1) . The larger is n, the higher the degree of political fragmentation. When it comes to making the notion operational one has to solve two issues. First, identifying the relevant political agents. In this respect, political parties are usually regarded as the basic cohesive entities representing specific interest groups. Second, how to weight them according to their actual bargaining power. Comparative political science has developed suitable measures that help address this issue. The Effective Number of Parties (*ENP*) introduced by Laakso and Taagepera (1979) parallels the Herfindahl index in evaluating political fragmentation according to the distribution of seats held by all parties:

$$ENP = \left[\sum_{j\in P}\left(\frac{s_j}{\sum\limits_{k\in P} s_k}\right)^2\right]^{-1}, \tag{9}$$

where s_j is the number of seats in the parliament held by the j-th party and P is some set of parties. Expression (9) says that if there are N parties, the *ENP* will take the value N if they all have the same number of seats, otherwise it will take lower values, in order to "discount" parties that are weaker in terms of parliamentary seats. As the number of parties increases, the single shares decrease on average and the *ENP* increases.

Since in any political system most of the veto power is held by the government, we first compute the index over parties forming the executive coalition; we call this variable *GENP*. At the same time, a highly fragmented parliament could also delay the implementation of reform policies which may require broader consensus than simple majority.[6] Thus, we will compute the index as well over all parties represented in the parliament; we will refer to this second measure as *PENP*.

Finally, a third measure of political fragmentation considers the barriers to entry imposed by different electoral systems. In particular, majoritarian systems tend to reduce the number of political parties (and thus veto players) gaining access to the parliament, as opposed to proportional systems (Persson, Roland and Tabellini (2007)). In empirical political economy studies, the electoral system is usually characterized in terms of binary variables. We refer instead to previous work by Gallagher (1991), who computed a continuous measure of the disproportionality of the electoral system:

$$DISPR = 100\sqrt{\frac{1}{2}\sum_{j\in P}\left[\left(\frac{s_j}{\sum\limits_{k\in P}s_k}\right)-\left(\frac{v_j}{\sum\limits_{k\in P}v_k}\right)\right]^2}, \tag{10}$$

where v_j is the number of electoral votes got by the j-th party and s_j and P are defined as in (9) . The index equals 0 if there is perfect proportionality between seats and votes. It increases, on average, as the electoral rule moves toward the majoritarian system; it is maximum for presidential elections, when the only seat at stake goes to the winner, in which case the index equals the percentage of votes obtained by the defeated candidate.

All the three indices, *GENP*, *PENP* and *DISPR* are continuous and defined for each country-year in the sample. As such, they account better than binary or discrete indices for the extreme heterogeneity observed at the institutional level.

The main source for the electoral data needed to compute the political variables was

[6] Notice that, in several countries (for example France, Belgium, etc.) the implementation of privatization entailed constitutional reforms, which in turn required a qualified majority in the Parliament (for instance, 2/3 of the votes).

Lijphart (1999). We have used his series for *DISPR* and *PENP*, updating to our end year. The other index, *GENP*, has been developed independently; as such, the relative series is compiled ex novo from various sources listed in Appendix A.[7]

To cross-check the reliability of our dataset, labeled FEEM Political Database (FPD), we have compared it with the World Bank Database on Political Institutions (DPI) by Beck et al. (2001), one of the most widely used sources in empirical political economy studies. Then we have compared FPD and DPI pair-wise with a third data base compiled by an independent source (Tsebelis (2002)) in the country years when the three overlap. Results of the cross-check are shown in Appendix B. Indeed, FPD and Tsebelis appear similar in several respects. The average difference between the number of seats is very low for each of the three main parties and for the government's coalition as a whole. Moreover, the percentage of "perfectly matched" cases is above 80% for each of the parties, and quite high for the government's coalition. On the contrary, the World Bank DPI does not seem to be related to any of the other two databases. First, the number of observations is much lower, which means that many electoral results are missing. Second, the pair-wise comparison yields a very high average difference in terms of reported seats (about 30 seats each election for the first party and for the government as a whole). Finally, the percentage of matched data is dramatically low, always under the 5% for the first party and for the government as a whole.

Table 1 reports cross-country averages of *GENP*, *PENP* and *DISPR*, and Figure 2 plots them on two- and three-dimensional scatters, along with the fitted OLS linear regressions.[8]

The slope of the regressions is consistent with the expected pair-wise relationship between the three variables. Persson, Roland and Tabellini (2007) have recently shown that electoral rules determine the number of parties gaining access to the parliament (higher in proportional systems, lower in majoritarian ones), which in turn shapes the fragmentation of the executive. The preliminary inspection of the data presented here is in line with their results. Most importantly, the three indices together univocally characterize the countries in the sample according to their political fragmentation. In particular, sticking to the terminology of Spolaore (2004), the cluster of Anglo-Saxon countries on

[7] The dataset is available at http://www.feem.it/fpd

[8] Three countries implemented institutional reforms in our sample period: Italy modified its electoral system in 1992, New Zealand and Japan in 1993. The two averages presented for these countries are computed over the two sub-periods before and after the first election taking place under the new regime.

the bottom-left of Figure 2d provides a reasonable empirical counterpart to the cabinet theoretical model. At the opposite, the countries on the top-right (Switzerland, Italy, the Lowlands and part of Scandinavia) resemble well the features of the consensus system.

We next turn to the description of the control variables that enter the vector x in the estimating equation (2) .

4.3 Control variables

While we investigate the role of political fragmentation on the timing of privatization, we may want to control for other possible determinants of privatization. Two of them deserve special attention.

First, initial conditions matter. In particular, privatization could be simply affected by the initial size of the SOE sector and/or the fiscal imbalance. Notice further that initial conditions determine the size of adjustment θ in equation (1) .[9] We control for one country's initial size of the *SOE* sector by the average of the SOE value added as a percentage of GDP in the three years preceding the first privatization reported in SDC. Similarly, we measure fiscal pressure by one country's average budget *DEFICIT* in the same pre-privatization period.[10]

Second, the strong distributive effects of privatization suggest that the ideological orientation of the executive matters in explaining the timing of the reform. Measuring partisanship of the government faces methodological issues analogous to those described above for political fragmentation[11]. We refer to the study by Huber and Inglehart (1995), who, by means of expert interviews (over 800 for 42 countries, including the 21 in our sample), have produced a comprehensive classification of political parties according to a score ranging between 1 (extreme left) to 10 (extreme right). We computed a weighted average of the scores obtained by all parties forming the executive in each country-year, with weights equal to the number of parliamentary seats held by each party over the

[9] In general, initial conditions are key in almost any political economy model of reform; see, for instance, the discussion in chapter 13 of Drazen (2002).

[10] Stock variables such as the value of State-owned assets or central government debt would certainly provide better proxies for initial conditions. Unfortunately, complete time series on debt and comparable data on State assets in OECD countries are still missing.

[11] Empirical studies of partisan political economy usually rely upon dummy or discrete variables, with very limited methodological refinement since the seminal work by Hibbs (1977).

number of seats held by the executive as a whole:

$$PARTISAN = \sum_{j \in G} \left(\frac{s_j}{\sum_{k \in G} s_k} \right) HI_j, \tag{11}$$

where HI_j is the score attached by Huber and Inglehart (1995) to the j-th party, G is the set of parties forming the government and s_j is defined as in (9) .

The GDP per capita is included as well in all the specifications since, even within our sample of OECD countries, it is still present some heterogeneity in terms of economic development, which could possibly play a role in the start of the reform. Finally, we will check the robustness of the results to the inclusion of further variables that previous work has found to be relevant for privatization: the stage of financial market development, which plays an important role since deep and liquid stock markets, as measured by $MKTCAP$ and $TURNOVER$ respectively, facilitate the flotation of large companies; and the set of legal origins ($COMMON$, $GERMAN$, $FRENCH$ and $SCANDINAVIAN$ LAW) by La Porta et al. (1998).[12]

4.4 Descriptive analysis

Table 1 presents summary statistics which are useful for a first account of the role of political institutions in the timing of privatization. The column $POLFRAG$ reports the cross country average of the three measures of political fragmentation, standardized for the whole sample.[13] It takes the lowest values in the Anglo-Saxon countries and France. Interestingly, almost all of the countries in this group (with the exception of Australia) were among the few ones raising median revenues within the 80's (the only other one, out of the group, is Japan). On the contrary, privatization has been long delayed in highly fragmented political systems such as Switzerland, Belgium and Finland.

Table 2 provides more systematic evidence. The first two columns report the average values of the explanatory variables for early and late reformers, defined as the first and last five countries, respectively, to raise revenues above the median. The third and fourth columns report the difference between the two and its t-statistic, respectively. The results

[12] For econometric evidence about the role of financial markets and legal origins, see Bortolotti, Fantini and Siniscalco (2003) and Bortolotti and Siniscalco (2004).

[13] The index $DISPR$ enters with a negative sign, in order to be consistent with $GENP$ and $PENP$, which are increasing political fragmentation.

reported for the control variables resemble those obtained in previous empirical studies of privatization, indicating a role for macroeconomic variables, legal origins, ideology and (to lesser extent) financial markets indicators in explaining also the timing of privatization. Political fragmentation appears to be the novelty: early privatizing are less politically fragmented democracies. The difference is highly statistically significant for all the three measures. This preliminary evidence suggests the potential explanatory power of political fragmentation, which we test extensively in the next section by estimating the econometric model (2) -(6) .

5 Results

Tables 3, 4 and 5 show the estimation results for the Cox, Weibull and Lognormal models, respectively. The specifications we present are the same for all the three models. A benchmark equation in column [1] of all tables includes *SOE*, *DEFICIT*, *GDP* and *PARTISAN*. Columns [2]-[4] add the political fragmentation variables *GENP*, *PENP* and *DISPR*. They are never included together in the same regression since they all proxy for the same theoretical variable, namely political fragmentation, which would make it hard to disentangle their distinct effects. Finally, columns [5]-[7] and [8]-[10] check the robustness of the results to the inclusion of the financial markets variables and legal origins, respectively. We start by discussing the information conveyed by the proportional hazard models (Cox and Weibull), since it is most easily interpretable, especially in terms of marginal effects of changes in the explanatory variables. Then, we will check the robustness of the results as we relax the proportionality assumption.

The first conclusion we can draw by looking at tables 3 and 4 is that well-established economic determinants of the extent of privatization (for instance in terms of total revenues) fail instead to account for the timing of the reform. The univariate correlations found in the descriptive statistics do not survive in the multivariate analysis, which yields unstable and statistically not significant point estimates of the coefficients of all the economic controls. On the other hand, the *PARTISAN* index is strongly significant in any specification and apparently controls for an important effect that the ideology of the executive exerts on the start of the reform.

Turning to the measures of political fragmentation, they show considerable explanatory power. First, they are always statistically significant at conventional levels in the benchmark specifications and the significance is robust to the inclusion of the financial markets variables. When we additionally control for the legal origins, the t-ratios lower considerably and in one case *DISPR* falls slightly short of the 10% significance level (its t-ratio in the Weibull model is 1.50). Notice that the inclusion of legal origins represents an important (and severe) robustness check for our variables of interest, since legal origins and political fragmentation go hand by hand for most of the countries in our sample. Yet, point estimates for the coefficients of legal origins are extremely unstable across specifications. Moreover, the null that they equal 0 can not be rejected in most of the equations. On the contrary, estimates of the effect of political fragmentation, while made somewhat noisier by the inclusion of legal origins, remain overall statistically significant.

Second, the absolute values of the estimated coefficients of *GENP*, *PENP* and *DISPR* are reasonably stable among the semi-parametric (Cox) and the parametric (Weibull) model (once again, the only exceptions come with the inclusion of the legal origins, which seems to reinforce considerably, in absolute value, the effect of *GENP* and *PENP* in the Cox model). This is reassuring about the specification of the functional form for the parametric model. We re-estimated the model for different specifications of the dynamics as well, by introducing lags and leads of potentially endogenous variables (namely *MKTCAP* and *TURNOVER*). Such changes do not affect results at all; this leads us to exclude both the existence of significant simultaneity bias and possible misspecifications of the dynamics.

Third, the estimated effect of our measures of political fragmentation is economically, other than statistically, significant. In particular, let us consider the effect of adding one (effective) party either to the government or to the parliament. We focus on such a unit increase in GENP and PENP because it is close to the sample standard deviation of both variables (0.848 and 1.247 respectively) and, further, because it relates to some very concrete feature of the political equilibrium (i.e. how many "important" parties enter the government or the parliament). These changes imply a reduction in the hazard rate of between 52 and 70 percent (according to Weibull and Cox estimates, respectively) if the additional party enters the executive, and of something more (62 to 72 percent) if it gains representation in the parliament.[14]

[14] The percentage change in the hazard rate is computed as $\Delta(\beta^k) = 100\left(\exp\beta^k - 1\right)$. To see this, consider two vectors of covariates x and $x' = x + e^k$, where e^k is the unit vector having its k-th element equal to 1 and all the other ones equal to 0. Recalling the proportionality assumption (3) and equation (5) $\frac{\lambda'}{\lambda} = \exp\left[(x' - x)\prime\beta\right] = \exp\beta^k$, which implies $100\left(\frac{\lambda'-\lambda}{\lambda}\right) = 100\left(\exp\beta^k - 1\right)$, where $100\left(\frac{\hat{\lambda}-\lambda}{\lambda}\right)$ is

Finally, to get more a sense of what these numbers imply, in Figure 3 the hazard rate and the survival function $S(T \mid x) = 1 - F(T \mid x)$ estimated for the parametric model are plotted against the delay to reform. Figure 3a plots the hazard rates when *GENP* is equal to 1 and 2.85, which are the two averages for the clusters of cabinet and consensus countries, respectively, identified in Figure 2; Figure 3b does the same for the survival functions; finally, bottom figures 3c and 3d repeat the exercise by considering average values of *PENP*.

The qualitative and quantitative differences between the two cases, cabinet and consensus, are striking. The hazard rate is always increasing because of $\widehat{\alpha} > 1$. However, in the cabinet system the increase in the slope is noticeable already after about 5 years, which is at the beginning of the 80s, and reform becomes almost sure in the early 90s, according to both *GENP* and *PENP* estimates of political fragmentation. On the contrary, in the consensus system privatization never gains momentum. Indeed, according to *GENP* estimates, by the end of the sample period the predicted hazard rate is still below 40%; as a consequence, 25% of consensus countries should have not reformed yet. Measuring political fragmentation by *PENP* entails even greater difference between the two cases.

The results of both Cox and Weibull estimates strongly support the empirical implications of the war of attrition model. The data show that indeed greater political fragmentation entails longer delays to implement large scale divestiture. We now check further the robustness of these results by relaxing the proportionality assumption (3) . Table 5 shows the results for the Lognormal model. The estimates for the coefficients of the variables of interest remain always statistically significant at conventional levels. Moreover, their absolute value is always close to the ratio $-(\beta/\alpha)$ estimated by the Weibull model. Thus, while proportionality of the hazard rate is convenient both in terms of tractability and interpretation of the results, the qualitative and quantitative results we discussed above do not rely heavily on such assumption.

nothing else than the percentage change in the hazard rate implied by increasing the k-th element of x by one unit (one party, in our case).

6 Conclusions

Political economy has recently provided several models to understand the determinants of economic reform. Yet empirical analysis on this topic faces severe measurement problems in finding suitable variables to gauge the economic relevance of reform processes and to link political-economic equilibria to factual institutional settings.

The big privatization wave that started in the United Kingdom in the late 70s, swept the world in the last two decades, and declined abruptly right after the turn of the century provides an ideal experiment to analyze empirically the timing of large-scale reform. Importantly, research is not limited by data availability given that reliable sources provide comprehensive information on privatization processes both across countries and overtime.

War of attrition models suggest that political fragmentation is a fundamental factor in explaining the timing of reform. In particular, these models posit a positive relationship between the delay of reform and the number of agents with veto power in a given political system.

In this paper, we first identify a formal link between the theoretical war of attrition model and the statistical duration model, then we study the delay of privatization in a large sample of developed countries over the 1977-2002 period. Our results confirm the empirical validity of war of attrition model: large scale divestitures is delayed longer the larger the number of parties and the greater the proportionality of the electoral rule.

The estimated coefficients of these variables are significant and robust across different specifications. Moreover, the hazard rates predicted by the model, conditional on our proxies for political fragmentation, generate expected delays of privatization that are consistent with those observed in reality in more versus less fragmented democracies.

References

[1] Alesina, A. and A. Drazen (1991). Why are stabilizations delayed?, *American Economic Review*, 81, 1170-88.

[2] Alesina, A., R. Perotti and J. Tavares (1998). The Political Economy of Fiscal Adjustments. *Brookings Papers on Economic Activity*, Vol. 1998, No. 1., 197-266.

[3] Banks, A.S., T.C. Day and W.R. Muller (2002), *Political Handbook of the World 2000-2002* - CSA publications, State University of New York.

[4] Beck, T., G. Clarke, A. Groff, P. Keefer, and P. Walsh (2001). New tools in comparative political economy: The Database of Political Institutions, *World Bank Economic Review*, 15, 165-176.

[5] Beck, T., A. Demirgüç-Kunt, and R. Levine (1999) A New Database on Financial Development and Structure (updated 2003), World Bank Policy Research Working Paper n. 2146.

[6] Boehmer, E., R. C. Nash and J. M. Netter (2005) Bank privatization in developing and developed countries: Cross-sectional evidence on the impact of economic and political factors, *Journal of Banking and Finance*, 29, 1981-2013.

[7] Bortolotti, B., M. Fantini and D. Siniscalco (2003). Privatization around the world: evidence from panel data, *Journal of Public Economics*, 88, 305-32.

[8] Bortolotti, B. and D Siniscalco (2004). *The challenges of privatization: an international analysis.* Oxford, New York, Oxford University Press

[9] Cox, D. R. (1972). Regression Models and Life Tables, *Journal of the Royal Statistical Society*, Series B 24, 187-220.

[10] Clarke G. and R. Cull (2002). Political and economic determinants of the likelihood of privatizing argentine public banks, *Journal of Law and Economics*, 45, 165-197.

[11] Dinc, S. and N. Gupta (2005). The Decision to Privatize: The Role of Political Competition and Patronage, mimeo, Northwestern University

[12] Drazen, A. (2002). *Political Economy in Macroeconomics*, Cambridge (MA), The MIT Press.

[13] Gallagher, M. (1991). Proportionality, disproportionality and electoral systems. *Electoral Studies*, 10, 33-51.

[14] Haltinwanger, J. And M. Singh (1999). Cross-country evidence on public sector retrenchement, *World Bank Economic Review*, 13:23-66

[15] Hibbs, D. (1977). Political parties and macroeconomic policy. *American Political Science Review*, 71, 1467-78

[16] Huber, G., M. Kocher and M. Sutter (2003). Government Strength, Power Dispersion in Governments and Budget Deficits in OECD-Countries. A Voting Power Approach, *Public Choice*, 116, 333 - 350

[17] Huber, J. and R. Inglehart (1995). Expert interpretations of party space and party locations in 42 societies, *Party Politics*, 1, 73-111.

[18] Kiefer, N.M. (1988) Economic duration data and hazard functions, *Journal of Economic Literature*, 26, 646–679.

[19] Laakso, M. and R. Taagepera (1979). Effective number of parties: a measure with application to West Europe. *Comparative Political Studies* 12: 3-27

[20] La Porta, R., F. Lopez-de-Silanes, A. Shleifer, R. W. Vishny (1998). Law and Finance, *Journal of Political Economy*, 106, 1113-1155.

[21] Lijphart, A. (1999). *Patterns of democracy.* New Haven and London, Yale University Press

[22] Megginson, W. L., Nash, R. C. and M. van Randenborgh (1994) The financial and operating performance of newly privatized firms: An international empirical analysis, *Journal of Finance*, 49, 403-452

[23] Megginson, W.L. and J. Netter (2001). From state to market: a survey of empirical studies on privatization, *Jounal of Economic Literature*, 39, 321-389.

[24] Padovano, F. and L. Venturi (2001). Wars of Attrition in Italian Government Coalitions and Fiscal Performance: 1948–1994, *Public Choice*, 109, 15 - 54

[25] Perotti, R., Y. Kontopoulos (2002) Fragmented Fiscal Policy, *Journal of Public Economics*, 82, pp. 191-222.

[26] Persson, T, Roland, G., and G. Tabellini (2007). How do electoral rules shape party structures, government coalitions, and economic policies?, *Quarterly Journal of Political Science*, forthcoming.

[27] Persson, T. and G. Tabellini (2000). *Political Economics*, Cambridge (MA), The MIT Press.

[28] Roubini, N., and J, Sachs (1989) Political and economic determinants of budget deficits in the industrial democracies, *European Economic Review*, 33, 903-33.

[29] Spolaore, E. (2004). Adjustments in Different Government Systems, *Economics and Politics*, 16(2).

[30] Tsebelis, G. (2002). *Veto Players: How Political Institutions Work*, Princeton University Press

[31] Van den Berg, G.J. (2001). Duration models: Specification, identification, and multiple durations, in J.J. Heckman and E. Leamer, editors, *Handbook of Econometrics*, Volume V, North Holland, Amsterdam.

[32] Velasco, A. (1997). Debts and Deficits with Fragmented Fiscal Policymaking. *NBER* Working Paper No. W6286.

[33] Vickers, J., and Yarrow, G. (1988). *Privatization: an economic analysis*, Cambridge (MA), MIT Press.

[34] Weibull (1951) Statistical distribution function of wide applicability. *Journal of Applied Mechanisms*,18, 293-302

[35] Woo, J. (2003). Economic, political, and institutional determinants of public deficits, *Journal of Public Economics*, 87, 387-426.

[36] World Bank (1995). *Bureaucrats in business*, Oxford, Oxford University Press.

Appendix A. Description of Variables and Sources

Variable	Definition	Source
DEFICIT	Central government deficit as percentage of GDP	International Financial Statistics
DISPR	Disproportionality index computed over the difference between the shares of votes and seats held by each party. Mathematical formulation of the index is presented in section 3.	Original dataset from Lijphart (1999) updated using the review *Electoral Studies*, various issues; Banks et al. (2002); Elections Around the World (www.electionworld.org)
GDP	Ratio of Gross Domestic Product in constant 1996 US Dollars to population. Total population counts all residents regardless of legal status or citizenship.	World Development Indicators, World Bank, International Financial Statistics
GENP	Concentration index computed over government parties' seats in the legislative chamber. Mathematical formulation of the index is presented in section 3.	*Electoral Studies*, various issues; Banks et al. (2002); Elections Around the World (www.electionworld.org);
MKTCAP	Stock market capitalization to Gross Domestic Product. Stock market capitalization in year *t* is calculated as the average between the end-of-year market capitalization deflated by the end-of-year Consumer Price Index in year *t* and *t-1*. Stock market capitalization refers to a country's main stock exchange.	Beck, Demirgüç-Kunt, and Levine (1999)
PARTISAN	Government's ideology. It is computed as the weighted average of the ideology attached by Huber and Inglehart to parties forming the government coalition. Mathematical formulation of the index is presented in section 3.	(same as *GENP*)
PENP	Concentration index computed over distribution of parties' seats in the parliament. Mathematical formulation of the index is presented in section 3.	(same as *DISPR*)
REVGDP	Total revenues from privatization (*Public Offers* and *Private Sales*) as a fraction of GDP.	Securities Data Corporation, Privatization Barometer database (http://www.privatizationbarometer.net /database.php)
SOE	Average of the SOE value added as a percentage of GDP in the three years preceding the first privatization reported in SDC.	World Bank (1995), World Development Indicators
T	Delay to reform. Defined for each country as T=t-1977, where $t=\min\{t: REVGDP_t \geq median[REVGDP_s]\ s=1977,\ldots,2002\}$	
TURNOVER	Stock market total value traded to total market capitalization. Total market value in year *t* is deflated by the Consumer Price Index in year *t*. Market capitalization in year *t* is calculated as the average between the end-of-year market capitalization deflated by the end-of-year Consumer Price Index in year *t* and *t-1*. Trading value and market capitalization refer to a country's main stock exchange.	Beck, Demirgüç-Kunt, and Levine (1999)
COMMON LAW	Dummy taking value 1 for Common Law countries	La Porta et al. (1998)
FRENCH LAW	Dummy taking value 1 for French Law countries	
GERMAN LAW	Dummy taking value 1 for German Law countries	
SCAND. LAW	Dummy taking value 1 for Scandinavian Law countries	

Appendix B. Comparison of Political Datasets

This table presents a comparisons between the electoral data reported in the FEEM Political Dataset (http://www.feem.it/fpd), the World Bank DPI by Beck et al. (2001) and Tsebelis (2002). Column OBS reports the number of common observations (i.e., elections) between two datasets. SEATS DIFF is the average difference between the number of seats reported for, respectively, the first, second and third party forming the executive, and for the government as a whole. % MATCHED SEATS is the percentage of cases in which the number of seats coincides exactly for two datasets.

			Tsebelis									**FPD**								
			OBS	SEATS DIFF				% MATCHED SEATS				OBS	SEATS DIFF				% MATCHED SEATS			
				P1	P2	P3	GOV	P1	P2	P3	GOV		P1	P2	P3	GOV	P1	P2	P3	GOV
FPD	OBS		126																	
	SEATS DIFF	P1		3.73																
		P2			2.76															
		P3				1.81														
		GOV					6.36													
	% MATCHED SEATS	P1						80.16												
		P2							81.75											
		P3								84.13										
		GOV									66.67									
World Bank DPI	OBS		103									109								
	SEATS DIFF	P1		28.58									29.64							
		P2			13.85									13.30						
		P3				4.21									5.01					
		GOV					30.79									30.46				
	% MATCHED SEATS	P1						4.85									4.59			
		P2							43.69									44.95		
		P3								66.99									66.06	
		GOV									2.91									4.59

Table 1. Political Data and Timing

This table presents cross-country averages of the political fragmentation indices over the period 1977-2002. The column *POLFRAG* reports the standardized average of the three measures (*DISPR* enters with negative sign in order to be increasing in the degree of political fragmentation, like *GENP* and *PENP*). In countries where an institutional reform occurred, averages are computed over the two sample periods defined by the first election under the new regime. Finally, the last column reports *T*, defined as the number of years between the first privatization ever (in 1977) and the year in which median (per country) privatization revenues are observed.

Countries	*GENP*	*PENP*	*DISPR*	*POLFRAG*	*T*
Australia	1.249	2.427	10.803	-0.759	18
Austria	1.636	2.800	1.679	0.169	14
Belgium	2.456	4.793	3.721	1.104	16
Canada	1.000	2.350	13.642	-1.101	10
Denmark	1.776	4.870	1.495	0.935	17
Finland	2.959	5.111	3.354	1.486	16
France	1.519	3.330	24.749	-1.259	10
Germany	1.357	2.661	2.204	-0.053	17
Greece	1.028	2.231	7.699	-0.728	15
Ireland	1.309	2.882	4.370	-0.149	22
Italy (-94)	1.898	3.955	3.505	0.560	17
Italy (94-)	3.278	6.267	7.111	1.777	
Japan (-96)	1.146	2.990	6.087	-0.310	11
Japan (96-)	1.084	3.147	8.779	-0.470	
Netherlands	2.221	4.321	1.316	0.990	14
New Zealand (-96)	1.000	1.965	14.858	-1.309	12
New Zealand (96-)	1.333	3.744	4.440	0.143	
Norway	1.467	3.432	6.889	-0.056	17
Portugal	1.103	2.993	4.589	-0.231	19
Spain	1.000	2.723	7.750	-0.583	16
Sweden	1.524	3.666	1.841	0.387	17
Switzerland	3.779	5.562	3.081	2.068	21
United Kingdom	1.000	2.173	14.968	-1.248	9
United States	1.000	1.940	15.538	-1.363	11
Average	1.630	3.431	7.270		15.19
Std. Dev.	0.776	1.187	5.848		3.614

Table 2. Univariate Tests

This table reports: the average of each explanatory variable over all observations for the countries at the bottom and top quartile of the distribution of *T (Delay of Privatization)*; the difference between the two averages; finally, the t-statistic of the null hypothesis of the difference being significantly different from 0. ***, ** and * denote statistical significance at 1, 5 and 10 % level, respectively.

	T (Delay of Privatization)			
	Bottom 25%	Top 25%	Difference	t-statistic
SOE	11.248	14.083	-2.835 ***	-3.70
DEFICIT	3.885	7.256	-3.370 ***	-8.60
GDP	23192	19059	4133 ***	5.54
PARTISAN	6.403	5.670	0.733 ***	4.65
GENP	1.130	1.963	-0.834 ***	-7.85
PENP	2.565	3.724	-1.159 ***	-8.17
DISPR	15.142	5.519	9.623 ***	13.92
TURNOVER	0.531	0.607	0.077	-1.07
CAPMKT	0.694	0.464	0.230 ***	4.89
COMMON LAW	0.600	0.400	0.200 ***	3.04
FRENCH LAW	0.200	0.400	-0.200 ***	3.04
GERMAN LAW	0.200	0.200	0.000	0.00

Table 3. Duration analysis: Cox model

This table reports estimated coefficients and associated t-statistics (in parentheses) for the Cox model. The dependent variable is the hazard rate of observing median (per country) privatization revenues. The maximum value of the log-likelihood function is reported below. The Wald Test refers to the null hypothesis of all the coefficients being jointly equal to 0. ***, ** and * denote statistical significance at 1, 5 and 10% confidence level, respectively.

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
SOE	-0.065	-0.074	-0.078	-0.082	-0.075	-0.089	-0.104*	0.093	0.048	-0.090
	(-0.96)	(-1.05)	(-1.13)	(-1.3)	(-1.16)	(-1.36)	(-1.69)	(1.41)	(0.43)	(-0.71)
DEFICIT	0.027	0.019	0.067	0.124**	-0.012	0.002	0.173**	-0.055	0.008	0.184
	(0.42)	(0.26)	(0.8)	(2.02)	(-0.07)	(0.01)	2.23	(-0.45)	(0.05)	0.80
GDP	1e-05	4.1e-05	7.8e-05*	4.7e-05*	1.4e-05	8.2e-05*	2.2e-05	0.0004***	0.0004***	2.1E-05
	(0.29)	(1.11)	(1.83)	(1.55)	(0.38)	(1.79)	(0.62)	(3.48)	(2.76)	(0.17)
PARTISAN	0.330**	0.442***	0.350***	0.575***	0.344***	0.283**	0.672***	0.917**	0.653*	0.661***
	(2.48)	(4.45)	(2.87)	(5.22)	(2.69)	(2.18)	(2.69)	(2.38)	(1.91)	(4.72)
GENP		-0.910**			-0.856**			-1.930***		
		(-2.2)			(-2.09)			(-3.8)		
PENP			-0.958**			-1.125**			-1.742*	
			(-2.15)			(-2.4)			(-1.85)	
DISPR				0.275***			0.430***			0.433*
				(4.42)			(2.73)			(1.76)
TURNOVER					-0.098	-0.900	1.148**	-1.112	-0.371	-0.779
					(-0.2)	(-1.45)	(2.29)	(-1.12)	(-0.32)	(-0.30)
MKTCAP					2.081	1.625	3.559*	-3.263*	-2.108	1.543
					(1.16)	(0.93)	(1.7)	(-1.65)	(-1.14)	(0.85)
COMMON LAW								1.706	0.523	3.732
								(1.53)	(0.27)	(1.44)
FRENCH LAW								-1.386	-2.865*	2.225
								(-1.51)	(-1.68)	(0.55)
SCAND. LAW								-5.609***	-4.678**	0.239
								(-4.46)	(-2.03)	(0.07)
Obs.	209	209	209	209	192	192	192	192	192	192
Log likelihood	-30.59	-29.1	-28.41	-25.27	-25.15	-24.31	-20.34	-19.59	-20.56	-18.21
Wald Test	6.86	36.59***	22.42***	42.65***	24.25***	24.26***	19.56***	67.22***	63.12***	208.55***

Table 4. Duration analysis: Weibull model

This table reports estimated coefficients and associated t-statistics (in parentheses) for the Weibull model. The dependent variable is the hazard rate of observing median (per country) privatization revenues. The maximum value of the log-likelihood function is reported below. The Wald Test refers to the null hypothesis of all the coefficients being jointly equal to 0. ***, ** and * denote statistical significance at 1, 5 and 10% confidence level, respectively.

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
CONSTANT	-18.570***	-18.802***	-18.662***	-29.793***	-19.272***	-18.498***	-30.800***	-30.684***	-29.355***	-57.558**
	(-3.95)	(-3.68)	(-3.34)	(-3.90)	(-3.31)	(-3.05)	(-3.55)	(-6.53)	(-5.85)	(-2.20)
SOE	-0.054	-0.045	-0.055	-0.078	-0.069	-0.086*	-0.107*	-0.007	-0.077	-0.338
	(-1.28)	(-1.19)	(-1.36)	(-1.61)	(-1.43)	(-1.67)	(-1.93)	(-0.05)	(-0.43)	(-1.40)
DEFICIT	0.027	0.028	0.069	0.141*	-0.008	0.008	0.140	0.083	0.138	0.565
	(0.38)	(0.38)	(0.87)	(1.86)	(-0.05)	(0.04)	(0.77)	(0.43)	(0.59)	(1.24)
GDP	7.5e-06	3.7e-05	7.7e-05*	4.6e-05	-8e-07	6.1e-05	9.2e-07	0.0001(7	0.00019	-0.0001
	(-0.21)	(0.98)	(1.7)	(1.41)	(-0.03)	(1.11)	(0.02)	(1.08)	(1.37)	(-0.73)
PARTISAN	0.312**	0.399***	0.330***	0.561***	0.282**	0.206*	0.468***	0.723**	0.622**	1.042**
	(2.07)	(3.17)	(2.58)	(3.52)	(2.37)	(1.66)	(2.54)	(2.42)	(2.38)	(2.14)
GENP		-0.688*			-0.720**			-0.802*		
		(-1.85)			(-2.16)			(-1.73)		
PENP			-0.846**			-1.043**			-1.056*	
			(-2.19)			(-2.05)			(-1.73)	
DISPR				0.283***			0.303***			0.654
				(4.05)			(4.04)			(1.50)
TURNOVER					0.090	-0.646	1.023	1.306	1.484	2.306
					(0.16)	(-0.73)	(1.16)	(0.83)	(1.04)	(1.04)
MKTCAP					3.146**	3.163**	2.490*	1.047	1.596	1.502
					(2.34)	(2.19)	(1.83)	(0.49)	(0.66)	(0.71)
COMMON LAW								2.900	2.590	7.427*
								(1.34)	(1.20)	(1.67)
FRENCH LAW								-1.960	-2.656	2.222
								(-0.71)	(-1.08)	(0.57)
SCAND. LAW								-2.837	-2.231	3.436
								(-0.99)	(-0.75)	(0.85)
α	6.430	6.455	6.867	9.112	7.034	7.473	9.718	8.948	9.446	18.208
Obs.	209	209	209	209	192	192	192	192	192	192
Log likelihood	4.36	5.37	6.19	10.99	6.46	7.46	11.96	11.40	11.50	18.57
Wald Test	5.57	10.80**	15.64***	28.380***	33.04***	36.66***	63.300***	677.15***	183.61***	54.770***

Table 5. Duration analysis: Lognormal model

This table reports estimated coefficients and associated t-statistics (in parentheses) for the Lognormal model. The dependent variable is the log of the delay between the first privatization ever (in 1977) and the year in which median (per country) privatization revenues are observed. The maximum value of the log-likelihood function is reported below. The Wald Test refers to the null hypothesis of all the coefficients being jointly equal to 0. ***, ** and * denote statistical significance at 1, 5 and 10% confidence level, respectively.

	[1]	[2]	[3]	[4]	[5]	[6]	[7]	[8]	[9]	[10]
CONSTANT	3.001***	2.853***	2.660***	3.311***	2.740***	2.549***	3.164***	3.454*	3.234***	3.151***
	(11.83)	(12.80)	(10.31)	(13.15)	(9.70)	(9.51)	(10.67)	(11.17)	(9.19)	(20.99)
SOE	0.011	0.013*	0.012	0.011	0.017*	0.015*	0.013*	0.001	0.003	0.014**
	(0.90)	(1.67)	(1.52)	(1.62)	(1.80)	(1.85)	(1.75)	(0.12)	(0.29)	(2.29)
DEFICIT	0.007	0.006	-0.001	-0.101	0.003	-0.002	-0.015	-0.010	-0.014	-0.023***
	(0.54)	(0.39)	(-0.10)	(-1.13)	(0.17)	(-0.08)	(-1.55)	(-1.31)	(-1.58)	(-3.32)
GDP	-0.00001	-7.72e-06	-0.00001	-5.37e-06	-2.01e-06	0.00001	8.97e-07	-2e-05	-3e-05*	2.43e-06
	(-0.42)	(-0.72)	(-1.30)	(-0.85)	(-0.23)	(-1.24)	(0.21)	(-1.57)	(-1.76)	(0.28)
PARTISAN	-0.088**	-0.109***	-0.091***	-0.083***	-0.091**	-0.085**	0.051**	-0.102***	-0.091***	-0.062***
	(-2.54)	(-2.89)	(-3.24)	(-3.16)	(-1.96)	(-2.32)	(-2.03)	(-2.98)	(-2.94)	(-3.58)
GENP		0.216**			0.182*			0.105**		
		(1.96)			(1.80)			(2.42)		
PENP			0.183**			0.183***			0.127*	
			(2.32)			(2.66)			(1.88)	
DISPR				-0.037***			-0.035***			-0.027***
				(-3.39)			(-4.12)			(-3.89)
TURNOVER					0.102	0.196	-0.081	-0.099	-0.116	-0.115
					(0.72)	(1.50)	(-1.03)	(-0.70)	(-0.92)	(-1.31)
MKTCAP					-0.467	-0.403	-0.469**	0.027	0.027	-0.180
					(-1.53)	(-1.57)	(-2.09)	(0.17)	(0.16)	(-1.51)
COMMON LAW								-0.397***	-0.283**	-0.356***
								(-3.21)	(-1.96)	(-4.51)
FRENCH LAW								0.253	0.362**	-0.012
								(1.84)*	(2.21)	(-0.09)
SCAND. LAW								0.397	0.376*	-0.020
								(2.00)**	(1.74)	(-0.12)
Obs.	209	209	209	209	192	192	192	192	192	192
Log likelihood	2.41	4.59	5.76	9.93	5.25	6.83	11.52	12.52	12.38	16.64
Wald Test	7.55	9.40*	14.41**	16.92***	7.99	17.33**	26.02***	508.86***	251.19***	618.44***

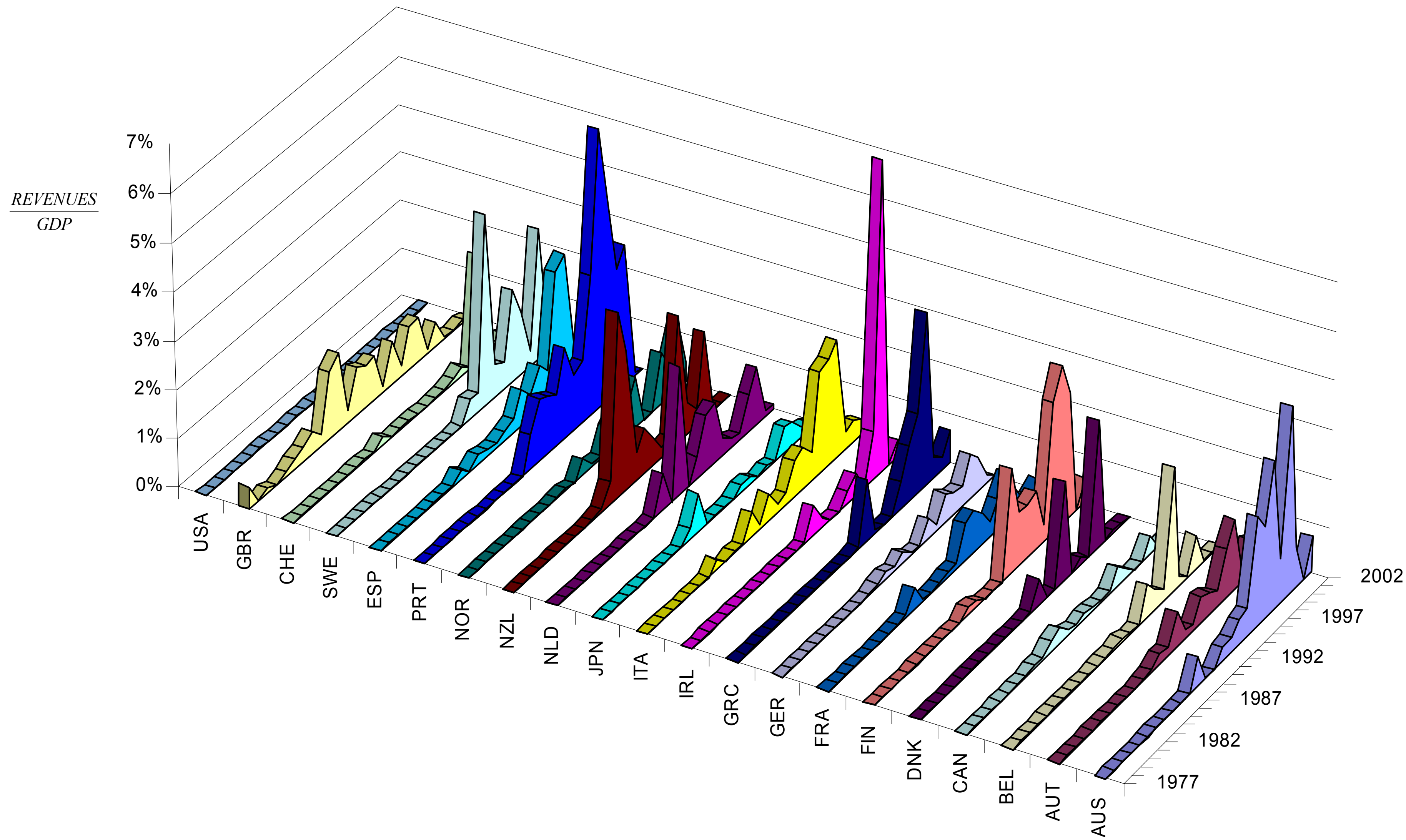

Figure 1. The Big Privatization Wave, OECD Countries

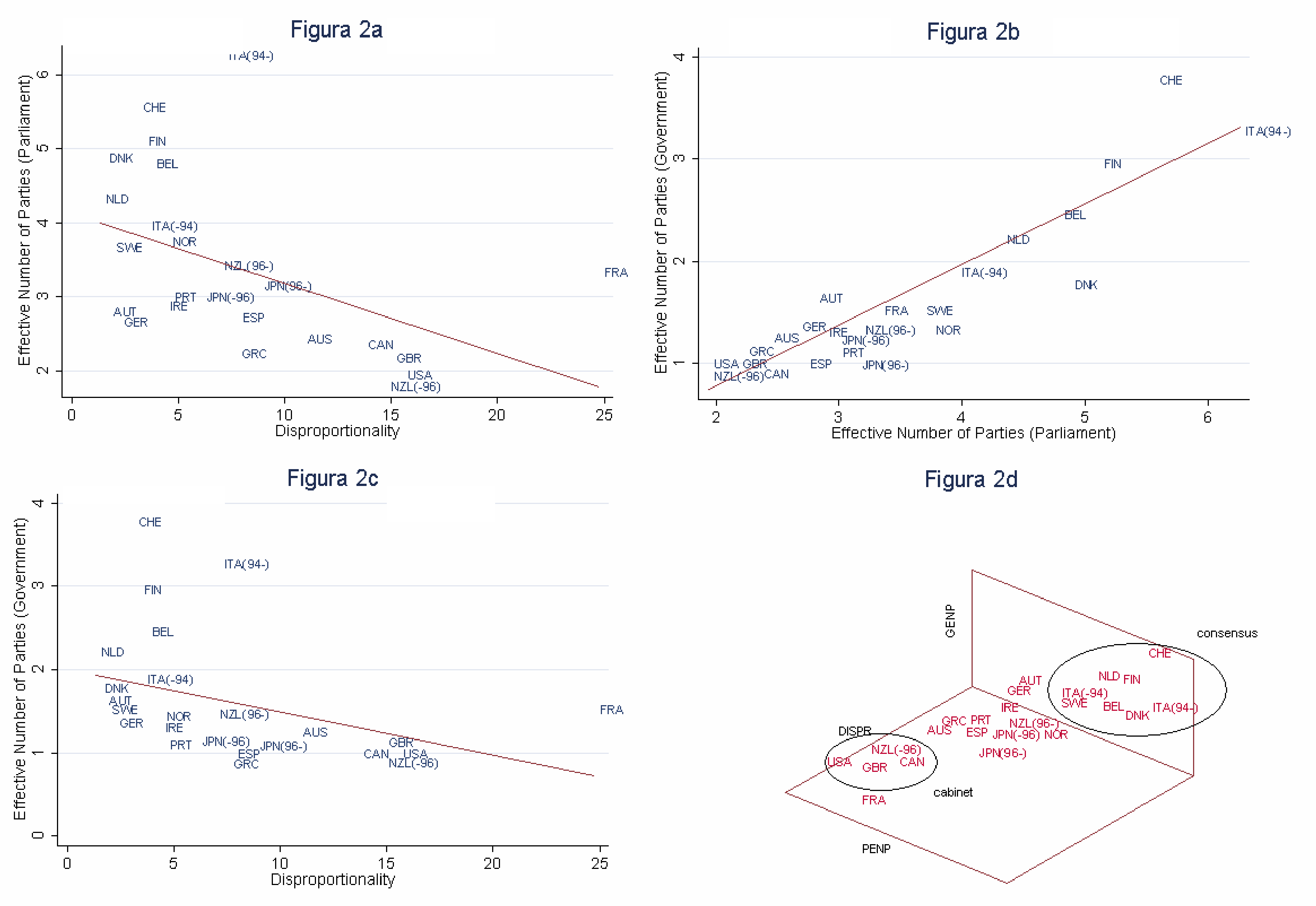

Figure 2. Measures of Political Fragmentation

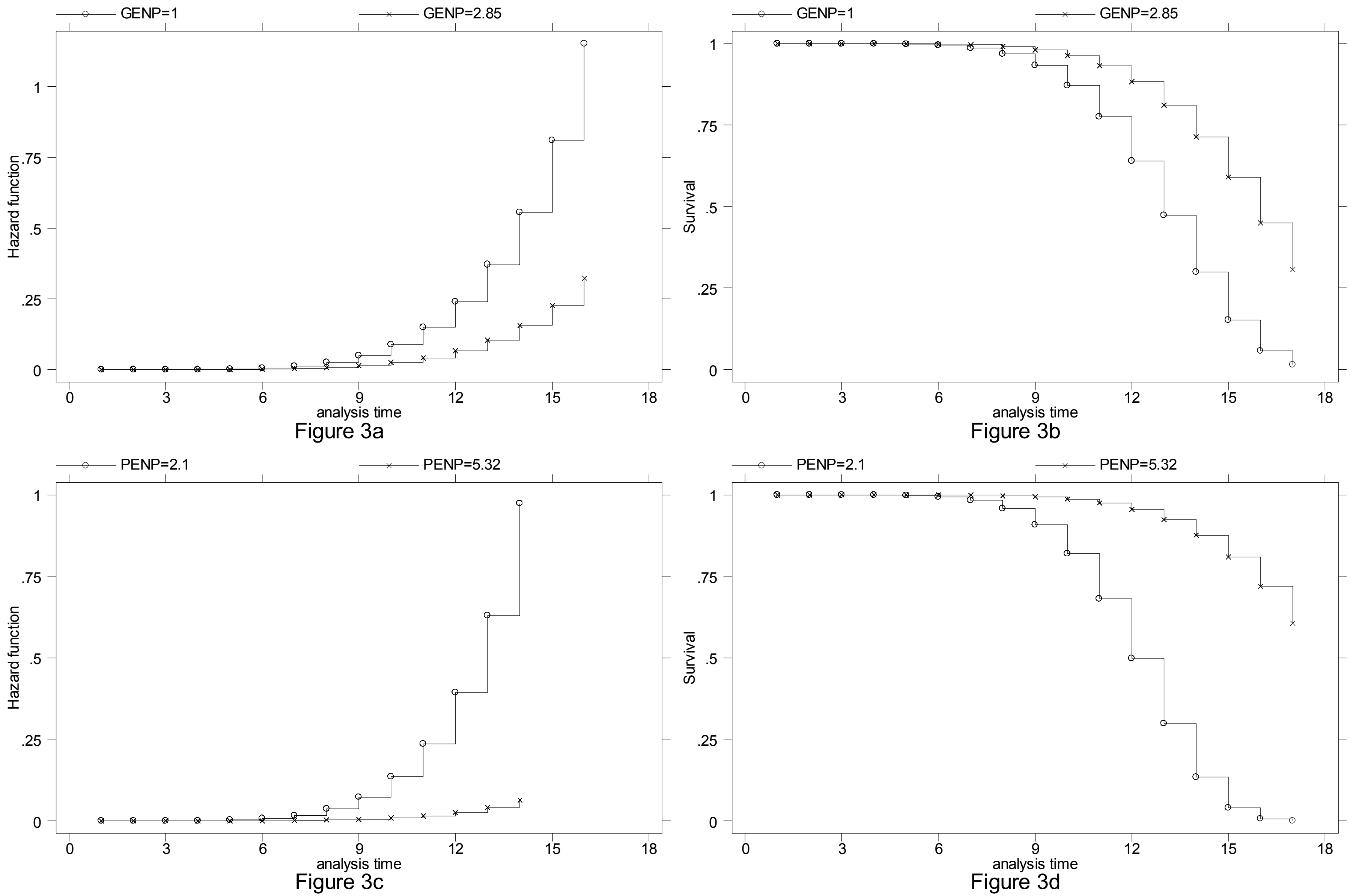

Figure 3. Hazard Rates and Survival Functions

RECENTLY PUBLISHED "TEMI" (*)

N. 637 – *Emerging markets spreads and global financial conditions*, by Alessio Ciarlone, Paolo Piselli and Giorgio Trebeschi (June 2007).

N. 638 – *Comparative advantage patterns and domestic determinants in emerging countries: An analysis with a focus on technology*, by Daniela Marconi and Valeria Rolli (September 2007).

N. 639 – *The generation gap: Relative earnings of young and old workers in Italy*, by Alfonso Rosolia and Roberto Torrini (September 2007).

N. 640 – *The financing of small innovative firms: The Italian case*, by Silvia Magri (September 2007).

N. 641 – *Assessing financial contagion in the interbank market: Maximum entropy versus observed interbank lending patterns*, by Paolo Emilio Mistrulli (September 2007).

N. 642 – *Detecting long memory co-movements in macroeconomic time series, by* Gianluca Moretti (September 2007).

N. 643 – *The producer service sector in Italy: Long-term growth and its local determinants*, by Valter Di Giacinto and Giacinto Micucci (September 2007).

N. 644 – *Aggregazioni bancarie e specializzazione nel credito alle PMI: peculiarità per area geografica*, by Enrico Beretta and Silvia Del Prete (November 2007).

N. 645 – *Costs and benefits of creditor concentration: An empirical approach*, by Amanda Carmignani and Massimo Omiccioli (November 2007).

N. 646 – *Does the underground economy hold back financial deepening? Evidence from the Italian credit market*, by Giorgio Gobbi and Roberta Zizza (November 2007).

N. 647 – *Optimal monetary policy under low trend inflation*, by Guido Ascari and Tiziano Ropele (November 2007).

N. 648 – *Indici di bilancio e rendimenti di borsa: un'analisi per le banche italiane*, by Angela Romagnoli (November 2007).

N. 649 – *Bank profitability and taxation,* by Ugo Albertazzi and Leonardo Gambacorta (November 2007).

N. 650 – *Modelling bank lending in the euro area: A non-linear approach,* by Leonardo Gambacorta and Carlotta Rossi (November 2007).

N. 651 – *Revisiting poverty and welfare dominance,* by Gian Maria Tomat (November 2007).

N. 652 – *The general equilibrium effects of fiscal policy: Estimates for the euro area,* by Lorenzo Forni, Libero Monteforte and Luca Sessa (November 2007).

N. 653 – *Securitisation and the bank lending channel,* by Yener Altunbas, Leonardo Gambacorta and David Marqués (November 2007).

N. 654 – *The cyclical response of fiscal policies in the euro area. Why do results of empirical research differ so strongly?*, by Roberto Golinelli and Sandro Momigliano (January 2008).

N. 655 – *What's behind "inflation perceptions"? A survey-based analysis of Italian consumers*, by Paolo Del Giovane, Silvia Fabiani and Roberto Sabbatini (January 2008).

N. 656 – *The effects of fiscal policy in Italy: Evidence from a VAR model*, by Raffaela Giordano, Sandro Momigliano, Stefano Neri and Roberto Perotti (January 2008).

N. 657 – *Excess money growth and inflation dynamics*, by Barbara Roffia and Andrea Zaghini (January 2008).

N. 658 – *R&D and market structure in a horizontal differentiation framework*, by Davide Fantino (January 2008).

N. 659 – *Housing market spillovers: Evidence from an estimated DSGE model*, by Matteo Iacoviello and Stefano Neri (January 2008).

N. 660 – *Real exchange rate volatility and disconnect: An empirical investigation*, by Riccardo Cristadoro, Andrea Gerali, Stefano Neri and Massimiliano Pisani (April 2008).

N. 661 – *The effect of investment tax credit: Evidence from an atypical programme in Italy*, by Raffaello Bronzini, Guido de Blasio, Guido Pellegrini and Alessandro Scognamiglio (April 2008).

N. 662 – *Accounting for sampling design in the SHIW*, by Ivan Faiella (April 2008).

(*) Requests for copies should be sent to:
Banca d'Italia – Servizio Studi di struttura economica e finanziaria – Divisione Biblioteca e Archivio storico – Via Nazionale, 91 – 00184 Rome – (fax 0039 06 47922059). They are available on the Internet www.bancaditalia.it.

2005

L. DEDOLA and F. LIPPI, *The monetary transmission mechanism: Evidence from the industries of 5 OECD countries,* European Economic Review, 2005, Vol. 49, 6, pp. 1543-1569, **TD No. 389 (December 2000)**.

D. Jr. MARCHETTI and F. NUCCI, *Price stickiness and the contractionary effects of technology shocks*. European Economic Review, Vol. 49, 5, pp. 1137-1164, **TD No. 392 (February 2001)**.

G. CORSETTI, M. PERICOLI and M. SBRACIA, *Some contagion, some interdependence: More pitfalls in tests of financial contagion,* Journal of International Money and Finance, Vol. 24, 8, pp. 1177-1199, **TD No. 408 (June 2001)**.

GUISO L., L. PISTAFERRI and F. SCHIVARDI, *Insurance within the firm*. Journal of Political Economy, Vol. 113, 5, pp. 1054-1087, **TD No. 414** (**August 2001**)

R. CRISTADORO, M. FORNI, L. REICHLIN and G. VERONESE, *A core inflation indicator for the euro area,* Journal of Money, Credit, and Banking, Vol. 37, 3, pp. 539-560, **TD No. 435 (December 2001)**.

F. ALTISSIMO, E. GAIOTTI and A. LOCARNO, *Is money informative? Evidence from a large model used for policy analysis,* Economic & Financial Modelling, Vol. 22, 2, pp. 285-304, **TD No. 445 (July 2002)**.

G. DE BLASIO and S. DI ADDARIO, *Do workers benefit from industrial agglomeration?* Journal of regional Science, Vol. 45, (4), pp. 797-827, **TD No. 453 (October 2002).**

G. DE BLASIO and S. DI ADDARIO, *Salari, imprenditorialità e mobilità nei distretti industriali italiani,* in L. F. Signorini, M. Omiccioli (eds.), *Economie locali e competizione globale: il localismo industriale italiano di fronte a nuove sfide*, Bologna, il Mulino, **TD No. 453 (October 2002).**

R. TORRINI, *Cross-country differences in self-employment rates: The role of institutions*, Labour Economics, Vol. 12, 5, pp. 661-683, **TD No. 459 (December 2002)**.

A. CUKIERMAN and F. LIPPI, *Endogenous monetary policy with unobserved potential output,* Journal of Economic Dynamics and Control, Vol. 29, 11, pp. 1951-1983, **TD No. 493 (June 2004)**.

M. OMICCIOLI, *Il credito commerciale: problemi e teorie*, in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 494 (June 2004)**.

L. CANNARI, S. CHIRI and M. OMICCIOLI, *Condizioni di pagamento e differenziazione della clientela,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 495 (June 2004)**.

P. FINALDI RUSSO and L. LEVA, *Il debito commerciale in Italia: quanto contano le motivazioni finanziarie?,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 496 (June 2004)**.

A. CARMIGNANI, *Funzionamento della giustizia civile e struttura finanziaria delle imprese: il ruolo del credito commerciale,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 497 (June 2004)**.

G. DE BLASIO, *Credito commerciale e politica monetaria: una verifica basata sull'investimento in scorte,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia,* Bologna, Il Mulino, **TD No. 498 (June 2004)**.

G. DE BLASIO, *Does trade credit substitute bank credit? Evidence from firm-level data.* Economic notes, Vol. 34, 1, pp. 85-112, **TD No. 498 (June 2004).**

A. DI CESARE, *Estimating expectations of shocks using option prices,* The ICFAI Journal of Derivatives Markets, Vol. 2, 1, pp. 42-53, **TD No. 506 (July 2004).**

M. BENVENUTI and M. GALLO, *Il ricorso al "factoring" da parte delle imprese italiane,* in L. Cannari, S. Chiri e M. Omiccioli (eds.), *Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia*, Bologna, Il Mulino, **TD No. 518 (October 2004)**.

L. CASOLARO and L. GAMBACORTA, *Redditività bancaria e ciclo economico,* Bancaria, Vol. 61, 3, pp. 19-27, **TD No. 519 (October 2004)**.

F. PANETTA, F. SCHIVARDI and M. SHUM, *Do mergers improve information? Evidence from the loan market*, CEPR Discussion Paper, 4961, **TD No. 521 (October 2004)**.

P. DEL GIOVANE and R. SABBATINI, *La divergenza tra inflazione rilevata e percepita in Italia,* in P. Del Giovane, F. Lippi e R. Sabbatini (eds.), *L'euro e l'inflazione: percezioni, fatti e analisi,* Bologna, Il Mulino, **TD No. 532 (December 2004).**

R. TORRINI, *Quota dei profitti e redditività del capitale in Italia: un tentativo di interpretazione*, Politica economica, Vol. 21, 1, pp. 7-41, **TD No. 551 (June 2005)**.

M. OMICCIOLI, *Il credito commerciale come "collateral"*, in L. Cannari, S. Chiri, M. Omiccioli (eds.), Imprese o intermediari? Aspetti finanziari e commerciali del credito tra imprese in Italia, Bologna, il Mulino, **TD No. 553 (June 2005)**.

L. CASOLARO, L. GAMBACORTA and L. GUISO, *Regulation, formal and informal enforcement and the development of the household loan market. Lessons from Italy*, in Bertola G., Grant C. and Disney R. (eds.) The Economics of Consumer Credit: European Experience and Lessons from the US, Boston, MIT Press, **TD No. 560 (September 2005)**.

S. DI ADDARIO and E. PATACCHINI, *Lavorare in una grande città paga, ma poco*, in Brucchi Luchino (ed.), *Per un'analisi critica del mercato del lavoro*, Bologna , Il Mulino, **TD No. 570 (January 2006)**.

P. ANGELINI and F. LIPPI, *Did inflation really soar after the euro changeover? Indirect evidence from ATM withdrawals,* CEPR Discussion Paper, 4950, **TD No. 581 (March 2006)**.

S. FEDERICO, *Internazionalizzazione produttiva, distretti industriali e investimenti diretti all'estero,* in L. F. Signorini, M. Omiccioli (eds.), *Economie locali e competizione globale: il localismo industriale italiano di fronte a nuove sfide*, Bologna, il Mulino, **TD No. 592 (October 2002).**

S. DI ADDARIO, *Job search in thick markets: Evidence from Italy*, Oxford Discussion Paper 235, Department of Economics Series, **TD No. 605 (December 2006)**.

2006

F. BUSETTI, *Tests of seasonal integration and cointegration in multivariate unobserved component models,* Journal of Applied Econometrics, Vol. 21, 4, pp. 419-438, **TD No. 476 (June 2003).**

C. BIANCOTTI, *A polarization of inequality? The distribution of national Gini coefficients 1970-1996*, Journal of Economic Inequality, Vol. 4, 1, pp. 1-32, **TD No. 487 (March 2004)**.

L. CANNARI and S. CHIRI, *La bilancia dei pagamenti di parte corrente Nord-Sud (1998-2000),* in L. Cannari, F. Panetta (a cura di), Il sistema finanziario e il Mezzogiorno: squilibri strutturali e divari finanziari, Bari, Cacucci, **TD No. 490 (March 2004).**

M. BOFONDI and G. GOBBI, *Information barriers to entry into credit markets,* Review of Finance, Vol. 10, 1, pp. 39-67, **TD No. 509 (July 2004).**

FUCHS W. and LIPPI F., *Monetary union with voluntary participation,* Review of Economic Studies, Vol. 73, pp. 437-457 **TD No. 512 (July 2004).**

GAIOTTI E. and A. SECCHI, *Is there a cost channel of monetary transmission? An investigation into the pricing behaviour of 2000 firms*, Journal of Money, Credit and Banking, Vol. 38, 8, pp. 2013-2038 **TD No. 525 (December 2004)**.

A. BRANDOLINI, P. CIPOLLONE and E. VIVIANO, *Does the ILO definition capture all unemployment?,* Journal of the European Economic Association, Vol. 4, 1, pp. 153-179, **TD No. 529 (December 2004)**.

A. BRANDOLINI, L. CANNARI, G. D'ALESSIO and I. FAIELLA, *Household wealth distribution in Italy in the 1990s,* in E. N. Wolff (ed.) International Perspectives on Household Wealth, Cheltenham, Edward Elgar, **TD No. 530 (December 2004)**.

P. DEL GIOVANE and R. SABBATINI, *Perceived and measured inflation after the launch of the Euro: Explaining the gap in Italy,* Giornale degli economisti e annali di economia, Vol. 65, 2 , pp. 155-192, **TD No. 532 (December 2004).**

M. CARUSO, *Monetary policy impulses, local output and the transmission mechanism*, Giornale degli economisti e annali di economia, Vol. 65, 1, pp. 1-30, **TD No. 537 (December 2004).**

A. NOBILI, *Assessing the predictive power of financial spreads in the euro area: does parameters instability matter?,* Empirical Economics, Vol. 31, 1, pp. 177-195, **TD No. 544 (February 2005)**.

L. GUISO and M. PAIELLA, *The role of risk aversion in predicting individual behavior, In P. A. Chiappori e C. Gollier (eds.) Competitive Failures in Insurance Markets: Theory and Policy Implications, Monaco, CESifo*, **TD No. 546 (February 2005).**

G. M. TOMAT, *Prices product differentiation and quality measurement: A comparison between hedonic and matched model methods*, Research in Economics, Vol. 60, 1, pp. 54-68, **TD No. 547 (February 2005)**.

F. LOTTI, E. SANTARELLI and M. VIVARELLI, *Gibrat's law in a medium-technology industry: Empirical evidence for Italy*, in E. Santarelli (ed.), Entrepreneurship, Growth, and Innovation: the Dynamics of Firms and Industries, New York, Springer, **TD No. 555 (June 2005).**

F. BUSETTI, S. FABIANI and A. HARVEY, *Convergence of prices and rates of inflation*, Oxford Bulletin of Economics and Statistics, Vol. 68, 1, pp. 863-878, **TD No. 575 (February 2006).**

M. CARUSO, *Stock market fluctuations and money demand in Italy, 1913 - 2003*, Economic Notes, Vol. 35, 1, pp. 1-47, **TD No. 576 (February 2006)**.

S. IRANZO, F. SCHIVARDI and E. TOSETTI, *Skill dispersion and productivity: An analysis with matched data*, CEPR Discussion Paper, 5539, **TD No. 577 (February 2006).**

R. BRONZINI and G. DE BLASIO, *Evaluating the impact of investment incentives: The case of Italy's Law 488/92.* Journal of Urban Economics, Vol. 60, 2, pp. 327-349, **TD No. 582 (March 2006).**

R. BRONZINI and G. DE BLASIO, *Una valutazione degli incentivi pubblici agli investimenti*, Rivista Italiana degli Economisti , Vol. 11, 3, pp. 331-362, **TD No. 582 (March 2006).**

A. DI CESARE, *Do market-based indicators anticipate rating agencies? Evidence for international banks*, Economic Notes, Vol. 35, pp. 121-150, **TD No. 593 (May 2006).**

L. DEDOLA and S. NERI, *What does a technology shock do? A VAR analysis with model-based sign restrictions*, Journal of Monetary Economics, Vol. 54, 2, pp. 512-549, **TD No. 607 (December 2006)**.

R. GOLINELLI and S. MOMIGLIANO, *Real-time determinants of fiscal policies in the euro area*, Journal of Policy Modeling, Vol. 28, 9, pp. 943-964, **TD No. 609 (December 2006).**

P. ANGELINI, S. GERLACH, G. GRANDE, A. LEVY, F. PANETTA, R. PERLI,S. RAMASWAMY, M. SCATIGNA and P. YESIN, *The recent behaviour of financial market volatility*, BIS Papers, 29, **QEF No. 2 (August 2006)**.

2007

L. CASOLARO. and G. GOBBI, *Information technology and productivity changes in the banking industry*, Economic Notes, Vol. 36, 1, pp. 43-76, **TD No. 489 (March 2004)**.

M. PAIELLA, *Does wealth affect consumption? Evidence for Italy*, Journal of Macroeconomics, Vol. 29, 1, pp. 189-205, **TD No. 510 (July 2004).**

F. LIPPI. and S. NERI, *Information variables for monetary policy in a small structural model of the euro area*, Journal of Monetary Economics, Vol. 54, 4, pp. 1256-1270, **TD No. 511 (July 2004)**.

A. ANZUINI and A. LEVY, *Monetary policy shocks in the new EU members: A VAR approach*, Applied Economics, Vol. 39, 9, pp. 1147-1161, **TD No. 514 (July 2004)**.

R. BRONZINI, *FDI Inflows, agglomeration and host country firms' size: Evidence from Italy*, Regional Studies, Vol. 41, 7, pp. 963-978, **TD No. 526 (December 2004).**

L. MONTEFORTE, *Aggregation bias in macro models: Does it matter for the euro area?*, Economic Modelling, 24, pp. 236-261, **TD No. 534 (December 2004)**.

A. DALMAZZO and G. DE BLASIO, *Production and consumption externalities of human capital: An empirical study for Italy*, Journal of Population Economics, Vol. 20, 2, pp. 359-382, **TD No. 554 (June 2005).**

M. BUGAMELLI and R. TEDESCHI, *Le strategie di prezzo delle imprese esportatrici italiane*, Politica Economica, v. 3, pp. 321-350, **TD No. 563 (November 2005)**.

L. GAMBACORTA and S. IANNOTTI, *Are there asymmetries in the response of bank interest rates to monetary shocks?*, Applied Economics, v. 39, 19, pp. 2503-2517, **TD No. 566 (November 2005).**

S. DI ADDARIO and E. PATACCHINI, *Wages and the city. Evidence from Italy*, Development Studies Working Papers 231, Centro Studi Luca d'Agliano, **TD No. 570 (January 2006)**.

P. ANGELINI and F. LIPPI, *Did prices really soar after the euro cash changeover? Evidence from ATM withdrawals*, International Journal of Central Banking, Vol. 3, 4, pp. 1-22, **TD No. 581 (March 2006)**.

A. LOCARNO, *Imperfect knowledge, adaptive learning and the bias against activist monetary policies*, International Journal of Central Banking, v. 3, 3, pp. 47-85, **TD No. 590 (May 2006)**.

F. LOTTI and J. MARCUCCI, *Revisiting the empirical evidence on firms' money demand*, Journal of Economics and Business, Vol. 59, 1, pp. 51-73, **TD No. 595 (May 2006)**.

P. CIPOLLONE and A. ROSOLIA, *Social interactions in high school: Lessons from an earthquake*, American Economic Review, Vol. 97, 3, pp. 948-965, **TD No. 596 (September 2006).**

A. BRANDOLINI, *Measurement of income distribution in supranational entities: The case of the European Union*, in S. P. Jenkins e J. Micklewright (eds.), Inequality and Poverty Re-examined, Oxford, Oxford University Press, **TD No. 623 (April 2007).**

M. PAIELLA, *The foregone gains of incomplete portfolios*, Review of Financial Studies, Vol. 20, 5, pp. 1623-1646, **TD No. 625 (April 2007).**

K. BEHRENS, A. R. LAMORGESE, G.I.P. OTTAVIANO and T. TABUCHI, *Changes in transport and non transport costs: local vs. global impacts in a spatial network*, Regional Science and Urban Economics, Vol. 37, 6, pp. 625-648, **TD No. 628 (April 2007).**

G. ASCARI and T. ROPELE, *Optimal monetary policy under low trend inflation*, Journal of Monetary Economics, v. 54, 8, pp. 2568-2583, **TD No. 647 (November 2007).**

R. GIORDANO, S. MOMIGLIANO, S. NERI and R. PEROTTI, *The Effects of Fiscal Policy in Italy: Evidence from a VAR Model*, European Journal of Political Economy, Vol. 23, 3, pp. 707-733, **TD No. 656 (December 2007).**

2008

S. MOMIGLIANO, J. Henry and P. Hernández de Cos, *The impact of government budget on prices: Evidence from macroeconometric models*, Journal of Policy Modelling, v. 30, 1, pp. 123-143 **TD No. 523 (October 2004).**

P. DEL GIOVANE, S. FABIANI and R. SABATINI, *What's behind "inflation perceptions"? A survey-based analysis of Italian consumers, i*n P. Del Giovane e R. Sabbatini (eds.), *The Euro Inflation and Consumers' Perceptions. Lessons from Italy*, Berlin-Heidelberg, Springer, **TD No. 655 (January 2008).**

FORTHCOMING

S. SIVIERO and D. TERLIZZESE, *Macroeconomic forecasting: Debunking a few old wives' tales*, Journal of Business Cycle Measurement and Analysis, **TD No. 395 (February 2001)**.

P. ANGELINI, *Liquidity and announcement effects in the euro area*, Giornale degli economisti e annali di economia, **TD No. 451 (October 2002).**

S. MAGRI, *Italian households' debt: The participation to the debt market and the size of the loan*, Empirical Economics, **TD No. 454 (October 2002)**.

P. ANGELINI, P. DEL GIOVANE, S. SIVIERO and D. TERLIZZESE, *Monetary policy in a monetary union: What role for regional information?*, International Journal of Central Banking, **TD No. 457 (December 2002)**.

L. MONTEFORTE and S. SIVIERO, *The Economic Consequences of Euro Area Modelling Shortcuts*, Applied Economics, **TD No. 458 (December 2002).**

L. GUISO and M. PAIELLA,, *Risk aversion, wealth and background risk*, Journal of the European Economic Association, **TD No. 483 (September 2003).**

G. FERRERO, *Monetary policy, learning and the speed of convergence*, Journal of Economic Dynamics and Control, **TD No. 499 (June 2004).**

F. SCHIVARDI e R. TORRINI, *Identifying the effects of firing restrictions through size-contingent Differences in regulation*, Labour Economics, **TD No. 504 (giugno 2004)**.

C. BIANCOTTI, G. D'ALESSIO and A. NERI, *Measurement errors in the Bank of Italy's survey of household income and wealth*, Review of Income and Wealth, **TD No. 520 (October 2004)**.

D. Jr. MARCHETTI and F. Nucci, *Pricing behavior and the response of hours to productivity shocks,* Journal of Money Credit and Banking, **TD No. 524 (December 2004).**

L. GAMBACORTA, *How do banks set interest rates?,* European Economic Review, **TD No. 542 (February 2005).**

P. ANGELINI and A. Generale, *On the evolution of firm size distributions*, American Economic Review, **TD No. 549 (June 2005).**

R. FELICI and M. PAGNINI,, *Distance, bank heterogeneity and entry in local banking markets*, The Journal of Industrial Economics, **TD No. 557 (June 2005).**

M. BUGAMELLI and R. TEDESCHI, *Le strategie di prezzo delle imprese esportatrici italiane*, Politica Economica, **TD No. 563 (November 2005).**

S. DI ADDARIO and E. PATACCHINI, *Wages and the city. Evidence from Italy*, Labour Economics, **TD No. 570 (January 2006)**.

M. BUGAMELLI and A. ROSOLIA, *Produttività e concorrenza estera*, Rivista di politica economica, **TD No. 578 (February 2006).**

PERICOLI M. and M. TABOGA, *Canonical term-structure models with observable factors and the dynamics of bond risk premia*, **TD No. 580 (February 2006).**

E. VIVIANO, *Entry regulations and labour market outcomes. Evidence from the Italian retail trade sector*, Labour Economics, **TD No. 594 (May 2006)**.

S. FEDERICO and G. A. MINERVA, *Outward FDI and local employment growth in Italy*, Review of World Economics, Journal of Money, Credit and Banking, **TD No. 613 (February 2007).**

F. BUSETTI and A. HARVEY, *Testing for trend*, Econometric Theory **TD No. 614 (February 2007).**

V. CESTARI, P. DEL GIOVANE and C. ROSSI-ARNAUD, *Memory for Prices and the Euro Cash Changeover: An Analysis for Cinema Prices in Italy*, In P. Del Giovane e R. Sabbatini (eds.), The Euro Inflation and Consumers' Perceptions. Lessons from Italy, Berlin-Heidelberg, Springer, **TD No. 619 (February 2007).**

B. ROFFIA and A. ZAGHINI, *Excess money growth and inflation dynamics*, International Finance, **TD No. 629 (June 2007)**.

M. DEL GATTO, GIANMARCO I. P. OTTAVIANO and M. PAGNINI, *Openness to trade and industry cost dispersion: Evidence from a panel of Italian firms*, Journal of Regional Science, **TD No. 635 (June 2007)**.

A. CIARLONE, P. PISELLI and G. TREBESCHI, *Emerging Markets' Spreads and Global Financial Conditions*, Journal of International Financial Markets, Institutions & Money, **TD No. 637 (June 2007).**

S. MAGRI, *The financing of small innovative firms: The Italian case*, Economics of Innovation and New Technology, **TD No. 640 (September 2007)**.