Electoral rules and agricultural protection: A missing link?

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Abstract

This paper analyzes the role of electoral rules and bargaining within legislatures of industrialized parliamentary systems in determining the political success of agriculture in attracting government transfers based on a probabilistic voting environment. Assuming voters expect a pro-agrarian policy, rural districts are pivotal in determining if the coalition obtains a majority, whereas urban districts are pivotal within the majority itself. In bargaining at the legislature, this generates a conflict between the prime minister, who will tend to favor rural districts, and the parliamentary majority, which will be dominated by urban concerns. As district size grows and the electoral system converges to a pure proportional system, both of these biases are attenuated. Overall, the result is an inverse u-shaped relationship between district size and agricultural subsidies. However, when voter beliefs tend toward a liberal agricultural policy a u-shaped relationship results. Based on a dynamic econometric panel model using time-series cross-country data for 23 parliamentary democracies since 1962 our theory is empirically validated. The findings remain stable under various robustness checks including a test of potential endogeneity of electoral rules.

Keywords comparative political economy, agricultural protection, electoral rules, endogeneity of political institutions, time-series cross-national data

1. Introduction

Since the seminal papers of Persson and Tabellini (1999, 2000, 2003) the question how constitutional rules influence economic policies and hence economic performance is on top of the research agenda in comparative political economy. In particular, Persson and Tabellini are interested in identifying the causal effects of formal political institutions on economic and political outcomes. However, Acemoglu and Johnson (2005) demonstrate that identifying causal effects of formal constitutional rules is a complex undertaking. For example, disentangling the impact of formal constitutional rules from the impact of informal institutions is often plagued by the problem of "clustered" institutions. "Clustered" institutions describe the fact that a combination of mutually reinforcing formal and informal institutions evolve jointly (Acemoglu and Johnson, 2005). Thus, oberserved political outcomes are the result of informal and formal rules of political games. In this context identifying true causal effects of constitutional rules demands for a comprehensive theory that reflects

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the interaction of formal and informal political institutions. Additionally, adequate econometric techniques must be used to guarantee a valid empirical analysis of the causal effects of formal constitutional rules.

A case in point to analyze the effects of constitutional rules are special interest politics, i.e. policy biases in favor of a specific voter group at the expense of the general public. *Pars pro toto* this analysis focuses on political success of agricultural voter groups in attracting government transfers. Reviewing the literature to date important questions on how and why constitutional rules determine agricultural protection, are still unsolved. While classical political economy models explain observed differences in agricultural protection comparing industrialized and developing countries (Gardner, 1987; Swinnen, 1994; Tyers and Anderson, 1992; Miller, 1991; Zusman, 1976), these approaches fail to shed light on observed large cross-country differences in agricultural protection among industrialized or developing countries, respectively. As these models neglect political institutions, they might be the missing link explaining observed variation of agricultural protection even among industrialized countries.

Based on well-knwon works of Beghin and Kherallah (1994); Beghin et al. (1996); Swinnen et al. (2000), Thies and Porche (2007) and Olper and Raimondi (2008) provide a comprehensive econometric analysis of the political determinants of agricultural protection, including socio-economic factors as control variables. But Thies and Porche (2007) and Olper and Raimondi (2008) do not yet provide a political economy theory of agricultural protection that explains the observed effects of political determinants on agricultural protection. They derive their hypotheses rather ad hoc applying various existing political economy theories on protection.

Hence this paper develops a micro-political founded theory to understand the interaction of formal and informal political institutions in determining the level of agricultural protection. We explicitly derive legislators' policy preferences from electoral competition and final policy outcomes from postelection bargaining in legislatures. Since recent theoretical and empirical work analyzes the impact of the electoral system as constitutional rule on policy outcomes, this work also focuses on the impact of the different electoral systems in combination with informal rules of the policy game on policy outcomes. However, apparently conflicting theories exist in the literature regarding special interest politics and electoral rules[1]. Scholars such as Persson and Tabellini (2003) or Grossman and Helpman (2005) argue that special interest politics, occur more frequently in majoritarian than in proportional representation systems. Other scholars such as Milesi-Ferretti et al. (2002) or Rogowski and Kayser (2002) state that distributional policies are less likely in majoritarian than in proportional systems. Moreover, Hee Park and Jensen (2007) criticize the use of a simple majoritarian-proportional dichotomy to explain distributive politics and suggest the Cox-threshold as the relevant indicator to measure the impact of electoral rules on distributive policy outcomes (see Cox, 1987; Myerson, 1993). The Cox-Myerson theory, however, does not provide a complete model of political decision making. The latter necessarily incorporates a model of postelection legislative bargaining among legislators representing different constituencies with heterogenous interests.

Assuming voters expect a pro-agrarian policy, rural districts are pivotal in determining the coalition obtaining a majority, whereas urban districts are pivotal within the majority itself. In bargaining at the legislature, this generates a conflict between the prime minister, who will tend to favor rural districts, and the parliamentary majority, which will be dominated by urban concerns. As district size grows and the electoral system converges to a

[1] In particular, the effects of two archetypical electoral systems, labeled "majoritarian" and "proportional", on general economic policy outcomes are contrasted. Scholars mostly neglect mixed electoral systems in their analysis

pure proportional system, both of these biases are attenuated. Overall, an inverse u-shaped relationship between district size and agricultural subsidies results. Assuming, however, voter beliefs tend toward a liberal agricultural policy, the prime minister tends to favor urban concerns and a rural legislator becomes decisive within his parliamentary majority. Accordingly, a u-shaped relationship results. Hence, in contrast to classical approaches, our theory is able to explain observed large cross-country differences in agricultural protection among industrialized countries.

Furthermore, these hypotheses are tested empirically. Based on difference-in-difference estimation and instrument variable estimation via two stage least squares (2SLS), empirical results confirm the suggested nonlinear relationship between the electoral system and agricultural protection. However, although our estimation results imply that we are able to valid instruments, interpretation of estimated coefficients as causal effects is still problematic. The idea of clustered institutions put forward by Acemoglu and Johnson (2005) well captured within the developed theory implies an interaction between formal electoral rules and coalition discipline, as well as the influence of interest as informal institutions. Thus the documented link may not be interpreted as a causal relationship.

This paper starts in section 2 introducing the theoretical model, while section 3 provides our empirical analysis, including the derivation of the applied econometric estimation strategy and description of used data. Further, we test potential endogeneity of electoral rules in section 4. Finally, section 5 summarizes our main results and gives further perspectives on future research.

2. Theoretical model

2.1. The population and economy

Consider a society is divided into two sectors: agricultural and non-agricultural. The group of voters economically active in the agricultural sector are denoted with A. M represents the group of voters economically active in the non-agricultural sector. If government does not engange in agricultural policy, the equilibrium per capita income of the agricultural and non-agricultural population is I_0^A and I_0^M, respectively. Each group has unit mass, and the share of each group in total population is denoted by α^A or α^M, respectively.

Agricultural policy is characterized by redistributive transfers from the non-agricultural to the agricultural sector. For simplicity we assume that income redistribution occurs via subsidization of agricultural and taxation of non-agricultural sectors. Let s denote the resulting per capita subsidization of the agricultural population, while t denotes the per capita taxation of the non-agricultural population. Any feasible policy must satisfy the following budget constraint:

$$\alpha^A \tilde{\Gamma}(s) = \alpha^M t \quad \Leftrightarrow \quad t = \frac{\alpha^A}{\alpha^M}\tilde{\Gamma}(s) = \Gamma(s) \tag{1}$$

The function Γ includes deadweight costs. In particular, we assume Γ to be strictly convex and increasing in the level of subsidization, i.e., $\Gamma' > 0$ and $\Gamma'' > 0$ [2]. Assuming identical individuals for both groups implies the following welfare function of each member

[2]Deadweight costs significantly vary across various agricultural policy instruments. However, we do not focus on the choice of economically efficient redistribution instruments, although discussion on agricultural policy is to a large extent concerned about this issue (see e.g. de Gorter and Swinnen, 2002; Becker, 1983; Lohmann, 1998).

given the policy s:

$$\begin{aligned} W^A &= I_0^A + s; \\ W^M &= I_0^M + b^M(PG - \Gamma(s)), \end{aligned} \tag{2}$$

where $PG - \Gamma(s)$ denotes the per capita public good expenditure resulting for the non-agricultural population, where b^M denotes the relative preference of non-agricultural households for public goods. Thus, we assume that agricultural subsidization does not directly reduce private income of non-agricultural households (I_0^M), but implies a reduction of per capita public good expenditures for the latter.

2.2. Legislative decision making

For our theoretical model of legislative bargaining in parliamentary systems, we suggest a rather simple legislative majority bargaining game that is based on the existence of a stable ex ante majority coalition and on the principle of proposal power of the government. As has been demonstrated by Huber (1996) and Diermeier and Feddersen (1998), stable ex ante majority coalition built among legislators essentially characterize parliamentary systems. Legislators who are members of this majority coalition make legislative decisions exclusively. The rational of ex ante majority coalition building corresponds to the fact that this coalition at least weakly increases the utility of all majority members when compared to their utilities derived under a default outcome $\bar{s}$ that would result from non-cooperative behavior of legislators. In particular, ex ante fixed parliamentary majorities are able to guarantee their members higher utilities due to additional rent legislators realize from being part of a stable majority (Huber, 1996).

We formally define a legislative system as a finite set of political agents, N, where $i = 1, \ldots, n$ denotes a generic element of the legislative system. Within the political system specific institutions are defined as specific subsets of N: the prime minister (PM), the majority (P_M) and the opposition (P_O). In general, P_M could correspond to a multi-party coalition or a single majority party. However, to simplify following analyses at the election stage, we assume a two-party set-up, i.e., P_M corresponds to the majority party P_M, where P_O denotes the opposition party. Further, P_M is a finite subset of legislators $g \in N$ and g is a generic element of P_M. Moreover, we assume that PM is also the party leader of the majority party. Following Huber (1996) as well as Diermeier and Feddersen (1998), we can concentrate on the prime minister, PM, and her majority in the parliament P_M that is ex ante identifiable.

The model has two stages. At the first stage, we model the default policy outcome $\bar{s}$. For simplicity, we assume that agricultural policy is one-dimensional and that parliament decides about agricultural policy by simple majority voting[3]. We denote the unidimensional policy space by S. Further, we assume that agents' policy preferences can be represented by a single-peaked function $U_i(s)$. Let Y_i denote the ideal point of legislator i. Obviously, under these assumptions the well-known median voter theorem applies, i.e., the unique equilibrium outcome of the non-cooperative legislative decision-making game neglecting any ex ante coalition building is the ideal point of the floor median (Black, 1958).

At the second stage, legislators, who are members of the majority P_M, and the PM bargain over policy to improve their utility derived under the default outcome. In detail, they proceed in two steps. First, the PM proposes a policy, s_{PM}, to her parliamentary

[3] Of course we could also assume more complex legislative decision-making procedures including agenda setting power of the parliamentary committees or the government. However, this would not change our major results and therefore we keep analyses as simple as possible at this point and leave the analysis of more complex legislative institutions for future work.

majority and announces side payments γ being paid to the majority in case it admits the governmental proposal. Regarding content, we interpret these side payments as rent the PM can pay to the majority due to specific formal legislative procedures, e.g., issuing a confidence vote, or informal procedures, i.e., the possibility to generate favors in terms of political career for party members. In this paper, we are not specifically interested in modeling exactly how the PM can generate rent valuable to her majority, but generally subsume this under the term party or coalition discipline that is exerted by the PM[4].

At the second step each individual majority member can decide whether or not to accept the proposal of the PM. If all majority members agree to the proposal, the proposed policy, s_{PM}, passes parliament and all majority members receive the announced rent. Otherwise, the default policy $\bar{s}$ becomes the legislative decision and no rent is paid. We assume that legislators maximize the sum of actual rent, γ, and the utility derived from policy, $U_g(s)$, while making their decision to agree or not to agree:

$$u_g = U_g(s) + \gamma \tag{3}$$

Proposition 1. *Assuming a one-dimensional agricultural policy choice s, there exists a unique subgame perfect Nash equilibrium for our legislative majority bargaining game defined above. The equilibrium outcome, s^*, depends on the rent, γ, the default policy outcome, $\bar{s}$, and the policy preferences of the PM and the majority members, g.*

In equilibrium agricultural policy choice, s^* results from the following maximization[5]:

$$s^* = \arg\max_s \quad U_{PM}(s) \quad s.t. \quad s \in \bigcap_g S_g \tag{4}$$
$$with \quad S_g = \{ s \in S | \, U_g(s) + \gamma \geq U_g(\bar{s}) \}$$

Defining $s^- = \min \bigcap_g S_g$ and $s^+ = \max \bigcap_g S_g$ the outcome of the legislative bargaining game corresponds to the minimal distance between the ideal point of the PM and the interval $[s^-, s^+]$.

Interestingly, if the rent, γ, is sufficiently large or if legislators' preferences are sufficiently homogeneous, the final agricultural policy outcome corresponds to the ideal point of the prime minister. Hence, under this condition our model corresponds to pre-election political models, which generally assume that governmental policy simply corresponds to political preferences of the party leader (becoming the omnipotent head of government after elections). However, if party discipline, i.e., the rent γ, is not sufficiently high or analogous, policy preferences of the PM and her parliamentary majority are sufficiently heterogeneous, agricultural policy outcome is no more fully determined by the PM's policy preferences. In contrast, under this assumption policy outcome is solely determined by the intersection set of the subsets S_g, i.e., the policy preferences of the majority member, the majority rent, γ, and the default policy, $\bar{s}$.

2.3. Election stage

Probabilistic voting models are generally applied to derive policy preferences of legislators from electoral competition (see Persson and Tabellini, 2000). For explaining special interest politics, these approaches basically argue that specific groups, such as farmers, are less

[4]Note further that we assume that at this stage the PM can commit to paying the rent. However, this assumption is not necessary; in a richer modeling set-up including the specific procedures it is possible to get essentially the same result without assuming this kind of commitment.

[5]Note that the maximization problem always has a unique solution, as long as the utility functions of legislators are strictly concave. Note that all sets S_g are compact and convex subsets of S.

ideologically biased relative to other groups and therefore become a natural target for politicians who vie for electoral support. In particular, we assume that legislators are rent seeking, i.e., legislators' behavior can be derived from the maximization of actual rent and future rent. Future rent depends on the probability of being re-elected. Obviously, this probability depends on voters' electoral response to observed policies.

In general, voters elect legislators in electoral districts. Every electoral district d_k contains the same share of voter population, α_{d_k}, and the sum of voter population over all districts covers total population eligible to vote. Usually, proportional representation (PR) and a majoritarian election system (MS) are distinguished as ideal-typical electoral systems if electoral systems are characterized by the number of legislators elected in a constituency, i.e. by the district magnitude. In PR systems candidates are elected in one multiple-member national electoral district, while candidates are elected in a one-member constituency in pure majoritarian systems. Denoting the total number of parliamentary seats by n, the district magnitude of PR systems is n and of pure MS systems 1, respectively. In general, the district magnitude of a specific electoral system k ranges from n to 1.

2.3.1. Voter behaviour

An individual incumbent $g \in P_M$ is re-elected in a generic voting district d. In principle, we assume that a voter votes for an incumbent if the utility she has derived under the implemented policy, s^*, is higher than her specific reservation utility. However, beyond economic welfare derived under observed policies, $W^J(s^*)$, voters care for another dimension, which generally is referred to as ideological preferences for parties, although this dimension could include other characteristics of parties or candidates, e.g., competence or appearance. The crucial point is that ideological preferences are exogenous in the sense that ideology is a permanent attribute of parties, i.e., cannot be changed at will during election campaign.

In this paper we do not further analyze ideological preferences of voters; we only assume that ideological preferences can be subdivided into three components: a group-specific relative importance of ideology compared to economic well-being, K^J; a regional component μ_{jd}; and a national component, δ. Thus, a voter $j \in J$ votes for the incumbent g if the utility she observes under the agricultural policy s^* is higher than a specific reservation utility, $W^J(s^0)$, corrected by the ideological preferences for the incumbent party P_M:

$$W^J(s^*) > W^J(s^0) + K^J(\mu_{jd} + \delta) \tag{5}$$

Parameters μ_{jd} and δ can take negative and positive values and measure the ideological bias of voter j toward the opposition party P_O. Thus, a positive value implies that voter j has a bias in district d in favor of party P_O.

The ideological preferences are uncertain at the time political agents have to make their policy decisions. In detail, we assume that the parameter μ_{jd} has region-specific uniform distributions on $\left[\bar{\mu}_d - \frac{1}{2\chi}, \bar{\mu}_d + \frac{1}{2\chi}\right]$. Thus, two parameters, $\bar{\mu}_d$ and χ, fully characterize the regional distribution of ideological preferences.

Moreover, we assume that the relative importance of ideology, K^J, differs across groups. Note that assuming a different relative importance of ideological preferences implies that groups generally differ in their effective ideological homogeneity, i.e., have different effective densities $\phi^J = \frac{\chi}{K^J}$. We make specific assumptions about the differences in these distributions. In particular, we assume that the agricultural population has less relative interest in ideology, i.e. $K^A < K^M$. Thus, it results that the agricultural population is more ideologically homogeneous than the non-agricultural population, i.e., $\phi^A > \phi^M$.

Political agents know the regional and group-specific distribution, $\bar{\mu}_d$ and ϕ^J, when they decide on agricultural policy, while the electoral uncertainty derives from the uncertainty of the national components, δ. The parameter δ measures the average popularity of party P_O in comparison to party P_M. Here, we assume a uniform distribution on $\left[-\frac{1}{2\psi}, +\frac{1}{2\psi}\right]$. Thus, on average, the national ideological shock is unbiased.

Given the assumption above the total vote share an incumbent g receives in district d after regional and national ideological shocks have been realized follows as:

$$\Pi_d = \sum_J \alpha_d^J \phi^J \omega^J - \chi \left[\bar{\mu}_d + \delta\right] + \frac{1}{2}, \tag{6}$$

where $\omega^J = \left(W^J(s^*) - W^J(s^0)\right)$. Given the electoral support incumbents receive in a generic district d, the re-election probability of a political agent g crucially depends on district characteristics that in turn are determined by the organization of the electoral system. We will derive the re-election probabilities of a political agent g and of the PM, respectively, subject to the organization of the electoral system in the following paragraphs.

2.3.2. The impact of the electoral system on legislators' preferences

To derive legislator's preferences we characterize precisely how electoral districts d_k differ in their demographic and ideological composition within and across electoral systems k. We differentiate two types of districts, rural districts (D^R) and urban districts (D^U), respectively. Rural districts are characterized by a higher share of agricultural population compared to urban districts. Furthermore, electoral districts differ regarding their ideological bias $\bar{\mu}_{d_k}$. We assume that three different clusters D_k^1, D_k^2 and D_k^3, respectively, exist. The first cluster is biased in favor of party P_M, while the third is biased in favor of the opposition party P_O, and the second cluster is unbiased, i.e., it holds:

$$\begin{aligned} \bar{\mu}_{d_k} &= \bar{\mu}_k^{P_M} < 0 \quad \forall \quad d \in D_k^1 \\ \bar{\mu}_{d_k} &= \bar{\mu}_k^{P_O} > 0 \quad \forall \quad d \in D_k^3 \end{aligned} \tag{7}$$

Overall, both rural and urban population is unbiased, that is it holds:

$$\bar{\mu}_k^{P_O} + \bar{\mu}_k^{P_M} = 0 \tag{8}$$

$$\sum_{d_k \in D^R} \alpha_{d_k} \bar{\mu}_{d_k} = \sum_{d_k \in D^U} \alpha_{d_k} \bar{\mu}_{d_k} = 0. \tag{9}$$

Moreover, we assume that none of the regional clusters D_k^1, D_k^2 and D_k^3 includes the majority of voter population, while any two clusters together include the majority of voters.

The basic idea behind these assumptions is that both the agricultural and the ideologically biased voter populations are clustered regionally. Empirically, such clusters can be found in most countries. Thus, our assumptions correspond approximately to real world societies, although real structures are certainly more heterogeneous. To simplify our analyses, though, we abstract from real world heterogeneity at this point.

Finally, larger electoral districts are demographically and ideologically more homogeneous. To cover this increasing homogeneity of larger electoral districts, we assume the following for $k = 1, ..., n$:

$$\alpha_{d_k}^A \leq \alpha_{d_{(k-1)}}^A \quad \forall d \in D^R \quad \text{and} \quad \alpha_{d_k}^A \geq \alpha_{d_{(k-1)}}^A \quad \forall d \in D^U \tag{10}$$

$$(\bar{\mu}_k^{P_O} - \bar{\mu}_k^{P_M}) \leq (\bar{\mu}_{k-1}^{P_O} - \bar{\mu}_{k-1}^{P_M}) \tag{11}$$

Assuming that all k candidates of party P_M running for election in the k-member district d_k have the same chances, $\frac{1}{k}$, to get a parliamentary seat won by party P_M in this district, the re-election probability of a majority member $g \in P_M$ in an electoral system k, $\tilde{\pi}_{rg}^k$, conditional on the national shock, δ, is given by:

$$\tilde{\pi}_{rg}^k(\delta) = \frac{1}{k}\Pi_{d_k} = \frac{1}{k}\left[\sum_J \alpha_{d_k}^J \phi^J \omega^J - \chi\left[\bar{\mu}_{d_k} + \delta\right] + \frac{1}{2}\right] \tag{12}$$

Analogously, the expected re-election probability, π_{rg}^k, results in:

$$\pi_{rg}^k = \int_{-\frac{1}{2\psi}}^{\frac{1}{2\psi}} \tilde{\pi}_{rg}^k(\delta)\psi d\delta = \frac{1}{k}\left[\sum_J \alpha_{d_k}^J \phi^J \omega^J - \chi\bar{\mu}_{d_k} + \frac{1}{2}\right] \tag{13}$$

Overall, maximizing the expected probability of re-election taking the groups' reservation utilities as given corresponds to maximizing an additive social welfare function (SWF_{d_k}), where the weight of group J, $\bar{\beta}_{d_k}^J$, results as:

$$\bar{\beta}_{d_k}^J = \alpha_{d_k}^J \phi^J \tag{14}$$

Thus, obviously legislators have different policy preferences as long as electoral districts are demographically heterogeneous. In particular, we can define rural and urban policy preferences depending on legislators' re-election in a rural or urban district, respectively. For any electoral systems, k, rural in comparison to urban preferences are characterized by a higher relative SWF weight for the agricultural population, implying a higher preferred subsidization level for the former, where s_k^u and s_k^r denote the preferred subsidization level of urban and rural legislators, respectively. However, because demographic homogeneity increases with district size k, subsidization levels preferred by rural and urban legislators, respectively, converge toward a common national level with increasing district size k.

2.3.3. Deriving the re-election probability of the PM

In contrast to a majority member, the PM is only re-elected if party P_M wins the election, thus only if party P_M wins the majority of total seats. To formally derive the probability of re-election of party P_M as the governmental party, we define the following stochastic variable Λ_{d_k} for each electoral district:

$$\Lambda_{d_k} = \begin{cases} k \quad \text{with} \quad \text{probability} \quad \mathrm{Pr^k}_{d_k} \\ k-1 \quad \text{with} \quad \text{probability} \quad \mathrm{Pr_{d_k}^{k-1}} \\ \vdots \\ 1 \quad \text{with} \quad \text{probability} \quad \mathrm{Pr_{d_k}^1} \\ 0 \quad \text{with} \quad \text{probability} \quad (1 - \sum\limits_{\mathrm{k}} \mathrm{Pr_{d_k}^k}) \end{cases} \tag{15}$$

Given the definition of Λ_{d_k} the probability that party P_M wins the election or rather that the PM is re-elected in an electoral system k, π^k_{rPM}, results:

$$\pi^k_{rPM} = \Pr\left[\sum_{d_k} \Lambda_{d_k} \geq 0.5\ n\ +1\right] \tag{16}$$

Formally, the ideal position of the party leader, Y^{PM}, is the policy position that maximizes the re-election probability of the PM and results from the following mixed-integer maximization problem:

$$\begin{aligned} Y^{PM} &= \arg\max_{s,\Lambda_{d_k},\delta} \delta \quad \text{s.t.} \\ &\left[\frac{\psi}{\chi}\sum_J \alpha^J_{d_k}\phi^J\omega^J - \bar{\mu}_{d_k} - \delta\right]\ k \geq \Lambda_{d_k} \\ &\sum_{d_k} \Lambda_{d_k} \geq 0.5\ n + 1. \end{aligned} \tag{17}$$

Despite assuming perfectly homogeneous electoral districts or a pure proportional system with only one national district (i.e., $k = n$), it is generally difficult to characterize the solution of eq. 17, i.e., the preferred policy position of the PM. Obviously, Y^{PM} will always lie in the interval $[s^u_k, s^r_k]$, as outside of this interval the expected vote shares of all districts monotonically increase or decrease with s. But beyond this interval it is tedious to characterize Y^{PM}. However, especially regarding the final outcome of legislative bargaining, it is crucial if the PM prefers a higher or a lower subsidization level when compared to the default outcome $\bar{s}_k$.

Therefore, we basically follow Persson and Tabellini (2000) and introduce additional assumptions that allow a further characterization of the equilibrium for our simple electoral competition set-up. In particular, we assume that the ideological biases toward party P_M in district type D^1, $\bar{\mu}^{P_M}$, and toward party P_O in district type D^3, $\bar{\mu}^{P_O}$, are sufficiently large that electoral competition only takes place in the unbiased district type D^2, i.e., we assume that party P_M wins the election if and only if it wins the majority of parliamentary seats in the unbiased electoral districts.

However, generally the cluster of unbiased districts D^2 comprises both rural and urban districts, respectively. Thus, the PM only wins the elections if his party P_M wins the majority of seats in both unbiased rural and urban districts. Denoting by Π^l, $\quad l = u, r$, the probability that the party P_M wins the majority of seats in the unbiased district $l = r, u$, the ideal position of the PM results form the maximization of his winning probability as follows:

$$\begin{aligned} Y^{PM} &= \arg\max_s \{\min\{\Pi^u(s), \Pi^r(s)\}\} \\ \Pi^l &= \frac{\psi}{\chi}\sum_J \alpha^l_J\phi^J\omega^J + 0.5 \quad l = u, r \end{aligned} \tag{18}$$

2.3.4. Voter belief formation and PM's policy position

The solution of this minimax problem crucially depends on the reservation utilities of agricultural and non-agricultural voters, $W^0(\tilde{s})$, where the latter depends on the policy outcome expected by voters. In particular, it can be shown that as long as voter beliefs fall in the interval $[s_k^u, s_k^r]$ there always exist a $s^{\#}$ with the following properties:

$$\begin{array}{rcll} \Pi^u(s^{\#}) & = & \Pi^r(s^{\#}) & \\ \Pi^u(s) & > & \Pi^u(s^{\#}) \quad \text{and} \quad \Pi^r(s) < \Pi^r(s^{\#}) & \forall s < s^{\#} \\ \Pi^u(s) & < & \Pi^u(s^{\#}) \quad \text{and} \quad \Pi^r(s) > \Pi^r(s^{\#}) & \forall s > s^{\#} \\ & \Rightarrow & s^{\#} = \arg\max_s \{\min\{\Pi^u, \Pi^r\}\} & \end{array} \tag{19}$$

Obviously, $s^{\#}$ is the solution of the maximin problem eq. (18). Denoting by $\tilde{s}$ the vector of voters' policy beliefs, the PM position results ceteris paribus as a function of voters' beliefs, i.e., $Y^{PM} = Y^{PM}(\tilde{s})$. Moreover, the final policy outcome s^* results from legislative bargaining and, hence, is also determined by voters' policy beliefs, i.e., $s^* = s^*(\tilde{s})$. Thus, assuming rational expectations implies that voters' beliefs $\tilde{s}$ result in a final policy outcome that corresponds with voters' initial beliefs, i.e., $s^*(\tilde{s}) = \tilde{s}$.

Now, for any position of the PM the outcome of legislative bargaining is restricted to the interval $\left[s_k^-, s_k^+\right]$, where s_k^- and s_k^+ are solely determined by the demographic composition of the electoral system and the party discipline γ (see above). Therefore, assuming voters form rational expectations implies that voters form common beliefs that lie in the interval $\tilde{s} \in \left[s_k^-, s_k^+\right]$. Moreover, it is easy to see that for any common voter belief $\tilde{s} \in \left[s_k^-, s_k^+\right]$ it holds:

$$\begin{array}{rcll} \Pi^u(\tilde{s}) & = & \Pi^r(\tilde{s}) = 0 & \\ \Pi^u(s) & < & 0.5 \quad \text{and} \quad \Pi^r(s) > 0.5, & s > \tilde{s} \\ \Pi^u(s) & > & 0.5 \quad \text{and} \quad \Pi^r(s) < 0.5, & s < \tilde{s} \end{array} \tag{20}$$

Accordingly, for any common voter belief $\tilde{s} \in \left[s_k^-, s_k^+\right]$ the unique solution to the minimax problem in eq.(18) is $Y^{PM} = \tilde{s}$. Hence, any commonly believed outcome $\tilde{s} \in \left[s_k^-, s_k^+\right]$ corresponds to rational expectations. The question that arises is how voters form their common beliefs. Belief formation of voters is a very interesting subject in itself. Nevertheless, we leave this interesting topic for future work and provide only an intuitive model of voters' belief formation. Following the interesting work of Golub and Jackson (2009), we assume belief formation results from a social communication process among voters. Belief formation might be biased to the extent that communication is dominated by specific central actors. In this regard interest groups are often central players dominating stakeholder communication.

Given the fact that the PM's preferred policy position resulting from the rationale in eq.(18) is increasing in voters' beliefs, agrarian interest groups have an interest to influence voters' beliefs toward high agricultural subsidization, while non-agrarian interest groups have an interest to influence voters' beliefs toward low agricultural subsidization levels. Taken into account that at least in industrialized democracies political communication is strongly dominated by agrarian interest groups, it seems plausible that voters' initial beliefs correspond to high agricultural protection levels in these countries. Formally, we assume that voters' common beliefs result as a weighted mean of preferred policy outcomes of rural and urban legislators, respectively, i.e., $\tilde{s} = C_A s_k^r + (1 - C_A) s_k^u$, with $0 \leq C_A \leq 1$. Thus, C_A measures the relative influence of agrarian interest groups.

Overall, it follows from our theory that legislators' policy preferences systematically change with the electoral system, in which the preferred agricultural subsidization levels decrease with the district size k for rural and increase for urban legislators. For pure proportional representation systems (k=n), all legislators and the PM have identical political preferences. (See also figure 1 below.) We denote $\bar{s}_n$ as the common ideal point of all legislators under proportional representation that trivially becomes also the unique policy outcome. Assuming further that initial voter beliefs in industrialized countries correspond to high agricultural support implies that under mixed and majority rules (k < n) the PM also tends toward rural preferences [6]. Next, we are able to determine the overall equilibrium of our legislative bargaining game under different electoral systems.

2.4. Policy outcomes under different electoral systems

Given the impact of the electoral system on legislators' policy preferences, we can now summarize the impact of the electoral system on the equilibrium policy outcome of our legislative bargaining game in *proposition 2.* (The proof is given in the appendix.)

Proposition 2. *Let s_k^* and $\bar{s}_k$ denote the equilibrium and default policy outcome, respectively, of the majority bargaining game defined in proposition 1 assuming an electoral system $k = 1, .., n$. Then the following holds:*

1. *The equilibrium policy outcome is defined by:*

$$\begin{aligned} s_k^* &= \arg\min_s \left\| Y_k^{PM} - s \right\| \quad s.t. \quad s \in \left[s_k^-, s_k^+ \right] \\ s_k^+ &= \max \left\{ s \in S_k \,\middle|\, \sum_J \beta_{u_k}^J W^J(s) + \gamma \geq \sum_J \beta_{u_k}^J W^J(\bar{s}_k) \right\} \\ s_k^- &= \min \left\{ s \in S_k \,\middle|\, \sum_J \beta_{r_k}^J W^J(s) + \gamma \geq \sum_J \beta_{r_k}^J W^J(\bar{s}_k) \right\} \end{aligned} \tag{21}$$

   where $\beta_{u_k}^J$ and $\beta_{r_k}^J$ denote the group weights of an additive SWF corresponding to the electoral competition equilibrium in urban and rural districts, respectively, defined by the electoral system k.

2. *In particular, it holds for the equilibrium outcome s_k^*:*

$$\begin{aligned} s_k^* &= \max \left\{ s_k^-, Y_k^{PM} \right\} \leq \bar{s}_k \quad if \quad Y_k^{PM} \leq \bar{s_k} \\ s_k^* &= \min \left\{ s_k^+, Y_k^{PM} \right\} \geq \bar{s}_k \quad if \quad Y_k^{PM} \geq \bar{s}_k. \end{aligned} \tag{22}$$

3. *Let $\tilde{s} = \max \left[\min \left[C_A * s_k^r + (1 - C_A) s_k^u, s_k^+ \right], s_k^- \right]$ denote the common beliefs of voters regarding the agricultural policy outcome, then it holds:*

$$\begin{aligned} Y_k^{PM} &= \tilde{s} \leq \bar{s}_n \quad for \quad C_A \quad sufficiently\ close\ to\ 0 \\ Y_k^{PM} &= \tilde{s} \geq \bar{s}_n \quad for \quad C_A \quad sufficiently\ close\ to\ 1, \end{aligned} \tag{23}$$

   where $\bar{s}_n$ is the unique common ideal position of all legislators under proportional representation, i.e., k=n.

4. *There always exists a k^* with $1 \leq k^* \leq n$ and it holds:*

   (i) *For C_A sufficiently close to 1:*

$$\left[s_k^* \leq s_{k+1}^* \quad \forall k < k^* \quad and \quad s_k^* \geq s_{k+1}^* \quad \forall k \geq k^* \right]$$

[6] However, please note that even if we would assume other voter beliefs implying urban preferences of the PM our main theoretical implication that agricultural policy outcomes systematically change with the electoral system would still result.

(*ii*) *For C_A sufficiently close to 0:*

$$\left[s_k^* \geq s_{k+1}^* \quad \forall k < k^* \quad and \quad s_k^* \leq s_{k+1}^* \quad \forall k \geq k^*\right]$$

Three things are worth noting. First, assuming the non-agrarian population is sufficiently more ideologically biased when compared to the agrarian population, i.e., $\phi^M << \phi^A$, both rural and urban legislators prefer a subsidization of the farm sector for any electoral system k. Second, in extreme cases of perfect party discipline, the restriction of the decisive (urban or rural) majority member is never binding, i.e., the equilibrium outcome is solely determined by the preferences of the PM. In this case k^* equals 1. Third, if this restriction is binding, the equilibrium is solely determined by the preferences of the decisive majority member being re-elected in an urban or rural district and the rent γ. Note, in particular, that under this condition the equilibrium would not change with changed preferences of the PM as long as the PM prefers a sufficiently high (low) subsidization level, i.e., a level that is higher (lower) than the maximum (minimum) level the decisive urban (rural) majority member is willing to accept in exchange for the rent, γ. This last point is crucial regarding the impact of the election system on agricultural policy. In contrast to existing pre-election models, in our approach the impact of an increased district size on agricultural protection is ambiguous and depends on the interplay of formal constitutional rules and informal institutions, i.e., coalition discipline and relative strength of agrarian interest groups.

In particular, if political communication is dominated by agrarian interest groups, voters believe in protectionist agricultural policy and electoral competition pushes the PM to take pro-agrarian policy positions. In this case legislative bargaining occurs between a pro-agrarian PM and an urban legislator that tends to be biased against agrarian interest. A contrario, if political communication is dominated by non-agrarian interest groups, voter beliefs tend toward a liberal agricultural policy. Hence, the PM tends to favor non-agrarian interests and legislative bargaining occurs among a pro-urban PM and a pro-agrarian rural legislator as a decisive coalition member. As district size grows and the electoral system converges to a pure proportional system, policy biases among rural and urban districts are attenuated. Accordingly, in the first case there is an inverse u-shaped relation, while in the second case a u-shaped relation between district size and agricultural protection results as shown in figure 1.

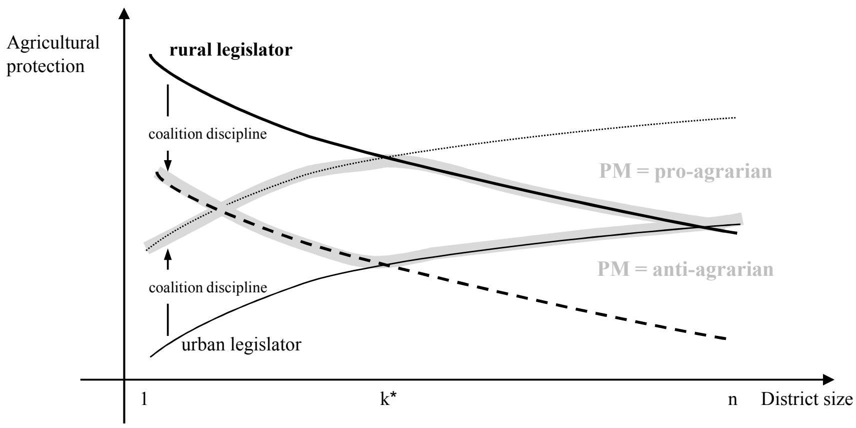

Figure 1: Ideal points (black thin lines) and policy outcomes (grey bold lines) under different electoral rules

3. Empirical Analysis

3.1. Data and estimation strategy

Given the fact that our theory emphasizes the impact of electoral rules on agricultural protection within parliamentary democracies, we restrict the sample of countries for the empirical analysis to fully consolidated parliamentary democracies. Basically, we follow this conservative approach of selecting countries to avoid the problem of "comparing the incomparable". In detail, we defined a country as a fully consolidated parliamentary democracy if the country has a polity score provided by the Polity IV data set above 0 (see database of Marshall et al., 2008) and if the constitution provides the presence of a vote of confidence and the separation of the head of state and the head of government (see database of Lundell and Karvonen (2003) and of Beck et al. (2001)[7]. All country-years where both requirements were fulfilled are included into our analysis. Furthermore, we exclude countries belonging to the European Union (EU) from our analysis, because the European Common Agricultural Policy (CAP) is negotiated at the supranational level by EU institutions. In particular, countries are dropped from our sample one year before their EU accession to take the impact of anticipated CAP into account.

As endogenous variable that characterizes redistributive transfers to the agricultural sector, we use an advanced measure of agricultural protection provided by Anderson and Valenzuela (2008), the Nominal Rate of Assistance to Agriculture (*NRA*). The *NRA* is calculated as a weighted average of commodity-specific *NRAs* using the undistorted production values of the commodities as weights. Analogously to the commonly used Producer Support Estimate (PSE) published by the OECD (2001), the NRA includes indirect market interventions, e.g., direct transfer payments as well as exchange rates distortions[8].

To test our theoretical model, we define the following three electoral systems based on the principle of district size: (1) a majoritarian system where only one legislator gets elected in a district, (2) a mixed system where on average 2 up to 9.9 legislators are elected per district and (3) a proportional representation system where 10 or more legislators get elected per district on average in a country. Such classification particularly allows to analyze the impact of an intermediate system between a pure majoritarian and a pure proportional representation system on special interest politics. Furthermore this results in two binary indicator variables as set of institutional variables I , maj and $prop$, where maj=1 indicates a majoritarian and $prop$=1 indicates a proportional electoral rule, while $maj = prop$=0 indicates a mixed system. We take information on district size from Lundell and Karvonen (2003) and Beck et al. (2001).

Following existing studies the set of macroeconomic controls denoted as X includes the initial gross domestic product (GDP) per capita (*initialgdppc*) and the real GDP per capita growth (*gdppcgrowth*) to capture economic development; the logarithm of agricultural share in employment (*emplln*) to account for differences in economic structure and industrialization; the ratio of the agricultural share in value-added and the agricultural share in employment (*compad*) to proxy comparative advantages in agriculture; and arable land per farm worker (*factorend*) to take the relative incomes of agricultural farmers into account (see for information on this variables Tyers and Anderson, 1992; Beghin and Kherallah, 1994; Swinnen, 1994; Olper, 2001; Anderson, 2008). Following Beghin and Kherallah

[7] If these data sets differ in their classification or if countries are not clearly coded as parliamentary or presidential, we use further information by Ismayr (2002a),Ismayr (2002b), Lijphart (1999) and Armingeon et al. (2008).

[8] In contrast to the concept of the PSE, the unit value difference of production between the world and domestic market is expressed as a fraction of the undistorted product value and not as a fraction of the distorted product value.

(1994) we define *budget* as the net agricultural export per capita in order to account for governmental budget constraints. In particular, budget costs due to agricultural trade policy crucially depend on the country's agricultural net trade position. All used economic and demographic control variables were calculated using data available from the World Bank database of development indicators and the database of the Food and Agriculture Organization of the United Nations (FAO, 2008; World Bank, 2008).

Overall, our sample corresponds to time-series cross-section data including 23 countries with an average time period of ???? years per country.[9]. Therefore, several issues endemic to these data type need to be addressed by the econometric specifications in order to provide a valid assessment of the influence of political institutions on the level of agricultural protection. First is the modeling of country specific heterogeneity and second is the consideration of model dynamics. Note that the issue of possible endogeneity of institutional variables is discussed in detail in section 4.

To empirically test our theory we start with the following baseline *difference-in-differences* specification, which is closely related to that used in Besley and Case (1995):

$$NRA_{it} = \kappa + \upsilon I_{it} + \varrho X_{it} + \varphi_t + \xi_i + \epsilon_{it}, \tag{24}$$

where NRA_{it} denotes the measure of agricultural protection, I_{it} denotes the set of institutional indicators, X_{it} refers to the set of macroeconomic control variables, and φ_t represents a time specific effect.[10]

Given the fact that our used institutional indicators are time-variant, though rarely changing variables, a fixed effect regression is applicable, though not necessarily optimal. While the baseline specification in eq.(24) addresses the issues of country specific heterogeneity via consideration of country specific effects ξ_i, the use of fixed effects hinders the direct incorporation of country specific variables (see Greene, 2010). Therefore, we follow an alternative modeling of country specific heterogeneity providing as a byproduct possibly more efficient estimation results.[11] Based on a cross validation experiment discussed by Stone (1974) and adapted to time series cross-section data by Beck et al. (2001), we test whether heterogeneity is related to groups of countries:

$$NRA_{it} = \kappa + \upsilon I_{it} + \varrho X_{it} + \varphi_t + \xi Z_i + \epsilon_{it}, \tag{25}$$

where Z_i incorporates country specific information as well as a group indicator. This specification puts more emphasis on the cross-section variation, which may be preferable to analyze the impact of rarely changing political institutions.

To test for heterogeneity related to groups, we start with an estimation of a pooled model specification of eq. (25) neglecting fixed effects, but including initial GDP per capita as a time invariant variable. Based on this pooled estimation we fit regression with $N-1$ countries and predict the NRA of the left out country to identify country specific heterogeneity via computed mean squared forecast errors (MSFE). Based on the comparison of

[9]Summary statistics of all variables are presented in table 3 in the appendix. A complete list of countries and considered time periods is presented in table 4 in the appendix.

[10]Note that specifying time specific variable captures the effect of the Uruguay negotiations in 1994. A model without time fixed effects and with a dummy for the effect of the Uruguay negotiations revealed similar results.

[11]Plümper and Troeger (2007) suggest a three-step procedure to estimate time invariant and slowly changing variables in the presence of country heterogeneity. Several authors have criticized this approach, recommending instead the application of a standard Hausman-Taylor model or a shrinkage estimator combining the Hausman-Taylor and the Plümper and Tröger approach (Greene, 2010; Mitze, 2009; Breusch et al., 2010). Therefore, we suggest an alternative estimation strategy in this paper to incorporate country specific heterogeneity.

country specific MSFE's with average MSFE, we were able to identify countries that are less well predicted via pooled regression. In a next step we introduce group specific dummy variables into eq. (25) to capture identified unobserved heterogeneity.

After controlling for country specific heterogeneity we deal with model dynamics. It is well-known in the literature that significance testing in specifications as given in model 24-25 is possibly subject to substantially inflated t-values, when the issue of model dynamics and autocorrelation is not properly addressed (see Bertrand et al., 2004). To deal with this issue, we check the robustness of results via the inclusion of the lagged dependent variable, $lNRA_{it}$, as an explaining factor.Note that inference for these specifications is based on cluster robust standard errors defining countries as clusters as suggested by White (1980). However, as noted by Cameron et al. (2008) these standard errors although controlling for both heteroscedasticity and general correlation patterns within clusters will generally still be biased downwards. We hence adapt the proposed wild cluster residual based bootstrap-se of Cameron et al. (2008) in order to perform a robust test of the hypothesis that electoral rules influence agricultural protection as suggested by theory.[12] Note that this strategy is accompanied by a careful check of the adequate model specification incorporating features of latent heterogeneity and serial correlation as outlined above.

3.2. Results

Following our estimation strategy, we present in table 1 five models corresponding to the different approaches discussed above. In detail, model 1 corresponds to the standard fixed effects model (eq.(24)), while model 2 is a standard pooled estimation of eq.(25). Based on model 2 we were able to undertake out-of-sample predictions as described above. By comparing the country specific MSFEs, we see that Norway, Iceland and Switzerland are less well predicted by a pooled model with country specific time invariant variables than other countries. These three countries can be seen as some kind of outstanding European countries with high GDP per capita and a very specialized agricultural sector. Thus, we employ a dummy $Protec^{+}$ in model 3 to tackle the group related unobserved heterogeneity. The positive coefficient of this country group dummy reveals that these countries tend to more highly subsidize their agricultural sector compared to other parliamentary democracies. To compare non-nested models 1-3, we refer to the standard information criteria, the Akaike (AIC) and Schwarz (BIC) criterion. Both criteria indicate that model 1 is preferable when compared to model 2. However, according to the BIC criterion, model 3 is preferred when compared to the standard fixed effect model 1.

After testing for the correct modeling of country heterogeneity, we next deal with model dynamics, i.e., serially correlated errors endemic to our type of data.

In Table 1 model 4 reports estimation results from a dynamic specification including a lagged dependent variable into model 3. We also examined a dynamic specification of model 1, but both information criteria reveal that a specification with modeling group related heterogeneity and a lagged dependent variable is preferred. Thus, overall model 4 is our preferred model[13]. The coefficient of the lagged dependent variable turns out to be highly significant. Intuitively, the use of a lagged Nominal Rate of Assistance can be

[12] In detail, we follow the strategy outlined in Appendix B of Cameron et al. (2008) based on the Wald type test statistic for assessing the hypothesis $\nu = 0$, i.e.

$$W_{\nu} = \hat{\nu}' \tilde{\Sigma_{\nu}}^{-1} \hat{\nu},$$

where $\hat{\nu}$ denote the estimate of ν and $\tilde{\Sigma_{\nu}}$ the corresponding covariance matrix.

[13] Please note, however, that our central results are unaffected by model specifications as proposed and discussed by Greene (2010). Estimation results for all model specifications are available from the authors upon request.

explained by a slow speed of adjustment of agricultural policies to changing socio-economic framework conditions.

Finally, we perform in model 5 a first robustness check of our results replicating model 4 with the continuous measure *distr* of the electoral system. More advanced robustness checks are postponed to section 4, where we discuss the problem of endogenous institutions in more detail.

Overall, the standard control variables exhibit the expected pattern if they are significant. In particular, we find a negative and highly significant impact of the agricultural employment share *emplln* on agricultural protection for all models except Model 2. Thus, following Olson's theory the former indicates lower cost of collective action due to a decreasing free-riding problem for smaller farm groups and, thus, implies c.p. higher agricultural protection, while the latter captures the negative effects of agreed WTO restrictions on agricultural protection levels.

Furthermore, the negative sign for the variable *compad* is in line with the theory and empirical finding of Honma and Hayami (1986), i.e., lower comparative advantages in agriculture increase the demand for agricultural protection. However, this parameter is only slightly statistically significant for models 1 and 3. Analogously, the negative coefficient of *factorend* estimated for all model specifications corresponds to the relative income hypothesis of Tyers and Anderson (1992) and de Gorter and Tsur (1991), predicting decreasing rates of assistance with increasing relative income of the agricultural sector. This variable turns out to be statistically significant in all specifications.

For the control variables *inititalgdppc*, *GDPpcgrowth* and *budget*, however, some variations in the significance and sign of estimated parameters can be observed across model specifications. In particular, while in model 2 the expected positive and significant impact on the level of protection results, which is in line with the so-called development paradox by Tyers and Anderson (1992), an unexpected negative impact is found in model 3-5. A possible explanation for this unexpected result could be the fact that different model specifications exploit *between* and *within* variation at a different degree, where parameter estimations are quite sensitive to the concrete specification of heterogeneity. In particular, our modeling of group-related heterogeneity ($Protec^{+}$) partly eliminates the *between* variation of variables, while the pooled regression model without fixed effects fully exploits *between* variations of variables. Hence, because GDP per capita is extraordinarily high for the three high protection countries, namely Norway, Iceland and Switzerland, remaining *between* variation after the group dummy ($Protec^{+}$) has been introduced is rather low. Analogously, because all three high protection countries are large net importers of agricultural commodities remaining *between* variance of the variable *budget* is also rather low once the group dummy has been introduced. Analogously to *inititalgdppc* the estimated *budget* parameter is only significant and displays the correct negative sign for the pooled regression (model 2) excluding any fixed effects. Beghin and Kherallah (1994) state that increasing budget expenditures to finance agricultural protection c.p. reduce protection levels.

A contrario, for the variable *gdppcgrowth* a high within and comparatively low between variance is observed. Accordingly, a significant and expected positive parameter estimation for *gdppcgrowth* is observed when the econometric model specification focuses on *within* variation, i.e., for the standard fixed effect model 1, while for all other specifications the expected positive sign, but no statistical significance is found.

Now we turn to our central explaining variables, *maj* and *prop*, describing electoral rules. For all model specifications we find highly significant and negative coefficients for both variables *maj* and *prop*. The evidence is supported for all specifications except specification (1) by the performed Wald test assessing the hypothesis $\nu = 0$. As noted above, the high multicollinearity between fixed effects and the rarely changing institutional variables may cause the failure to reject the null hypothesis in case of the fixed effects specification. To investigate this issue further, we perform a second stage cross section analysis regressing thereby the estimated fixed effects on country specific variables and the institutional variables.[14] The corresponding regression results clearly indicate that cross country variation within fixed effects can be explained via the considered institutions.[15]

In particular, these main estimation results clearly support our theory. Agricultural protection first increases and than decreases with district size and, hence, protection is c.p. the highest for mixed electoral systems when compared to both pure majoritarian and pure proportional representation systems. Moreover, the simple majoritarian-proportional dichotomy is not sufficient to explain agricultural protection as a special case of redistributive politics.

Summarizing results we first conclude from model 1-3 that the effect of the electoral dummies survives while controlling for country specific or group related specific effects. With respect to the inclusion of the lagged dependent variable in model 4 all coefficients decrease. However, to generate parameters comparable to model 3 estimated coefficients of model 4 have to be divided by one minus the estimated parameter of the lagged dependent variable. Please note that after transformation all estimated parameters nearly correspond for models 3 and 4.

Furthermore, model 5 presents a robustness check against the specification of the electoral system. We substitute our dummy variables by a continuous measure of district size and include the squared district size in our specification. Again, estimation results support our theory, i.e., a non-linear (inverse u-shaped) relationship between district size and agricultural protection levels results.

Overall, we can conclude that by controlling for standard economic and demographic variables we find a statistically significant impact of electoral rules on agricultural protection levels. In particular, a non-linear inverse u-shaped relationship between district size and agricultural protection results as predicted by our theory. This relationship remains remarkably robust when addressing different model specifications concerning country heterogeneity and dynamic issues. To what extent our estimation results can be interpreted as causal effects of electoral rules, i.e., answering the counterfactual questions of what would be the short and long term impact of a change in electoral rules on agricultural policy, will be discussed in more depth below.

4. Endogeneity and causal effects of electoral rules

In the context of comparative politics, it is widely perceived that political institutions and political performance variables might be simultaneously affected by unobservable factors, e.g., political culture, history and geography. Accordingly, the literature on economic comparative politics stresses the importance to control for potential endogeneity of political institutions (see Persson and Tabellini, 2003; Acemoglu et al., 2001; Acemoglu and Johnson, 2005). The main approach to control for potential endogeneity is an instrument variable estimation (IV). Beyond eliminating estimation bias, Persson and Tabellini (2003) use

[14] To deal with the occurring time variation within institutions, we classified a country according to its modal institutional setting.

[15] Regressing estimated country specific effects on a constant, *inititalgdppc*, *maj* and *prop* revealed the following results (overall $R^2 = 0.1905$)

$$\hat{\xi}_i = -0.9029 + 0.0292 initital gdppc_i + -0.4182 maj_i + -0.0842 prop_i + e_i.$$

Table 1: Political Institutions and Agricultural Protection: Basic Results

	(1)	(2)	(3)	(4)	(5)
constant	$--$	-0.2402 0.2981	0.1451 0.1376	0.0296 0.0386	-0.0796 0.0744
lNRA	$--$	$--$	$--$	0.6836^{***} 0.0357	0.6849^{***} 0.0345
initialgdppc	$--$	0.0800^{***} 0.0194	-0.0196 0.0144	-0.0065 0.0048	-0.0066 0.0048
gdppcgrowth	0.0130^{***} 0.0043	0.0089 0.0098	0.0046 0.0052	0.0071^{***} 0.0026	0.0070^{***} 0.0027
factorend	-0.3959^{**} 0.1946	-0.6291^{***} 0.1929	-0.3645^{***} 0.1470	-0.1036^{**} 0.0495	-0.1041^{**} 0.0514
budget	0.4941^{*} 0.2792	-0.6945^{***} 0.1632	-0.0349 0.1090	-0.0123 0.0354	-0.0129 0.0356
compad	-0.3418^{*} 0.2074	-0.0300 0.3595	-0.2134^{*} 0.1289	-0.0450 0.0452	-0.0369 0.0428
emplln	-1.0575^{***} 0.2818	-0.0406 0.1991	-0.3988^{***} 0.1010	-0.1178^{***} 0.0324	-0.1175^{***} 0.0320
maj	-0.9982^{***} 0.3034	-0.3857^{*} 0.2058	-0.3149^{**} 0.1437	-0.1143^{**} 0.0531	$--$
prop	-0.9269^{***} 0.1902	-0.6789^{***} 0.2891	-0.2145^{**} 0.1014	-0.0702^{*} 0.0378	$--$
distr	$--$	$--$	$--$	$--$	0.5966^{*} 0.3473
distr^2	$--$	$--$	$--$	$--$	-0.6142^{*} 0.3199
PROTEC$^+$	$--$	$--$	2.0791^{***} 0.2008	0.6322^{***} 0.0735	0.6601^{***} 0.0785
Wald test $\nu = 0$	26.5915	21.8805^{**}	12.1923^{*}	14.3666^{*}	14.2690^{*}
AIC	-2.4531	-1.6978	-2.2713	-2.9642	-2.9618
BIC	-1.7636	-1.2151	-1.7788	-2.4619	-2.4595
Country fixed effects	yes	no	no	no	no
Time fixed effects	yes	yes	yes	yes	yes
# of regressors	69	48	49	50	50
# of observations	407	407	407	407	407

Notes: Bootstrapped cluster robust standard errors with countries defined as clusters are given in parentheses. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, and *** indicates significance at the 1 percent level.

instrument variable estimation to attempt to estimate the causal effects of constitutional rules. Especially the latter point is of interest for the new comparative political economy, because "[T]he ultimate goal is to draw conclusions about causal effects of constitutions on specific policy outcomes. We would like to answer questions like the following: if United Kingdom were to switch its electoral rule from majoritarian to proportional, how would this affect the size of its welfare state or its budget deficits?" (Persson and Tabellini, 2003, page 7). Hence, it is certainly interesting to analyze causal effects of electoral rules on the level of agricultural protection. Interpreting estimated parameters as causal effects implies that the maximal increase in agricultural protection induced by a shift from a majoritarian to a mixed system amounts to 0.12 in the short run and 0.31 in the long run. Analogously, a change from a pure proportional representation to a mixed system induces an increase in the NRA measure of 0.07 in the short run and 0.21 in the long run.

Technically, several alternative estimation procedures have been suggested in the literature to solve endogeneity problems. See Persson and Tabellini (2003) for an overview in the framework of endogeneity of political institutions. Here, an IV estimation approach as suggested by Angrist (2001) is adapted to check our results against possible endogeneity of political institutions. In particular, based on the preferred model 4 three IV approaches are conducted. First, a linear probability model is used for instrumenting the political institutions, second a multinomial logit model addressing explicitly the discrete nature of the possibly endogenous electoral rule, and finally a 2SLS estimation is performed for the continuous measurement of electoral rules. For all specifications, the exogeneity of our electoral rule variables is assessed via Wu-Hausmann tests. Within the first stage of estimation, the instruments are used to provide a prediction of the political institutional variables, via linear and nonlinear regressions.[16]

As has been manifold discussed in the literature, the critical part within an IV estimation is to find valid instruments, e.g., variables that on the one hand are sufficiently correlated with the endogenous variable, but not with the error term of the explained variable. Otherwise the IV strategy will not solve the endogeneity problem. In the context of endogenous political institutions, Persson and Tabellini (2003) propose a set of instrument variables for constitutional rules that includes variables dating the origin of the current constitution, a variable indicating a British colonial history of a country, two variables as measures of geography, as well as two indicators of cultural heritage. Further endogeneity of political institutions is discussed by Aghion et al. (2004), who focus on ethnic fragmentation as an important determinant of constitutional choices.

We choose instrument variables out of the set that are relevant for our sample and that can be theoretically declared as exogenous. Because the two cultural and geographic variables show no variance in our sample, we refrain from using these variables. Moreover, we exclude colonial history from our set of instruments, as theoretical considerations imply reasonable doubt on the exogeneity of this variable. It is conceivable that colonial history is correlated with other factors than the electoral system influencing agricultural protection. For example, a British colonial history might imply a preference for liberal economic policies. In this case colonial history would not be a valid instrument. In contrast to colonial history, exogeneity of the constitutional timing variables including age of democracy is pretty straightforward. Accordingly, we include ethnic fragmentation, the set of constitutional timing dummies, and age of democracy into our model. Age of democracy (*age*) is directly calculated from the Polity IV data set, while the following constitutional tim-

[16]Note that the possible endogeneity of macroeconomic controls has also been subject to robustness checks. We adapt the common approach in macroeconometrics which uses lagged variables as instruments. With lagged macroeconomic controls no changes with respect to significance and direction in parameter estimates occurred. Results are provided upon request by the authors.

ing variables indicating whether a country adopted its electoral rule after 1981 ($CON81$), between 1951-80 ($CON5180$), between 1921-50 ($CON21$), with before 1921 as an omitted category, are taken from data provided by Persson and Tabellini (2003). Additionally, we use the ethnic fractionalization index (*ethnic*) provided by Alesina et al. (2003) to consider the theory of Aghion et al. (2004).

Table 2 presents the estimation results. Overall, results on the effect of political institutions are robust when endogeneity is taken into account. Furthermore, no evidence is found for endogeneity of electoral rules for our panel of countries. This is in line with the results reported by Persson and Tabellini (2003). Therefore, we overall conclude that our estimation results implying an inverse u-shape relation of district size and agricultural protection are quite robust, implying empirical support of our theory.

Table 2: Political Institutions and Agricultural Protection: Robustness

	(I)	(II)
constant	0.0385	−0.1057
	0.0389	0.1151
lNRA	0.6817	0.6902
	0.0347	0.0358
initialgdppc	−0.0098	−0.0059
	0.0051	0.0051
gdppcgrowth	0.0069	0.0073
	0.0027	0.0028
factorend	−0.1202	−0.1202
	0.0441	0.0370
budget	−0.0321	−0.0629
	0.0396	0.0455
compad	−0.0556	−0.0185
	0.0477	0.0435
emplln	−0.1315	−0.1083
	0.0361	0.0351
maj	−0.0732	0.8325
	0.0517	0.7723
prop	−0.1082	−0.7951
	0.0551	0.6837
distr		0.8325
		0.7390
distr^2		−0.7951
		0.6837
PROTEC$^+$	0.6858	0.6514
	0.0814	0.0758
Wald test $\nu = 0$	22.8235**	4.2408
AIC	-2.9642	-2.9553
BIC	-2.4619	-2.4530
Hausman Wu test	0.59	0.074
overid-test	0.412	1.771

	maj	pr	distr	distr^2
con	5.2160	3.7996	0.4306	0.4332
	7.2551	5.3733	0.3317	0.3634
age	−11.7194	−7.3828	−0.1441	−0.2771
	12.1089	11.6018	0.3151	0.3617
ethnic	9.4215	−2.3631	0.8922	1.0204
	6.7143	24.7827	0.2990	0.3285
con2150	−4.9326	−1.4873	−0.2990	−0.3533
	4.5511	3.8707	0.2493	0.2686
con5180	−4.7731	−1.4890	−0.1041	−0.1783
	6.7132	4.6266	0.2017	0.2209
con81	−19.0316	−1.9684	−0.5982	−0.7049
	64.9286	7.7039	0.2216	0.2395
$\overline{R}^2$/			0.5142	0.5254

Notes: Standard errors are given in parentheses. * indicates significance at the 10 percent level, ** indicates significance at the 5 percent level, and *** indicates significance at the 1 percent level.

But can we interpret estimated parameters as causal effects of electoral rules? In this regard an interesting argument refers again to Acemoglu and Johnson (2005), raising the point of clustered institutions. In particular, Acemoglu and Johnson (2005) point out that political, economic, social and legal institutions are likely evolving jointly, i.e., reinforcing each other and thus can only be observed in specific clusters. Accordingly, even if instruments are able to identify correct parameters, an interpretation of these parameters as a causal effect of a specific institution might still be problematic due to the fact of "clustered" institutions.

Interestingly, it follows directly from our theoretical model that the specific impact of electoral rules on agricultural protectionism crucially depends on the relative strength of agrarian when compared to non-agrarian interest groups. In specific terms only if agrarian interest groups are sufficiently strong to dominate political communication inducing voters to expect high protection outcomes a switch from majoritarian to a mixed system implies an increase in protection. In contrast, assuming agrarian interest groups are comparatively weak implies voters expect low agricultural protection levels and, hence, the same switch implies a decrease in protection. Thus, our theory of agricultural protection is a good example of Acemoglu's "clustered institutions" argument. Therefore, although our IV estimations imply that we are able to find valid instruments for electoral rule variables, based on our theory estimated parameters are causal effects of clustered institutions combining electoral rules, coalition discipline and interest groups. Accordingly, these can be at best interpreted as conditional causal effects of electoral rules.

5. Conclusion

Despite numerous bilateral and multilateral agreements on trade liberalization, agriculture remains highly protected in many but not all industrialized countries. Thus, inspired by new comparative political economy studies of Persson and Tabellini (2003) and others, our main aim in this article is to explain how agricultural protection levels are systematically tied to different modes of political representation in industrialized countries.

To this end, this paper suggests a political economy model that explicitly takes the role of electoral rules on agricultural policy outcomes into account. In detail, our model derives legislators' policy preferences within a probabilistic voting environment where agrarian voters are less ideologically committed when compared to non-agrarian voters. As a consequence our theory implies that, while all electoral systems tend to protect agricultural interests at the expense of the general public, especially mixed systems when compared to majoritarian or proportional systems, respectively, induce high governmental incentives to protect agriculture.

In particular, our theory implies for parliamentary systems an interaction effect between formal electoral rules, coalition discipline and the influence of interest groups on voters' beliefs as informal political institutions. Assuming agrarian interest groups have a strong influence on voters' beliefs implies voters expect a pro-agrarian policy. In this case electoral competition drives the prime minister to favor high agricultural protection, whereas urban districts are pivotal within the parliamentary majority. In bargaining within the legislature, this generates a conflict between the prime minister and the parliamentary majority, where the bargaining result mainly depends on coalition discipline.

As district size grows and the electoral system converges to a pure proportional system, both of these biases are attenuated. Overall, an inverse u-shaped relationship results between district size and agricultural subsidies. Assuming, however, voter beliefs tend toward a liberal agricultural policy the prime minister tends to favor urban concerns and a rural legislator becomes decisive within his parliamentary majority. Accordingly, a u-shaped relationship results. Hence, in contrast to classical political economy approaches, our theory can explain observed large cross-country variation of agricultural protection levels even within industrialized countries.

Using a dynamic panel estimation based on time-series cross-country data for 23 parliamentary democracies since 1962, our theory is confirmed empirically. Interpreting estimated parameters as causal effects would imply that electoral rules have an important impact on agricultural protection levels, which is comparable with the impact of WTO restrictions. However, we consider our theory of agricultural protection as a good example of Acemoglu's "clustered institutions" argument. Although our IV estimations imply that we are able to find valid instruments for electoral rule variables, estimated parameters have to be interpreted as causal effects of clustered institutions combining electoral rules, coalition discipline and interest groups. Therefore, we conclude that identified effects can be at best interpreted as conditional causal effects of formal electoral rules. Please note further that homogeneity of countries included in the data sample in regard to unobserved informal institutions is crucial for the identification of causal effects of clustered institutions. We tried to take care of this problem by following a conservative strategy that just allows to include industrialized countries protecting agriculture into our country sample.

However, future research might focus on more advanced estimation techniques, e.g., applying switching regression models with latent regimes that allow identification of causal effects even within a more heterogeneous sample comprising of industrialized and developing countries.

A. Proof of proposition 2

Part (1) and (2) follow directly from proposition 1 given the assumption that only two types of legislators' preferences, rural and urban, exist. These preferences correspond to an additive SWF characterized by specific relative weights of the agricultural population, $\beta^A_{u_k}$ and $\beta^A_{r_k}$, which depends on the relative shares of the agricultural population in the rural and urban district, respectively.

Proof of Part (3): Note that for any common voter belief $\tilde{s} \in S_k$ it holds:

$$\Pi^u(\tilde{s}) = \Pi^r(\tilde{s}) = 0. \tag{26}$$

Note further that it holds:

$$\begin{aligned} \Pi^u(s) > 0 \quad \Pi^r(s) < 0 \quad \text{for } s^-_k \leq s < \tilde{s} \\ \Pi^u(s) < 0 \quad \Pi^r(s) > 0 \quad \text{for } s^+_k \geq s > \tilde{s} \end{aligned} \tag{27}$$

Therefore, it follows that $Y^{PM}_k = \tilde{s}$ delivering directly part 3. Finally, to prove part (4) assume electoral competition implies that $Y^{PM}_k > \bar{s_k}$. For simplicity we assume $C_A = 1$ [17], then it follows:

$$Y^{PM}_k = \min\left\{s^+_k, s^r_k\right\} \tag{28}$$

Then using part (2) the final policy outcome results as:

$$s^*_k = \min\{Y^{PM}_k, \tilde{s}^+_k\} \tag{29}$$

By assumption it holds:

$$\alpha^A_{d_k} \leq \alpha^A_{d_{(k-1)}} \quad \forall d \in D^R \quad \text{and} \quad \alpha^A_{d_k} \geq \alpha^A_{d_{(k-1)}} \quad \forall d \in D^U \tag{30}$$

Thus, it directly follows:

$$s^r_k \geq s^r_{k+1}, \quad s^u_k \leq s^u_{k+1}, \quad s^-_k \geq s^-_{k+1}, \quad s^+_k \leq s^+_{k+1} \tag{31}$$

Therefore, it follows that if there exists a $k^+ = 1, .., N$ such that it holds: $s^r_{k^+} \leq s^+_{k^+}$, then it already holds:

$$s^r_k \leq s^+_k \quad \forall k \geq k^+ \tag{32}$$

Obviously, there always exists such a k^+, i.e., equation 32 holds for $k^+ = N$. We define k^* as the minimum of all k^+'s for which equation 32 holds. Trivially, k* always exists and it follows:

$$\begin{aligned} Y^{PM}_k = \min\left\{s^+_k, s^r_k\right\} = s^+_k \quad \forall k < k^* \\ Y^{PM}_k = \min\left\{s^+_k, s^r_k\right\} = s^r_k \quad \forall k \geq k^* \end{aligned} \tag{33}$$

Therefore, it follows that s^*_k equals s^+_k for all $k < k^*$ and s^*_k equals s^r_k for all $k \geq k^*$. Thus, the first statement of part (4) is proven. The proof of the second statement is perfectly analogous (assuming $C_A = 0$) and therewith *proposition 2* is proven.

Q.E.D.

B. Data description

[17] Please note that the proof will not change substantially if we assume that C_A is lower, but sufficiently close to 1.

Table 3: Summary statistics

Variable	Mean	Standard Deviaton	Minimum	Maximum
NRA	0.713	0.938	-0.244	4.239
initialgdppc	14.094	6.956	1.583	33.295
gdppcgrowth	2.570	3.181	-15.840	12.507
compad	0.652	0.277	0.219	1.991
factorend	0.209	0.337	0.004	1.255
budget	0.024	0.390	-0.926	1.547
emplln	-2.340	0.729	-4.142	-0.502
distr	0.383	0.411	0.007	1.000

Table 4: Sample characteristics

Country	first year	last year	# obs.	øNRA	ødistr
Australia	1971	2005	35	0.05	1.000
Austria	1971	1993	23	0.31	0.049
Canada	1962	2005	42	0.27	1.000
Czech Rep.	1993	2002	10	0.24	0.055
Denmark	1966	1971	6	0.37	0.075
Estonia	1993	2002	10	0.11	0.109
Hungary	1993	2002	10	0.24	0.053
Iceland	1980	2005	23	2.49	0.135
India	1998	2004	7	0.14	1.000
Japan	1965	2005	41	0.93	0.226
Latvia	1993	2002	10	0.22	0.050
Lithuania	1993	2002	10	0.17	0.028
Malaysia	1989	2004	16	0.01	1.000
New Zealand	1971	2005	32	0.11	0.790
Norway	1971	2005	26	2.67	0.117
Poland	1993	2002	10	0.17	0.066
Portugal	1976	1984	9	0.19	0.089
Romania	1993	2005	13	0.43	0.124
Slovak Rep.	1998	2002	5	0.43	0.007
Slovenia	1993	2002	10	0.73	0.163
Sweden	1971	1993	23	0.84	0.070
Switzerland	1990	2005	16	2.69	0.110
Turkey	1984	2005	22	0.17	0.184
Total			412	0.713	0.393

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