The Political Economy of Privatization: Evidence from OECD Countries*

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Abstract

This paper presents a new political dataset for 21 industrialized countries in the 1977-2002 period, and tests the effect of political institutions and partisan orientation on privatization policy. First, we show that the degree of political fragmentation affects significantly the timing of privatization. As theory predicts, in more fragmented democracies the "war of attrition" among different veto players delays large-scale privatization. Second, privatization methods seem strongly affected by partisan politics. Particularly, right-wing executives with re-election concerns design privatization to spread share ownership among domestic voters.

Keywords: Political institutions, partisan politics, privatization

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1. Introduction

In the last two decades, privatization policy has swept the world. Up to end 2004, governments sold assets worth US$1,360bn in more than 4,053 privatization deals. Not surprisingly, industrialized economies got the lion's share of total revenues (76 per cent) (Securities Data Corporation); however, developing countries have also privatized large chunks of their State-owned enterprise (SOE) sector under the pressure of international lending agencies (Figure 1).

Even if privatization processes seem to follow a common global trend, the extent of divestiture varies greatly across countries. In some countries, governments have pursued a consistent and sustained privatization policy as a part of wider reform packages, while in others ambitious programs have been blocked on their way by adverse interest groups, so that privatization has been sporadic and small-scaled.

The United Kingdom provides an illustrative example of the first kind of privatization policy. Although not explicitly mentioned in the Conservatives' program at the 1979 elections which brought Mrs. Thatcher to power, privatization started in (reasonably) competitive industries with the sale of British Aerospace, Britoil, and Cable and Wireless, and then gained momentum after the 1983 re-election with the privatization of sizeable companies with market power (such as British Telecommunications, British Gas, and British Airport Authority). The complete divestiture of the SOE sector was instead a top priority of the political manifesto which allowed the Conservatives to obtain re-election in 1987. Privatization continued apace with important sales in the newly liberalized electricity market and in the water industry. At the end of the third Conservative legislature, the annual proceeds approached £5bn (i.e. ten times the initial level) and virtually all State-owned corporations have been sold out, with SOE value added accounting for a marginal share of domestic GDP (Vickers and Yarrow, 1991; World Bank, 1995).

Importantly, the privatization process in the United Kingdom was initially fiercely ousted by the trade unions. In 1985-86, the National Union of Mineworkers went on a two-year strike against the restructuring of the to-be-privatized coal industry. The engineers of BT also called a strike to oppose the major reductions in the staff numbers that privatization foresaw. However, the power enjoyed by the Conservatives in Parliament and the strength of the cabinet allowed to push back the opposition and to accomplish the announced program.

Some countries' privatization history is also fraught with failed attempts. The Belgian case is certainly interesting in that respect. A significant attempt to restructure and denationalize the public sector was made at the beginning of the 80s under various weak coalitional governments led by Prime Minister Martens. This attempt was thwarted by trade unions in 1983, with a general strike lasting several weeks. This strong reaction forced the governments to postpone this first reforming effort.

In 1986, Martens tried to launch an austerity program which also included privatization. In this direction, a public commission was established with the aim to study the rationalization of state-owned enterprises, eventually recommending the partial sale of Sabena, Belgacom, Societé Nationale d'Investissement (SNI) and Caisse Générale d'Epargne et de Retraite (CGER) (Spinnewyn, 2000). Again, this program was deeply ousted even within the coalition members and did not result in any actual privatization. In the beginning of the 90s, the sales recommended by the 1986 public commission were finally launched, amid strong political and social resistance leading to a new wave of strikes by public sector employees.

The worsening of public finance and the urgent need to meet Maastricht convergence criteria called again for fiscal discipline and privatization. In order to overcome the political stalemate that characterized the previous stabilization attempts, in 1995 the Prime Minister Deahene asked and obtained by the Parliament a special authorization to legislate by decree on certain economic matters, including divestiture. Only under these exceptional rules, has Belgium been able to float in the stock market a large number of shares of two important SOEs (i.e. Distrigaz and Dexia), to generate in a two-year period three quarters of total proceeds raised to date (end 2002), and to implement privatization sales amid a wide social protest.

We claim that different political institutions matter in explaining one country's ability of implementing policies with significant distributional consequences, such as privatization.

Particularly, countries characterized by a lower degree of political fragmentation should be more likely to privatize a sizeable fraction of their SOE sector. These countries are typically endowed with political institutions which tend to reduce the number of veto players, which in turn provide higher executive stability. Greater political cohesion allows incumbent governments to privatize 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 using an original dataset with continuous and time varying measures of political institutions for 21 OECD economies for the 1977-2002 period. Duration analysis provides a statistical model that matches perfectly the theoretical framework of the war of attrition model. Our results show that indeed political institutions matter to explain the timing of the privatization: countries characterized by lower political fragmentation tend to privatize sooner while divided political systems lag behind.

A political economy approach seems therefore useful in understanding the determinants of privatization. This fact begs a natural question: does ideology or political orientation also matter?

According to a largely held view, governments supported by right-wing coalitions are more prone to favor market economy, and therefore privatization, than leftist governments, traditionally more inclined to reduce social inequalities and to broaden the size of government.

In the privatization history large scale processes implemented by right wing executives abound. Privatization in the United Kingdom is again a notable example, as the process in its entirety has been implemented in the course of three consecutive legislatures with Mrs. Thatcher in office (1979-1991). The French case is also deeply shaped by partisan politics. In the beginning of the 80s, the newly elected socialist government undertakes a massive nationalization plan involving 5 industrial firms (Compagnie Générale d'Electricité (CGE), Rhône Poulenc, Saint Gobain, Péchiney, and Thomson Brandt), 2 financial firms (Paribas and Suez), and 39 banks. Following the electoral defeat of the socialists in 1986, the conservative government led by Chirac decided to re-privatize 13 firms and financial institutions. The privatization wave stopped with the return to power of the socialists between 1988 and 1992. Privatization resumed in 1993 when the socialists lost the presidential elections, and continued under conservative governments led by Balladur and Juppé. At the end of the 90s most of the companies that were nationalized in 1982 were again (partially) private (Dumontier and Laurin, 2002).

However, center-left governments have also embarked privatization especially when fiscal conditions deteriorate. In Italy, proceeds worth more than US$135bn (the third value in the global ranking by proceeds after the UK and Japan) have been raised almost exclusively by center-left governments. At a smaller scale, the timing of Danish privatizations has coincided exactly with the tenure of a social-democratic cabinet led by Rasmussen.

The logic that privatization policy is *a priori* adopted on the grounds of ideological preferences is not completely satisfactory. Indeed, privatization might be a consistent policy also for left wing governments if revenues are used for redistribution.

Theoretically, political preferences should instead matter in the choice of privatization method. Even if governments of all political stripes may privatize, only market-oriented (right-wing) governments design privatization to spread share ownership and foster popular capitalism. The rationale for this policy is re-election: by selling underpriced shares in the domestic retail market, right wing governments make equity investment attractive for the median voter, and create a constituency interested in the maximization of the value of financial assets and averse to the redistribution policies of the left. Strategic privatization can therefore be a rational strategy for raising the probability of success of market-oriented coalitions at future elections (Biais and Perotti, 2002).

To assess this theory, we first construct a continuous and time-varying measure of the governments' ideological orientation, and then use it as an explanatory variable in several privatization regressions.

Our results indicate that - in the context of advanced economies - the partisan orientation of the government has a significant impact on the methods of privatization. Governments leaning towards the right of the political spectrum tend to sell shares in the domestic retail market rather than selling them to strategic investors or abroad, as theory suggests. This evidence confirms that privatization is politically motivated, and that a political economy approach is particularly useful in understanding why and how divestiture takes place around the world.

This study is related to the growing field of empirical literature analyzing the political-institutional determinants of economic policy and reform (surveyed by Alesina and Perotti, 1995; Tommasi and Velasco, 1996; Persson and Tabellini, 2003).

The paper addresses and improves upon numerous methodological problems in existing privatization and political economy studies. First, we present some new continuous and time-varying political variables. These measures parallel the variables widely used in other fields of economics, such as Herfindal indexes, and simple weighted averages. They represent a major improvement on the dummy indicators and discrete variables typifying empirical studies on political economy by enhancing descriptive power of data on ideological orientation and political fragmentation (Alesina, Roubini and Cohen, 1997; Woldendrop, Keman and Budge, 1998; Perotti and Kontopoulos, 2002). Importantly, these new data allow a proper test of partisan theories of privatization which mainly used political dummy variable or proxies (Bortolotti, Fantini and Siniscalco, 2003; Jones et al., 2001; Megginson et al., 2002).

Second, several empirical studies in political economy have made widespread use of the Database of Political Institutions (DPI) assembled by Beck, Clarke, Groff, Keefer and Walsh (2001) on behalf of World Bank, which provides a wide array of political variables based on electoral data for a large panel of countries. This paper shows that DPI data are flawed by severe systematic mistakes on reported electoral outcomes compromising in turn the reliability of the political and institutional indicators available in the database. Although compiled for a more limited number of countries, the data presented in this paper survive to cross-checking with existing independent political data, providing a reliable source for empirical work in political economy.

The paper is organized as follows: section 2 reviews the relevant political economy models and states the theoretical hypotheses being tested; section 3 describes our political measures; section 4 presents the data; section 5 describes the empirical methodology and the econometric results; section 6 concludes.

2. Theoretical framework

A formal theory about the effects of political institutions on privatization is not available in the literature. However, some contributions in modern political economy can be suitably adapted to the context of privatization. In what follows, we review these theories and draw our empirical implications from these models. A partisan theory of privatization has instead been developed by Biais and Perotti (2002). In this section, we will also present this model and its main predictions.

2.1 War of attrition models and delayed privatization

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 cost of waiting equals the expected benefit from waiting. Importantly, the model predicts that if stabilization costs are unequally distributed between "winners" and "losers" stabilization is delayed. The intuition is clear: the more skewed is the distribution of the tax burden, the higher are the benefits from waiting. Alesina and Drazen note that large coalition cabinets made of diverse parties may hardly reach an agreement on how to allocate tax increase among the different constituencies. Therefore delayed stabilization should be associated with lower political cohesion.

However, the empirical implications of this model are too far fetched to allow a proper test of the role of the political system in explaining delays in stabilizations. Indeed, it is not straightforward to link the asymmetries of the distribution of stabilization costs to different political-institutional settings.

Spolaore (2004) makes an important step in this direction, by developing a model which 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, and how this affects the relative performance of different political systems in terms of efficient adjustment. 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 can be simply defined 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 consider 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 shocks for which immediate adjustment would be efficient, 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 in efficient adjustment (i.e. stabilization) is increasing in the number of political agents. The degree of political fragmentation therefore affects the delay of adjustment.

War of attrition models have been tested empirically in the context of fiscal stabilization (Roubini and Sachs, 1989; Alesina and Perotti, 1995). However, privatization is another important example of 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 (managers and 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 the context of privatization, Spolaore (2004) model yields therefore the following empirical implication.

H1. *Ceteris paribus, the expected delay in implementing privatization is increasing in the degree of political fragmentation.*

2.2 The partisan dimension of privatization

During the last fifteen years political economy has witnessed a growing interest in the positive analysis of the consequences of the political conflict between partisan politicians on economic policy. Within this strand of literature, some contributions analyzed the possibility of strategic manipulation of economic variables by politicians in order to achieve reappointment (Aghion and Bolton, 1990).

Biais and Perotti (2002) develop a model of privatization where right wing politicians privatize in order to gain future support from the constituencies of shareholders of newly privatized firms. They assume that the right wing party maximizes the utility of the rich, the left the utility of the poor, and each party needs the vote of the median class to win the elections. They show that by allocating a substantial amount of shares of privatized companies to the middle class, the right makes the median voter averse to the redistribution policies of the left, and more prone to vote with the right at future elections. A large-scale privatization program may therefore represent a strategy for switching to forms of "popular capitalism", by creating a constituency of voters interested in the maximization of the value of their financial assets. Importantly, Bias and Perotti show that the left can also strategically design privatization to obtain re-election. However, the privatization objectives of the two parties would be different, as the left wing does not have any incentive to underprice shares, but instead to maximize revenues available for redistribution.

This theoretical argument suggests that while privatization can be a bi-partisan policy, its implementation will be affected by political preferences. On the one hand, right wing governments will tend to privatize by public offer, earmarking (underpriced) shares to domestic investors. On the other hand, left wing governments will opt more frequently for private placements (i.e. direct sales of control blocks to strategic investors) or share issues in international (and more liquid) exchanges as both strategies allow to generate higher privatization revenues (Megginson et al. 2004; Ellul and Pagano, 2002).

The partisan model of privatization yields the following empirical implication:

H2. *Ceteris paribus, governments leaning towards the right of the political spectrum wing should privatize by spreading share ownership among domestic voters.*

The next sections will describe how we assess the empirical validity of these theoretical predictions.

3. Measuring political institutions and partisan orientation

In order to test empirically the above mentioned theories, quantitative indicators on the relevant features of political systems are needed. First, the test of war of attrition models should rely on a measure of political fragmentation capturing the number of political agents with veto power (H1). Second, a proper test of partisan model predicates on objective measures for the ideological orientation of the executive (H2).

3.1 Political fragmentation

Conceptually, political fragmentation relates to the presence of political agents enjoying veto power. The larger the number of these veto players, the higher the degree of political fragmentation. When it comes to make the notion operational, one has first to decide who the relevant political agents are, and second to provide an objectively quantifiable measure of their veto power in a given political system.

As to the first issue, the political economy literature conventionally identifies political parties as the main decision making units being (more or less) cohesive entities representing specific interest groups.

Classifications of political fragmentation often used by economists are based on the number of parties in government and distinguish between one-party executives, two-party coalitions and broader ones (Roubini and Sachs, 1989; Persson, Roland and Tabellini, 2003). However, this approach does not allow for an adequate account of the real veto power enjoyed by political parties in different types of government.

First, the relative strength of the veto players within the executive is not captured by a simple count of the number of parties supporting the government. Second, this method does not take into account the bargaining power of the executive in dealing with the parliament, which in turn affects the power of veto players. For example, the power enjoyed by a given party within a minimal coalition cabinet (including only parties whose support is necessary to achieve majority) is stronger than in an oversized one, even if the number of parties in government is the same. Similarly, a single-party minority government may not be a powerful political agent, being typically exposed to threats by the parties of the opposition.

As to the first issue, we improve on conventional proxies of political fragmentation by using the *Effective Number of Parties* (*ENP*), a measure developed in comparative political science. This measure, which parallels the Herfindal concentration index commonly used in industrial economics, puts more weight to those parties which either hold "coalition potential" or "blackmail potential", i.e. substantial bargaining power in terms of parliamentary seats (Sartori, 1976; Laakso and Taagepera, 1979). The novelty here is that, in order to focus on the veto players within the executive, we compute it only over the parties forming the government coalition. Thus, our first index *GENP* is given by the following formula:

$$GENP = \left[\sum_{j \in G} \left(\frac{n_j}{\sum_{j \in G} n_j} \right)^2 \right]^{-1} \tag{1}$$

where n_j is the number of seats in the parliament held by the j-th party and G is the set of parties forming the coalition. Expression (1) says that if there are N_G parties in the government, the *GENP* will take the value N_G if their shares are exactly equal, otherwise it will take lower values, in order to "discount" less powerful parties within the coalition. As the number of parties increases, the single shares decrease on average and the *GENP* increases.

We address our second concern about the effect of structure of the government on veto power by using the *Type Of Government* (*TOG*) indicator, developed by Lijphart (1999). This characterization takes into account both the transaction costs of political bargaining within the executive and the transaction costs of government's dealing with the parliament. Accordingly, the variable *TOG* characterizes governments along two dimensions: (i) one-party governments from coalition ones; (ii) minimal coalition governments from minority and oversized ones. Minimal coalition governments include only parties whose support is necessary to achieve parliamentary majority, while oversized ones do not. Using this classification, governments are attributed scores according to the following matrix:

Type of Government (*TOG*)	One-Party	Coalition
Minimal Winning	1	0,5
Minority or Oversized	0,5	0

The minimal winning – one party executive obtains the maximum score, as it enjoys considerable bargaining power ensuring executive stability. As such, it fits well with the the "cabinet" system as labeled by Spolaore (2004). On the contrary, the leadership in coalition minority governments is typically exposed to threats of turnover both by coalition allies and by the opposition; such possibility in turn fosters political bargaining and compromise. Furthermore, oversized coalitions tend to accord decision-making power to parties other than those strictly necessary for the coalition to stay in office; in this aspect they clearly fit in with Spolaore's "consensus" system.

Governments deviating from these two benchmark models along one dimension receive middle scores; adjusted scores provide classification for particular systems, such as presidential ones.[1] Clearly, this discrete index increases as political fragmentation decreases.

[1] In particular, presidential cabinets are, in a way, one party – minimal winning by definition; thus, they should always receive score 1. However, the score ranges from 0,5 to 1 to take into account whether or not the president faces a hostile legislative assembly (the so-called divided government).

In comparative political science, political fragmentation is usually referred to the type of government but also to the party structure in the legislature. While we are particularly interested in measuring veto power within the executive, a highly fragmented party structure could also delay the implementation of reform policies which may require broader consensus than simple majority. Thus, we include as a measure of political fragmentation the *Effective Number of Parties* (*ENP*) which is computed using exactly the same formula of *GENP* but over the total number of parties represented in the parliament.

3.2 Partisan orientation

Empirical analyses on partisan politics usually predicate upon dummy variables which crudely distinguish between left and right wing governments, with very limited methodological refinement since the seminal work by Hibbs (1977). This classification is suitable for the small sample of countries where political competition results in a strong and clearly marked bi-polarism. However, political dummy variables are certainly unsatisfactory to measure government's partisan orientation in countries where the party system is highly fragmented and/or there exists a significant "center" block. For example, they fail to discriminate between a left wing government and a center-left one with strong representation by Christian-Democrats, a typical case in several Continental European countries. Moreover, dummy variables assign the same score to moderate or extremist parties. Discrete variables scaling more than two values (see Perotti and Kontopoulos, 2002 and Alesina et al., 1997, 3 and 4 values indexes, respectively or Woldendrop et al., 1998, 5 values index) represent only a partial solution to this problem, because they still arbitrarily weigh the role played within the ruling coalition by extremist or moderate parties.

We argue instead that in order to assess precisely the partisan orientation of the executive, first a score has to be assigned to each party entering in the ruling coalition, and *not* to the coalition as a whole; second, the values obtained by the different parties of the coalition need to be aggregated into a single score. In this direction, we proceed as follows: (i) by locating different parties of several countries on a left-right scale of political orientation; (ii) by weighting the relative importance of each party within the coalition. We will address these issues in turn.

During the last 20 years, several methodologies have been established to locate different parties on a left-right spectrum of political orientation. All these approaches trade inaccuracy of dummy indicators for arbitrariness of continuous measures. Among them, expert survey methodology has proven itself a reliable tool in limiting researchers' discretion. Huber and Inglehart (1995) have produced a comprehensive dataset for a very large sample both in terms of countries considered (42) and of experts interviewed (over 800). As far as we know, this is the most recent attempt to provide such a partisan classification, and one of the broadest ones in terms of coverage. Therefore, in the construction of our partisan index, we will use their classification, assigning to each party a score ranging between 1 (extreme left) to 10 (extreme right).

Our index is a weighted average of the scores obtained by parties forming the executive, according to the Huber and Inglehart survey, where the weights are given by the power enjoyed by each party within the government coalition. For majority governments, parties whose support is not essential for the coalition to obtain 50%+1 of seats in the parliament have been excluded in the computation of the index, as in principle they cannot exert any veto power on the decision making of the coalition. All the parties are instead accounted for in minority governments.

We proxy political power by the number of parliamentary seats held by each party over the total held by the ruling coalition as a whole. Our *PARTISAN* index of political orientation is thus formally defined as follows:

$$PARTISAN = \frac{\sum_{j \in G} \Phi\left\{\left[\left(\underset{i \in G}{S} > 50\%\right) \wedge \left(\underset{i \in G \wedge i \neq j}{S} \leq 50\%\right)\right] \vee \left(\underset{i \in G}{S} \leq 50\%\right)\right\} n_j HI_j}{\sum_{j \in G} \Phi\left\{\left[\left(\underset{i \in G}{S} > 50\%\right) \wedge \left(\underset{i \in G \wedge i \neq j}{S} \leq 50\%\right)\right] \vee \left(\underset{i \in G}{S} \leq 50\%\right)\right\} n_j} \qquad (2)$$

$$\text{with } \underset{i}{S} = \sum_i \frac{n_i}{\sum_{k \in P} n_k}$$

where $\Phi(.)$ is an indicator function taking value 1 when its argument holds true and 0 otherwise, $\wedge$ and $\vee$ denote the "AND" "OR" logic operators respectively, HI_j is the score attached by Huber and Inglehart (1995) to the j-th party, n_j is the number of seats in the parliament held by the *j*-th party and G and P are the set of parties forming the government and the parliament, respectively.

An alternative methodology could have used the number of ministries held by each party as weights. However, as Laver and Shepsle (1996) have pointed out, the two criteria are strongly related, since the percentage of parliamentary seats and of ministries held by parties are on average very similar. Nevertheless, at least in principle the method based on the percentage of parliamentary seats seems more convincing, as the alternative criterion would have implied to assign the same importance to all the executive posts, despite the obvious differences in terms of prestige and power.

An important feature of all the politico-institutional variables just described is that they are time varying, as they change around election years in a given country[2]. Moreover, three of them are continuous as well, the *TOG* discrete index being the exception. As such, they account for the extreme heterogeneity observed at the ideological and institutional level better than dummy or discrete indexes adopted so far by most of the political economics literature.

4. Data

In the previous section, we have presented the methodology that we will use in the construction of our political indexes. We now describe precisely our rules for sampling and our sources; we present the FEEM Political Database (FPD) and compare it with existing databases; finally, we describe our privatization measures and control variables.

4.1 Political variables

The first issue to address is the selection of countries. In this direction, we strictly follow the political science literature by choosing sound democracies with established political institutions enabling an orderly succession of powers. Our sample covers: most of Western Continental Europe (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.

Our main source for the data on political institutions is Lijphart (1999). As we mentioned in the previous section, Lijphart (1999) has developed a series of country

[2] In electoral years these variables are the weighted averages of the data in the pre and post election periods, and the weights are given by the proportion of months before and after the elections.

indicators along several dimensions of the political system using electoral data. We have used his series for two of our indexes (*TOG* and *ENP*) for the 21 countries in our sample for the period 1977-2002, updating the original dataset to our end year (2002). The other two indices, *GENP* and *PARTISAN* have been developed independently; as such, the relative series are compiled *ex novo* from various sources listed in Table 1.

Table 2 shows the average values for these variables for the countries of our sample. 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 on the two sub-periods around the first election taking place under the new regime. Figure 2 plots the same cross-country averages for the three institutional measures *GENP*, *TOG* and *ENP* on two and three-dimensional graphs, along with the fitted values from an OLS regression.

The slope of the regression lines is consistent with the expected pair-wise relationship between the three variables. The number of parties in the parliament, as measured by the *ENP*, shows a clear relationship with both the effective number of parties in the government and the incidence of strong minimal winning – one party majority governments, reflected in the *TOG* index (Figure 2.a and 2.b, respectively). Persson, Roland and Tabellini (2003) have recently shown that political fragmentation is affected by the electoral rule. Indeed, electoral rules determine the number of parties gaining access to the parliament (higher in proportional systems, lower in majoritarian ones); party structure shapes political fragmentation, and thus indirectly fiscal policy.

In line with previous results, our dimensions of political fragmentation identify a cluster of strongly majoritarian (Anglo-Saxon) countries such as Australia, Canada, United Kingdom, United States and New Zealand (before 1993 reform) (see Figure 2d).

At the opposite, proportional-consensual democracies such as the Low Countries (Belgium and Netherlands), the Scandinavian Countries (Denmark, Finland, Norway and Sweden), Italy and, finally, Switzerland for a second cluster characterized on average by higher political fragmentation.

As we already mentioned, three cases of electoral systems' reform are reported. Since they are extremely rare events (3 out of 483 country-years in our sample), it may be interesting to evaluate their impact on our political indicators.

In New Zealand the 1993 reform from majoritarian to proportional electoral system resulted in an increased number of parties in the parliament and in enlarged government coalitions,[3] according to the basic mechanisms linking electoral rules, effective number of parties and government fractionalization mentioned above.

Japan and Italy, attempting to curb corruption and improve government stability, moved instead in the opposite direction, shifting from proportional to majoritarian systems. However, these reforms did not pay off as expected. In Japan stability of the government, as measured by the *TOG* index, increased, but the effective number of parties in the parliament increased as well, even if only slightly. In Italy, the introduction of a mixed proportional/majoritarian system resulted into a sharp increase in the number of parties, both within the parliament and the government coalition (in all the graphs, Italy moves towards the right of the spectrum), contrary to the intentions of the reform; moreover the average score for the type of government was left unaffected.

We now briefly describe our *PARTISAN* index. Even it may theoretically range between 1 and 10, the country averages reported in Table 2 show a convergence to

[3] Before 1993, New Zealand was the only country of the sample that has been ruled by majority one-party cabinets for the whole period; after the reform its index TOG has always been around 0.5.

middle values, the only outlier in this respect being Japan. Such a tendency is confirmed by the whole sample mean being about 6 and a standard deviation lower than 1. We can interpret this as a tendency of established democracies to converge to moderate ideological positions, at least in the medium-long run.

4.2 Comparison with existing political databases

The World Bank Database on Political Institutions (DPI) by Beck et al. (2001) reports the number of seats obtained by each of the three main parties of the ruling coalition and the number of seats held by the coalition as a whole. These data could be an important starting point in the construction of our political-institutional indexes. However, the DPI data do not appear accurate as one can realize at a first sight (for instance, by noticing that in many cases the total number of parliamentary seats in a given country changes across elections). We have therefore constructed a new database using the primary sources listed in Table 1.

To cross-check the precision of our database and to identify the presence of any systematic error in the World Bank DPI, we compared them pair-wise to a third data base compiled by an independent source (Tsebelis, 2001) in the country years when the three databases overlap. Table 3 presents the results of this comparison.

The column OBS refers to the number of observations (i.e., the number of elections) reported in two databases. SEATS DIFF is the average difference between the number of seats reported in two databases for the first, the second, and the third largest party in the government coalition, respectively, and for the government as a whole. Finally, % MATCHED SEATS reports the percentage of cases in which the number of seats exactly coincides.

Indeed, the FPD and Tsebelis' appear similar in several respects. The average difference between the reported number of seats is very low for each of the three main parties and for the government 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 coalition.

On the contrary, the World Bank DPI does not seem to be related to any of the other two databases. First, a lower number of observations is reported, so that several electoral results are missing. Second, the pair-wise comparison yields a very high average difference in terms of reported seats (around 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.

The way the FPD and Tsebelis' dataset fit in with each other and their pair-wise divergence from the World Bank DPI allows us to conclude that some systematic measurement error must affect the last one. The comparison casts serious doubts about the reliability of the "PARTIES" section of the World Bank DPI, at least as far as the variables mentioned above are concerned. Moreover, these mistakes are very likely to affect the computation of fractionalization indexes for the government and for the legislative assembly, also included in the database.

4.3 Privatization variables

Our source for privatization data is *Securities Data Corporation*, certainly one of the most comprehensive sources of information at the transaction level. The database contains detailed information about Public Offers of shares (i.e. privatization on public equity markets) and also about private equity placements. Clearly, the first ones refer to large-scale operations that often involve the targeting of shares to different classes of investors (retail, institutional) in different marketplaces (domestic, or international). The second ones refer instead to the sale of a large stake (often a control stake) to strategic investors.

We have aggregated transaction data to construct a panel database for our 21 countries with the following variables. Total privatization revenues (i.e. from public offers and private sales) to GDP in country *i* in year *t* *(REV/GDP)*; privatization revenues from the domestic retail market as a percentage of total revenues in country *i* in year *t* *(DOM/REV)*. The variable *REV/GDP* will be employed to measure the delay in privatization, which is given by the time elapsing from the first privatization reported in SDC (the sale of British Petroleum in 1977) to the year corresponding to the median value of *REV/GDP* for a given country. Median revenues are used to set the length of this process rather than the first transaction in the country because the initial privatizations are often sporadic and small-scaled. The second variable *(DOM/REV)* captures government's intentions to tap domestic citizens and to diffuse share ownership among domestic voters. This variable is used to verify whether the choice of the privatization method is politically motivated.[4]

4. 4 Control variables

In order to isolate the role of political institutions, we have to control for the other possible determinants of privatization. First, privatization can be simply affected by the initial size of the SOE sector. Second, fiscal conditions should be considered, given that privatization revenues are typically used to square public finance. Third, the stage of financial market development plays an important role, as deep and liquid stock markets facilitate the flotation of large companies and allow governments to maximize proceeds. Finally, the current economic outlook matters, as it has been empirically documented that large privatization programs have been typically implemented in rich countries during times of declining economic activity (Bortolotti et al. 2003). We will control for these factors by use of the variables described below.

Comprehensive datasets on SOE activity in developed countries suitable to control for initial conditions are not available. Complete time series about the SOE value added as a percentage of GDP (or about the share of employment in SOE) are missing for countries such as Austria, Canada, Denmark, Finland, New Zealand, Spain, Switzerland, and the Netherlands. Data availability then forces us to use a proxy for the size of the SOE sector pre-privatization given by the ratio of taxes over GPD, averaged for the three years before one country's first privatization (*AVGTAX*). Despite its limitations, such a measure captures one of the main effects of large state ownership in the economy, such as higher average tax rates to finance subsidies to SOEs. We complement such a measure with the value of total (domestic and foreign) debt as a percentage of GDP (*DEBT*), which mirrors the current outlook of public finances in a country-year. We use two conventional measures of a country's financial development: the ratio of stock market capitalization to GDP, (*MKTCAP*), and the turnover ratio, given by the stock market total value traded to market capitalization (*TURNOVER*). Finally, we include GDP per capita (in constant dollars 1996). With the exception of *AVGTAX*, all the controls are time-varying covariates.

[4] Revenues raised by Public Offer could be an alternative measure. However, *DOM/REV* allows a more precise empirical test of H2, as floating shares of privatized firms on the domestic stock market is a

4.5 Descriptive analysis

The data presented in Table 2 are useful for a first account of the role of political institutions and government's political orientation in privatization. The average of the (standardized) values of the three measures of political fragmentation takes the highest values in New Zealand (before 1993 electoral reform), the United Kingdom, the United States, Canada, and Australia. Interestingly, some of these countries have also been involved earlier in large-scale privatization. The United Kingdom, the US, and Canada raised their median revenues in 1977, 1983, and 1987, respectively. On the contrary, privatization has been delayed longer in highly fragmented countries such as such as Belgium, Sweden, Norway, and Finland.

Table 4 provides a more systematic test based on univariate statistics. The statistics reported for the control variables yield similar results with respect to those obtained in previous work, indicating a possible role of traditional macroeconomic variables and financial market indicators in explaining also in explaining the timing of privatization. But a new factor also appears to be relevant: political fragmentation. Interestingly, countries in the first quartile in terms of privatization delay (therefore privatizing the median revenues sooner) appear less politically fragmented then countries in the last quartile, with highly statistically significant differences in all three measures. Partisan politics does not seem to explain why privatization is delayed, but it appears to matter in the choice of the privatization method. In Table 4, we find the fraction of revenues raised in the domestic retail market to be associated with right wing market-oriented governments in office. This preliminary evidence suggests that our political-institutional

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necessary condition for shifting the political preferences of the median voter. Similar results are obtained when data on public offers are used.

measures might have some explanatory power, and indicate the need of a thorough econometric test.

5. Econometric analysis

Our main goal is to analyze the political and institutional determinants of the timing and methods of privatization. We describe in turn the way we deal with these two issues.

5.1 Methodology

We first estimate a duration model with time-varying covariates to investigate how long it takes a country to privatize a sizeable part of its SOE sector. War of attrition models surveyed in Section 2.1 establish a monotone relationship between the concession hazard rate, i.e. the conditional probability of observing the adjustment/privatization, and political fragmentation. Hazard-rate models provide an exact translation of this concept into a statistical model. The dependent variable is the hazard rate $\lambda(t \mid x)$ defined as

$$\lambda(t \mid x) = -\frac{\partial \ln\left[1 - F(t \mid x)\right]}{\partial t} = \frac{f(t \mid x)}{1 - F(t \mid x)} \tag{3}$$

where t represent the duration of some process; x is a vector of covariates; finally, $F(t \mid x)$ and $f(t \mid x)$ are, respectively, the cumulative and the density function of the duration t. Thus, according to equation (3) we are modelling the probability of observing some event (or, as it is usually referred to, "failure") at date t, given that it has not been observed for any period τ<t. In our context, the event is observing the median value for the ratio of privatization revenues over GDP.

Following the literature on survival-transition analysis (Cox, 1972, Kiefer, 1988, and Van Den Berg, 2001), we restrict our attention to proportional hazard models, so that equation (3) takes the form

$$\lambda(t \mid x) = k(x)\lambda_0(t) \tag{4}$$

where a nonnegative $k(.)$ function of the vector of covariates x multiplies the baseline hazard $\lambda_0(t)$. In particular, we follow two different estimation strategies in turn. First, we follow Cox (1972) and adopt a semi-parametric procedure that leaves the baseline hazard function $\lambda_0(t)$ unspecified, while assuming an exponential form for $k(.)$:

$$k(x) = \exp(x'\beta)$$

It then follows that for the particular failure at time $t_{(j)}$, conditionally on the risk set R_j[5], the probability that the failure is on the k-th unit as observed is given by

$$f(t_{(j)} \mid x_k) = \frac{\exp(x_k{}'\beta)}{\sum_{i \in R_j} \exp(x_i{}'\beta)}$$

The vector of coefficients β is estimated by partial maximum likelihood over the objective function:

$$\ln L(\beta) = \sum_{j=1}^{J} \left[\sum_{k \in t_{(j)}} x_k{}'\beta - m_j \ln\left\{ \sum_{i \in R_j} \exp(x_i{}'\beta) \right\} \right] \tag{5}$$

[5] The risk set R_j is the set of all the units (countries, in our case) for whom the failure happens to be observed at $t \geq t_j$

where j indexes the ordered failure times $t_{(j)}$ (j=1,…,J), k is any of the m_j observations that fail at $t_{(j)}$, R_j is the risk set at time $t_{(j)}$ and x_k is the vector of covariates referring to the k-th unit as observed at time $t_{(j)}$. This semi-parametric approach is well suited to study how the political-institutional covariates shift the hazard function. We additionally perform a robustness check by estimating a fully parametric model specifying a functional form for the baseline hazard. In particular, we refer to the conventional Weibull specification given by

$$\lambda(t \mid x) = \gamma\alpha t^{\alpha-1}\exp(x'\beta) \tag{6}$$

Notice that now the ancillary nonnegative parameter α provides a measure of the so-called *duration dependence*. For α>1 the process shows positive duration dependence, i.e. the probability of failure increases as time goes by; of course, the opposite holds true as α<1, while for α=1 the hazard rate is independent of time. In the last case the Weibull model collapses to the simpler exponential form. The vector of coefficients β and the ancillary parameter α in (6) are estimated by ordinary maximum likelihood.

We next turn to the issue of testing our H2 hypothesis using panel data on the ratio of revenues raised by public offer to the domestic retail market over total privatization revenues per country-year (*DOMREV*). This measure is both left and right censored, at 0 (for all the country-years in which no operation was issued to the domestic retail market) and at 1 (for the years in which all the sales took place by domestic public offer) respectively. As a result, conventional regression methods would fail to account for the qualitative difference between limit (0 and 1) observations and non-limit (continuous) observations. Tobit analysis, instead, being based on a new random variable that infers the missing tail in the distribution of the observed variable, allows for estimation by conventional maximum likelihood methods (Amemiya, 1985).

In particular, consider a the following linear relationship between variable $\tilde{y}$ and the vector of covariates x

$$\tilde{y}_{it} = x_{it}'\delta + \nu_i + \varepsilon_{it} \tag{7}$$

where i and t denote values for the variables referring to the i-th country in the t-th year, ν_i is an i.i.d. random variable and ε_{it} is i.i.d. as well and independent of ν_i.[6] Panel data methods would in general consistently and efficiently estimate the vector of coefficients δ in (7) by Feasible Generalized Least Squares.

However, consider now the case when we do not actually observe $\tilde{y}$. Instead, we observe y according to the following censoring rule:

$$y = \begin{cases} 1 & \Leftrightarrow \quad 1 \leq \tilde{y} \\ \tilde{y} & \Leftrightarrow \quad 0 < \tilde{y} < 1 \\ 0 & \Leftrightarrow \quad \tilde{y} \leq 0 \end{cases}$$

Within the weaker hypotheses of the Least Squares framework there is no hope to recover some point estimate or confidence interval for δ. Thus, we have to impose some distribution for the error terms ν_i and ε_{it} in order to move to a Maximum Likelihood framework. Assuming normality, we get the following log-likelihood function:

$$L(\delta) = \sum_{i=1}^{n} \ln\left[\int_{-\infty}^{\infty} \frac{\exp(-\nu_i^2 / 2\sigma_\nu^2)}{\sqrt{2\pi\sigma_\nu^2}} \left\{\prod_{t=1}^{T} \Gamma\left(y_{it}, x_{it}'\delta + \nu_i\right)\right\} d\nu_i\right] \tag{8}$$

where

$$\Gamma\left(y_{it}, x_{it}'\delta + \nu_i\right) = \begin{cases} \Phi\left(\dfrac{y_{it} - x_{it}'\delta - \nu_i}{\sigma_\varepsilon}\right) & \Leftrightarrow \quad y_{it} = 0 \\ \left(\sqrt{2\pi\sigma_\varepsilon^2}\right)^{-1} \exp\left\{-\left[\dfrac{y_{it} - x_{it}'\delta - \nu_i}{\sigma_\varepsilon}\right]^2\right\} & \Leftrightarrow \quad 0 < y_{it} < 1 \\ 1 - \Phi\left(\dfrac{y_{it} - x_{it}'\delta - \nu_i}{\sigma_\varepsilon}\right) & \Leftrightarrow \quad y_{it} = 1 \end{cases}$$

Maximization over the objective function $L(\delta)$ (usually approximated by quadrature) provides consistent and asymptotically efficient estimates for the vector of coefficients δ.

5.2 Results

Tables 5 and 6 show the results of our econometric analysis. We present an analogous specification for both models, the only difference being the inclusion of time dummies per year only in the tobit model.[7] We follow previous empirical studies about economic determinants of privatization[8] and include in all regressions controls for the level of real GDP per capita (*GDP*), the initial conditions of public finance (*AVGTAX* and *DEBT*), the financial markets (*MKTCAP* and *TURNOVER*) and the ideological orientation of the government (*PARTISAN*); column (1) presents such baseline specification. In columns (2), (3) and (4) we then add in turn the three institutional indexes: *GENP*, *TOG* and *ENP*. We never include more than one indicator in the same regression because both from a theoretical and empirical point of view (as we discussed in section 4) we would run into serious multicollinearity problems.

[6] Given that one of the regressors (*AVGTAX*) is time-invariant, a random effect specification is assumed about the error term.

In Table 5 we present both the un-exponentiated coefficients (δ) and the hazard ratios[9] resulting from estimation of duration models. The first conclusion we can draw is that well-established economic explanations for the likelihood and extent of privatization fail to account for the time it takes for the process to take off. The univariate correlations found in the descriptive analysis do not survive in the multivariate analysis which yields not significant and unstable point estimates of the coefficients of all the economic controls. The same conclusion holds true for the *PARTISAN* index: ideological orientation of the government does not seem to matter in explaining the exit from the political stalemate.

On the other hand, the three institutional measures of political fragmentation have explanatory power. First of all, they are found to be statistically significant at conventional levels in almost any specification (the only exception being the coefficient on *GENP* for the Weibull model, which is borderline). Second, both the absolute values and the t-ratios of the coefficients are extremely stable among the semi-parametric (Cox) and the parametric (Weibull) model; this is reassuring about the robustness of the estimated effects for the variables of interest and 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 (mainly *MKTCAP*, *TURNOVER* and *DEBT*). Such changes do

[7] While it is important to include time trends in the estimation of the privatization method, the inclusion of time dummies in the duration model would be redundant and inefficient, given its peculiar specification of time dependence.
[8] See e.g. Bortolotti et al. (2003).
[9] The hazard ratios are nothing else than the exponentiated coefficients. While the simple coefficient provides immediate information about the sign of the effect, the hazard ratios are on the other hand more useful in order to perform any sensitive analysis. A change in one explanatory variable implies a proportionate change in the hazard rate at any point in time of 100*[hazard rate – 1] percent. See e.g. chapter 20 in Woolridge (2002).

not affect at all the results; this leads us to exclude both the existence of significant simultaneity bias and possible misspecifications of the statistical model.

Finally, the effect of our measures of political fragmentation appears economically significant. In particular, let us consider the effect of adding one (effective) party either to the government or to the parliament. Such increase in *GENP* and *ENP* is close to the sample standard deviation of both variables (0.776 and 1.186 respectively) and it relates to some observable of political equilibria. This change implies a reduction in the hazard rate of about 45 to 50 percent (according to Cox and Weibull estimates, respectively) at any point in time if the additional party enters the executive (i.e. if we increase the *GENP*) and by something more (50% to 55%) if it gains representation in the parliament (i.e. if we increase the *ENP*).

Changes in the type of government *(TOG)* index have no direct counterpart in observable political equilibria. Nevertheless, we can interpret a standard deviation increase in the index (0.324) as a shift from the "consensus" to the "cabinet" theoretical benchmark described by Spolaore by moving away from weak coalition and/or minority cabinets to stable majority and single-party ones. According to our estimates such a shift would approximately double the hazard rate in any period. Notice that such a coefficient is the most precisely estimated one.

Finally, in figure 3 hazard rates and survival functions estimated by the parametric model are plotted against years since 1977, the initial year of our sample. The survival function is given by $S(t \mid x) = 1 - F(t \mid x)$, thus being the predicted risk set. Figure 3.a and 3.b plot the two functions assuming *TOG*=0, which in our metric identifies the pure "consensus" benchmark; the sample analogues are Switzerland and, to lesser extent, Italy and the Scandinavian countries. Figure 3c and Figure 3d plot the same functions for pure "cabinet" model (assuming TOG=1); this model fits in with the data of New Zealand (before of the institutional reform occurred in 1996) and other Anglo-Saxon countries (with the exception of Ireland).

The qualitative and quantitative differences among the two benchmark models are striking. While the hazard rate is increasing through time in both cases (according to our estimated coefficient $\hat{\alpha}$ being significantly greater than 1 in all the regressions), in the cabinet system the increase in the slope (see figure 3c) is noticeable already after 6 years (namely at the beginning of the 80s) and the event becomes almost sure at the beginning of the 90s. On the contrary, in the consensus system the privatization process never gains momentum. Indeed, by the end of the sample period the implied hazard rate is still about 50% (see Figure 3a). According to the estimated survival function (Figure 3b) consensus countries face a 40% probability of not raising significant privatization revenues even by the end of the period.

The empirical evidence presented so far allows us to conclude that the timing of privatization is strongly affected by different institutional arrangements and political equilibria. Indeed, our first theoretical hypothesis finds strong support in the data: political fragmentation delays large scale divestiture.

In the remaining part of this section we present the results of the test of the Biais-Perotti hypothesis (Table 6). As far as control variables are concerned, the bottom line is the same that emerged from duration analysis. While explaining much about the variation of the extent of privatization across countries and over time, economic variables fail broadly to explain the choice of the privatization method when the appropriate econometric techniques are used, the only exception being the initial size of the SOE sector, proxied by the average level of taxes before the first privatization (*AVGTAX*).

We now turn to comment the results obtained for the main variables of interest. The conclusions drawn in the previous paragraph about the relative importance of ideological vs. institutional determinants affecting the evolution of the privatization process are now reversed when the privatization method is analyzed. While the coefficients of the measures of political fragmentation are insignificant and negligible in absolute value, the government's ideological orientation seems a key variable to explain how they privatize. First, the estimated coefficients of the variable *PARTISAN* are always significant at conventional levels. Second, the magnitude of the effect seems precisely estimated as it remains extremely stable across specifications and robust to the inclusion of the other politico-institutional indexes. Similarly to duration analysis, we ran several regressions allowing for different dynamic specifications of the model. Once again, the results are not affected by such changes.[10]

A quantitative assessment of the effect of our political variable suggests that again the economic effect of partisan orientation is not only statistically but also economically significant. A standard deviation (0.85) shift form left to right in the government's partisan orientation induces on average a 5 percent increase of the ratio of revenues raised from the domestic retail market. The more the government is leaning to the right to the political spectrum the more he is willing to earmark shares to domestic voters. The bulk evidence that we amassed so far does not allow us to reject neither the conjecture that a war of attrition among veto players delays privatization nor that re-election concerns affect the way governments choose the privatization method. The results we get highlight the statistical and economic relevance of our political and institutional variables to understand the issues at stake.

6. Conclusions

This paper has tried to explore empirically how political institutions and the partisan orientation of the government affect the timing and methods of privatization in developed economies.

Our results show that a political economy approach is particularly useful in understanding why and how governments privatize. The timing of divestiture is affected by the existence of political institutions curbing the bargaining power of veto players and enhancing executive stability. Privatization methods seem shaped by political preferences, with market oriented governments involved in spreading share ownership among domestic voters.

It would certainly be interesting to use the data collected to provide a final test of the political economy of privatization, i.e. to assess whether strategic privatization by right wing governments has indeed contributed to shift political preferences and

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[10] For the tobit/probit analysis, two stage maximum likelihood estimators for simultaneous equations models have been developed in order to check for the existence of significant simultaneity bias and to control for that. In particular, as far as tobit analysis is concerned, the main reference is Vella and Verbeek (1999), which extends to panel data the methodology developed by Smith and Blundell (1987) for cross sectional analysis. We performed the same analysis in order to check for the robustness of our results to possible endogeneity issues. While considerably raising the computational requirements, the two stage estimator did not convey additional information with respect to the simpler one stage ML tobit estimates we present, thus excluding the existence of significant endogeneity bias.

increase the likelihood of success of market-oriented platforms. It would also be interesting to test the power of our political dataset in other areas of structural reform. We leave this to future research.

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Table 1. Description of the Variables

Variable	Definition	Source
AVGTAX	Average tax revenues as a percentage of Gross Domestic Product of country *i* in the three years before the first privatization reported in the country by Securities Data Corporation. Tax revenues comprises compulsory transfers to the central government for public purposes.	International Monetary Fund, Government Finance Statistics Yearbook
DEBT	Total debt as a percentage of Gross Domestic Product of country *i* in year *t*. Total debt is expressed as the whole stock of direct, government, fixed term contractual obligations to others outstanding at a particular date. It includes domestic debt (such as debt held by monetary authorities, deposit money banks, non financial public enterprises, and households) and foreign debt (such as debt to international development institutions and foreign governments).	International Financial Statistics, World Development Indicators 2004, World Bank
DOM/REV	Ratio of privatization revenues raised in the domestic retail market to total privatization revenues in country *i* in year *t*.	Securities Data Corporation, Privatization Barometer database (http://www.privatizationbarometer.net/database.php)
ENP	Concentration index computed over parties seats shares in the legislative chamber. Mathematical formulation of the index is presented in the text.	Original dataset from Lijphart, updated using Electoral Studies, various years; Banks et al. (1997); Elections Around the World (www.electionworld.org); Parties and Elections in Europe (www.parties-and-elections.de/indexe.html), Political Reference Almanac (http://www.polisci.com/almanac/nations.htm)
GDP	Ratio of Gross Domestic Product in constant 1996 US Dollars to population in country *i* in year *t*. 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 shares in the legislative chamber. Mathematical formulation of the index is presented in the text.	Electoral Studies, various years; Banks et al. (1997); Elections Around the World (www.electionworld.org); Parties and Elections in Europe (www.parties-and-elections.de/indexe.html), Political Reference Almanac (http://www.polisci.com/almanac/nations.htm)
MKTCAP	Stock market capitalization to Gross Domestic Product in country *i* in year *t*. 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).
PARTISAN	Indicator for the government's partisanship. It is computed as the weighted average of the score attached to parties forming the government coalition, according to Huber and Inglehart (1995) and it ranges from 0 to 10 as well. Weight i-th equal the number of seats held by party i-th in the legislative chamber over the total held by the government coalition. Null weight is assigned to parties whose seats are not essential for the government coalition to hold the absolute majority.	Electoral Studies, various years, Banks et al. (1997), Zaratè's World Political Leaders since 1945 (www.terra.es/personal2/monolith), Library of Congress Country Studies (http://lcweb2.loc.gov/frd/cs/cshome.html), Administration and Cost of Elections (www.aceproject.org), Elections Around the World (www.electionworld.org) Parties and Elections in Europè (www.parties-and-elections.de/indexe.html), Political Reference Almanac

Table 1. (continued)

REV/GDP	Total revenues from privatization to Gross Domestic Product in country i in year t. Total revenues are revenues in current US dollars from total privatization deals (*Public Offers* and *Private Sales*). Gross Domestic Product is expressed in current US dollars.	Securities Data Corporation Privatisation, World Development Indicators, Privatization Barometer database (http://www.privatizationbarometer.net/database.php)
TOG	Discrete measure which accounts for the type of government in office: one party, minimal winning, minimal winning – one party, or neither of them. See matrix in the text	Original dataset from Lijphart, updated using Electoral Studies, various years; Banks et al. (1997); Elections Around the World (www.electionworld.org); Parties and Elections in Europe (www.parties-and-elections.de/indexe.html). Political Reference Almanac (http://www.polisci.com/almanac/nations.htm)
TURNOVER	Stock market total value traded to total market capitalization in a country in year t. 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).

Table 2. Political and Privatization Data

This table presents country averages of our politico-institutional indices, economic controls and privatization variables over the period 1977-2002. GENP is the effective number of parties in the government, TOG is the type of government and ENP is the effective number of parties in the parliament. PARTISAN is the measure of ideological orientation, ranging from extreme left (1) to extreme right (10) of the political spectrum. Finally, MEDIAN is the year when the country raised its median total privatization revenues as a percentage of GDP and DOM/REV is the average privatization revenues raised from the domestic retail market as a fraction of total revenues. In countries where an institutional reform occurred during the sample period, politico-institutional data are split in two sub-periods by considering the first electoral year under the new regime.

COUNTRIES	GENP	TOG	ENP	AVERAGE	PARTISAN	MEDIAN	DOM REV
Australia	1,249	0,813	2,427	0,841	6,012	1988	0.102
Austria	1,636	0,546	2,800	0,263	5,476	1991	0.178
Belgium	2,456	0,287	4,793	-0,989	5,499	1993	0.012
Canada	1,000	0,986	2,350	1,166	5,939	1987	0.094
Denmark	1,776	0,119	4,870	-0,885	5,822	1991	0.075
Finland	2,959	0,016	5,111	-1,608	5,692	1993	0.196
France	1,519	0,628	3,330	0,248	5,514	1987	0.206
Germany	1,357	0,463	2,661	0,341	5,620	1991	0.155
Greece	1,028	0,974	2,231	1,176	5,861	1992	0.242
Ireland	1,309	0,433	2,882	0,264	5,931	1991	0.153
Italy (-94)	1,898	0,048	3,955	-0,745	6,054	1989	0.242
Italy (94-)	3,278	0,037	6,267	-2,076	5,440		
Japan (-96)	1,146	0,184	2,990	0,035	8,286	1987	0.654
Japan (96-)	1,084	0,440	3,147	0,296	8,187		
Netherlands	2,221	0,385	4,321	-0,634	5,957	1989	0.084
New Zealand (-96)	1,000	1,000	1,965	1,297	6,800	1989	0.000
New Zealand (96-)	1,467	0,315	3,432	-0,100	6,486		
Norway	1,333	0,397	3,744	-0,042	5,221	1993	0.120
Portugal	1,103	0,447	2,993	0,340	5,925	1990	0.407
Spain	1,000	0,723	2,723	0,768	5,563	1992	0.216
Sweden	1,524	0,415	3,666	-0,087	4,960	1994	0.075
Switzerland	3,779	0,000	5,562	-2,134	4,642	1987	0.081
United Kingdom	1,000	0,955	2,173	1,185	6,701	1977	0.457
United States	1,000	0,795	1,940	1,080	5,604	1983	0.733
Average	1,630	0.475	3.431	0	5.966	1988	0.213
Std. Dev.	0.776	0.324	1.187	1	0.850	6.160	0.196

Table 3. Comparison of Political Datasets

This table presents a comparison between the FEEM FPD, the World Bank DPI, and Tsebelis (2001) database. Data about electoral observations are considered. Column OBS reports the number of common observations (i.e., the number of elections) between two databases. SEATS DIFF is the average difference between the number of seats reported for, respectively, the first, second and third party forming the government coalition, and for the government as a whole. % MATCHED SEATS is the percentage of cases in which the number of seats coincides exactly in two databases.

			Tsebelis									FEEM 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
FEEM 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 4. Univariate Tests

This table presents the test of significance of the differences in means of the explanatory variables. Column (1) reports the differences a their statistical significance between the average values of the explanatory variables among the countries that were the first 5 (bottom quartile) and the last 5 (top quartile) to raise median revenues (t-statistics are in parentheses). Column (2) reports the differences between the average values of the explanatory variables in the top and bottom quartile of the distribution of the values of the variable DOMREV. a, b, c bold characters denote statistical significance at 1, 5 and 10 percent level, respectively.

	MEDIAN	**MEDIAN**	(1)	**DOM/REV**	**DOM/REV**	(2)
	bottom 25%	top 25%	**Difference**	top 25%	bottom 25%	**difference**
GDP	20986.36	20218.92	767.43 (1.52)	20619.03	24306.21	**-3687.18**[a] (-3.11)
AVGTAX	25.91.	32.44	**-6.53**[a] (-6.08)	25.42	27.92	**-2.50**[b]. (-2.17)
DEBT	0.53	0.72	**-0.19**[a] (-6.40)	0.46	0.53	-0.07 (-1.34)
MKTCAP	0.49	0.34	**0.15**[a] (5.98)	0.54	0.60	-0.06 (-0.84)
TURNOVER	0.50	0.30	**0.20**[a] (9.56)	0.63	0.60	0.03 (0.45)
GENP	1.38	1.62	**-0.24**[b] (-2.25)	1.42	1.57	-0.15 (-1.39)
TOG	0.68	0.45	**0.24**[a] (4.18)	0.53	0.54	-0.01 (-0.23)
ENP	2.91	3.69	**-0.78**[a] (-4.47)	3.06	3.42	**-0.36**[b] (-2.06)
PARTISAN	5.92	5.61	**0.31**[b] (1.84)	6.14	5.72	**0.42**[b] (2.14)

Table 5. Duration Analysis

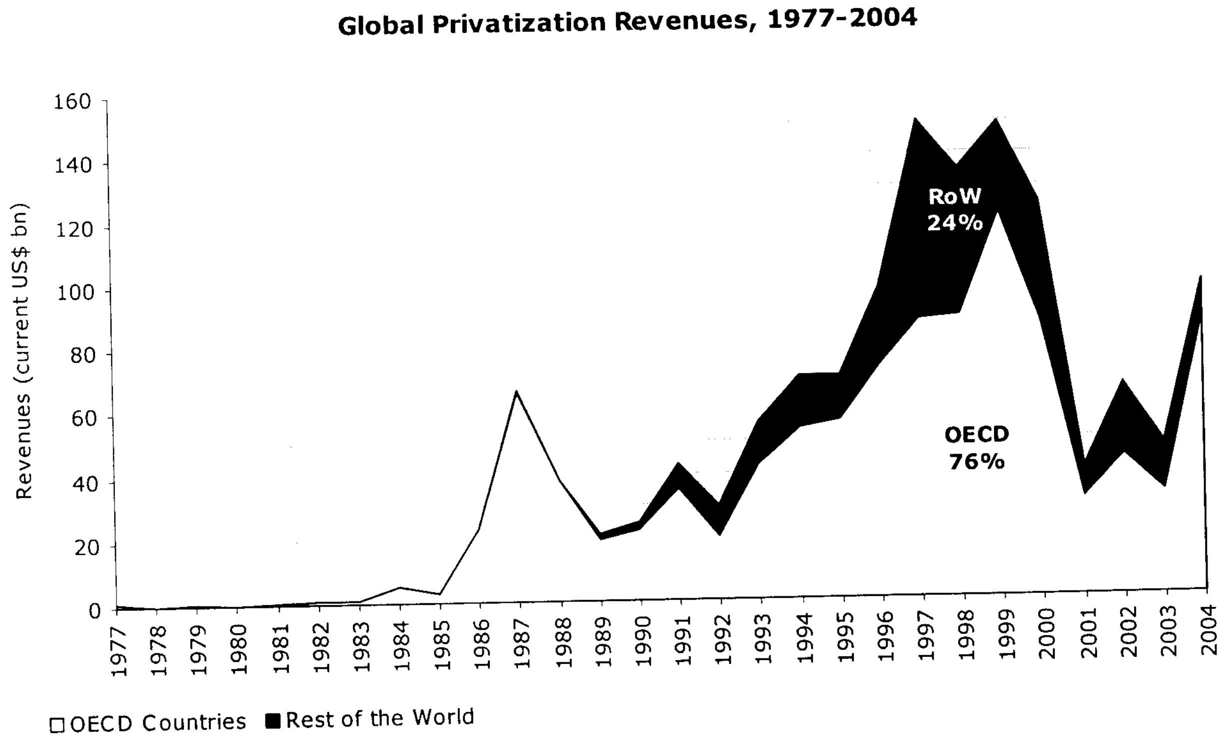
This table reports the coefficients and associated t-statistics (in parenthesis) of a duration model with time-varying covariates. We model the duration between the first-ever privatization (the sale of BP in 1977) and the year when median (per country) value of REV/GDP have been raised. Both Cox (odd columns) and Weibull (even columns) estimates results are presented. For the Weibull results, the estimate of the additional time-dependence parameter α is presented. a, b, c bold characters denote statistical significance at 1, 5 and 10 percent level, respectively.

Explanatory variables	(1A) Cox	(1B) Weibull	(2A) Cox	(2B) Weibull	(3A) Cox	(3B) Weibull	(4A) Cox	(4B) Weibull
GDP	0.00001	-0.00001	0.00004	-0.00001	0.0001	0.00007	0.0001	0.00004
	[1.000]	[1.000]	[1.000]	[1.000]	[1.000]	[1.000]	[1.000]	[1.000]
	(0.17)	(-0.21)	(0.61)	(-0.05)	(1.43)	(1.01)	(1.12)	(0.57)
AVGTAX	-0.061	-0.056	-0.035	-0.015	-0.059	-0.035	-0.052	-0.032
	[0.940]	[0.946]	[0.966]	[0.985]	[0.943]	[0.965]	[0.949]	[0.969]
	(-1.15)	(-1.10)	(-0.60)	(-0.28)	(-1.00)	(-0.67)	(-0.87)	(-0.60)
DEBT	0.633	0.417	0.427	-0.021	1.048	0.544	0.837	0.356
	[1.883]	[1.517]	[1.533]	[0.979]	[2.853]	[1.723]	[2.309]	[1.427]
	(0.48)	(0.31)	(0.30)	(-0.01)	(0.73)	(0.39)	(0.58)	(0.25)
MKTCAP	0.848	1.837	0.249	2.246	-0.040	2.037	-0.250	1.705
	[2.336]	[6.280]	[1.283]	[9.454]	[0.960]	[7.671]	[0.779]	[5.503]
	(0.44)	(1.11)	(0.12)	(1.32)	(-0.02)	(1.20)	(-0.12)	(1.05)
TURNOVER	-0.100	0.007	-0.213	0.059	0.151	0.408	-0.364	-0.191
	[0.905]	[1.007]	[0.808]	[1.061]	[1.163]	[1.504]	[0.695]	[0.826]
	(-0.14)	(0.01)	(-0.26)	(0.08)	(0.21)	(0.56)	(-0.42)	(-0.24)
PARTISAN	0.187	0.155	0.400	0.304	0.423	0.309	0.252	0.191
	[1.206]	[1.167]	[1.492]	[1.356]	[1.527]	[1.362]	[1.287]	[1.210]
	(0.87)	(0.78)	(1.49)	(1.38)	(1.59)	(1.41)	(1.05)	(0.93)
GENP			**-0.812**[c]	-0.717				
			[0.444[c]]	[0.488]				
			(-1.67)	(-1.56)				
TOG					**2.198**[b]	**2.172**[b]		
					[9.003[b]]	**[8.780[b]]**		
					(2.09)	(1.96)		
ENP							**-0.792**[b]	**-0.689**[c]
							[0.453[b]]	**[0.502[c]]**
							(-1.86)	(-1.64)
$\hat{\alpha}$		**5.881**[a]		**6.729**[a]		**6.578**[a]		**6.346**[a]
		(8.66)		(8.70)		(8.14)		(8.72)
Obs.	191	191	191	191	191	191	191	191
Log likelihood	-31.389	6.821	-29.934	8.452	-28.96	9.110	-29.446	8.136

Table 6. Privatization in Domestic Equity Markets: Tobit Regressions

This table reports the estimated coefficients and associated standard errors (in parenthesis) of tobit estimation. The dependent variable is given by the ratio of revenues from privatization by domestic public offer to total revenues from privatization in country *i* in year *t*. The dependent variable is left censored in 0 for the years in which all privatization operations occurred by private sales; it is right censored for all the years in which only public offers were issued; finally, it is not defined when no privatization occurred. Normality of the individual effects is assumed (random-effects model). a, b, c bold characters denote statistical significance at 1, 5 and 10 percent level, respectively.

Explanatory variables	(1)	(2)	(3)	(4)
GDP	7.92e-06	6.92e-06	1.83e-06	6.59e-06
	(0.89)	(0.75)	(0.18)	(0.68)
AVGTAX	**-0.018**[b]	**-0.019**[b]	**-0.019**[b]	**-0.019**[b]
	(-2.07)	(-1.99)	(-2.04)	(-2.01)
DEBT	-0.162	-0.191	-0.209	-0.189
	(-0.81)	(-0.90)	(-0.98)	(-0.88)
MKTCAP	-0.218	-0.230	-0.170	-0.216
	(-1.25)	(-1.26)	(-0.97)	(-1.20)
TURNOVER	**0.225**[c]	0.221	0.186	0.218
	(1.74)	(1.62)	(1.41)	(1.61)
PARTISAN	**0.065**[b]	**0.061**[c]	**0.062**[b]	**0.066**[b]
	(2.08)	(1.91)	(2.02)	(2.12)
GENP		0.032		
		(0.59)		
TOG			-0.209	
			(-1.38)	
ENP				0.020
				(0.39)
Time dummies	Yes	Yes	Yes	Yes
Obs.	198	198	198	198
Left censored	88	88	88	88
Right censored	11	11	11	11
Log likelihood	-108.002	-107.825	-107.058	-107.928



Source: Elaborations on *Securities Data Corporation*

Figure 1

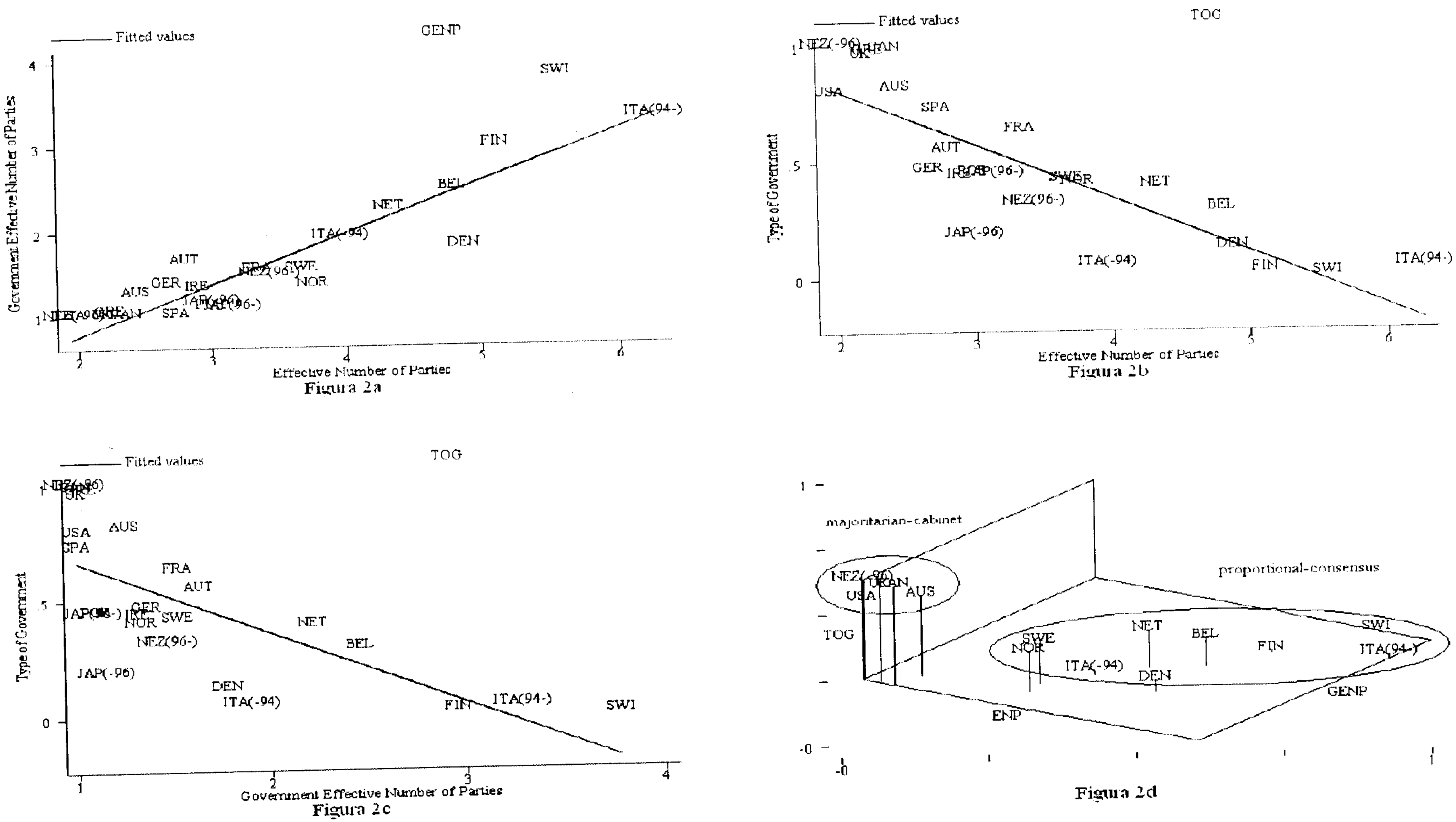


Figura 2 - The Geography of Institutions

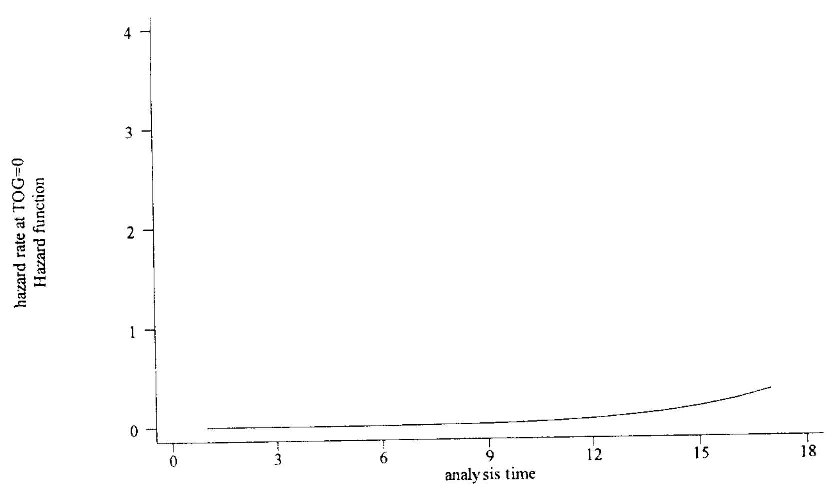


Figure 3a

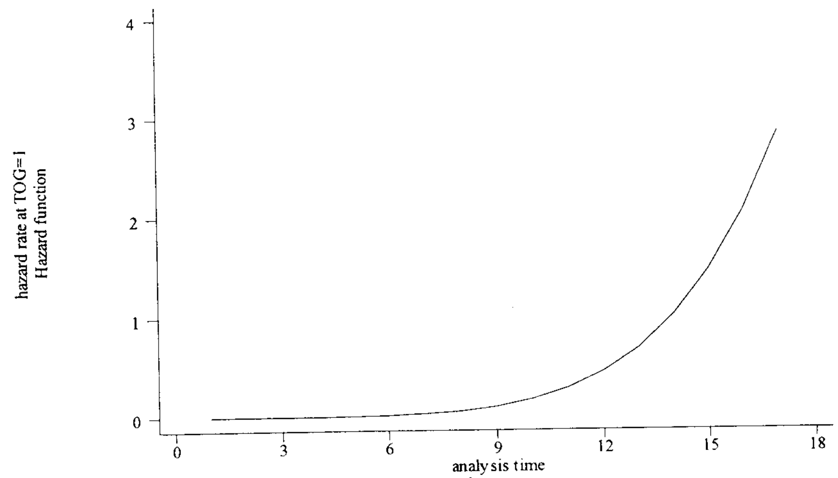


Figure 3b

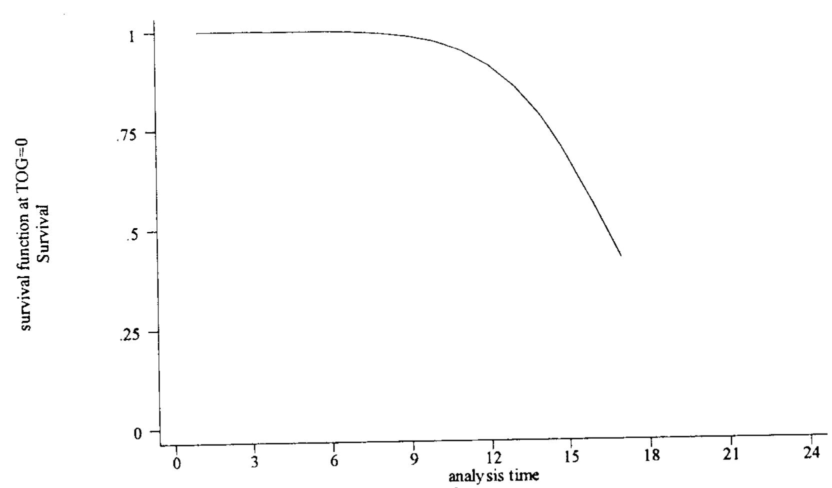


Figure 3c

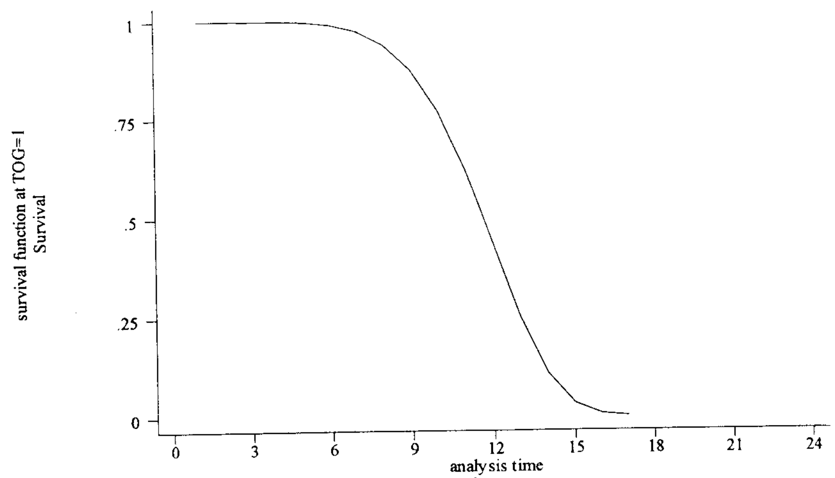


Figure 3d

Figure 3 - Hazard Rates and Survival Functions