Essays in Political Economics

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Contents

I. It's a matter of confidence: Institutions, government stability and economic outcomes 4

1. Introduction 5
2. Theoretical background 6
3. Empirical strategy 8
   3.1. Data 8
   3.2. Model 10
4. Results 13
   4.1. Robustness checks and possible transmission channels 16
5. Conclusions 19

A. Appendix 22

II. Voters' preferences and electoral systems: The EuroVotePlus experiment in Italy 31

1. Introduction 32
2. The EuroVotePlus experiment 34
3. Italians are indeed Europeans: consistency of the empirical findings with previous studies 35
   3.1. Home candidate bias 36
   3.2. Gender bias 39
4. Preferences over institutions: the importance of experience 42
   4.1. Empirical analysis 44
   4.2. Robustness checks 46
5. Conclusions 47

A. Appendix 49

III. From revolution to election: A comparative analysis of Tunisia and Egypt 55

1. Introduction and motivations 56
2. Data and descriptive statistics 59
3. Methodology 62
4. Results 65
5. Conclusions 69

A. Appendix 74

It's a matter of confidence.

Institutions, government stability and economic outcomes.*

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In this paper, we analyze the effect of constitutional structures over policy outcomes. In particular, we exploit the heterogeneity in parliamentary systems deriving from the presence and the use of the confidence vote to investigate whether stable and unstable parliamentary systems behave differently in terms of the policy they implement. This finer partition of parliamentary systems allows us to identify effects that are more robust than those in the literature. We show that the difference between presidential and parliamentary systems documented in previous works is driven by a difference between presidential and stable parliamentary systems. We suggest that possible transmission channels are legislative cohesion and (the absence of) selection.

Keywords: presidential system, parliamentary system, confidence vote, government stability

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1 Introduction

Over the last decade the political economy literature has focused on the impact that political institutions have on national policies (Persson, 2002). The seminal work of Persson and Tabellini (2003) has shown that institutions, namely political regimes, do matter in shaping size and composition of goverment spending. Since these findings, a rich literature (e.g. Blume *et al.*, 2009) has highlighted how those results are not robust to changes such as, for example, the set of countries and the time span. Furthermore, some authors (e.g. Acemoglu, 2015; Voigt, 2011b) have suggested that the distinction between parliamentary and presidential systems may simply be too coarse and that possible extensions include the use of more fine grained variables to classify constitutional systems.

The purpose of this paper is to better analyse the dichotomy between presidential and parliamentary regimes. In particular, we consider the presence and effective use of the confidence vote as the key variable to distinguish parliamentary from presidential systems.[1] This characterizing constitutional feature operates through different mechanisms that may also depend from several underlying aspects of the political environment. In some cases, the confidence vote does indeed generate frequent changes of government, thus replacing possibly bad politicians and generating a different government composition (*selection effect*). In other countries, the confidence vote acts as a credible threat and may induce either the executive to behave better (*disciplining effect*) or the parliament to accept more frequently the executive's misbehavior (*legislative cohesion*). Hence, the performance of parliamentary systems may depend on politicians' characteristics such as, for example, the quality of the information available and/or the alignement of their interests with the citizens. Given this complexity, we investigate more deeply the characteristics of countries that adopt a parliamentary constitution by considering the stability of the government as a proxy to distinguish different parliamentary systems (Lijphart, 2004). We measure stability as inversely related to the frequency of government changes, which is clearly correlated with the effective use of the confidence vote.

The issue is to understand if institutions and their actual implementation, picked up by our finer partition of political regimes, do matter in terms of implemented policies. The main result of this paper is that this classification of constitutional systems (presidential, stable parliamentary, unstable parliamentary) delivers more robust results than those in the literature. In detail, we find that stable parliamentary systems are significantly different both from presidential and unstable parliamentary ones. On the other side, unstable parliamentary systems and presidential systems behave alike in terms of the policy they implement. This result is robust to changes in the set of countries included in the dataset and in the definition of stability.

[1] For a detailed review of the relevance of the confidence vote, see Lijphart (1999).

Hence, we contribute to the literature by refining the standard classification of constitutions (see Persson and Tabellini, 2003) introducing stable and unstable parliamentary systems. In that pathbreaking work the authors compare constitutional systems - presidentialism *vs.* parliamentarism - and electoral rules - majoritarian *vs.* proportional - in order to identify the differences, if any, in a number of relevant social and economic indicators.[2]

A large literature followed Persson and Tabellini (2003), with the aim of extending, testing or questioning their results. First, Blume et al. (2009) find that, while the results on the effect of the electoral rules are robust, the effects of the parliamentary *vs.* presidential constitutional choice are sensitive to an enlargement of the dataset and the updating of the economic indicators used as regressors.

Over time the profession has felt the need of an extension of the Persson and Tabellini analysis. Acemoglu (2005) and Voigt (2011b), among others, question the presidential/parliamentary classification of constitutional structure in a twofold manner: on one side they advocate the endogenous nature of the constitutional form of government, noting that it is an equilibrium outcome rather than an exogenous characteristic, on the other side they ask for a finer partition of countries, taking into account the heterogeneity within each group. Moreover Voigt (2011b) highlights the lack of an analysis of possible transmission channels. While recent work, such as Hayo and Voigt (2013) and Robinson and Torvik (2008), investigate the determinants of the choice or change of the constitutional structure, not much has been done to refine the classification of constitutional structures when studying their effect on policy.[3]

In Section 2 we suggest how our findings may be explained by two effects that are consistent with previous theoretical literature: a selection effect, as in Cella et al. (2015), and a legislative cohesion effect, as in Baron (1998), Diermeier and Feddersen (1998a, 1998b) and Diermeier and Vlaicu (2011). Section 3 presents the data and the model, Section 4 discusses the results and Section 5 concludes.

2 Theoretical background

The idea that further partitioning parliamentary systems may provide insights about the link between insitutions and economic policy finds its support in several theoretical papers. In fact while the interplay of uncertainty and incentives operates unambigously

[2]Persson and Tabellini identify presidential/parliamentary regimes according to the legal existence of the confidence vote, so that presidential countries where the government is subject to a confidence vote - as for instance France - are classified as parliamentary. Authors find that presidential systems systematically spend the five percent less than parliamentary systems.

[3]An excpetion is the work of Ardanaz and Scartascini (2014) who sustain that the degree of separation of power within presidential system is heterogeneous. They replicate the analysis of Persson and Tabellini (2003) but interacting presidential systems with a dummy indicating the executive budget discretion. They find that presidentialism has a larger negative impact on government size only when executive discretion regarding budget allocation is low.

in presidential systems and generally give rise to unique equilibria, parliamentary systems are more heterogeneous and can support multiple equilibrium strategies and induce multiple equilibrium outcomes (see for example Cella et al. (2015)). For example, parliamentary systems outcomes may differ depending on whether the confidence vote is mostly used as a threat or whether it effectively replaces politicians. This refers to the *de facto* behaviour of politicians that in turn depends on their (short- or long-term) incentives structure. In this context, the confidence vote may affect government duration given that a legislative defeat would lead to new elections that will replace the executive and the legislative body with positive probability. On the contrary, in presidential systems both bodies have fixed terms and government stability is not affected by the working of the policy process.[4] In other words, under presidentialism politicians face undistorted incentives and vote according to policy preferences in order to increase their reputation and thus the probability to be re-voted in next scheduled elections.

In particular, the literature identifies three channels through which the confidence vote may affect the policy-making process in parliamentary systems. First of all, if the confidence vote is actively used to replace politicians, it improves their expected quality (*selection effect*, see Cella et al., 2015; Huber and Gallardo, 2008). If instead the confidence vote acts as a threat, it may either reduce the distortions to the executive's behavior (*disciplining effect*, see Cella et al., 2015; Huber, 1996), or induce the voting cohesion in parliament (*legislative cohesion*, see Baron, 1998; Diermeier and Feddersen, 1998b; Diermeier and Vlaicu, 2011).

Cella et al. (2015) highlight this twofold nature of the confidence vote. They model an executive and a legislative body in a parliamentary system where politicians may face early elections if the parliament does not approve the executive's proposed policy. Politicians can be of two types, they either care about implementing the efficient policy or they only care about being in office. The authors show that in such a setting two equilibria may arise, depending on the parameters that describe politicans' quality (type distribution) and information. The confidence vote may act as a threat and induce an office oriented executive to propose the efficient policy in order to prevent early termination of the legislature. In presence of this *disciplining effect*, stable systems are characterised by a low level of inefficiency. Alternatively, the confidence vote may be used in equililbrium to replace possibly bad politicians. As, on average, office oriented politicians are replaced more often, the expected quality of the executive improves. Hence the selection effect operates more efficiently in unstable systems in which we should observe a better alignment of the executive's and voters preferences.

[4] As noted by Diermeir and Vlaicu (2011): "Under presidentialism the policy process is driven by short-term issue-by-issue incentives because there are only short-term consequences of an unsuccessful proposal. In parliamentary systems, on the other hand, the failure of a policy proposal can lead to a change in the composition of the governing coalition. This injects political incentives in the policy process whereby coalition members consider both their short-term policy interests and their long-term political interest".

In a model of parliamentary democracy, where the government controls the legislative agenda, Baron (1998) instead shows that government members may change policies regarding government spending to preserve the government and may also seek support from the minority in parliament. In other words, the legislative cohesion effects contributes to the stability of parliamentary systems through the approval of a larger fraction of policies aimed at keeping current politicians in power.

The existence of all these effects implies that parliamentary systems are more heterogeneous than presidential ones, as, depending on their underlying characteristics, they may have a different response to the policy implementation process and a different degree of stability even for a given set of constitutional rules. Moreover, stability and policy response are theoretically correlated, so that it is meaningful to use stability as a proxy to refine the classification of parliamentary systems. The focus of our paper is to exploit empirically this de facto heterogeneity.

Our model then moves from these considerations to investigate the complex mechanisms that link constitutional features to economic outcomes. We compare the effects that constitutional structures have on the policy-making process, adopting a partition of parliamentary systems that takes into account their degree of stability. In particular, the disciplining effect would imply that the difference between presidential and parliamentary structures should derive from unstable parliamentary systems. On the opposite view, both legislative cohesion and selection effects would induce stable systems to be the ones that differ more from presidential ones. Hence our approach will allow a better understanding not only of the existence of the link between institutions and policies but also on the underlying mechanisms that generate it.

3 Empirical strategy

We start by replicating the standard empirical setting so to ease comparison with the previous literature. Then, we introduce some modifications to handle the identification issue. We now present the datasets and the empirical specifications.

3.1 Data

Data are taken from three main sources. We start with the same dataset as in Persson and Tabellini (2003) (PT). The dataset is composed by 85 countries, including data on a set of economic and social indicators.[5] The main dependent variables are central government expenditure (*cgexp*) and central government revenues (*cgrev*). These variables are computed as percentage of the GDP and are averaged between 1990 and

[5]Countries are classified as follows: OECD (*oecd*); Central, Latin America and Caribbeans (*laam*); Africa (*africa*); South and Central Asia (*asiae*). There is a prevalence of OECD and LAAM countries that jointly represent the 60% of the sample. For a detailed list of variables and sources, see Persson and Tabellini (2003).

1998.[6] The set of covariates includes indicators for the continental location and colonial history that always enter the regression equations,[7] dummies for the origin of the constitution,[8] age of democracy (*age*), distance from equator (*lat01*), percentage of people having either English or other European languages as mother-tongue (*engfrac* and *eurfrac*, respectively), democracy level (*gastil*), per-capita income (*lyp*), proportion of people between the age 15-64 (*prop1564*) and over 65 (*prop65*), population size (*lpop*) and a dummy indicating a federalism system (*federal*).

We then extend the dataset (to obtain what we call the BCIM dataset) to 116 countries, using data from Blume et al. (2009), thus updating the following variables: output per worker (*logyl*) from 1988 in the PT dataset to 2000 and perception of corruption (*cpi*) from 1995-2000 to 2000-2005. We include data on additional dependent variables to provide robustness checks: central government expenditure on social services and welfare as percentage of the GDP (*ssw*) for the period 1990-1998, social protection as percentage of the GDP (*socprot*) for the period 1995-2012, expenditure on education as percentage of the GDP (*edspend*) for the period 1995-2012 and general government expenditure as percentage of the GDP (*ggexp*) for the period 1990-2014.[9] We also consider the executive's ideological position (*right_left*), and the district magnitude (*magn*) to perform additional robustness checks in Section 4.2.[10]

Finally, in order to partition parliamentary countries according to the stability distribution, we create several stability indexes from a set of political indicators drawn from the World Bank Database of Political Institutions (DPI, 2012). The dataset covers the period 1975-2012. Our main indicator is $gov\ life$, defined as follows:

$$gov\ life = \frac{\sum_i D_i / \sum_i E_i}{L_i}, \qquad (1)$$

where D_i represents the number of years a government has been in office between two elections, E_i is a dummy which indicates elections, and L_i, is the legal length of any electoral term according to country-specific constitutional rules.[11] Thus, $gov\ life$ is the average length of any electoral cycle computed for each country i, normalized by

[6]Other possible dependent variables included in the PT dataset are central government expenditure on social services and welfare as percentage of the GDP (*ssw*), log of the output per worker (*logyl*) and perception of corruption (*cpi*).

[7]The continental location variables are reported in Footnote 4. As for the colonial history, variables include: *col_espa* if a country is a former colony of Spain or Portugal; *col_uka* if a country is a former English colony and *col_otha* if a country is a former colony of a country other than England, Spain and Portugal. All the variables are weighted for the years of independence as follows: $col\_uka = col\_uk * (250 - t\_indep)/250$, where $col\_uk = 1$ is a dummy indicating a former English colony, $t\_indep \in [0, 250]$ are the years of independence and 250 is used as the standard value for all non-colonized countries. The same exercise holds for *col_espa* and *col_otha*.

[8]The variables *con20, con2150, con5180*, respectively dating the constitution's origin before 1920, between 1921-1950, and between 1951-1980.

[9]Sources: IMF/GFS Yearbook.

[10]Source: DPI, 2012.

[11]The index $gov\ life$ is built using the indicator *yrcurnt* from the DPI dataset which is coded zero in the election year, and $X_i - 1$ in the year after the election.

the legal length of the electoral cycle. The index ranges from zero to one, with higher values corresponding to higher stability.

We classify parliamentary countries in *parl stab* if their value of $gov\ life$ is above the median of the stability distribution, and *parl unstab* if their value of $gov\ life$ is below the median.[12] We introduce separate measures of stability in Section 4.2 for the robustness checks. Table 15 in the Appendix A.1 reports some descriptive statistics.

3.2 Model

We first replicate results by Persson and Tabellini (2003) and Blume et al. (2009), considering the effects of a twofold classification of countries in presidential and parliamentary systems, according to the legal existence of the confidence vote. The empirical equation is estimated through OLS:

$$Y_i = \alpha + \beta_1 pres_i + \gamma maj_i + \delta X_i + \varepsilon_i, \tag{2}$$

where X_i is the set of observable country-specific covariates, and ε_i is the error term which is assumed to be normally distributed.

Then, following the categorization of parliamentary systems in terms of stability, we introduce our finer classification of countries, that further partitions parliamentary systems into stable and unstable ones. We apply the dummy coding technique to account for the heterogeneity in the subgroup of parliamentary systems, thus generating three categories: *pres*, *parl stab* and *parl unstab*. We estimate the model with the following multiple regression procedure:

$$Y_i = \alpha + \beta_2 parlstab_i + \beta_3 parlunstab_i + \gamma maj_i + \delta X_i + \varepsilon_i, \tag{3}$$

where *pres* is the baseline category that represents the control group in our setting.[13] We are interested in testing whether presidential systems differ from stable parliamentary ones ($\beta_2 \neq 0$), whether presidential systems differ from unstable ones ($\beta_3 \neq 0$) and whether stable and unstable parliamentary systems differ from each other ($\beta_2 \neq \beta_3$).

Finally, we re-estimate the model using the instrumental variable (IV) strategy. We

[12] More precisely, we first drop three countries which are within 0.05 points from the median of the stability distribution in order to avoid a random assignment of countries due to measurement errors. Note results hold even when we make the threshold move along the stability distribution. In detail, results remain significant until stable parliamentary countries are in the 75th percentile of the stability distribution. After that, results are no longer significant. This is consistent with the disciplining effect discussed in Section 2 according to which fully stable parliamentary countries should not be significantly different from any other constitutional category.

[13] We always perform the *Variance Inflation Factor* (VIF) to control for problems of multicollinearity. A VIF above 5 may indicate multicollinearity problems. None of our regressions has an average VIF above 3, and the explanatory variables always show a VIF between 1 and 2. Therefore, we explore the sensitivity of our key estimates by starting from a very parsimonious specification to then progressively include additional covariates. Results are available under request.

do so because we aknowledge that the choice of a constitution may be an equilibrium outcome therefore determined by agent's preferences. In other words there may exist variables that simultaneously affect both the choice of constitutions and the dependent variable, thus potentially biasing the OLS coefficents due an omitted variable problem.[14] In order to address this issue, we simultaneously estimate a nonlinear system of equations by a maximum likelihood estimator. The instrumenting equation (4) is a multinomial probit model where the constitutional categories are regressed over a set of instruments and additional controls. The headline equation (5) follows the same structure of equation (3):

$$P_c(j) = Prob[U_j > U_\kappa; \kappa \in C, \kappa \neq j], \quad K = 0, 1, 2$$
$$U_{ij} = \zeta_j X_i + \varphi Z_i + \eta_{ij}; \tag{4}$$
$$Y_i = \alpha + \beta_2 parlstab_i + \beta_3 parlunstab_i + \gamma maj_i + \delta X_i + \varepsilon_i. \tag{5}$$

We jointly estimate equations (4) and (5) to allow error terms (η_{ij}, ε_i) to be correlated, thus taking into account the full covariance structure of the model (Roodman, 2011).[15] Equation (4) defines the choice probability of a given constitutional alternative j from the choice set C containing K elements, where $K = 0, 1, 2$ represent the presidential system, the unstable parliamentary system and the stable parliamentary system, respectively. Such probability depends on the relative utility that country i obtains from adopting the constitutional type j instead of a different one, where X_i is the same set of covariates that enter equation (5), and Z_i is the vector of instruments. The error term η_{ij} is assumed to have multivariate normal distribution and is not necessarily independent across choices.[16]

The instrumental variables are chosen to ensure the *exclusion restriction*, that is some of the variables entering equation (4) have not direct effects on the policy outcome, but the effect on the constitutional choice, once we control for other regressors $Cov(Z_i, \varepsilon_i \mid X_i = 0)$. As noted by Persson and Tabellini (2003), the exclusion restriction is guaranteed by three variables, i.e. *con2150, con5180* and *con81*, respectively dating the adoption of a constitution before 1920, between 1921-1950, and between 1951-1980. These variables may be used as instruments as they are clearly exogenous to recent policy outcomes but correlated with the consitutional choice, as historically there have been waves of adoptions of specific types of constitutions. However, the pre-

[14]The identifying assumption in equations (2) and (3) is that, conditional on the vector of controls, the type of constitution and the error term are orthogonal. If this is not true, then the OLS estimator is no longer consistent. See Acemoglu (2005) for a detailed discussion.

[15]We estimate a likelihood function having more components, one for the linear equation and N-1 for the multinomial equation, where N is the number of options. This method is more efficient than the traditional two steps procedure proposed by Heckman (1979) even in the case of weak instruments (Perez and Sanz, 2005).

[16]See Appendix A.2 for a detailed analysis of the empirical model.

dictive power of the constitutional dating variables is somewhat weak.[17] Thus, Persson and Tabellini include three more instruments in the first-stage estimation: the fraction of the population speaking major European languages as native tongues (*engfrac, eurfrac*), and the distance from the equator (*lat01*). These variables are proxies for the European influence on the constitutional decision, following the argument of Hall and Jones (1999) who suggest that the extent to which countries have been influenced by Europe has a considerable impact on the quality and type of institutions. The languages predictors and latitude are indeed highly correlated with the form of government, but their validity as intruments has been widely questioned in the literature (see Acemoglu, 2005; Rockey, 2012), and by Persson and Tabellini themselves.[18] To account for this, we include the Hall-Jones instruments as covariates both in the instrumenting equation and in the headline regression.

Finally we introduce an additional intrument to improve the identification strategy. Such instrument, which we label *confl mean*, is defined as the proportion of years a country has been involved in internal and external violent conflicts between years 1816 and 1900, thus representing the country-level degree of social and political conflict during the XIX century.[19] The index *confl mean* may significantly affect the probability of a country to fall into a particular constitutional category. Indeed, the degree of conflict may strongly impact the choice of the constitutional system itself. High values of the index may indicate deeply divided societies in which political decision making needs to rely on power-sharing rules. Parliamentary systems offer the ideal environment for a broad power-sharing executive, given that the cabinet is a collegial decision-making body (Lijphart, 2004). On the contrary, presidential systems introduce rules that favour a winner-takes-all outcome, also by facilitating the adoption of a majoritarian electoral rule (Linz, 1994). As additional evidence in favour of this correlation, Jung and Deering (2015) show that unstable conditions at the time of the constitutional choice increase the likelihood of adoption of a parliamentary system. This effect should be slighly larger in the case of unstable parliamentary systems given that the degree of stability of a country's political environment may be persistent, i.e. a more unstable environment is more likely to arise in presence of past political instability (Alesina *et al.*, 1996).

[17] Nonetheless, the F-test of the joint significance of these three variables significantly rejects the null hypothesis. See Table 6 in the Appendix.

[18] "We think that [...] the three constitutional dating variables [...] are uncorrelated with the remaining unobserved determinants of fiscal policy, while we are less certain about the remaining instruments [i.e. the language variables and the latitude]. Assuming that the first three instruments are valid, the validity of the remaining [intruments] can be tested via the implied overidentifying restrictions."(Persson and Tabellini, 2004, p. 37). However, the overidentifying restriction is not convincing given the little predictive power of the constitutional timing variables. Thus, as noted by Acemoglu (2005, p.1041), "there are good reasons to suspect that they may not be excludable from the regression of interest."

[19] Source: Correlates of War Project. The index is computed as the sum of the years between 1816-1900 a country has been involved in violent conflicts over the reference period. Violent conflicts include both intra and inter states' conflicts. For a detailed review of the definition and categorization of conflicts, see Sarkees (2010).

Therefore, we argue that the new instrument has a robust predictive power for the endogenous regressor, i.e. the constitutional choice, and it is orthogonal to the error term of the headline equation, given the time span elapsed between the instrument and the dependent variable.[20]

4 Results

We start our empirical analysis by comparing our new framework with the standard analysis by Persson and Tabellini (2003) and Blume et al. (2009). We first run regressions (2) and (3) on the PT dataset using central government expenditure and central government revenues as dependent variables (Table 1). A look at Column (1) and (3) reminds us of the standard results in the literature, namely that presidential systems spend systematically less than parliamentary ones, regardless of the chosen measure of the government size.

Table 1: Constitutions, Central Government Expenditure and Central Government Revenues. OLS estimations. PT dataset

Dep.Var.	cgexp (1)	cgexp (2)	cgrev (3)	cgrev (4)
pres	-5.181^{***}		-5.001^{**}	
	(1.93)		(2.02)	
parl stab		6.298^{***}		7.541^{***}
		(2.28)		(3.20)
parl unstab		1.383		-0.104
		(1.87)		(1.05)
F-test		6.97^{**}		10.82^{***}
Observations	80	80	76	76
Adjusted R^2	0.631	0.643	0.586	0.640

Notes: White heteroskedasticity-consistent standard errors in parentheses. All the regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*. *F-test* (columns (2), (4)) refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_2 = \beta_3$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

In Table 1, Column (2) and (4) show that the difference between constitutional systems is driven by the subgroup of stable parliamentary countries. Indeed β_2 is statistically significant in both columns and seems slightly larger in magnitude when compared to previous results. Coefficient β_3 is never significantly different from zero so that we cannot reject the hypothesis that unstable parliamentary systems behave like presidential ones.

If we test whether parliamentary countries can be treated as an homogeneous group ($\beta_2 = \beta_3$), we find that we can always reject the null hypothesis (p-values are 0.013

[20]Results from the instrumenting equations are reported in the Appendix A.1, Table 6.

and 0.002, respectively).

We then run the same set of regressions on the extended BCIM dataset to check the robustness of our approach (Table 2). As shown by Blume et al. (2009) and reported in Column (1) and (4) the difference between constitutional systems in the traditional classification is no longer significant, even though the coefficients retain the same sign.

The significancy of the coefficients of the finer partition is instead preserved. Columns (2) and (5) show that β_2 is still significantly different from zero, that β_3 is not significantly different from zero and that we can reject the hypothesis that $\beta_2 = \beta_3$ (p-values are 0.0495 and 0.009, respectively). In other words extending the dataset does not change any of the empirical facts observed in the original dataset, thus suggesting that not treating parliamentary systems as a homogeneous group is a modelling improvement.

Table 2: Constitutions, Central Government Expenditure, Central Government Revenues. OLS estimations. BCIM dataset

Dep.Var.	cgexp (1)	cgexp (2)	cgexp (3)	cgrev (4)	cgrev (5)	cgrev (6)
pres	-3.755			-2.701		
	(2.42)			(2.36)		
parl stab		5.206^{**}	5.128^{**}		6.882^{***}	6.929^{***}
		(2.11)	(2.18)		(2.32)	(2.49)
parl unstab		1.734	1.442		0.302	0.052
		(1.98)	(1.96)		(1.95)	(1.96)
PPI			4.486			5.186
			(4.92)			(5.02)
F-test		2.30^{**}	2.92^{**}		9.82^{***}	6.74^{***}
Observations	91	89	82	87	85	82
Adjusted R^2	0.67	0.69	0.70	0.66	0.71	0.73

Notes: White heteroskedasticity-consistent standard errors in parentheses. All regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*. *F-test* (columns (2)-(3) and (5)-(6)) refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_2 = \beta_3$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Our partition of parliamentary systems is based on the observed behaviour of the country and not on the details of the constitutional rules. We argued in Section 2 that we intend to capture the heterogeneity generated by the different equilibria that may arise in a given constitutional setup. To further validate this interpretation of the results we include as additional control the Parliamentary Power Index (PPI) proposed by Fish and Kroenig (2009). This index reports the strength of the legislative body by measuring the fraction of "powers" that the national legislature held out of 32 listed ones in 2007. The introduction of PPI as a regressor allows us to control for differences in the constitutional features of parliamentary countries. Columns (3) and

(6) show that results do not change when we include the index, thus supporting our interpretation.

Finally, we replicate the analysis but adopting the IV approach as described in Section 3.2 both on the PT dataset (Table 3) and on the BCIM dataset (Table 4). Columns (2) and (4) of both tables show that the empirical findings are confirmed even when we instrument the possibly endogenous constitutional decision, further supporting our modelling choice.

Table 3: Constitutions, Central Government Expenditure and Central Government Revenues. IV estimations. PT dataset

Dep.Var.	cgexp (1)	cgexp (2)	cgrev (3)	cgrev (4)
pres	-6.51^{*}		-6.47^{*}	
	(3.16)		(3.72)	
parl stab		5.774^{**}		6.231^{**}
		(2.92)		(2.29)
parl unstab		2.346		0.571
		(1.50)		(1.36)
F-test		2.97^{**}		9.82^{***}
Headline covariates	Yes	Yes	Yes	Yes
Observations	83	82	82	81
Adjusted R^2	0.68	0.68	0.67	0.66

Notes: White heteroskedasticity-consistent standard errors in parentheses. The IV approach follows the strategy presented in Section 3.2. In Columns (1) and (3), the instrumenting equation is estimated using a probit model. In Columns (2) and (4), the instrumenting equation is a multinomial probit model. Results from the instrumenting equations are reported in the Appendix, Table 6. Columns (1)-(4) include the following controls: *engfrac, eurfrac, lat01, age, lyp, trade, prop1564, prop65, gastil, maj, federal, lpop, oecd*. *F-test* (columns (2) and (4)) refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_2 = \beta_3$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Constitutions, Central Government Expenditure and Central Government Revenues. IV estimations. BCIM dataset

Dep.Var.	cgexp (1)	cgexp (2)	cgrev (3)	cgrev (4)
pres	-3.29 (3.91)		-4.38 (3.93)	
parl stab		5.702^{**} (2.23)		7.243^{***} (2.75)
parl unstab		2.981 (2.33)		2.185 (2.16)
F-test		2.85^{**}		6.42^{***}
Headline covariates	Yes	Yes	Yes	Yes
Observations	100	86	99	86
Adjusted R^2	0.59	0.61	0.60	0.63

Notes: White heteroskedasticity-consistent standard errors in parentheses. The IV approach follows the strategy presented in Section 3.2. In Columns (1) and (3), the instrumenting equation is estimated using a probit model. In Columns (2) and (4), the instrumenting equation is a multinomial probit model. Results from the instrumenting equations are reported in the Appendix, Table 6. Columns (1)-(4) include the following controls: *engfrac, eurfrac, lat01, age, lyp, trade, prop1564, prop65, gastil, maj, federal, lpop, oecd*. *F-test* (columns (2) and (4)) refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_2 = \beta_3$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

4.1 Robustness checks and possible transmission channels

We perform several robustness checks to test the validity of the results. First of all, we test the model on dependent variables different from those analysed so far: central government expenditure on social services and welfare as percentage of the GDP, output per worker, perception of corruption, social protection expenditure as percentage of the GDP, expenditure on education as percentage of the GDP and general government expenditure as percentage of the GDP. We run regressions (2) and (3) on both the original and enlarged datasets whenever possible. Results are reported in the Appendix A.1, Tables 7-11. The only exception is *ssw* which is only available for a subgroup of countries even in the PT dataset. The change in dependent variables does not alter our results. Moreover, when estimating the effect of the constitutional design on the share of social welfare spending, Persson and Tabellini (2003) slightly modify the original specification by dropping three control variables - *lpop*, *pro1564* and *trade*. Our specification remains significant even when all controls are included.

We then estimate the model including the executive's ideological position and the district magnitude as additional regressors. The underlying idea is that a leftist executive should implement higher public expenditure and that district magnitude may have a positive impact on the size of fiscal policies (Milesi-Ferretti et al., 2002). Table 12 in the Appendix A.1 shows that results do not change.

A third set of robustness checks is performed on the partition of parliamentary countries in stable and unstable ones, as this partition is the distinguishing feature of the empirical strategy. In the main analysis countries are partitioned according to the index *gov life*. We consider three alternative stability indexes: *gov end*, defined as the fraction of governments that are successful in reaching the legal term of the mandate;[21] *year exec*, defined as the average tenure of the head of the executive weighted by the legal length of any electoral term;[22] *year party*, defined as the average number of years the governing party has been in office weighted by the legal length of the electoral term.[23] Results reported in Table 13 of the Appendix A.1 show that the empirical findings are not sensitive to the choice of the stability index.

A further concern is related to the possible reverse causality between stability and policy outcomes. Indeed, public expenditure may be increased by the government in order to remain in power, thus increasing stability. As a robustness check we classify parliamentary countries in stable and unstable ones based on the index *gov life* in the time interval 1975-1989 (instead of 1975-2012). In this way, the dependent variables (which are averages of the expenditure/revenues in the period 1990-1998) cannot have a direct effect on the classification. Results are reported in Table 5. The sign and magnitude of the coefficients in Table 2 and Table 5 supports the validity of our approach.

[21]The index is built using the indicator *yrcurnt* from the DPI dataset. Higher values of the index correspond to higher stability.

[22]This index is built using the indicator $yearoff$ from the DPI dataset, which collects information about the number of years the head of the executive has been in office. Higher values of the index correspond to higher stability. Note that this index may provide different results with respect to the previous ones, since it keeps counting the number of years a government has been in power even if an election occurs, if the incumbent government wins the election.

[23]This index is built using the indicator *prtyin* from the DPI dataset, which counts the number of years the chief executive party has been in office. Higher values of the index correspond to higher stability. Note that the index accounts for the possibility that a single party holds the power for a long time span.

Table 5: Constitutions, Central Government Expenditure and Central Government Revenues. OLS estimations. BCIM dataset

Dep.Var.	cgexp (1)	cgrev (2)
parl stab	6.201**	6.403**
	(2.87)	(3.05)
parl unstab	1.26	0.646
	(1.66)	(2.49)
F-test	3.43*	5.51**
Observations	89	82
Adjusted R^2	0.735	0.708

Notes: White heteroskedasticity-consistent standard errors in parentheses. All the regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*. *F-test* refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_1 = \beta_2$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The overall empirical analysis is consistent with the theoretical intuitions stated in Section 2. In particular in a model as in Cella et al. (2015), the performance of parliamentary systems closely approaches the performance of presidential ones if the confidence vote entails a better ability of legislators to reject bad policy proposals and thus a more unstable political environment. On the opposite, if the *selection effect* is not effective then the *de facto* behaviour of politicians will be driven by political incentives. This mechanism drives stable parliamentary systems far from the performance of presidential systems, but also from the one of unstable parliamentary systems. The same effect is consistent with the presence of the *legislative cohesion* effect in models as in Baron (1998), according to which the executive and the parliament coordinate to keep politicians in power and to avoid a no confidence motion. Coordination indeed leads either the executive to formulate policy proposals that please the majority of the veto-players or legislators to accept a larger fraction of executive's proposals to avoid early elections.

The main findings described above do not find evidence in favour of the *disciplining effect* that predicts a similar performance between presidential and stable parliamentary systems. However, this may depend on the difficulty to empirically isolate the subgroup of fully stable parliamentary countries. This leads us to suppose that the difference between presidential and parliamentary system is not monotonically increasing in the level of stability of the latter system. To test this insight, we try to further split parliamentary countries in more than two categories according to their stability distribution. Indeed, even if the analysis is sensitive to the small number of countries included in each category, we find that the difference between constitutional systems is increasing in the stability of the parliamentary constitutional design, but it drops when we consider fully stable parliamentary countries. Results are reported in the Appendix

A.1, Table 14.

5 Conclusion

This paper analyses the effect of constitutional structures on policy outcomes with a specific attention to the role of the confidence vote. In particular, the novelty of the paper rests with the understanding of the link between government stability and economic outcomes, particularly for parliamentary systems. Hence, the empirical analysis we perform introduces finer partitions of parliamentary countries according to their degree of stability. We find that stable parliamentary systems behave differently both from presidential and from unstable parliamentary ones with respect to every dependent variables we consider.

We also provide some novel insights about the transmission channels that may generate our empirical results. When the executive is disciplined by the threat of the confidence vote (*disciplining effect*), then it will always formulate a congruent policy proposals. In this case, the confidence vote is never actively used and the performance of fully stable parliamentary systems and presidential systems will tend to coincide. Indeed, when the confidence vote is actively used by the parliament, then the difference in performance between the two constitutional systems will be increasing in the stability of the parliamentary system. That is either when the parliament is for the majority office-motivated (*selection effect*) or when the executive and the parliament coordinate to stay in office till the end of the term (*legislative cohesion effect*).

We also introduce a novel empirical approach by proposing a new instrument that better predicts the constitutional choice of countries, thus tackling the problems of endogeneity.

Thus, we contribute to the growing body of literature of empirical constitutional economics by dealing with some of the critiques that have been moved to the previous works in particular by offering a method of analysis that generates results that are more robust and that shed some light on the possible transmission channels.

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Appendix

A.1 Tables

Table 6: IV instrumenting equations. BCIM dataset

Model	Probit	Multinomial (base=pres)	
Dep-Var.	pres	stable	unstable
	(1)	(2)	(3)
confl mean	-1.71^{***}	0.65^{**}	0.87^{***}
	(0.07)	(0.30)	(0.31)
con2150	-0.003	-0.24	0.28
	(0.08)	(0.21)	(0.23)
con5180	0.37^{**}	-0.79^{***}	0.27
	(0.09)	(0.27)	(0.23)
con81	0.78^{***}	-0.97^{***}	-0.03
	(0.13)	(0.33)	(0.26)
engfrac	-0.46^{**}	0.27	0.28
	(0.16)	(0.21)	(0.23)
eurfrac	0.82^{***}	-0.46^{***}	-0.49^{***}
	(0.15)	(0.18)	(0.23)
lat01	-0.87^{**}	0.89^{*}	0.02
	(0.39)	(0.48)	(0.51)
age	1.14^{***}	-2.03^{***}	-0.49^{*}
	(0.29)	(0.67)	(0.43)
F-TEST on confl mean		13.6^{***}	
F-TEST on constitution variables	21.4^{***}	28.98^{***}	27.39^{***}
Observations	83	80	80
Adjusted R^2	0.735		

Notes: White heteroskedasticity-consistent standard errors in parentheses. We estimate Column (1) using a probit model and Columns (2) and (3) using a multinomial probit. Entries are average marginal effects. All columns include (but do not report) the following controls: *maj, gastil, lyp, lpop, trade, prop1564, prop65, federal, oecd*. F-test on *confl mean* refers to the joint significance of confl mean in the stable and unstable categories. F-test on *constitution variables* refers to the joint significance of the constitution dating variables, i.e. *con2150, con5180, con81.*
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Constitutions and Social Welfare Expenditure. OLS estimations. PT dataset

Specification	PT-modified	BCIM-modified	PT	BCIM
Dep.Var.	ssw	ssw	ssw	ssw
	(1)	(2)	(3)	(4)
pres	-2.244^{**}		-1.927	
	(2.03)		(1.69)	
parl stab		2.707^{**}		3.366^{***}
		(2.03)		(2.25)
parl unstab		-0.543		-0.487
		(1.48)		(2.46)
F-test		5.79^{**}		6.83^{**}
Observations	69	69	69	69
Adjusted R^2	0.632	0.699	0.523	0.683

Notes: White heteroskedasticity-consistent standard errors in parentheses. PT-modified and BCIM-modified refer to Persson and Tabellini (2003) specification of the model where the authors include all the standard controls - *age, lyp, prop65, gastil, federal, oecd, africa, asiae, laam, col_uka, col_espa, col_otha* - except that *lpop, prop1564* and *trade* are missing. Then, we re-estimate the model using the same specification as in the previous table. *F-test* (columns (2), (4)) refers to the hypothesis that the coefficients for *parl stab* and *parl unstab* are equal ($\beta_1 = \beta_2$).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Constitutions and Output per Worker. OLS estimations.

Dataset	PT		BCIM	
Dep.Var.	logyl	logyl	logyl	logyl
	(1)	(2)	(3)	(4)
pres	-0.294^{*}		-0.157	
	(1.84)		(1.01)	
parl stab		0.325^{*}		0.392^{**}
		(1.78)		(2.04)
parl unstab		0.115		-0.0364
		(1.55)		(2.20)
Observations	74	73	84	83
Adjusted R^2	0.731	0.695	0.753	0.721

Notes: White heteroskedasticity-consistent standard errors in parentheses. *logyl* is the productivity level as in Persson and Tabellini (2003) (Columns (1) and (2)), and in Blume et al. (2009) (Columns (3) and (4)). The regressions include the following controls: *age, lyp, trade, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha, avelf, prot80, catho80, confu.*
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 9: Constitutions and Perception of corruption. OLS estimations.

Dataset	PT		BCIM	
Dep.Var.	cpi (1)	cpi (2)	cpi (3)	cpi (4)
pres	-0.620^{*} (1.76)		-0.326 (-1.05)	
parl stab		0.627^{*} (1.80)		0.559^{*} (1.72)
parl unstab		0.491 (1.40)		0.362 (1.19)
avelf	1.274^{**} (2.09)	1.567^{**} (2.42)	0.987^{*} (1.83)	1.432^{**} (2.49)
Observations	78	78	88	88
Adjusted R^2	0.829	0.833	0.806	0.820

Notes: White heteroskedasticity-consistent standard errors in parentheses. *cpi* is the perception of corruption as in Persson and Tabellini (2003) (Columns (1) and (2)), and in Blume et al. (2009) (Columns (3) and (4)). The regressions include: *age, lyp, trade, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha, avelf, prot80, catho80, confu.* The addtional control *avelf* is included and reported in the table. *avelf* is the index of ethnolinguistic fractionalization, as in La Porta et al. (1998).
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 10: Constitutions and Social Protection. OLS estimations.

Dataset	PT		BCIM	
Dep.Var.	socprot (1)	socprot (2)	socprot (3)	socprot (4)
pres	-2.184 (2.10)		-2.028 (2.01)	
parl stab		2.924^{**} (2.05)		2.149^{**} (1.09)
parl unstab		-0.346 (2.16)		0.270 (1.71)
Observations	76	76	86	86
Adjusted R^2	0.758	0.761	0.714	0.729

Notes: White heteroskedasticity-consistent standard errors in parentheses. *socprot* is the central government social protection expenditure as defined by the IMF-GFS dataset (averaged over years 1990-2012). The regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha.*
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 11: Constitutions and Education Expenditure. OLS estimations.

Dataset	PT		BCIM	
Dep.Var.	edspend (1)	edspend (2)	edspend (3)	edspend (4)
pres	−0.413		−0.512	
	(1.10)		(1.16)	
parl stab		0.832^{**}		1.10^{**}
		(0.65)		(0.89)
parl unstab		0.346		0.270
		(1.16)		(1.71)
Observations	75	75	84	84
Adjusted R^2	0.508	0.521	0.514	0.529

Notes: White heteroskedasticity-consistent standard errors in parentheses. *edspend* is the central government expenditure on education as percentage of the GDP as defined by the IMF-GFS (averaged over years 1995-2012). The regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha.*
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 12: Constitutions, Central Government Expenditure, District Magnitude and Ideology. OLS Estimations. PT and BCIM dataset.

Dataset	PT		BCIM	
Dep.Var.	cgexp (1)	cgexp (2)	cgexp (3)	cgexp (4)
pres	−3.327		−2.641	
	(1.54)		(2.05)	
parl stab		5.543^{**}		4.502^{**}
		(2.26)		(2.30)
parl unstab		−0.0559		1.484
		(2.03)		(1.64)
right_left	1.266	0.970	1.968	2.055
	(1.94)	(1.72)	(1.39)	(1.48)
magn	−0.228	−0.271	−0.333	−0.354
	(0.24)	(0.24)	(0.27)	(0.27)
Observations	75	75	85	85
Adjusted R^2	0.671	0.692	0.628	0.645

Notes: White heteroskedasticity-consistent standard errors in parentheses. The regressions include: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*. The addition controls *right_left* and *magn* are included. *right_left* reports the average ideological position of the executive from 1970 to 2012. Values are between 1 - right-oriented executive - to 3 - left-oriented executive. *magn* represents the district magnitude weighted by the country's population. Source: DPI dataset.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 13: Robustness checks with different stability indexes. OLS estimations

Dep.Var.	cgexp (1)	cgexp (2)	cgexp (3)	cgexp (4)
Dataset		PT		
parl stab	6.298***	5.932**	5.816***	6.064**
	(1.93)	(2.62)	(2.78)	(2.03)
parl unstab	1.383	−0.316	2.560	1.852
	(1.87)	(2.57)	(1.63)	(1.94)
Dataset		BCIM		
parl stab	5.206**	4.510**	4.656*	4.590**
	(2.11)	(2.08)	(1.98)	(2.17)
parl unstab	1.734	−1.700	2.522	2.827
	(1.98)	(1.81)	(2.07)	(1.50)

Notes: White heteroskedasticity-consistent standard errors in parentheses. A detailed explanation of the way in which the stability indexes have been assembled is reported in Section 3.1 and Section 4.2. Columns (1), (2), (3) and (4) report the stability indexes *gov life, gov end, year exec, year party*, respectively. The regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 14: Constitutions and Central Government Expenditure. OLS estimations. BCIM dataset

Dep.Var.	cgexp (1)	cgexp (2)
parl stab_1	2.67	1.96
	(3.01)	(2.93)
parl stab_2	4.91**	2.24
	(2.60)	(2.83)
parl stab_3	3.53	5.31**
	(3.01)	(2.99)
parl stab_4		3.26
		(3.33)
Observations	91	87
Adjusted R^2	0.67	0.71

Notes: White heteroskedasticity-consistent standard errors in parentheses. Parliamentary countries are split into 3 categories (columns (1)) and 4 categories (columns (2)) according to the index *gov life*. Note that in this specification *parl stab*(0)=*pres*. The regressions include the following controls: *age, lyp, trade, prop1564, prop65, gastil, federal, oecd, lpop, africa, asiae, laam, col_uka, col_espa, col_otha*.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 15: DESCRIPTIVE STATISTICS

	Pres (1)	Parl Unstab (2)	Parl Stab (3)	p(1,2) (4)	p(1,3) (5)	p(2,3) (6)	p(pt,bcim) (7)
CGEXP	22.9	32.1	32.5	0.00	0.00	0.32	0.43
CGREV	20.7	27.6	30.7	0.00	0.00	0.08	0.91
SSW	4.84	9.89	9.45	0.00	0.02	0.60	-
LYP	7.88	8.53	8.77	0.01	0.00	0.31	0.22
TRADE	68.8	79.5	92.5	0.27	0.05	0.31	0.05
GASTIL	2.92	2.49	2.12	0.09	0.07	0.08	0.06
PROP65	5.83	9.78	8.87	0.00	0.01	0.95	0.11
AGE	0.14	0.21	0.23	0.16	0.08	0.97	0.00
OECD	0.05	0.34	0.36	0.00	0.00	0.98	0.08
AFRICA	0.20	0.17	0.08	0.78	0.14	0.15	0.39
ASIA	0.10	0.14	0.21	0.63	0.19	0.58	0.08
LAAM	0.52	0.07	0.18	0.00	0.00	0.25	0.89
PPI	0.49	0.66	0.67	0.00	0.00	0.73	0.89

Notes: Entries in columns (1), (2) and (3) are mean values for constitutional categories when using the BCIM extended dataset. p(x,y) is the probability of falsely rejecting equal means across groups corresponding to columns x and y, under the assumption of equal variances. Column (7) is the probability of falsely rejecting equal means across the original PT dataset and the BCIM extended dataset, under the assumption of equal variances.

A.2 Technical notes

The basic framework as reported in equations (2) and (3) is based on the assumption that the explanatory variables are orthogonal to the error term, that is conditional on the set of regressors that enter the regression equation, the covariance between the type of constitution (S_{ij}, and $j = 1, ..., M$ indicating the constitution type) and the error is equal to zero:

$$Y_i = S_{ij}\gamma + X_i\beta + \nu_i \quad \text{and} \qquad\qquad cov(S_{ij} \mid X_i, \nu_i) = 0. \tag{6}$$

If the assumption is satisfied, then the OLS estimator of $\tilde{\gamma}$ is consistent. However, there are good enough reasons to suppose that the choice of the constitution is endogenous and determined by the same set of preferences that simultaneously affect both the dependent and independent variables (Acemoglu, 2005). Therefore, we introduce a selection equation where the constitutional choice depends on a set of observable characteristics, such that:

$$S_{ij} = W_{ij}\delta + \eta_{ij}; \tag{7}$$

where the subscript i and j indicate individual observations and outcome alternatives respectively (where $j = 0, 1, 2$ in our setting), and W_{ij} is a vector of covariates. Given the categorical nature of the dependent variable, equation (7) is estimated by a multinomial probit model that allows us to relax the independence of irrelevant alternatives assumption. Disturbances have multivariate normal distribution with positive-definite symmetric covariance matrix Σ. The alternative with the higher utility is chosen. This implies that the utilities of all other alternatives are negative relative to the chosen alternative. The likelihood is then:

$$L_i(\delta_j, \Sigma; S_i \mid W_i) = \Phi(-W_{ij}\delta; \Sigma_i); \tag{8}$$

where Φ is the multidimensional cumulative normal distribution. Note that Σ is indexed by i because it depends on which alternative is chosen in case i. However, this model makes necessary some identification restrictions given that not all elements of Σ can be identified. Thus, we take one alternative as reference altetnative (i.e. the presidential system in our empirical setting) and we reduce the selection process to two equations. For instance, if the reference category is $S_j = S_2$, then the transformed model becomes:

$$S_j^* = S_j - S_2 = W_j\delta^* + \eta_j^*,$$
$$\delta_j^* = \delta_j^* - \delta_2 \quad \text{and}$$
$$\eta_j^* = \eta_j - \eta_2.$$

The superindex $*$ differentiates the transformed model from the original three equations selection model. The covariance matrix reduces in its dimension to $(J-1) \times (J-1)$, i.e. in our setting it reduces to a bivariate distribution. The constraint $\sigma_{11} = 1$ is also needed to normalize the scale of the variances.

Instead of using the two-stages procedures, as the one proposed by Heckman (1997) that entails the inclusion of the predicted values from equation (7) into equation (6), we joint estimate the entire model through a maximum likelihood estimator. This allows us to take into account the full model's covariance structure of error terms that jointly have multivariate distribution. Given the nonlinear functional form of the multinomial equation, the parameters of this joint model can be identified even if the regressors that appear in the two equations are the same. However, to better qualify the identification strategy, we include a vector $(M \times 1)$ of excluded instruments (Z_i), under the assumption that the $Cov(Z_i, \nu_i \mid X_i) = 0$, such that $W_i = Z_i + X_i$. Note that the covariance between the vector of instruments and the error term of the headline equation (6) is equal to zero, conditional on the vector of controls X_i. This is the reason why we include the same X_i's in both equations (6) and (7). Otherwise, as noted by Acemoglu (2005), the residuals from the instrumenting equation that authomatically go into the headline equation are no longer orthogonal to the covariates, unless we assume that the X_i's omitted from equation (7) have no predictive power for the endogenous regressor, conditional on the other covariates.

The headline equation is estimated by a classical linear regression model of the following form:

$$Y_i = \beta_0 + \beta_1 S_{ij=0} + \beta_2 S_{ij=1} + \zeta X_i + \nu_i; \tag{9}$$

where $S_{ij=0}$ and $S_{ij=1}$ are two constitutional categories dummies (when category $j = 2$ represents the baseline group), X_i is the column vector of regressors and ε_i is the error term that we assume to have normal distribution. If we represent the zero-centered normal distribution by:

$$\phi(u; \sigma^2) = (1/\sqrt{2\pi\sigma^2})e^{-u^2/2\sigma^2};$$

then the likelihood function for observation i is:

$$L_i(\beta, \zeta, \sigma^2; y_i \mid S_i, X_i) = \phi(y_i - (\beta_0 + \beta_1 S_{ij=0} + \beta_2 S_{ij=1} + \zeta X_i); \sigma^2).$$

The likelihood function for the full model is thus composed by different components related to the headline linear regression and the transformed multinomial selection model. The parameters to be estimated are the following: the coefficients in the headline and selection equations (equations (10) and (7) respectively), the corresponding variances, the covariance between the error terms of the two equations, and the covariance within the selection equation. The resulting likelihood function has the following form:

$$
\begin{aligned}
& L(\beta_j, \delta_{ij}, \sigma^2_{\nu_i}, \sigma^2_{\eta_{ij}^*}, \sigma_{\nu_i \eta_{ij}^*}, \sigma_{\eta_{i0}^* \eta_{i1}^*}, Y, X, W, S^*) = \\
& = \prod_{\substack{S_0^*>0, \\ S_0^*-S_1^*>0}} [\phi(Y_i \mid S_j = S_0)\Phi(S_0^* > 0, S_0^* - S_1^* > 0)] \\
& \times \prod_{\substack{S_1^*>0, \\ S_0^*-S_1^*\leq 0}} [\phi(Y_i \mid S_j = S_1)\Phi(S_1^* > 0, S_0^* - S_1^* \leq 0)] \\
& \times \prod_{\substack{S_0^*\leq 0, \\ S_1^*\leq 0}} [\phi(Y_i \mid S_j = S_2)\Phi(S_0^* \leq 0, S_1^* \leq 0)];
\end{aligned} \tag{10}
$$

where $\phi(\cdot)$ describes the density function of the linear model and $\Phi(\cdot)$ the conditional cumulative distribution function of the bivariate selection process. The parameter of the covariance between the error terms of headline and selection equations indicates if the selection process is endogenous (covariance different from zero).

Voters' preferences and electoral systems.

The EuroVotePlus experiment in Italy.*

Luca Bettarelli†, Giovanna Iannantuoni‡, Elena Manzoni§ and Francesca Rossi¶

Motivated by the need to understand voting behaviour under different electoral rules, Laslier et al. (2015) have conducted an online experiment, the EuroVotePlus experiment, focusing on the effects of the different rules adopted to elect members of the European parliament on voters' behaviour. The experiment took place in several European countries in the three weeks before the 2014 elections for the European Parliament. This paper focuses on the Italian data. Firstly, we show that the behaviour of Italian respondents is consistent with the empirical findings at the European level. Then, we exploit the change from open list to closed list elections implemented in Italy in 1993 to investigate whether and how preferences over institutions are affected by experience. We find that respondents who voted using the open list system in Italy are more likely to prefer closed list systems, and that the effect is stronger the higher the number of open list elections the respondents have faced.

Keywords: European Parliament Election, Open list, Closed list, Voting rules.

JEL Classification: D7, C9.

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1 Introduction

The political economy literature (see, among many others, Cox 1997) has studied mass voting systems in modern democracies through the rational choice theory approach (since the seminal papers of Arrow, 1951; Downs, 1957). One of the most important issues is indeed the understanding of voting behaviour under different institutional setups. Scholars (for a review see Persson and Tabellini, 2000) have mainly focused on the electoral rules (electoral formula and district size), and regime types (presidential versus parliamentary systems). In this paper we exploit the data set of the EuroVotePlus (EVP) experiment (Laslier et al., 2015) to understand the voting behaviour of the Italian electorate when called to vote for the European Parliament under different electoral formulas, and the determinants of the electorate preferences over institutions. Italian data are particularly useful in addressing this last research question as national elections in Italy were characterised by open lists until 1993, and have been characterised by closed lists since 1993. Hence, discriminating by the age of the respondent, it is possible to correlate the preferences over electoral systems with the familiarity that voters have with the electoral system itself. Such analysis is of particular interest in addressing several issues, such as the European attitude of particular social groups (male/female, educated people, left/right oriented people) and the propensity to vote for pan-European parties.

The EVP experiment was conducted online during the three weeks before the 2014 elections for the European Parliament in the 28 European countries. An open-to-all multi-lingual website (www.eurovoteplus.eu) was created by a team of scholars (Laslier et al., 2015), where users were invited to learn more about European elections and rules used to elect MEPs, and then to participate in an online voting experiment. The website included a description of three electoral systems used at the time to elect MEPs in France (*closed list system*), Latvia (*open list system with preferential voting*), and Luxembourg (*open list system with cumulative voting and panachage*).[1] All visitors to the website were offered the opportunity to participate in a simulated vote for a pan-European election, where they were presented with ballots composed of seven lists corresponding to the seven political groups registered at the European Parliament at the time. Each list was made up of ten candidates randomly selected from the MEPs registered in the corresponding group. Each ballot provided the official photo of the candidate, name, nationality, group affiliation, and a link to the official page on the European Parliament website. Participants voted three times, once for every electoral system described. The experiment was covered by the press in Italy, where Giovanna

[1]A more extensive description of the experimental and electoral rules is provided in Section 2.

Iannantuoni was the national correspondent, giving Italy the second highest number of participants. In a restricted set of countries (which does not include Italy) respondents also voted on a simulated national election, where the ballot was composed of local candidates.

The EVP experiment was extensively analysed in several papers. Laslier et al. (2015) discuss the properties of the experiment by highlighting voters' behaviour under different electoral rules. The number of seats allocated to each party does not seem to vary substantially between closed and open lists, while the panachage rule seems to favour small parties over main ones. However, as somehow expected, more flexible electoral rules appear to have an impact on which candidates are elected within each party. In this respect, Laslier et al. (2015) suggest that voters tend to prefer candidates from their own country and that the effect seems larger for smaller countries. More specifically, Bol et al. (2016) show that the likelihood of assigning a positive vote to a co-national candidate is between seven and eight times higher than to a candidate of another nationality (controlling for all other characteristics). Moreover, with both closed and open lists the presence of at least one co-national candidate increases the probability of voting for that list by 48%. Harfst et al. (2015) analyze German data from EVP to investigate whether candidates' regional ties influence voting behaviour. The EVP in Germany had an extra feature that allows such analysis: only 50% of the respondents voted in a ballot that showed the candidate's Land as additional information. The authors find that preference votes are cast to support regional candidates. Moreover, they find the effect both on treated and untreated voters, suggesting that voters are interested in regional candidates and already know their identity. A second dimension in which elected candidates may differ under different electoral rules is gender. Laslier et al. (2015) highlight that flexible electoral systems appear to favour female candidates and that this effect is particularly strong when the sub-sample of female voters is considered. In more detail, Golder et al. (2015) find empirical evidence that voters favour female candidates under open list and panachage systems, while there seems to be no significant gender effect under a closed list system.

The contribution of this paper is twofold. First, we replicate the main results of the above literature related to the EVP experiment focusing on the population of the Italian respondents. We find that the Italian data are consistent with the empirical findings at the European level both in terms of home-candidate bias (Bol et al., 2016) and gender effect (Golder et al., 2015).

More interestingly, we address an original research question related to the Italian case, which is the effect the experience of an electoral rule has on its popularity. More precisely, we exploit the institutional change that occurred in 1993, which reformed the open list system to a closed list system. We show that respondents who experienced the pre-1993 system are less likely to prefer open list over closed list electoral rules. We also find that this effect is stronger the higher the number of elections experienced under the pre-1993 rule. The intuition of this latter result is related to the Italian electorate's perception of the political system in the early 90's, as well as to the relationship between open lists and political inefficiencies and corruption in Italy at that time (see Golden, 2003).

The paper is organized as follows: Section 2 describes the EVP experiment, Section 3 discusses home-candidate bias and gender effects for the Italian case, Section 4 focuses on the effect of experience over preferences for electoral rules, and Section 5 concludes.

2 The EuroVotePlus experiment

The EVP experiment focuses on the effects that electoral rules have on voters' behaviour. European member states implement several voting rules to elect the European Parliament. Among these rules, EVP focuses on the three rules which differ the most in the extent of voters' possible choices.

France (closed list). In the French system voters express their vote for one party list. Candidates in each party lists are ranked ex-ante and the ranking is used to determine who on the list is elected.

Latvia (open list with preferential voting). The electoral rule in Latvia allows voters to vote for one party list. Moreover, for each candidate on the chosen list, the voter can choose to express extra preference (with a +), or to cross out his/her name. The score of a candidate, which is used to determine who gets elected in the list, is equal to the number of votes for the list itself, plus the number of votes with a + for the candidate, minus the number of votes where the candidate's name is crossed out. Thus, the unique feature of this electoral system is the fact that the voter may cast a *positive* or *negative* vote for candidates on his chosen list by, respectively, assigning a + or crossing out the candidate's name. Otherwise, we say that the voter casts a *neutral* vote to those candidates which are neither endorsed with a + nor crossed out on the list.[2]

Luxembourg (open list with cumulative voting and panachage). The electoral rule in Luxembourg does not dictate that a voter votes for only one party list. The

[2] In the online experiment, the possibility of endorsing a candidate or crossing out his name was induced as follows: each of the ten candidates in the chosen list was given one point by the computer; the respondent could then endorse the candidate adding an extra point, or cross out the candidate removing the point already assigned.

voter has a number of votes equal to the number of positions to fill that can be given to candidates from different lists, giving up to two votes to the same candidate. The total number of votes received by each list determines the distribution of seats, and candidates in each list are elected on the basis of the number of votes received.

After a brief introduction and explanation of the three aforementioned systems, each participant in the EVP experiment has to vote in a simulated election for pan-European lists. More specifically, seven different lists (corresponding to the seven political groups that were actually registered at the European Parliament at the time of the election) containing ten candidates are shown to each participant in the experiment. The candidates on each list are randomly selected from the members of the European Parliament affiliated to the corresponding group at the time, and they are kept fixed for a given participant for the duration of the experiment. Each participant is then required to vote in this simulated election under the three different electoral systems. At the end of the experiment each respondent answers a questionnaire that elicits their opinions and preferences over the three electoral rules and over the possibility of electing MEPs with transnational lists, together with other personal characteristics.

3 Italians are indeed Europeans: consistency of the empirical findings with previous studies

We begin our analysis of the EVP Italian data showing that Italian respondents behave consistently with the empirical findings at the European level. The dataset is composed of 385 respondents who fully completed the experiment, 34% of which are female.

Table 1: DESCRIPTIVE STATISTICS

	Tot (1)	SD (2)	Male (3)	Female (4)	Min (5)	Max (6)
Education	25.76	3.35	25.64	25.98	16	35
Year of birth	1980	11.88	1981	1979	1939	1996
EuDemo	4.56	2.40	4.60	4.47	0	10
EuFeeling	2.69	0.98	2.73	2.62	0	5
EuTurnout	4.62	0.87	4.69	4.47	1	5
Political Interest	7.60	2.12	7.86	7.07	0	10
LeftRight	3.96	2.10	4.11	3.70	0	10
Travel	0.97	0.16	0.96	0.98	0	1
Language	1.79	0.69	1.70	1.98	0	3

Notes: EVP dataset. Entries in columns (1), (3) and (4) are mean values, while column (2) reports the standard deviation. Columns (5) and (6) indicate the minimum and maximum values respectively.

Among the independent variables of interest we include gender, education and age of respondents.[3] Education is defined as either the age at which the respondent finished their education, or his/her current age where the respondent is still at school. In addition, we focus on three indicators characterizing the aptitude of people toward Europe and European institutions. More precisely, we focus on the level of satisfaction with democracy at a European level (*EuDemo*), with the self-perception of the respondents as European (*EuFeeling*) and with the propensity to vote in the European elections (*EuTurnout*).[4]

We also include indicators for political orientation and interest in politics in our analysis. The former (*LeftRight*) is a categorical variable ranging from zero to ten, with higher values corresponding to right oriented respondents, while the latter (*Political Interest*) is a categorical variable ranging from zero to ten, with higher values corresponding to higher interest.

Descriptive statistics in Table 1 show a slight feeling of detachment from European institutions, together with a large intention to vote in the European Parliament election and a large interest in politics. The sample seems biased toward left-oriented people, especially for the subgroup of female respondents. As expected, the proportion of respondents having visited at least one foreign country at least once (*Travel*) is very large; the average number of foreign languages (*Language*) spoken by the respondents is close to two.

3.1 Home candidate bias

We replicate the analysis of Bol et al. (2016) where the authors test three hypotheses related to the co-nationality between voters and candidates, the interaction between the *co-nationality effect* and the electoral system in use. As mentioned in Section 2, the EVP experiment proposes pan-European lists in which each voter is faced with candidates with different nationalities. The authors focus on respondents from France, Germany and Sweden. The analysis suggests that voters have a positive feeling toward pan-European lists with 59% of respondents that like this form of election. The presence of candidates of the same nationality may however affect voters' behaviour, and such

[3] Gender=1 for female respondents.

[4] Such indicators are based on the three following questions:

i) From 0 to 10, how satisfied are you with democracy at European level?

ii) If 0 indicates you only feel Italian and 5 only European, where do you place yourself?

iii) From 0 to 5, how sure are you sure that you will vote in the next European Parliament election?

Responses to questions i), ii) and iii) are used to construct the variables *EuDemo*, *EuFeeling* and *EuTurnout*, respectively.

an effect may depend on the electoral system in use. Bol et al. (2016) find evidence in favour of the probability of a positive (negative) preference being higher when the candidate is (is not) a co-national of the voter. Secondly, voters seem to prefer lists which include a co-national candidate. Hence, at the moment of casting their votes, voters tend to endorse co-national candidates, thus creating a *home-candidate bias*. Therefore, the co-nationality effect may partially frustrate the effects of the pan-European lists.

We replicate the same analysis for Italian respondents.[5] First of all, Italian voters seem to appreciate pan-European lists more than their European fellows. Indeed, 69% of Italian respondents approve of the idea of having pan-European lists.[6]

To test the co-nationality effect on voting preferences, we exploit the design of the EVP experiment, and, in particular, the possibility in the open list system to either endorse (positive vote) or cross-out (negative vote) a specific candidate on the list.[7]

We investigate the effects of co-nationality on the probability of expressing negative/positive votes for a candidate with a Multinomial logit model. The categorical outcome is the probability of crossing out/endorsing a candidate under the open list system. The baseline category in our setting is represented by the neutral vote. The main predictor is the dummy variable *Co-national Candidate* which identifies Italian candidates. In addition, we include a set of covariates to control for participant and candidate characteristics. At the participant level we control for age, gender and education of respondents. Furthermore, we include indicators for political orientation (*LeftRight*), interest in politics (*Political Interest*) and personal feeling about European institutions (*EuFeeling*) Then, we also include the total number of points assigned by each respondent (*Voter's Total Points*). At the candidate level, we control for age and gender of each candidate. We account for the possibility of party-level bias by including party-specific dummy variables.

[5]We highlight that in the replication of Bol et al. (2016) we cannot introduce one of their main controls, which is the dummy *Consistent* indicating if the pan-European list vote is consistent with the vote cast by the respondents in the national ballot (i.e. same party group), as national ballots were not introduced in the Italian version of EVP.

[6]If we consider the full EVP sample including all European countries, the percentage of respondents approving of pan-European lists is almost 60%. The null hypothesis of equal means between the percentage of people that approve the pan-European lists in Italy and in the rest of Europe can be rejected at the 1% level.

[7]In our sample the votes in the open list systems are distributed as follows: 18.83% negative votes, 20.96% positive votes and 60.21% neutral votes. The candidate's position within the party-list seems to have no effect on the decision of the voters, as 48% of the endorsed candidates are located in the first five positions on the list.

Table 2: Predicting preference votes and home-candidate bias under open list

	Negative vote		Positive vote	
	Coeff.	RRR	Coeff.	RRR
Co-national Candidate	-0.72^{*}	0.44^{*}	1.73^{***}	5.63^{***}
	(0.28)	(0.14)	(0.15)	(0.85)
Candidate's Gender	-0.24^{*}	0.79^{*}	0.53^{***}	1.70^{***}
	(0.13)	(0.10)	(0.11)	(0.18)
Candidate's Age	-0.003	0.99	0.10^{*}	1.01^{**}
	(0.006)	(0.006)	(0.005)	(0.005)
Voter's Total Points	-0.45^{***}	0.64^{***}	0.19^{***}	1.22^{***}
	(0.02)	(0.02)	(0.03)	(0.34)
Gender	0.27^{*}	1.26^{*}	0.26^{**}	1.28^{**}
	(0.12)	(0.15)	(0.11)	(0.15)
Education	-0.01	1.00	-0.03	0.98
	(0.01)	(0.02)	(0.02)	(0.015)
Age	-0.007	-1.00	-0.005	-1.00
	(0.005)	(0.005)	(0.004)	(0.004)
LeftRight	-0.06	0.94	-0.05	0.95
	(0.04)	(0.04)	(0.03)	(0.03)
Political Interest	-0.04	0.96	-0.02	0.98
	(0.02)	(0.03)	(0.03)	(0.03)
EuFeeling	0.14^{**}	1.28^{**}	0.10^{*}	1.10^{*}
	(0.06)	(0.06)	(0.06)	(0.06)
Constant	-31.776^{***}	-4.86^{**}	7.07^{**}	$< 0.01^{**}$
CHI^2	544.66^{***}			
N	19124	19124	19124	19124

Notes: White heteroskedasticity-consistent standard errors in parentheses. Entries are coefficients and the relative risk ratios (RRR). Dummies indicating the party voted under the open list system are not reported in the table.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The results, reported in Table 2, are consistent with the findings of Bol et al. (2016). Italian respondents, when allowed to express their preferences, reward candidates of the same nationality. Indeed, co-nationality represents a highly significant predictor of the probability of endorsing a particular candidate under the open list system. Also, from Table 2, female candidates appear more likely to be rewarded, and this effect is highly significant. If we run the same model restricting the sample to female respondents, the relative risk ratio increases to 2.97 and it is significant at the 1% level.

Furthermore, we test whether lists proposing a higher number of co-national candidates are more likely to be voted for. This effect should be larger under the closed list system, given that under the open list system voters can directly reward a co-national candidate. As in Bol et al. (2016), co-nationality is measured in three ways. Firstly, we consider a discrete variable indicating the total number of co-national candidates on the list.[8] Secondly, we consider a categorical variable where the number of co-nationals is included in the model as a single covariate, to verify whether the co-nationality effect is linear or not. Finally, we consider a dummy equal to one if there is at least one co-national candidate on the list. We run conditional logit models predicting the probability of voting for each list under the closed and open list systems. At respondents' level, we introduce fixed effects to account for heterogenous individual characteristics. At the candidate level, we include the number of female candidates and the average age of the candidates on each list. Results are reported in Appendix (Table 5 (closed list system) and 6 (open list system)).

Results are consistent with expectations. The presence of co-national candidates positively affects the probability of voting for a specific list. Moreover, the co-nationality effect is higher between lists including zero co-nationals and those that include one or two co-national candidates. After a threshold of two, the inclusion of an additional co-national candidate does not affect the probability of voting for the list. Results are robust to the introduction of the binary classification of the dependent variable. Furthermore, the co-nationality effect seems to be slightly larger under the closed list system. This follows theoretical expectations.

Among the other covariates, we highlight that the presence of female candidates is positively related to the probability of voting for a list. This further evidence motivates the analysis of possible gender-related effects across different electoral rules.

3.2 Gender bias

The experimental design gives us the opportunity to isolate some gender-related effects. The presence of women in the party list is supposed to favour female participation in the electorate and to increase the share of votes to women candidates. This mechanism should be stronger in more open electoral systems. However, party leaders may promote the presence of female candidates even in the case of closed list systems if they expect women to favour lists with a higher percentage of female candidates.

The work by Golder et al. (2016) investigates the presence of gender effects by analysing the EVP data at the European level. We again focus on the case of Italy. In line with the previous literature, we expect that electoral systems which provide voters with the opportunity to express their own preferences would favour female candidates. Moreover, female voters, by voting for female candidates, should in theory increase female representation, regardless of the electoral system in use. Finally, the left/right orientation of voters may affect the probability of voting for female candidates.

[8] The number ranges from zero to six.

We test the effect that different electoral rules have on the proportion of female candidates and on the voting behaviour of respondents. As the ballot given to participants is the same under the three voting rules, we are able to evaluate the effect of the electoral system while keeping other cultural, historical, economic and contextual factors, which normally vary with the electoral rules and may generate biases in the causal estimation, constant. Women represent on average 32% of the proposed candidates, with a 5% standard deviation. The proportion of women in the lists varies depending on the party and on the party ideological position, with leftist and less extreme parties associated with a larger presence of female candidates.[9]

We evaluate the propensity to vote for female candidates by considering the proportion of votes gathered by women candidates under different electoral rules. Specifically, under the closed list system, we create a measure of female support from the proportion of female candidates in the chosen list, given that voters are not allowed to vote directly for individual candidates. Under both open list and panachage systems, we measure female support by computing the proportion of points assigned to female candidates over the total number of points distributed by each respondent. The main predictor is the gender of the respondents. A measure of the overall proportion of women in the ballot always enters the regression. Other controls include voter-related characteristics such as age, education, interest in politics, left/right ideology and the European feeling. We also include a set of covariates indicating the party chosen under the closed list system in order to control for the effect of the party ideology.[10]

[9]Note that party lists have been randomly assembled by the roster of existing MEPs; thus, parties with lower female representation would naturally provide fewer potential female candidates for the experiment.

[10]The Conservatives and Reformists party group is the control group.

Table 3: Proportion of votes to female candidates under closed, open and panachage systems

	Closed	**Open**	**Panachage**
Gender	-0.034^{*}	0.048^{***}	0.113^{***}
	(0.002)	(0.003)	(0.004)
Freedom and Democracy	-0.301^{***}	-0.28^{***}	-0.33^{***}
	(0.007)	(0.007)	(0.01)
Alliance Liberals and Democrats	-0.36^{***}	-0.25^{***}	-0.18^{***}
	(0.003)	(0.007)	(0.008)
Progressive Alliance of Socialists and Democrats	-0.073^{***}	0.006	-0.014^{**}
	(0.004)	(0.006)	(0.007)
Greens Eu Free Alliance	0.1^{***}	0.077^{***}	0.024^{***}
	(0.05)	(0.007)	(0.007)
United Left-Nordic Greeen Left	0.33^{***}	-0.003	0.0106^{*}
	(0.03)	(0.004)	(0.006)
Eu People Party	-0.11^{***}	-0.174^{***}	-0.038^{***}
	(0.04)	(0.005)	(0.008)
% WomenBallot	0.95^{***}	1.17^{***}	0.757^{***}
	(0.20)	(0.27)	(0.039)
Education	0.04^{***}	0.006^{***}	0.02^{***}
	(0.00)	(0.00)	(0.00)
Age	0.001^{***}	0.002^{***}	0.004^{***}
	(0.00)	(0.00)	(0.00)
LeftRight	-0.003^{***}	-0.019^{***}	-0.027^{***}
	(0.00)	(0.001)	(0.001)
Political Interest	0.05^{***}	0.001^{*}	0.005^{***}
	(0.00)	(0.00)	(0.00)
EuFeeling	0.005^{***}	0.001^{**}	0.034^{***}
	(0.00)	(0.00)	(0.00)
Constant	1.56^{***}	3.35^{***}	7.56^{***}
R2	0.51	0.39	0.27
N	19180	19040	19180

Notes: White heteroskedasticity-consistent standard errors in parentheses. Among party covariates, the *Conservatives and Reformists* party represents the baseline category.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Results are consistent with the empirical findings at the European level (Golder et al., 2016). Female respondents are more likely to reward candidates of the same gender and the gender effect is higher the more flexible the electoral system is; the voting decisions when voters are not allowed to express their preferences toward specific candidates seem largely to depend on party affiliation; left-oriented voters are more inclined to reward female candidates as well as voters that are more educated and those that are more interested in politics.

4 Preferences over institutions: the importance of experience

With its final questionnaire, the EVP experiment investigates respondents' preferences over electoral rules, and it allows us to perform an analysis of the determinants of such preferences. More precisely, every respondent is asked to rate his approval of open list, closed list and panachage rules. The ratings of open vs. closed lists are of particular interest in the Italian case, given the constitutional reform which occurred in 1993, which abolished preference voting in national elections.

In the literature there is an open debate on whether it is preferable to adopt open-list or closed-list systems in terms of several outcomes, including among others efficiency, government corruption, minority representation (see among others Cox, 1997; Persson and Tabellini, 2000; Blumenau et al., 2016). It is however commonly agreed (see among others Lepre, 1993; Zincone, 1995; Barbagallo, 1998; Ginsborg, 2001) that the pre-reform open list system in Italy generated distortions in politicians' behaviour, up to the point of creating incentives to significantly decrease the quality of public services. Golden (2003) argues that public administration performed poorly because politicians were more concerned about their re-election prospects than policy implementation. Indeed, evidence suggests that open lists somehow promoted a system of political patronage "where patronage consisted of concrete individual benefits (jobs, especially in public administration) and help in negotiating the complex legal regulations affecting daily life. The response of many voters to such a system was to offer their votes in exchange for patronage and constituency services" (Golden, 2003).

In order to remedy the distortion and misbehaviour arising from preference voting, the first referendum was called in 1991, in which voters were asked to remove the option of expressing multiple preferences. This was the first but important step away from the distortions occuring in Italy as a result of preference voting. As Ginsborg (2001) explains "Being able to choose up to four candidates of the same political party had been a historic vehicle of political clientelism, a means of tying local clients to patrons, and of building factional strengths in specific areas."[11] The proposed reduction in the number of preferences was opposed by a large number of politicians, in particular by those who relied on preference voting to be elected. The high turnout, together with the fact that 95.6% of voters approved of the change, confirmed that the majority of the population had a profound dislike of the open list system, and it was generally interpreted as a call for a general reform of the electoral rule (Mack Smith, 1997). After a second referendum, held in 1993, the electoral rule was modified. The 1993 Mattarella law introduced a system that elected legislators with a mix of plurality

[11] Ginsborg P. (2001), Italy and Its Discontents. Family, Civil Society, State 1980-2001, p.173.

rule and closed list proportional rule, completely abolishing the open list system from national elections. It is thus of particular interest to investigate whether having voted at least once under the open list system systematically changes the perception of what is preferable.

The literature on preferences over institutions is quite recent. Aldrich et al. (2014) analyse the 2004 ballot initiative for a constitutional amendment (Amendment 36) that aimed to change the electoral rules for the election of the US Electoral College in Colorado. The proposed reform was to switch to a proportional electoral rule from the status-quo winner-take-all rule. The amendment did not pass. The authors show that self-interest was a strong determinant of the vote on the ballot. A citizen was much more likely to be in favour of the amendment if his most preferred candidate was unlikely to be the frontrunner in the state. Fournier et al. (2011), instead, show how citizens' values may be relevant in determining their preferences over institutions. The authors analyse the work of citizen assemblies on electoral reforms in Canada and the Netherlands and discuss how narrow self-interest seems to have no impact on their decisions. Citizens formed their decisions on the basis of both what they valued as most important in a representative democracy, and what they believed were the best institutions that could achieve their objective. The closest work to our analysis is Blais et al. (2015). The authors analyse data from a large internet-based quasi-experiment associated with the first round of the French 2012 presidential election. Blais et al. (2015) confirm that self-interest is one of the determinants of preferences over electoral rules, as they find that respondents who prefer one of the two viable candidates (who will very likely reach the second round, and who will possibly win the election) are 20% more likely to prefer the status-quo two-round system to alternative systems. Moreover, they find a correlation between ideology and preferences over institutions, as they have evidence of a right-wing bias in favour of single-vote systems.

A characterisation of preferences over institutions that is motivated by self-interest or values requires citizens to have a deep understanding of the consequences of each institution. Such an assumption may not be very realistic. Karp (2006) provides evidence suggesting that citizens are able to understand the consequences of voting under the status quo. If this is the case, the analysis of Italian data may be even more interesting. A subgroup of the Italian respondents have experience of multiple electoral rules which may enable them to understand the consequences of multiple rules even better and to express preferences depending on the implied outcome.

4.1 Empirical analysis

We analyse whether respondents who experienced voting in national elections under open list systems have a different propensity towards them. Moreover, we investigate whether this effect becomes greater the greater the experience of voting under an open list system. Our hypothesis is that respondents who voted under the open list electoral rule should be more likely to prefer the closed list system, in light of their acquired awareness concerning the inefficiencies generated by preference voting in Italy. According to our data, and consistent with expectations, the percentage of people who strictly prefer open lists over the closed list system is somewhat low, with only 11% of the sample preferring the former open lists. The closed list system is preferred by 70% of respondents, while 19% are indifferent between the two systems.[12] The sample includes respondents born between 1939 and 1996. As the 1994 national election was the first election in Italy adopting the closed list system, respondents born after 1975 have no direct experience of voting under the open list proportional rule in national elections. Our main regression analyses whether having experienced at least one election with open list proportional rule affects the respondent's opinion. We create a dummy (*Experience*) which identifies respondents born before 1974, that is, the subgroup of respondents that have experienced at least one election with the open list system. They represent 20% of the sample.[13] Respondents in the two subgroups differ not only in terms of experience of open list systems, but also in terms of age. Hence, we expect to observe two effects: a general age effect, and a specific experience effect.

We first estimate a simple linear probability model where the dependent variable is a binary variable equal to one if the respondent strictly prefers the open list system to the closed list system, and zero otherwise. Results are reported in Table 4.

[12]The questionnaire asks respondents to express their opinion about the proposed electoral systems on a scale between 0 and 5. We exploit this information to build three binary variables indicating if the respondent strictly prefers open lists over closed list, or viceversa, or if the respondent is indifferent between the two.

[13]We drop respondents born in 1974 as we don't know their month of birth.

Table 4: Voters preferences over electoral systems and experience

	(1)	(2)	(3)	(4)
Experience		-0.103^{***}		
		(0.0106)		
Year of birth	-0.00249^{***}	-0.00581^{***}	-0.0047^{***}	-0.006^{***}
	(0.00023)	(0.000384)	(0.000414)	(0.0005)
Elections			-0.0184^{***}	
			(0.00317)	
Election1				-0.0905^{***}
				(0.0122)
Election2				-0.106^{***}
				(0.014)
Election3				-0.121^{***}
				(0.018)
Election4				-0.083^{***}
				(0.029)
Election5				-0.039
				(0.027)
Election6				0.034^{*}
				(0.021)
Election7				-0.303^{***}
				(0.041)
Election8				-0.393^{***}
				(0.026)
Election9				-0.406^{***}
				(0.0260)
Gender	0.00870	0.00740	0.00789	-0.0073
	(0.00562)	(0.00558)	(0.00569)	(0.0057)
Education	-0.0039^{***}	-0.0045^{***}	-0.00603^{***}	-0.0009
	(0.0011)	(0.0012)	(0.0012)	(0.0013)
EuDemo	0.0131^{***}	0.0135^{***}	0.0126^{***}	0.011^{***}
	(0.00114)	(0.00114)	(0.00114)	(0.0013)
Political Interest	0.0045^{***}	0.0044^{**}	0.00414^{***}	0.0039^{**}
	(0.00156)	(0.00156)	(0.00155)	(0.00156)
EuFeeling	0.012^{***}	0.013^{***}	0.00129^{***}	0.015^{***}
	(0.0032)	(0.00333)	(0.0031)	(0.0032)
Travel	0.349^{***}	0.360^{***}	0.358^{***}	0.347^{***}
	(0.0212)	(0.0205)	(0.0211)	(0.021)
Language	0.0339^{***}	0.0417^{***}	0.0398^{***}	0.044^{***}
	(0.00431)	(0.00427)	(0.00428)	(0.0040)
EuTurnout	0.0194^{***}	0.0195^{***}	0.0184^{***}	0.025^{***}
	(0.00330)	(0.00328)	(0.00329)	(0.0034)
LeftRight	0.0167^{***}	0.0170^{***}	0.0167^{***}	0.0156^{***}
	(0.00124)	(0.00124)	(0.00124)	0.00128)
Constant	4.59^{***}	11.19^{***}	9.08^{***}	11.60^{***}
R^2	0.49	0.51	0.54	0.49
N	17710	17710	17710	17710

Notes: White heteroskedasticity-consistent standard errors in parentheses. Elections(1)-(9) indicate the number of open list elections faced by the experiment participants. Election(0) is the baseline category representing the subgroup of people born after 1974, that are the ones with no direct experience of open list elections. The dependent variable is a dummy equal to one if respondents strictly prefer open lists over the closed lists.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Column (1) displays results of our baseline regression, where we include the respondent's year of birth in addition to standard controls. The effect of the year of birth is significant and negative, showing that younger respondents have a lower probability of preferring the open list system. We then introduce the dummy *Experience* in the set of regressors (see Table 4, Column (2)). Results support our conjecture about the link between preferences and experience: those who have directly experienced the open list system are less likely to prefer open lists compared to the group of respondents who have never voted under that system, controlling for other characteristics. The effect of the respondent's year of birth remains negative and highly significant, as in the baseline case.

Ideology and preferences over electoral rules. An interesting effect on the respondents' preferences over the choice between open and closed list systems is given by ideology. We find that right-wing respondents are more likely to prefer open list systems than left-wing respondents. This may seem in contrast with Blais et al. (2015) who find that left-wing respondents favour more complex electoral rules. They motivate their findings in a twofold way. First, they report evidence suggesting that left-wing voters are more open to innovation, thus more easily favouring alternatives to the status quo. It has to be noted that the status quo in France was clearly the two-round electoral rule, while it is less clear what the Italian respondents may perceive as innovative, as many of them voted in several elections under an open list system. Moreover the debate on possible further changes of the electoral rule in Italy was very important even after the 1993 reform, possibly also giving a flavour of uncertainty over the duration of the post-reform electoral rule. Secondly, they suggest that this may be due to a different perception of the role of elections according to the voter's ideology as, in their presidential election choice, right-wing voters may be more focused on choosing the person who will lead the country, while left-wing voters may be more concerned about the possibility of voters expressing their opinions in the electoral process (*voter choice*). This issue is however less relevant for our analysis, as we are comparing two proportional rules, so that right-wing voters cannot focus on voting only for the leader.

4.2 Robustness checks

To further investigate the effects of experience, we build a count variable - *Elections* - equal to the total number of elections that were held using preference voting when the respondent was of voting age. The variable ranges from zero (for people born after 1974) to nine (for people born before 1940). We expect a negative effect of the number of open list elections a respondent has experienced on the probability of preferring the open list system. Indeed, Table 4, Column (3), validates our hypothesis. As a robustness check, we also refine the latter analysis by introducing a set of dummy variables *Electionj* for $j = 1, ..., 9$ which identify respondents who experienced j elections under the open list system when they were of voting age. Table 4, Column (4) shows that aversion to the open list system generally increases with the number of such elections. The few inconsistencies in Column (4) may be explained by the small number of respondents belonging to some of the subgroups (specifically to those defined by *Election*4, *Election*5 and *Election*6). Our results are robust even when coefficients are estimated by a Logit rather than by a linear probability model. Results are reported in Table 7, in the Appendix.

As a further robustness check, we replicate the same analysis after restricting the sample to subjects who prefer either system, i.e. after taking out respondents who are indifferent to the closed and open list systems. Results can be found in Table 8, in the Appendix.

Finally, we perform a placebo test, replicating the same analysis on a different dataset. We show that *Experience*, which has a strong and robust effect in Italy, has no effect on the European data. Hence, we run the baseline regression, in which only *Year of birth* is included as an additional regressor, and the specification in which *Experience* is also included both for the set of all European countries except Italy (see the Appendix, Table 9, Columns (1) and (2)) and for France, which is the country with the largest number of observations (see the Appendix, Table 9, Columns (3) and (4)). We find that *Experience* is no longer statistically significant and, moreover, the effect of *Year of birth* is positive and significant, thus further supporting our interpretation.

5 Conclusions

Electoral rules influence voters' behaviour, as suggested by the theoretical and empirical literature. In this paper, we also show that experience of electoral rules affects voters' preferences over electoral systems.

The analysis is based on the Italian data from the EuroVotePlus experiment which took place in the weeks preceding the 2014 elections for the European Parliament. We find that the influence of electoral rules on the behaviour of the Italian respondents is consistent with the empirical findings at the European level. More precisely, we find that, although Italians are more in favour of pan-European elections than average Europeans, they nonetheless display the co-nationality effect and favour Italian candidates in elections. We also find evidence of a gender effect, showing that women receive more votes under open list rules.

Moreover, the institutional reform that occurred in Italy in 1993, which replaced an open list system with a closed list system, allowed us to investigate the role of experience of a system in determining the respondents' preferences over electoral rules. In a political environment where open list systems were (perceived as) leading to corruption, inefficiencies and sub-optimal behaviour of politicians, we find that respondents with greater experience of elections held under the open list system were significantly less likely to prefer open list over closed list electoral rules.

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6 Appendix

Table 5: Predicting list voting and home candidate bias (Closed list)

	Continuous		Categorical		Binary	
	Coef.	OR	Coef.	OR	Coef.	OR
Number of co-nationals	0.112^{**}	1.12^{**}				
	(0.052)	(0.058)				
Categories:						
One co-national			0.30^{**}	1.35^{***}		
			(0.14)	(0.19)		
Two co-nationals			0.50^{***}	1.65^{***}		
			(0.16)	(0.26)		
Three co-nationals			-0.01	0.99		
			(0.27)	(0.27)		
Four co-nationals			0.44	1.56		
			(0.38)	(0.60)		
Five co-nationals			0.08	1.08		
			(1.08)	(1.17)		
Six co-nationals			0.38	1.42		
			(0.45)	(0.53)		
At least one co-national					0.33^{***}	1.39^{***}
					(0.12)	(0.16)
Number of women	0.204^{***}	1.23^{***}	0.19^{***}	1.22^{***}	0.19^{***}	1.21^{***}
	(0.034)	(0.042)	(0.034)	(0.041)	(0.032)	(0.039)
Age (mean)	0.987^{***}	2.68^{***}	0.99^{***}	2.68^{***}	0.98^{***}	2.67^{***}
	(0.129)	(0.346)	(0.034)	(0.36)	(0.13)	(0.35)
CHI^2	90.3^{***}		217.2^{***}		90.1^{***}	
N	2695		2695		2695	

Notes: White heteroskedasticity-consistent standard errors in parentheses. Entries are coefficients and odd ratios (OR). The model is a multinomial conditional logit with subject level fixed effects.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Predicting list voting and home candidate bias (Open list)

	Continuous		Categorical		Binary	
	Coef.	OR	Coef.	OR	Coef.	OR
Number of co-nationals	0.086^{**}	1.09^{**}				
	(0.039)	(0.043)				
Categories:						
One co-national			0.30^{**}	1.36^{**}		
			(0.13)	(0.17)		
Two co-nationals			0.49^{***}	1.63^{***}		
			(0.15)	(0.25)		
Three co-nationals			-0.31	0.72		
			(0.27)	(0.20)		
Four co-nationals			0.55	1.74		
			(0.37)	(0.64)		
Five co-nationals			-0.26	0.77		
			(1.06)	(0.82)		
Six co-nationals			0.08	0.98		
			(0.34)	(0.53)		
At least one co-national					0.30^{**}	1.35^{***}
					(0.11)	(0.15)
Number of women	0.24^{***}	1.27^{***}	0.23^{***}	1.26^{***}	0.23^{***}	1.26^{***}
	(0.032)	(0.040)	(0.032)	(0.040)	(0.030)	(0.038)
Age (mean)	0.745^{***}	2.11^{***}	0.74^{***}	2.10^{***}	0.73^{***}	2.08^{***}
	(0.089)	(0.189)	(0.092)	(0.19)	(0.09)	(0.19)
CHI^2	115.6^{***}		253.4^{***}		116.2^{***}	
N	2695		2695		2695	

Notes: White heteroskedasticity-consistent standard errors in parentheses. Entries are coefficients and odd ratios (OR). The model is a multinomial conditional logit with subject level fixed effects.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Voters preferences over electoral systems and experience (robustness check).

	Logit	**Logit**
Experience		-0.0955^{***}
		(0.0103)
Year of birth	-0.00242^{***}	-0.00523^{***}
	(0.00021)	(0.00034)
Gender	0.00691	0.0045
	(0.00575)	(0.0057)
Education	-0.00364^{***}	-0.0041^{**}
	(0.0011)	(0.0010)
EuDemo	0.0127^{***}	0.0123^{***}
	(0.00112)	(0.0011)
Political Interest	0.00214	0.00243
	(0.00167)	(0.00171)
EuFeeling	0.0113^{***}	0.0119^{***}
	(0.00314)	(0.0031)
Travel	0.214^{***}	0.220^{***}
	(0.0099)	(0.0096)
Language	0.0409^{***}	0.0423^{***}
	(0.00401)	(0.0039)
EuTurnout	0.0153^{***}	0.0144^{***}
	(0.0026)	(0.0025)
LeftRight	0.0154^{***}	0.0153^{***}
	(0.00116)	(0.0012)
R^2	0.51	0.55
N	17710	17710

Notes: White heteroskedasticity-consistent standard errors in parentheses. Entries are average marginal effects. The dependent variable is a dummy equal to one if respondents strictly prefer open lists over closed lists.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Voters preferences over electoral systems and experience (robustness checks).

	OLS	**OLS**	**Logit**	**Logit**
Experience		-0.104^{***}		-0.0925^{***}
		(0.012)		(0.011)
Year of birth	-0.00255^{***}	-0.00249^{***}	-0.0047^{***}	-0.0052^{***}
	(0.00024)	(0.000415)	(0.000216)	(0.0004)
Gender	0.00764	0.00690	0.00568	0.00383
	(0.00606)	(0.00604)	(0.00609)	(0.0061)
Education	-0.00267^{**}	-0.00259^{**}	-0.00247^{**}	-0.00318^{***}
	(0.0012)	(0.00120)	(0.0011)	(0.0011)
EuDemo	0.0125^{***}	0.0127^{***}	0.0120^{***}	0.012^{***}
	(0.00124)	(0.00123)	(0.00118)	(0.0012)
Political Interest	0.00672^{***}	0.00645^{***}	0.00452^{**}	0.0046^{**}
	(0.00166)	(0.00166)	(0.00181)	(0.00184)
EuFeeling	0.015^{***}	0.015^{***}	0.00141^{***}	0.014^{***}
	(0.0035)	(0.00342)	(0.0033)	(0.0033)
Travel	0.329^{***}	0.337^{***}	0.209^{***}	0.216^{***}
	(0.0214)	(0.021)	(0.011)	(0.010)
Language	0.0380^{***}	0.0403^{***}	0.039^{***}	0.041^{***}
	(0.0042)	(0.0041)	(0.0041)	(0.0040)
EuTurnout	0.0212^{***}	0.0201^{***}	0.0167^{***}	0.016^{***}
	(0.0035)	(0.0034)	(0.0027)	(0.0027)
LeftRight	0.0184^{***}	0.0185^{***}	0.0168^{***}	0.0167^{***}
	(0.00132)	(0.00132)	(0.00122)	(0.0013)
Constant	4.69^{***}	10.92^{***}		
R^2	0.51	0.55	0.52	0.56
N	16450	16450	16450	16450

Notes: White heteroskedasticity-consistent standard errors in parentheses. Results obtained by a linear probability model (Columns (1) and (2)) and by a logit model (Columns (3) and (4)) when sample is restricted to people who are not indifferent between closed and open list systems. Entries for the Logistic regressions are average marginal effects.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 9: Voters preferences over electoral systems and experience (robustness check).

	(1)	**(2)**	**(3)**	**(4)**
Experience		-0.012		0.0075
		(0.074)		(0.021)
Year of birth	0.0041^{***}	0.0038^{***}	0.0022^{***}	0.0024^{***}
	(0.00013)	(0.00025)	(0.000216)	(0.0007)
Gender	-0.0128^{**}	-0.0129^{**}	-0.0894^{***}	-0.0896^{***}
	(0.0044)	(0.0044)	(0.0103)	(0.0103)
Education	-0.00161^{***}	-0.0016^{***}	0.0158^{***}	0.0158^{***}
	(0.00037)	(0.00037)	(0.0017)	(0.0017)
EuDemo	0.00067^{***}	0.00069^{***}	0.0027^{***}	0.0026^{***}
	(0.00016)	(0.00016)	(0.0003)	(0.0003)
Political Interest	-0.0039^{***}	-0.0039^{***}	-0.066^{***}	-0.066^{***}
	(0.0011)	(0.0011)	(0.0039)	(0.0039)
EuFeeling	0.0280^{***}	0.0283^{***}	0.0345^{***}	0.0344^{***}
	(0.0015)	(0.0015)	(.00538)	(.0054)
Travel	-0.219^{***}	-0.219^{***}	-0.598^{***}	-0.596^{***}
	(0.012)	(0.012)	(0.016)	(0.017)
Language	-0.0368^{***}	-0.0367^{***}	-0.0047	-0.0049
	(0.0026)	(0.0026)	(0.0072)	(0.0072)
EuTurnout	-0.0256^{***}	-0.0256^{***}	-0.035^{***}	-0.035^{***}
	(0.00233)	(0.00233)	(0.004)	(0.004)
LeftRight	0.00016	0.00017	0.0005^{*}	0.0005^{*}
	(0.00147)	(0.00143)	(0.0003)	(0.0003)
Constant	-7.37^{***}	-6.68^{***}	-3.01^{***}	-3.46^{***}
Sample	EU	EU	FR	FR
R^2	0.31	0.55	0.42	0.46
N	52430	52430	9940	9940

Notes: White heteroskedasticity-consistent standard errors in parentheses. The dependent variable is a dummy equal to one if respondents strictly prefer open lists over closed lists. Columns (1) and (2) are based on data for all European countries except Italy. Columns (3) and (4) are based on French data.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

From revolution to election.

A comparative analysis of Tunisia and Egypt.*

Luca Bettarelli†

Political participation is far from being a trivial issue. Indeed, it eludes traditional rational choice theory given the low individual impact over outcomes and the high costs of participation should discourage people from personally standing for political reasons. However, contrary to theoretical predictions, people do participate in political activities of various kind. In this paper, I concentrate on two dissimilar acts of political participation: revolution and elections. In detail, I first describe participants in revolution and elections - who they are and why they do participate - to then analyse the link (if any) between revolution and election participation. I compare the case of Tunisia and Egypt in the context of the Arab Spring. I find that Tunisian insurgents are more likely to vote in post-revolution election when compared to their Egyptian fellows and it is consistent with the outcomes of the two revolutions.

Keywords: Election; Revolution; Religiosity; Gender.
JEL Classification: D02; D74; H11

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1 Introduction and motivations

At the end of 2010, an unexpected cycle of events upset the Middle East and North Africa (MENA) region leading to the collapse of several authoritarian regimes that had lasted decades. The spark leading to first protests was the self-immolation by street vendor Mohammaed Bouazizi in Sidi Bouzid (Tunisia), after being insulted by police and failing to retrieve his confiscated wares. Riots rapidly spread to nearby towns and then the capital Tunis causing the resignation of President Zine el-Abidine Bel Ali on 14 January 2011, after 23 years of dictatorship. Following the example of Tunisia, a wave of unrest disseminated along the region.[1] At first glance, these revolts have been branded as bottom-up initiatives led by young people determined to fight for better economic conditions, equal rights, political independency and cultural freedom (Wright, 2011). International scholars immediately identified the similarities among countries, out of which: high unemployment rates mainly amid young educated people, widespread corruption, long-standing authocratic regimes, demographic structure with a large youth population cohort, lack of political and social freedom and the diffusion of communication technologies (Campante and Chor, 2012; Ishay, 2013). Indeed, the term Arab Spring has been coined to provide a overall view of this phenomenon. This interpretation is consistent with the so-called *deprivation literature*, according to which inequality, oppression, regime corruption are key determinants of revolutions and dictatorship-democracy transitions (Gurr, 1970; Acemoglu and Robinson, 2006; Persson and Tabellini, 2006; Acemoglu et al., 2014).

However, few years after the revolution many countries experienced extreme violence, military takeover and religious extremism. This seems a good enough reason to further investigate alternative narratives. Some authors (Hoffman and Jamal, 2013) pointed to the role of Islam as the leading engine enabling the coordination of insurgents, according to the so-called *mosque to square* narrative. Lynch (2012) noted that Fridays became frequently days of rage because of the convenience of organizing insurgents during Friday prayers. Indeed, authors as Israeli (2013) noted that Islam emerged as the real promoter of the Spring turbulance thus branding the post-revolution path known as Islamic Winter. Other authors such as Varol (2012) sustain the thesis of the *democratic coupe d'etat* according to which elites (often the army) first supported to popular opposition against the authoritarian regime in order to pave the way for their political climbing. In this context, the army, after siding with the protestors, behaves as a self-interested actor and sets up the transition process so that the resulting

[1] By the end of 2011, rulers had been forced from power in Tunisia, Egypt, Libya and Yemen; civil uprisings erupted in Bahrain and Syria; major protests occured in Algeria, Iraq, Jordan, Kuwait, Morocco and Sudan while minor protests in Mauritania, Oman, Saudi Arabia, Djibouti, Western Sahara and Palestine.

constitution favors its policy and institutional preferences.[2] Empirical analyses have mainly compared the Egyptian and Tunisian transition paths. This is due to the very similar initial conditions in terms of economic performance, socio-cultural features and political status that characterised the two countries at the time of the revolution. Moreover, the revolution resulted in the overthrow of the old regimes in both countries. All this made scholars expect similar outcomes in the revolution's aftermath. But, while the Egypth's democratic experiment ended up in a military takeover, Tunisia approved a consensual constitution and, although the consolidation of Tunisa's democracy is still in progress, its positive achievements make this comparison meaningful.[3]

This work belongs to the same stream of research, focusing on the comparison of Tunisia and Egypt. In particular, I analyse revolution and the first post-revolution elections that occured in both Tunisia and Egypt on 23 October and 28 November 2011 respectively.[4] I describe the preferences of people directly fighting authoritarian regimes and their propensity to vote in a symultaneous way, through a system of equations. In detail, the questions addressed are the following: who participated in revolution and election? Is there a connection between revolution and election participation? Are individuals fighting regimes more likely to vote in post-revolution election? Does the latter question inform us about revolution type and post-revolution path? Indeed, the interaction between these two dissimilar acts of political participation may reveal crucial insights about the typology of the revolution itself.[5] First, as noted by the seminal work of Tullock (1971), the decision to directly participate in revolutions is by no means trivial, bringing about the *paradox of revolution*. Indeed, the revolution outcome has public good features, that is once the revolution has been successful the outcome is shared by all individuals symphatizing with the revolutionary goals, irrespective of participation (Olson, 1971; Kavka, 1982). Thus, rational people would be better off by free-riding on the risk taken by other participants without directly assuming the high costs of participation (see among others Kuran, 1989, 1995; Lohman, 1994; Fearon, 2011). This is confirmed by several empirical analyses where the share of individuals who directly fight regimes is far smaller than those who agree with the revolutionary

[2] As noted by Apolte (2015b, p.3): "Coupe d'etats are hence the rule, even if they may sometimes be accompanied by some myth of broad participation of a deprived population."

[3] Tunsia revolution led longtime dictator Ben Ali to resign on 14 January 2011. Then, election to a Constituent Assembly was held on 23 October 2011 with the Ennahda Movemen that emerged as the leading party. After a period of political turmoil between 2012 and 2013, on 26 January 2014 a new Constitution was passed. The new Constitution recognizes Islam as the official state religion, but protects freedom of belief. As for Egypt, on 11 February 2011, Hosni Mubarak, Egypt *de facto* dictator since 1981, was forced to resign. Mubarak's fall was followed by a phase of military rule until election when the islamist Mohammed Mursi was elected president. Mursi's presidency was followed by a second phase of military rule.

[4] The revolution began in Tunisia on 18 December 2010 and ended on 14 January 2011, while the timing of the Egyptian revolution is from 25 January to 11 February 2011

[5] This paper considers political participation in a broad sense in the same way as Verba et al. (2000) who refer to acts "intended to have the consequence of influencing the choice of governing official or the policies they make and implement."

ideals (Finkel and Mller, 1998; Bozzoli and Bruck, 2011; Alianak, 2014). A similar reasoning applies to the decision to vote in election, according to the so-called *paradox of voting*: in large elections the chance that a single vote will change the outcome is so unlikely that the expected private benefit is close to zero and so even a small cost of voting should deter a rational individual from participating (Downs, 1957).

However, regardless of theoretical expectations, individuals do participate in political activities of different kind. Hence, it is relevant to analyse the characteristics of these individuals - who they are and why they do participate - and the link (if any) between revolution and election participation. Protest participation is defined in literature as *unconventional* form of political activism (Teorell et al., 2007). As noted by Ekman and Amna (2012) "unconventional activism [...] provides its members or supporters with a sense of doing something, an opportunity to personally take a stand and make a difference [...] The concept refers to deliberate attempts to influence the people in power".[6] However, this does not necessarily translate into other more conventional forms of political engagement, such as electoral participation.[7] Indeed, other conditions/opportunities may affect such decision. From this perspective, the analysis of revolutionaires' electoral turnout may provide an appealing interpretative lens: first, one would expect that direct participation in protests signals the intention of individuals to challenge the socio-political *status quo* of the country. This should translate in higher participation in more conventional political activities, mostly elections. Elections are indeed very important: they are not just a technical issue of the political stage but they represent the link between people preferences and political representation (Popper, 1966; Sartori, 2001). As noted by Besley and Kudamatsu (2008), election is a mean of removing poorly-performing leaders from office within a democratic setting as well as revolution is in dictatorship. These intuitions make reasonable to connect revolution and election participation. However, in a different perspective, non-participation in conventional political activities of people engaged in unconvential forms of political activism (such as riots, boycotting, political violence) may reveal either frustration and alienation or anti-political orientation which, in both cases, may proxy the detachment between revolutionary ideals and actual political stage.[8]

Results corroborate this preliminary hypothesis. Indeed, I find that Tunisian insurgents are more likely to vote in post-revolution election. In other words, they found a

[6]Ekman and Amna (2012), p.290.

[7]Other authors as Maki (2006) refer to formal/informal forms of political participation to distinguish between non anti-systemic and anti-systemic actions of individuals.

[8]This is also consistent with theoretical models which study revolution and revolution outcomes by focusing on the role of the *selectorate* (i.e. the elite binding the dictator decision-making process, as the power of politicians in a democracy is bound to the electorate) [see Bueno de Mosquita et al., 2003; Besley and Kudamatsu, 2008; Gilli, 2012]. The behaviour of the selectorate strongly affects the revolution aftermath. In the framework of this paper the intuition is the following: revolutionaires' ideals may be frustrated by a well-coordinated selectorate who exploits revolution to takeover the power. Indeed, Egyptian selectorate (the Army) was far more developed and organized than the one in Tunisia (Ishay, 2013).

fertile political ground to support their ideals in the revolution aftermath. This is not the case of Egyptian revolutionaires. The finding is consistent with the outcomes of the two revolutions with the Egyptian case that ended up in a military takeover while Tunisia, through strenuous negotiotaions between political leaders and civil society, approaved a consensual constitution.

The determinants of participation in revolution and election in the context of Arab Spring have been the focus of several studies. Beissinger et al. (2015) portrayed the participants in the Tunisian and Egyptian revolutions and they found that Tunisian insurgents were younger, more secular and significantly more diverse in social composition than their Egyptian fellows, but they do not find any signifincant evidence of the impact of religiosity in revolution participation. The authors motivate these sharp differences by the very different strategies pursued by the incumbent regimes on the eve of the revolutions. Hoffman and Jamal (2014) also examine the correlates of participation in these two protests but they focus on the link between religiosity and participation. The authors find a significant positive effect of religiosity, partially explained by the positive relationship between Quran reading and support for democracy. Doherty and Schraeder (2015) study the determinants of revolution participation in Tunisia and they find little evidence that participation was driven by pro-democracy dispositions. Instead, participants in the revolution were distinguished primarily by their exposure to direct and indirect social cues that encouraged participation and by their ability to bear the costs of participation. Authors also analyse participation in post-revolution election, as I do, but without focusing on the link between the two political participation acts.

This paper contributes to the literature by simultaneously analysing both revolution and election participation strategies and by controlling for problems of endogeniety through the use of a novel instrument. The paper is structured as follows: Section 2 presents data and descriptive statistics, Section 3 introduces the econometric methodology, Section 4 discusses results and Section 5 concludes.

2 Data and descriptive statistics

Data are taken from the second and third waves of the Arab Barometer. The second wave has been conducted between June and October 2011, while the third wave between February and April 2013. The sample includes 1,219 respondents in the second wave and 1,196 in the third wave in Egypt, while 1,196 and 1,199 individuals have been interviewed in Tunisia in the second and third wave respectively. The total number of respondents is then 4,810. The survey is a national probability sample design of adults 18 years and older, stratified by governorate and further stratified by urban-rural. Interviews were distributed proportional to population size.

The main dependent variables are two dummies - *Revol* and *Elect* - which are equal to

Table 1: DESCRIPTIVE STATISTICS

	Tot	Egypt		Tunisia				
	Mean (1)	Mean (2)	SD (3)	Mean (4)	SD (5)	Min (6)	Max (7)	p(2,4) (8)
Election	0.76	0.79	0.33	0.72	0.45	0	1	0.00
Protest	0.14	0.11	0.31	0.18	0.38	0	1	0.00
Gender	0.50	0.50	0.50	0.50	0.50	0	1	0.98
Age	41.95	39.42	13.65	41.94	15.82	18	87	0.00
Rural	0.45	0.56	0.50	0.35	0.48	0	1	0.00
Education	2.41	2.48	1.09	2.34	1.00	1	4	0.00
Unemployment	0.45	0.43	0.49	0.48	0.50	0	1	0.00
Wealth	1.96	1.94	0.88	1.97	0.97	1	4	0.29
Polinternet	1.02	0.92	0.53	1.11	0.61	1	3	0.00
Member	0.10	0.13	0.34	0.07	0.25	0	1	0.00
Polislam	3.97	4.41	1.09	3.47	1.18	0	5.5	0.00
Armytrust	3.59	3.65	0.72	3.53	0.87	1	4	0.00
Demobest	3.53	3.55	0.68	3.51	0.67	1	4	0.09

Notes: Arab Barometer dataset. Entries are mean values (*Mean*) and Standard Deviation (*SD*). p(x,y) is the probability of falsely rejecting equal means across groups corresponding to columns x and y, under the assumption of equal variances.

one if respondents participated in the revolution and voted in the first post-revolution election respectively, and zero otherwise. The percentage of individuals who declared they have voted is 75% in Egypt and 71% in Tunisia, which is larger than what is registered in the official data (54% and 52%).[9] Even if this figure appears to be inflated, data are consistent with other survey studies such as the AfroBarometer and the original survey by Doherty and Schraeder (2015). As for protest participation, 11% and 18% of respondents reported to be part of insurgents in Egypt and Tunisia respectively.[10] In both cases, the share of women is significantly lower than men. Interestingly, a very large fraction of respondents (85%) took the revolution side when asked whether they support the revolution or the dictator. This is consistent with theoretical expectations according to which individuals directly fighting dictators are far less than those who sympathize with the revolutionary goals. Among the set of covariates, I include age and gender of respondents, as well as a variable indicating if people live in urban/rural areas.[11] Then, I control for the level of education which may matter because the more educated is the society, the higher is the probability to participate

[9] Wave 2 has been conducted before the 2011 election. However, the survey asks respondents the intention to vote in next election, specifying the election at which the question refers to. Instead, Wave 3 directly asks respondents if they voted in the first post-revolution election. I combine these two questions to build the dichotomous variable *Elect*.

[10] Data are again consistent with other studies as Moaddel, 2012.

[11] *Gender*=1 for female respondents.

in revolution and to vote in the election (Campante and Chor, 2012).[12] Revolution participation has often been described as strictly linked to socio-economic grievances, thus indicators for unemployment and people wealth are included among the controls, as they may be correlated with the response variables.[13] Another likely determinant of political participation is the use of communication technologies and social media. Indeed, Twitter, Facebook and Youtube have been extensively exploited by protesters to spread their ideals, to attract potential fellows and to coordinate street mobilizations (Edmond, 2013; Acemoglu et al., 2014). Thus I create an additive scale index (*Polinternet*) ranging from 0 to 3 to indicate the proportion of people using internet for political issues.[14] I include a dummy variable labelled *Member* equal to one if respondent is member of at least one association with other than political goals.[15] Indeed, as noted by Doherty and Schraeder (2015), direct exposure to social cues may affect the decision to participate or not in revolutionary riots. Then, I aim to shed light on the impact of the so-called political Islam on political participation. As stated in Section 1, some authors identified Islamic movements as the promoters of revolutions. Indeed, after years of repression under authoritarian regimes, Islam got a momentum with the Arab Spring, thus gaining large consensus. This led Islamic parties to win the first post-revolution elections both in Egypt and Tunisia. However, the approach followed by Islamic parties about the relationship between Islam and politics may largely vary between countries (Torelli et al., 2012; Kienle, 2015). Thus, I construct a variable labelled *Polislam* to describe the preferences of respondents about the role of Islam and Shari'a in issues others than personal life (Berman, 2003).[16] In detail, I create a synthetic PCA-based index with higher values indicating a higher degree of individual level political Islam.[17] I make use of four questions asking respondents if Shari'a should

[12] *Education* is coded as follows: 1 (no-education), 2 (primary education), 3 (secondary education) and 4 (tertiary education or higher).

[13] *Wealth* is a count variable with higher values associated with higher economic prosperity. The rank is a self-evaluation of respondents among the following options: (1) Our household income does not cover our expenses and we face significant difficulties in meeting our needs; (2) Our household income does not cover our expenses and we face some difficulties in meeting our needs; (3) Our household income covers our expenses without notable difficulties; (4) Our household income covers our expenses well and we are able to save.

[14] I exploit the following questions: "Do you use the internet in order to: (1) Find out about political activities taking place in your country; (2) Express your opinion about political issues; (3) Find out about opposing political opinions in your country" Possible answers are No or Yes. The variable *Polinternet* is then coded 0 if respondent answers No to all the questions, 1 if he/she answers Yes to at least one question, 2 if he/she answers Yes to two questions, 3 if he/she answers Yes to all the above three questions.

[15] I use the following quastion "Are you a member of []? (i) A charitable society; (ii) A professional association; (iii) A youth/cultural/sports organization; (iv) a local development association". Possible answers are Yes or No. The variable *Member* is equal to one if respondent replays Yes to at least one of the above points.

[16] Berman (2003) defines Islamism "the belief that Islam should guide social and political as well as personal life".

[17] As robustness check, I try to build *Polislam* as an additive scale index instead of a PCA-based index but results remain qualitatively unchanged.

drive the formulation of (1) penal laws, (2) personal status laws, (3) inheritance laws and (4) divorce laws.[18] Then, I control for the opinion of respondents about a specific institution: the army. Indeed, if individuals trust the army to a great extent, then the likelihood of a democratic coupe d'etat, as defined in Section 1, should be higher.[19] Finally, I test the idea that one of the main protesters' objectives has been to enforce democratic institutions by looking at the individual level democracy attractiveness on a scale from 1 to 4 (*Demobest*) with higher values associated with people in favour of democratic institutions.[20]

3 Methodology

I first run two separate regressions where the outcomes are *Elect* and *Revol.* Equations are estimated through a probit model with wave fixed effect and robust standard errors:

$$Elect_i^* = X_i^{'}\beta_1 + \varepsilon_i; \tag{1}$$

$$Revol_i^* = X_i^{'}\beta_2 + \eta_i; \tag{2}$$

where the subscript i stands for individual observations and $Elect_i^*$ and $Revol_i^*$ are the underlying continuous latent variables. The vector (K×1) of covariates X_i is the same in both equations and ε_i and η_i are error terms. Even though the latent process is not observable, I do observe its realization, such that:

$$y_j = \begin{cases} 1 & \text{if} \quad y_j^* > 0, \\ 0 & \text{if} \quad y_j^* \leq 0, \end{cases} \tag{3}$$

where $j = Elect$, $Revol$. Given the latent-variable models in (1)-(3), then:

$$\Pr(y_j = 1) = \Phi(X_i^{'}\beta_j), \tag{4}$$

where Φ is the standard cumulative normal distribution. I run both regressions for Egypt and Tunisia to describe participants in revolution and elections.

[18] The questions are as follows: "To what extent do you agree with each of the following principles in the formulation of your country's laws and regulations? The government and parliament should enact penal laws in accordance with Islamic law (1), The government and parliament should enact personal status laws (marriage, divorce) in accordance with Islamic law (2), The government and parliament should enact inheritance laws in accordance with Islamic law (3), The government and parliament should enact inheritance laws in accordance with Islamic law (4)". Answers range from 1 (I strongly disagree) to 4 (I strongly agree).

[19] I use the following question: "I will name a number of institutions, and I would like you to tell me to what extent you trust each of them: The armed forces (the army)." Answers range from 1 (I absolutely do not trust) to 4 (I trust it to a great extent).

[20] The question is the following: "I will name a number of political systems to you, and I want to ask you about your opinion of them with regard to the country's governance [*A democratic political system*: public freedoms, guarantees equality in political and civil rights, alternation of power, and accountability and transparency of the executive authority]". Possible answers are: (1) very bad, (2) bad, (3) good, and (4) very good.

Then, I simultaneously model participation in revolution and elections. Let $V^*_{Elect,i}$ being the utility that individual i gets from voting. Such utility is described by the following behavioural equation:

$$V^*_{Elect,i} = \beta_{Elect} X_i^{'} + \gamma Revol_i + \delta_t + \varepsilon_i,$$

where X_i is the same vector of covariates as in (1)-(4), δ_t is the wave dummy and $Revol_i$ is the dummy indicating if individuals participated in revolutionary riots. Of course, individual utility is not directly observable but the actual turnout of respondents is and it is described as follows:

$$Elect_i = 1(V^*_{Elect,i} > 0) = 1(\beta_{Elect} X_i^{'} + \gamma Revol_i + \delta_t + \varepsilon_i \geq 0), \tag{5}$$

where $1(\cdot)$ denotes the indicator function. If we assume that participation in revolutions is exogenous then the parameters of equation (5) could be directly estimated specifying a distribution for ε_i. However, the decision to participate in a revolution is expected to be endogenous and determined by variables that may simultaneously impact on the individual probability of participation both in election and revolution. Thus, failing to take this into account would result in biased estimates. I model revolution participation in the same fashion as electoral turnout. Let $V^*_{Revol,i}$ be individual i's net benefit from being a revolutionaires. The reduced-form behavioural model is then given by:

$$V^*_{Revol,i} = \beta_{Revol} X_i^{'} + \lambda Z_i^{'} + \delta_t + \eta_i.$$

Again, I do not observe $V^*_{Revol,i}$ but rather a dummy variable $Revol_i$ which indicates whether individual i reports having participated in revolution or not. This is defined by:

$$Revol_i = 1(V^*_{Revol,i} > 0) = 1(\beta_{Revol} X_i^{'} + \lambda Z_i^{'} + \delta_t + \eta_i \geq 0); \tag{6}$$

where X_i is the same set of covariates that enter equations (1)-(5), Z_i is the excluded instrument and δ_t is the wave dummy. I jointly estimate equations (5) and (6) by a bivariate probit model where ε_i and η_i are jointly normally distributed, $E(\varepsilon_i) = E(\eta_i) = 0$, $Var(\varepsilon_i) = Var(\eta_i) = 1$ and $Cov(\varepsilon_i, \eta_i) = \rho$.[21] The t-statistic on parameter $\hat{\rho}$ ia a Wald test of the hypothesis that the cross-equation error term correlation is statistically significant. This provides information as to whether the maximum likelihood bivariate probit estimates should be used or if single equation estimates are adequate. In other words, $\hat{\rho}$ measures the endogeneity of $Revol_i$ in equation (3). To test the null hypothesis of bivariate normality I use the goodness-of-fit score test developed by Murphy (2007).

[21] As robustness check, I estimate equations (5) and (6) by a generalized structural equation model. Results remain qualitatively unchanged. Results and details are reported in the Appendix, Table 4.

The likelihood function for the bivariate model can be written as:

$$\log L(\theta) = \sum_{i=1}^{N} \log\Phi(t_{1i}, t_{2i}; \rho_i{}^*);$$

where θ contains the equations' coefficients and the unknown parameters of the joint distribution of ε_i and η_i that need to be estimated and $\Phi(t_{1i}, t_{21}; \rho^*)$ denotes the joint probability function of $Elect_i$ and $Revol_i$.

For $i = 1, ..., N$,

$$\begin{aligned} t_{1i} &= (2Elect_i - 1)(\beta_{Elect} X_i^{'} + \gamma Revol_i + \delta_t), \\ t_{2i} &= (2Revol_i - 1)(\beta_{Revol} X_i^{'} + \lambda Z_i{}^{'} + \delta_t) \quad \text{and} \\ \rho_i{}^* &= (2Elect_i - 1)(2Revol_i - 1)\rho, \end{aligned}$$

with the subscript i indicating the i_{th} unit observed in the sample.

The bivariate model is identified if the instrument Z_i in (6) has no effect on the outcome of equation (5) other than its impact through the endogenous regressor *Revol*, i.e. if the esclusion restriction is satisfied (Greene, 1996). The chosen instrument Z_i is *Proximity to square* which is a dummy equal to one if the respondent comes from a governorate sistematically involved in violent activities during the Arab Spring. The information about the respondent's governorate of origin is available in the Arab Barometer. Data on place and date of riots come from the ACLED dataset that records day-by-day violent political activities for a wide number of countries, including Tunisia and Egypt, and for the period 1997-present.[22] I restrict the sample to the Arab Spring time span.[23] The dataset includes information on place, date, type and actors of any event, thus enabling to isolote those events directly related to the Arab Spring. From these data, I can identify governorates where systematic (two or more) violent activities took place and therefore build the dichotomous instrument. There are 24 governorates in Tunisia and 27 in Egypt out of which 9 and 6 experienced systematic Arab Spring-related political activities respectively. In the empirical analysis, I use cluster-robust standard errors to control for intra-governorate correlation and, even if the number of governorates is not such small, I perform the wild bootstrap test as suggested by Cameron et al. (2008).[24]

The proximity to square may affect the likelihood of protest participation in differ-

[22] ACLED stands for Armed Conflict Location and Event Data Project that is an independent organization affiliated with the University of Sussex. ACLED has recorded over 90,000 individual events. See Raleigh et al. (2010) for more details.

[23] I consider Arab Spring related protests till one month after the resignation of the dictators in both Tunisia and Egypt.

[24] The rule of thumb is that when clusters are approximately less than 40 (Angrist and Pischke, 2008), then standard errors may be biased. Cameron et al. (2008) propose a wild cluster bootstrap-t procedure to deal with this issue. Technical details are reported in their paper.

ent ways. First, it reduces practical costs of participation thus making revolutionaires payoff (the difference between benefits and costs of participation) positive. Second, it facilitates the diffusion of information. Indeed, people living close to places of fight has not to entirely rely on information from external sources - such as television, newspapers, radio - that may be easily mainpulated by the central power. Third, it stimulates emulative behaviors and increases the psychological involvement of potential rebels.[25] On the opposite, the proximity to square impacts on election turnout only through the effect on revolution participation, thus ensuring the exclusion restriction. Results confirm these conjectures.[26]

4 Results

Table 2 reports estimations for models (1) and (2). The dependent variable in columns (1) and (2) is the dummy *Protest* indicating people who actively took part in revolutionary riots during the Arab Spring, in Egypt and Tunisia respectively. As noted in Section 2, the proportion of individuals fighting regimes over those symphatizing with the revolutionary goals is very small, so that it is relevant to investigate some of the determinants of protest participation. Contrary to theoretical expectations, the individual-level economic status, proxied by the variables *Income* and *Unempl*, does not play any role and the coefficients are indeed very close to zero and not statistically significant.[27] Consistently with expectations, males are largely more likely to participare in riots as well as people living in urban areas and this is true independently from the country. However, the negative coefficient for *Rural* in the case of Tunisia is half that of Egypt, thus emphasizing the different characteristics of Tunisian revolution that slowly moved from small villages to the capital Tunis while Egyptian protesters, by contrast, concentrated in major towns (namely Cairo and Alexandria). The idea that protests were led by young and educated individuals finds ground in the data even if the effect is remarkably feebler than what expected (in particular for Egypt). The coefficient for *Member* and *Polinternet* have a large and significant impact on protest participation indicating that social connections and participation in associations have been extensively exploited by protestets to spread revolutionary values. The coefficient for the variable *Demobest* mantains the same direction amid countries, even if the effect is sligtly larger in the case of Tunisia. These findings are consistent with previous studies (such as Doherty and Shraeder, 2015) and do not significantly vary between Egypt and Tunisia.

[25] This is consistent with the idea of *dynamic of mutual expectations* in models as Yin (1998).

[26] Staiger and Stock (1997) suggested the rule of thumb that instrument be deemed weak if the instrumenting equation F-statistic of the excluded instrument is less than 10. The F-statistic for *Proximity to square* is 14.8 and 21.5 in Tunisia and Egypt respectively.

[27] The test of the joint significance of the two economic status variables fails to reject the null in both Tunisia (p-value=0.99) and Egypt (p-value=0.63).

Table 2: Probit estimates.

Dep.Var.	**Revol**		**Elect**	
Country	**Egypt** (1)	**Tunisia** (2)	**Egypt** (3)	**Tunisia** (4)
Age	-0.001^{**}	-0.004^{***}	0.004^{***}	0.005^{***}
	[0.000]	[0.000]	[0.001]	[0.001]
Gender	-0.081^{***}	-0.188^{***}	-0.101^{***}	0.006
	[0.014]	[0.019]	[0.024]	[0.023]
Educ	0.042^{***}	0.029^{***}	0.053^{***}	0.027^{**}
	[0.006]	[0.011]	[0.009]	[0.013]
Wealth	-0.010	-0.004	0.001	0.004
	[0.006]	[0.010]	[0.009]	[0.012]
Unempl	-0.013	0.003	-0.028	-0.025
	[0.015]	[0.020]	[0.025]	[0.025]
Rural	-0.062^{***}	-0.035^{***}	0.026	0.04
	[0.012]	[0.008]	[0.021]	[0.023]
Polinternet	0.027^{***}	0.037^{***}	0.030^{**}	0.024^{**}
	[0.006]	[0.009]	[0.012]	[0.012]
Member	0.060^{***}	0.108^{***}	0.038	0.057
	[0.021]	[0.024]	[0.028]	[0.038]
Demobest	0.007^{*}	0.009^{*}	0.001^{*}	0.010^{**}
	[0.004]	[0.005]	[0.000]	[0.005]
Armytrust	0.019^{**}	-0.022^{**}	0.040^{***}	-0.006
	[0.006]	[0.009]	[0.011]	[0.013]
Polislam	0.038^{**}	-0.006	0.002	0.003
	[0.010]	[0.009]	[0.004]	[0.008]
Observations	2218	1929	2202	1922
Waves dummy	Yes	Yes	Yes	Yes

Notes: White heteroskedasticity-consistent standard errors in brackets. Entries are marginal effects.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

This is not the case for *Armytrust* and *Polislam*. These two covariates have opposed effects in the two countries, thus providing some evidence about the difference in the two revolutionary paths. Egypt protesters have a significant feeling of trust towards military institutions as well as a positive opinion about the involvement of religious values into politics. The same is not true in the case of Tunisia. This is consistent with the outcome of the two revolutions. It is now of great interest to analyse if protesters have higher probability to vote in the election following the revolution. First, I briefly analyse the determinants of voting turnout (Table 2, columns (3) and (4)). Results are generally consistent with expectations, where voting participation is positively related to the level of education and the interest in politics. The coefficient for *Demobest* is positive but the effect both in terms of size and statistical significance is smaller than what expected. This may indicate that people voted to sustain motivations other than democratic values. Again *Armytrust* impacts in opposite directions in the two countries with Egyptian voters that significantly trust the army. This evidence may motivate the bid for power of military leaders in the aftermath of the election. On the opposite, *Polislam* does not affect voters turnout independently from the country. The same is true for the variable *Rural* and this is consistent with the choice of the instrument I propose. Indeed, turnout seems to be unrelated to dwelling place of respondents. Interestingly, *Gender* is significant and negatively associated with voters turnout in Egypt but not in Tunisia. Indeed, I cannot reject the null hypothesis that voters turnout is the same for Tunisian males and females respondents. This may indicate a more comprehensive political participation of Tunisian respondents.[28]

Finally, I turn to the core objective of the paper that is the simultaneous analysis of revolution and electoral participation. In detail, I aim to test if protesters are also more likely to vote in the first election following the revolution. As noted in Section 1, revolution and election participation are expected to share common characteristics.

Table 3 shows results from the bivariate model reported in Section 3.

[28] Indeed, the Global Gender Gap Index ranks Tunisa 69th in terms of political empowerment while Egypt is ranked 136th (Sika and Khodary, 2012).

Table 3: Bivariate probit estimates.

Dep.Var.	Elect	
Country	**Egypt** (1)	**Tunisia** (2)
Revol	-0.045	0.063^{***}
	[0.035]	[0.019]
Age	0.004^{***}	0.005^{***}
	[0.000]	[0.001]
Gender	-0.104^{***}	0.017
	[0.021]	[0.021]
Educ	0.053^{***}	0.032^{**}
	[0.008]	[0.014]
Wealth	0.005	0.009
	[0.016]	[0.014]
Unempl	-0.031	-0.021
	[0.026]	[0.022]
Rural	0.029	0.019
	[0.021]	[0.022]
Polinternet	0.032^{**}	0.019^{**}
	[0.013]	[0.009]
Member	0.025	0.055
	[0.047]	[0.033]
Demobest	0.001^{*}	0.009^{*}
	[0.000]	[0.005]
Armytrust	0.033^{***}	-0.001
	[0.012]	[0.010]
Polislam	0.003	0.004
	[0.004]	[0.008]
Observations	2119	1914
Rho	0.62^{***}	0.74^{***}
Normality test (p-value)	0.154	0.198
F-test on Proximity to square	14.8^{***}	21.5^{***}
Waves dummy	Yes	Yes

Notes: Standard errors [in brackets] clustered by governorate. Entries are marginal effects. Results for the instrumenting equation are not reported in the Table.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Protesters are more likely to vote only in the case of Tunisia. In other words, Tunisia revolutionaries seem to have acquired a greater sense of political entitlement alongside greater optimism about their influence on the political change. This is not the case of Egyptian revolutionaires where indeed revolution led to a military takeover. It is hard to point out the specific reasons leading to these opposite behaviours. Two explanations are possible: either revolutionaires differ in their fundamental motivations (as it is suggested by the analysis of the determinants of protest participation) or the events next to the revolution made Egyptian revolutionaires withdraw from political participation. The former element is related to the expressive side - as opposed to the instrumental side - of revolution participation. Instrumental motivations reveal the will of revolutionaires to affect the long-term political change process of the country and thus the acquired political awareness of revolutionaires. In this context, revolution is meant as a mean to be entitled to other conventional forms of political participation such as free and fair elections, beyond a number of goals of different kind. Instead, as noted by Tullock (1971), mental motivations of revolutionaires can hardly represent serious threats to a dictator leadership that instead is threated by the people in the inner circle around him/her. From this viewpoint, individuals participate in mass protests to express dissatisfaction, to signal their preferences and/or to emulate similar events (as may be the case of Egyptian revolutionaires in the aftermath of the Tunisian revolution). However, collective action problems prevent them from seriously threatening the regime. Thus, revolutions mainly result from the action of a small and coordinated elite (the army) that possibly enjoys the trust of the people and exploits them to oust the unpopular incumbent but, at the same time, preserving its dominant position. This is consistent with the notion of *abused rebels* (see Apolte, 2015a). From this perspective, revolution is perceived as a change of ruler from dictator to a coordinated elite with revolutionaires only playing a marginal role. The second argument refers to the ability of a new government to calm down the rebelling part of the population and to generate accountability mechanisms. Once the revolution starts, the elite close to the dictator (e.g. the army/police) can either remain loyal to the regime and suppress rebels - the Tiananmen Square situation - or decide to not intervene. In the first case, the *status-quo* outcome is implemented and rebels fall back into line. In the latter situation, the new leader should succed in reintegrating rebels into a more conventional form of political activism (proxied by the electoral participation). If this is not case and revolutionaires' ideals are frustrated by the political state, then social conflict persists even in the revolution aftermath. This may be proxied by the revolutionaires' act of abstaining from the vote out of protest (Teorell et al., 2007). In this context, the army feels empowered to calm down the stage by taking over the power.

These conjectures are consistent with the results of this paper and the actual revolutions' outcomes in Egypt and Tunisia.

5 Conclusion

This paper analyses the determinants of revolution and election participation comparing two countries which both experienced the Arab-Spring waves: Tunisia and Egypt. The comparison is meaningful given that both countries went through similar revolutionary paths leading to the overthrow of longstanding dictators. However, while Egyptian revolution resulted in a military takeover and heavy issues of social and political conflict, Tunisia approved a consensual constitution in 2014 and is currently sailing along a full democratic transition, even though with various difficulties.[29]

[29]Tunisia has been recently ranked as the best democratic regime among Arab countries. Source: Democracy Index 2015, Economist Inteligence Unit.

Consistently with the literature, I find that some of the traditional aspects motivating the political participation of people - e.g. unemployment, people wealth - do not have strong effect on participation. On the other side, I find that other covariates such as trust in the army and the inclusion of religious values into politics better explain the determinants of participation and impact in different ways depending on the country. This is consistent with the outcome of the two revolutions.

The novelty of the paper lies in the direct correlation between the two political participation acts under the idea that people directly fighting regimes should be more likely to vote in election. The validation or rejection of this hypothesis can both inform us regarding the roots of revolution and its outcome.

Moreover, the paper uses the instrumental variable approach to control for endogeneity issues and proposes a novel instrument. Results prove to be robust to the use of different econometric models and shed light on a crucial aspect - the link between revolution and elections participation - that it is worth being further investigated also in different contexts.

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Appendix

Table 4: Generalized structural equation model estimates.

Dep.Var.	**Elect**	
Country	**Egypt** (1)	**Tunisia** (2)
Revol	-0.153	0.152^{**}
	[0.132]	[0.076]
Age	0.0146^{***}	0.0148^{***}
	[0.003]	[0.002]
Gender	-0.435^{***}	0.0631
	[0.070]	[0.066]
Educ	0.222^{***}	0.080
	[0.039]	[0.049]
Wealth	-0.002	0.008
	[0.063]	[0.043]
Unempl	-0.102	-0.0820
	[0.099]	[0.068]
Rural	0.171^{*}	0.020
	[0.103]	[0.068]
Polinternet	0.128^{**}	0.058^{**}
	[0.053]	[0.029]
Member	0.142	0.163
	[0.201]	[0.119]
Demobest	-0.086	0.014
	[0.074]	[0.056]
Armytrust	0.149^{***}	-0.014
	[0.051]	[0.036]
Polislam	-0.023	0.016
	[0.054]	[0.024]

Notes: Standard errors [in brackets] clustered by governorate. Entries are coefficients.
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Coefficients are estimated by a generalized structural equation model (Rabe-Hesketh and Pickles, 2004) with common unobserved component:

$$E\big[Elect_i \mid Revol, X, \eta\big] = F\big(\beta_0 + Revol_i\alpha + X_i\beta_1 + \delta_t + \eta_i\big);$$
$$Revol_i = \beta_0 + X_i\beta_2 + Z_i\gamma + \delta_t + \eta_i + \varepsilon_i;$$

where F() is a smooth nonlinear function, X_i is the same vector of covariates that enter equations (1)-(6) which are assumed to be exogenous, Z_i is the excluded instrument, δ_t is the wave dummy, η_i is the common unobserved component which resolves the issues of endogeneity and ε_i is an error term.