Delayed Privatization[*]

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Abstract

This paper studies the timing of privatization in 21 major developed economies in the 1977-2002 period. Duration analysis shows that political fragmentation plays a significant role in explaining 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.

keywords: Privatization, political economy, war of attrition

JEL ns.: D72, D78, L33

1   Introduction

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% 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 1980s or early 1990s, peaked in the late 1990s 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 1990s?

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 (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

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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 1980s 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.[1] After a long negotiation, the 1998 Telecommunications Act was eventually enacted yielding a timid liberalization of the sector, and the floatation 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 sets forth a plan to further the privatization of the company but the policy was immediately blocked by the opposition of the centre-left Social

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[1]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).

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 under state ownership. Boemer, 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 (2007) 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. In an in-depth analysis of the economic and institutional determinants of privatization in EU15 countries, Belke et al. (2007) study empirically the relation between the number of veto players and privatization revenues, finding weak support for the theory. Our approach is complementary to theirs and shows that veto players indeed matter when privatization delays are considered.

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 dis-

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entangle 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.[2]

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

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[2] For instance, Velasco (1997) presents a dynamic model in which greater 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).

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 modelled as a "war of attrition" between groups characterized by political stalemate until one group concedes (Alesina and Drazen (1991)). Spolaore (2004) extends this model to compare patterns of adjustment policies in different systems of government. 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 provides prompt adjustment, while the consensus system may fail to adjust even when adjustment is optimal. Interestingly, in the presence of large adjustments, like those following a broad privatization program, the only equilibrium in the consensus system is a war of attrition à la Alesina and Drazen, and 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

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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 obseved 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.[3]

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[3]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).

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 ML in Cox, full ML in 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) ; we discuss such measures in the next section.

4 Data

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, marking what is conventionally considered the first pri-

vatization in recent times, the IPO of British Petroleum.[4] 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 1980s, a time 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 that country. Thus, we have first collected revenue data for all privatizations (public offers and private

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[4]Some important historical antecedents were the sales of Volkswagen and VEBA implemented in the Federal Republic of Germany by Adenauer in the early 1960s. However, these companies quickly returned to public hands and were bailed-out under the pressure from disappointed investors.

sales) reported in Securities Data Corporation (see Appendix A), 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

$$t_i = \min \left\{ s : REVGDP_{is} \geq median\left[REVGDP_{ir}\right], \ r = 1977, ..., 2002 \right\}, \tag{6}$$

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

$$T_i = t_i - t_0. \tag{7}$$

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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:

$$ENP = \left[\sum_{j \in P} \left(\frac{s_j}{\sum_{k \in P} s_k} \right)^2 \right]^{-1} , \tag{8}$$

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 (8) 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 gov-

ernment, 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.[5] 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:

$$DISPR = 100 \sqrt{\frac{1}{2} \sum_{j \in P} \left[\left(\frac{s_j}{\sum_{k \in P} s_k} \right) - \left(\frac{v_j}{\sum_{k \in P} v_k} \right) \right]^2}, \qquad (9)$$

where v_j is the number of votes received by the j-th party and s_j and P are defined as in (8) . 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 at a maximum for presidential elections,

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

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 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 de novo from various sources listed in Appendix A.[6]

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. The 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

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

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.[7]

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

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[7]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.

Anglo-Saxon countries on 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. Actually, initial conditions enter equation (1) through the size of adjustment parameter θ, which in turn affects the hazard rate.[8] We control for initial conditions by including among the explanatory variables the average size of the *SOE* sector in 1977 (that is, at the beginning of the sample period). For analogous reason we may want to take into account the effect of fiscal imbalances by controlling also for the level of debt over GDP (*DEBT*).

Second, the strong distributive effects of privatization suggest that the ideological orientation of the executive matters in explaining the timing of

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[8]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)

the reform. Measuring partisanship of the government faces methodological issues analogous to those described above for political fragmentation.[9] We refer to the study by Huber and Inglehart (1995), who, by the 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 number of seats held by the executive as a whole:

$$PARTISAN = \sum_{j \in G} \left(\frac{s_j}{\sum_{k \in G} s_k} \right) HI_j, \tag{10}$$

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 (8).

The GDP per capita is included as well in all the specifications since, even within our sample of OECD countries, it still presents 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 (see Belke et al. (2007), Bortolotti, Fantini and Siniscalco

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[9]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).

(2003) and Bortolotti and Siniscalco (2004)): the decentralization of the decision making process (*FEDERAL*); the level of social conflict, reflected in the number of working days lost through *STRIKES*;[10] 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 quality of legal institutions, captured by the creditor and anti-director rights (*CREDIT* and *ANTID*, respectively) by La Porta et al. (1998).

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.[11] 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

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[10] For the relevance of social conflict in privatization see Robinson (1992).

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

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 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 privatizers 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 models (2) -(5) .

5   Results

Tables 3 and 4 show the estimation results for the Cox and Weibull models, respectively. The specifications we present are the same for both models. A benchmark equation in column [1] of all tables includes *SOE*, *DEBT*, *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. Then columns [5] through [13] check the robustness of the results to the inclusion the additional control variables, namely: the capitalization and liquidity of the financial markets variables (columns [5]-[7]); the degree of federalism and the number of days lost through strikes (columns [8]-[10]); and the quality of legal institutions (columns [10]-[13]).

The first conclusion we can draw by looking at the tables 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 variables. Initial conditions do not seem to matter. 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. A federal political structure appears also very relevant in explaining privatization delays. Indeed, the coefficient of the dummy *FEDERAL* is positive and statistically significant. This result suggest that federalism (which is often combined with a hard budget constraints for governments at the sub-national level) provides stronger incentives to pursue privatization and structural adjustment.

Turning to the measures of political fragmentation, they show considerable explanatory power. First, the indices for parliamentary fragmentation and electoral disproportionality are always statistically significant at conven-

tional levels. The evidence seems somewhat weaker for government fragmentation, at least in the benchmark specification (column [1]). On the other hand, when we control for other determinants of the delay of privatization (columns [5] through [13]), all three indices are always significant at conventional levels.

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. 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, *DEBT*, *MKTCAP* and *TURNOVER*). Such changes do not affect the 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 40 and 75% according to different specifications of

the model.[12]

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 $\hat{\alpha} > 1$. However, in the cabinet system the increase in the slope is noticeable already after 5 to 10 years, which is during the first half of 1980s, and reform occurs with probability 1 in the early 1990s, according to both $GENP$ and $PENP$ estimates of political fragmentation. On the contrary, in the consensus system privatization never gains momentum. Indeed, by the end of the sample period the predicted hazard rate is still below 25%; as a consequence, between 25 and 50% of consensus countries should have not

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[12]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) $\frac{\lambda'}{\lambda} = \exp\left[(x' - x)'\beta\right] = \exp \beta^k$, 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 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).

reformed yet (according to *GENP* and *PENP* estimates, respectively).[13]

Overall, the results of duration analysis seem strongly supportive of the empirical implications of war of attrition model. The data show that indeed greater political fragmentation entails longer delays to implement large scale divestiture.

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 1970s, 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 funda-

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[13]Notice that these predictions for hazard rates and survival functions are conservative about the importance of political fragmentation because they are based on specifications [5] and [6], in which the magnitude of the coefficients of political variables are smaller. Predictions based on specifications [8] to [13], including the full set of control variables, would entail even greater differences between consensus and cabinet countries in terms of probaility of reforming.

mental 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 the war of attrition model: large scale divestiture is delayed longer the larger the number of parties in the political system 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.

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Appendix A. Description of Variables and Sources

Variable	Definition	Source
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)
T	Delay to reform. Defined for each country as T=t-1977, where t=min{t:$REVGDP_t \geq$median[$REVGDP_s$] s =1977,...,2002}	
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)
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*)
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)
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*)
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
SOE	Average public ownership in 1977 across different sectors of economic activity.	Conway and Nicoletti (2006)
DEBT	Central government debt as a percentage of GDP	International Financial Statistics
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), updated (2003)
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), updated (2003)
FEDERAL	Dummy variable equal to 1 if the country has a federal political structure and 0 otherwise.	Persson and Tabellini (2003)
STRIKES	Number of working days lost through strikes	Arimngeon et al. (2006)
ANTID	Discrete index measuring the rights of the shareholders against the management.	La Porta et al. (1998)
CREDIT	Discrete index measuring the rights of the creditor against the debtor.	

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. [4] and Tsebelis [30]. 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																
	SEATS DIFF P2			2.76															
	SEATS DIFF P3				1.81														
	SEATS DIFF GOV					6.36													
	% MATCHED P1						80.16												
	% MATCHED P2							81.75											
	% MATCHED P3								84.13										
	% MATCHED GOV									66.67									
World Bank DPI	OBS	103									109								
	SEATS DIFF P1		28.58									29.64							
	SEATS DIFF P2			13.85									13.30						
	SEATS DIFF P3				4.21									5.01					
	SEATS DIFF GOV					30.79									30.46				
	% MATCHED P1						4.85									4.59			
	% MATCHED P2							43.69									44.95		
	% MATCHED P3								66.99									66.06	
	% MATCHED GOV									2.91									4.59

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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	3.860	5.180	-1.320 ***	-14.28
DEBT	0.449	0.314	0.135 ***	3.55
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
MKTCAP	0.694	0.464	0.230 ***	4.89
FEDERAL	0.400	0.400	0	
STRIKES	123.310	152.827	-29.517	-1.45
ANTID	4.400	2.800	1.600 ***	12.14
CREDIT	1.600	1.200	0.400 ***	3.35

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]
SOE	-0.467	-0.349	-0.170	-0.151	-0.374	0.013	0.117
	(-.1.58)	(-1.13)	(-0.60)	(-0.94)	(-0.94)	(0.04)	(0.46)
DEBT	0.203	0.166	0.225	0.338	0.378	0.156	0.783
	(0.33)	(0.28)	(0.38)	(0.62)	(0.36)	(0.17)	(1.01)
GDP	2.1e-5	3.71e-4	6e-5 **	4.7e-5	4.1e-5	6.2e-5	1.8e-05
	(0.66)	(1.57)	(2.09)	(1.49)	(0.95)	(1.60)	(0.68)
PARTISAN	0.300 ***	0.370 ***	0.283 ***	0.301 **	0.523 ***	0.315 **	0.383 ***
	(3.77)	(2.90)	(3.05)	(2.53)	(3.59)	(2.02)	(3.07)
GENP		-0.434			-0.892 ***		
		(-0.98)			(-2.94)		
PENP			-0.528 *			-0.883 ***	
			(-1.72)			(-3.47)	
DISPR				0.142 ***			0.274 ***
				(2.95)			(4.49)
TURNOVER					-0.154	-0.355	1.094 **
					(-0.28)	(-0.66)	(2.51)
MKTCAP					-0.357	0.536	2.668
					(-0.17)	(0.31)	(1.16)
Observations	211	211	211	211	191	191	191
Likelihood	-35.074	-31.511	-34.243	-32.484	-29.776	-29.854	-26.602
Wald χ^2	22.21 ***	18.44 ***	15.80 ***	19.46 ***	32.06 ***	36.37 ***	31.75 ***

Table 3. Duration analysis: Cox model (continued)

	[8]	[9]	[10]	[11]	[12]	[13]
SOE	0.098	0.176	0.409	0.382	0.457	1.073 **
	(0.24)	(0.45)	(1.59)	(1.12)	(1.34)	(2.28)
DEBT	0.036	0.110	0.515	2.429	2.502	2.868 *
	(0.04)	(0.12)	(0.75)	(1.49)	(1.46)	(1.80)
GDP	-1e-5	3.2e-5	1.13e-6	-4e-5	2.2e-5	-2.13-5
	(-0.26)	(0.75)	(0.02)	(-0.63)	(0.47)	(-0.48)
PARTISAN	0.630 ***	0.367 **	0.369 ***	0.635 **	0.353	0.356 ***
	(2.90)	(2.36)	(2.75)	(2.02)	(1.52)	(3.07)
GENP	-1.239 ***			-1.513 **		
	(-2.72)			(-2.05)		
PENP		-0.675 ***			-1.082 **	
		(-2.98)			(-1.97)	
DISPR			0.232 ***			0.252 ***
			(4.85)			(4.43)
TURNOVER	-1.418 **	-0.833	0.458	-0.394	0.396	1.221
	(-1.99)	(-1.34)	(0.80)	(-0.38)	(0.42)	(1.60)
MKTCAP	1.719	1.577	3.561 *	1.058	0.666	3.869 *
	(0.88)	(0.87)	(1.81)	(0.47)	(0.40)	(1.93)
STRIKES	-0.0001	0.0001	0.0001	-0.003	-0.002	-0.0001
	(-0.52)	(0.59)	(0.82)	(-1.27)	(-0.99)	(-0.18)
FEDERAL	1.604 ***	0.828 *	0.915 *	1.911 ***	0.693	1.422 ***
	(2.71)	(1.70)	(1.89)	(2.65)	(1.21)	(2.72)
ANTID				0.504	0.522	0.761 **
				(1.55)	(1.52)	(2.21)
CREDIT				-0.730	-0.927 *	0.278
				(-1.23)	(-1.67)	(0.62)
Observations	179	179	179	179	179	179
Likelihood	-26.692	-28.125	-25.346	-25.418	-26.442	-24.444
Wald χ^2	14.79 *	31.09 ***	41.98 ***	16.66	22.39 **	211.14 ***

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]
CONSTANT	-13.678 ***	-15.519 ***	-15.019 ***	-20.532 ***	-18.052 ***	-17.010 ***	-24.791 ***
	(-5.36)	(-4.43)	(-4.73)	(-4.18)	(-4.21)	(-3.87)	(-3.85)
SOE	-0.680 *	-0.642 *	-0.490	-0.235	-0.548	-0.215	0.032
	(-1.72)	(-1.75)	(-1.37)	(-0.80)	(-0.74)	(-0.34)	(0.07)
DEBT	0.284	0.373	0.410	0.333	0.387	0.106	0.483
	(0.44)	(0.60)	(0.68)	(0.54)	(0.28)	(0.09)	(0.46)
GDP	1.14e-5	3.14e-5	4.56e-5	3.43e-5	1.45e-5	2.78e-5	-1.48e-6
	(0.30)	(1.21)	(1.43)	(0.99)	(0.27)	(0.63)	(-0.04)
PARTISAN	0.316 ***	0.400 ***	0.311 ***	0.299 ***	0.433 **	0.279 *	0.265 *
	(3.34)	(3.48)	(3.70)	(2.60)	(2.39)	(1.88)	(1.94)
GENP		-0.473			-0.804 ***		
		(-1.22)			(-2.62)		
PENP			-0.466 **			-0.709 **	
			(-1.99)			(-2.17)	
DISPR				0.157 ***			0.202 ***
				(3.32)			(3.87)
TURNOVER					0.111	-0.181	0.886 *
					(0.18)	(-0.30)	(1.67)
MKTCAP					0.871	1.491	1.479
					(0.33)	(0.68)	(0.58)
α	5.550	6.108	6.036	6.850	7.007	6.723	7.873
Observations	211	211	211	211	191	191	191
Likelihood	5.968	6.734	6.875	9.265	8.639	8.352	11.438
Wald χ^2	14.78 ***	19.35 ***	19.00 ***	28.05 ***	29.46 ***	57.65 ***	38.140 ***

Table 4. Duration analysis: Weibull model (continued)

|	[8]	[9]	[10]	[11]	[12]	[13]
CONSTANT	-25.472 ***	-19.708 ***	-26.335 ***	-27.465 **	-19.214 **	-31.85 **
	(-2.97)	(-3.87)	(-2.95)	(-2.41)	(-2.27)	(-2.06)
SOE	-0.118	0.033	0.368	0.208	0.115	0.867
	(-0.20)	(0.05)	(0.54)	(0.23)	(0.13)	(0.75)
DEBT	-0.297	-0.267	0.004	1.569	1.257	0.348
	(-0.26)	(-0.22)	(0.00)	(1.45)	(1.15)	(0.29)
GDP	-8.74e-5	-3.50e-5	-5.78e-5	-8.55e-5	-6.94e-6	-7.97e-5
	(-1.06)	(-0.77)	-(1.26)	(-0.94)	(-0.16)	(-1.09)
PARTISAN	0.661 ***	0.384 ***	0.270 *	0.659 ***	0.354 *	0.298 **
	(3.64)	(3.19)	(1.81)	(2.78)	(1.90)	(2.51)
GENP	-1.572 **			-1.809 **		
	(-2.02)			(-2.45)		
PENP		-0.765 *			-1.626 **	
		(-1.68)			(-2.08)	
DISPR			0.178 ***			0.217 *
			(3.30)			(1.87)
TURNOVER	-2.228 *	-1.478 *	-0.216	-1.223	-0.554	-0.447
	(-1.81)	(-1.72)	(-0.32)	(-0.91)	(-.042)	(-0.43)
MKTCAP	3.337	2.635	2.936	1.897	-0.064	3.384
	(1.36)	(1.50)	(1.11)	(0.69)	(-0.03)	(1.00)
STRIKES	-0.003	-0.001	-0.001	-0.005 *	-0.005 *	-0.001
	(-1.42)	(-1.12)	(-0.91)	(-1.81)	(-1.80)	(-1.21)
FEDERAL	1.957 **	1.048 *	1.257 *	2.207 **	0.353	1.468 *
	(2.52)	(1.88)	(1.89)	(2.05)	(0.50)	(1.71)
ANTID				0.415	0.235	0.277
				(1.22)	(0.70)	(1.09)
CREDIT				-0.447	-1.097	0.545
				(-0.78)	(-1.50)	(0.78)
α	10.136	7.821	8.443	10.027	9.094	8.913
Observations	179	179	179	179	179	179
Likelihood	12.967	10.588	12.958	14.366	12.87	13.523
Wald χ^2	99.92 ***	48.00 ***	76.41 ***	107.55 ***	70.32 ***	113.89 ***
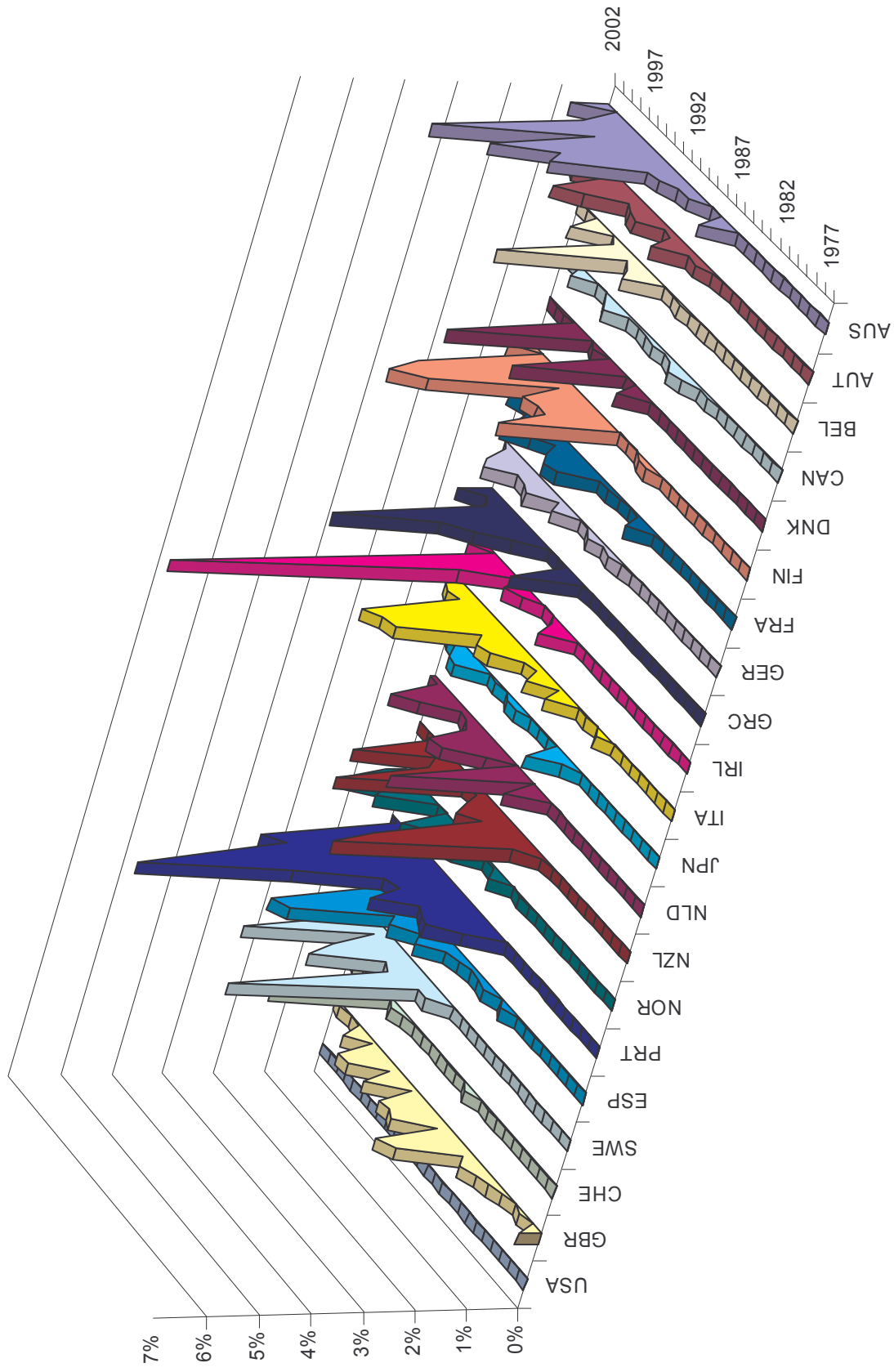
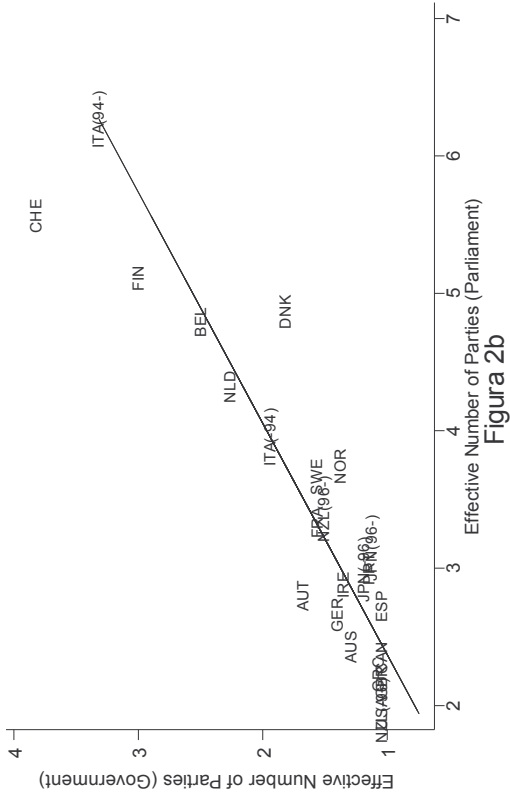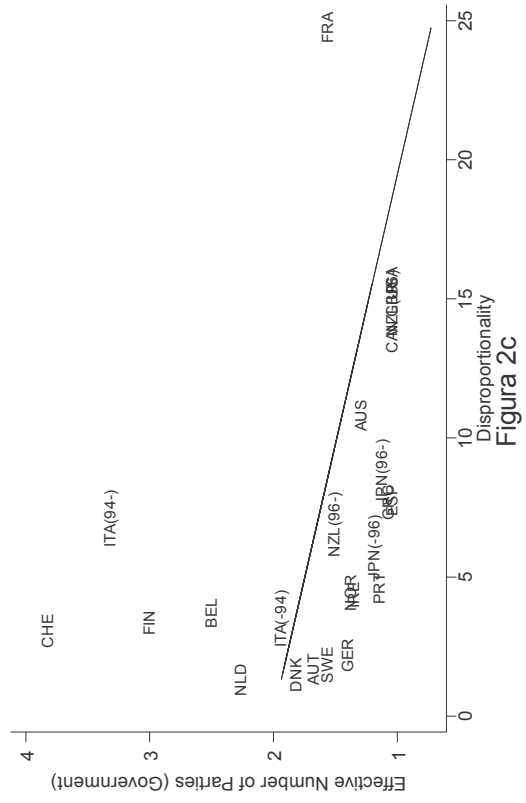Figure 1. The Big Privatization Wave, OECD Countries
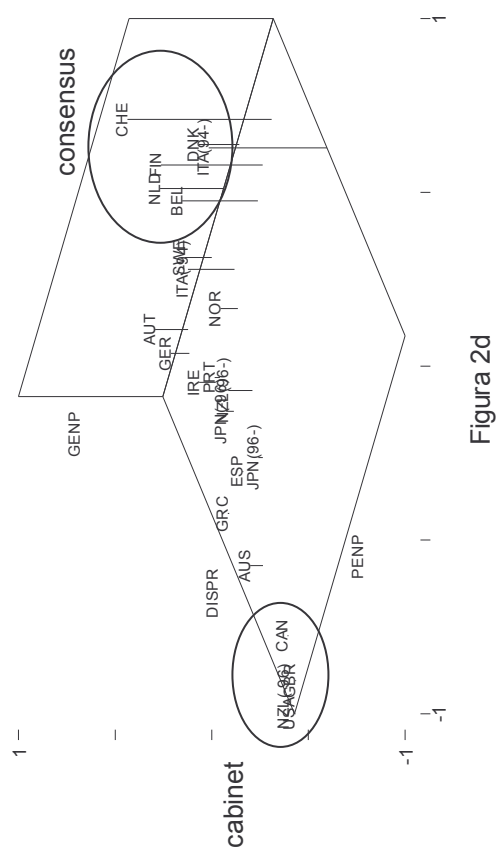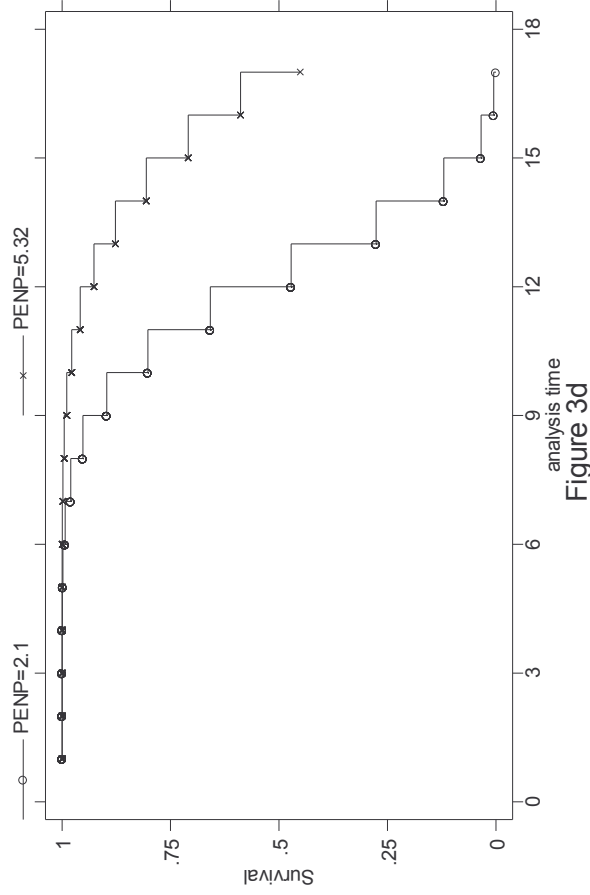
38
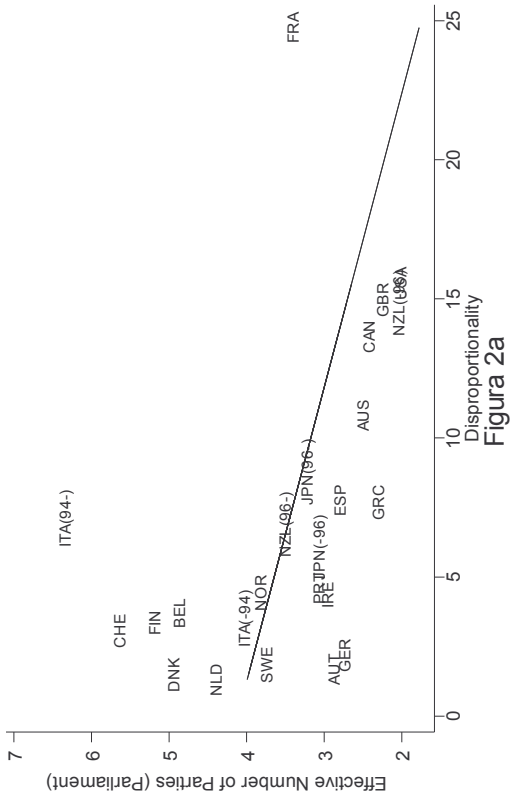
Figure 2. Measures of Political Fragmentation
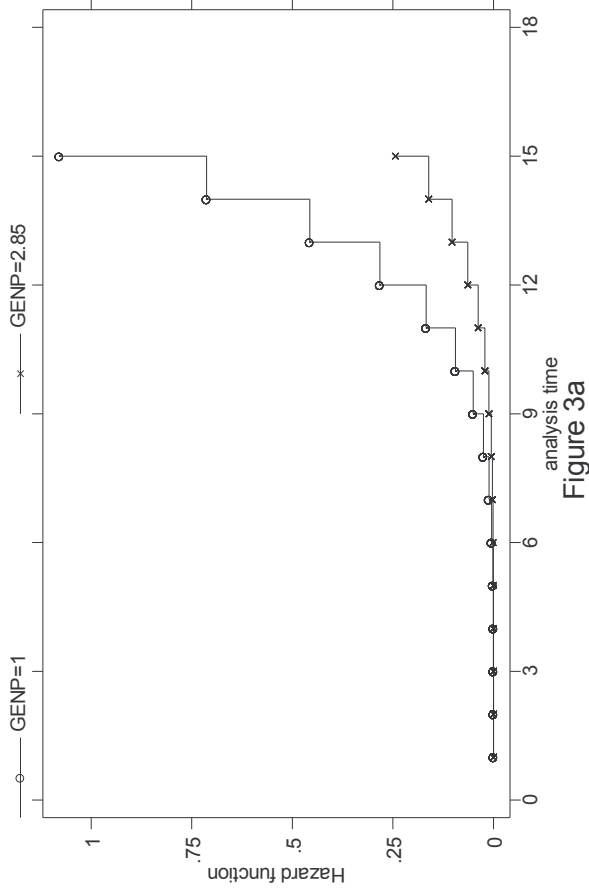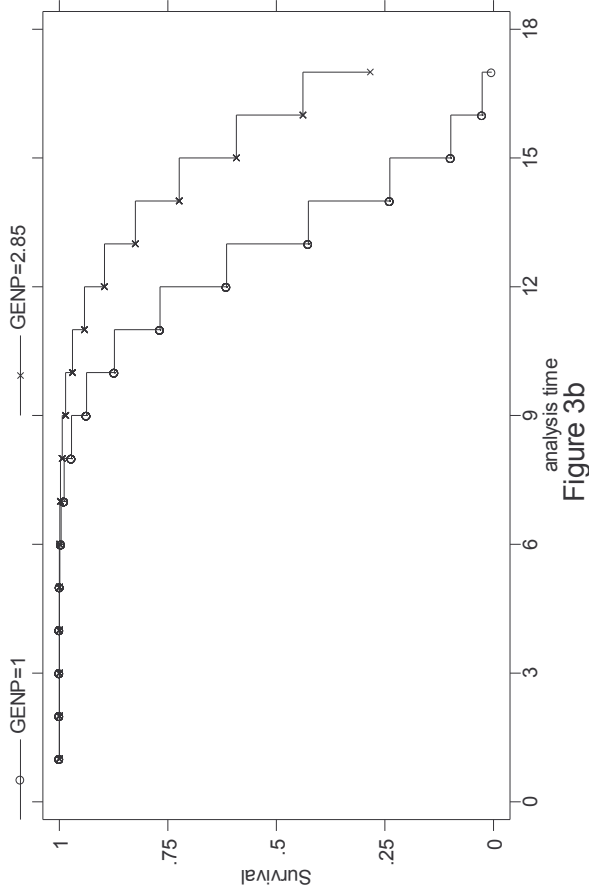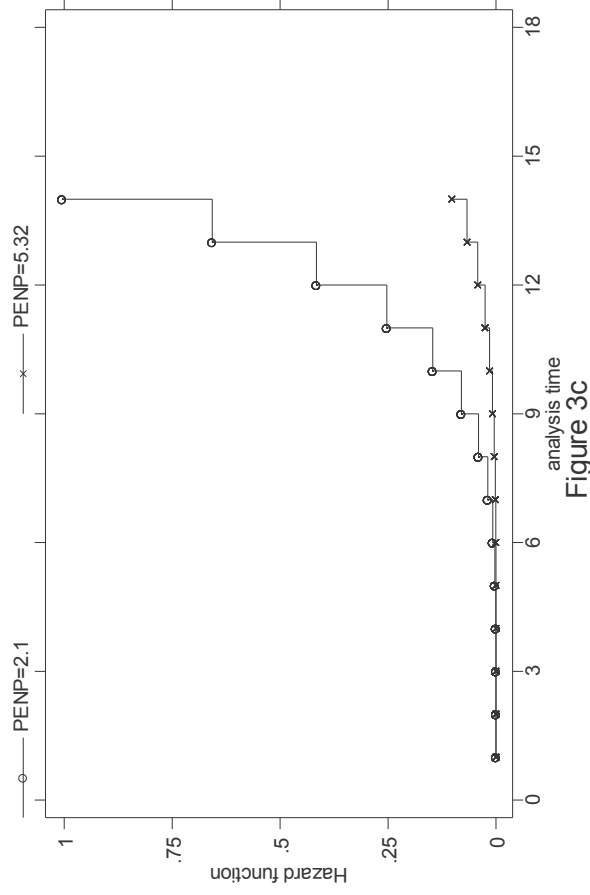
Figure 3. Hazard Rates and Survival Functions