Zur Spezifizierung von Risiko und Unsicherheit in räumlichen Modellen

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Zusammenfassung

Beiträge zur räumlichen Theorie des Wählens gehen zumeist von idealisierten Bedingungen aus: Wähler sind vollständig informiert und entscheiden sich strikt rational, Parteien beziehen klare und eindeutig identifizierbare Positionen im politischen Wettbewerbsraum, Wählereinstellungen zum Umgang mit Risiko und Unsicherheit sind à priori in theoretischen und statistischen Modellen fixiert. Dieser Beitrag hinterfragt diese Grundannahmen der „Neo-Downsianischen" Modelltradition. Er bestimmt empirisch, wie Wähler räumliche Distanzen in Nutzenfunktionen übersetzen und wie sie dabei mit Risiko und Unsicherheit umgehen. Ein wesentlicher Aspekt betrifft dabei die Angemessenheit von konkaven oder konvexen Nutzenfunktionen, also die Frage, ob theoretische und/ oder statistische Modelle Verlustfunktionen mit quadratischen oder mit linearen Metriken spezifizieren sollten. Die empirische Analyse verwendet das umfangreiche Datenmaterial desWahlforschungsprojekts „The **C**omparative **S**tudy of **E**lectoral **S**ystems" (CSES). Vergleichende Analysen des Wahlverhaltens zeigen dabei eindeutig, dass Wähler über neunzig heterogene Wahlkontexte hinweg wesentlich weniger risikoavers sind als von der großen Mehrheit theoretischer und empirischer Beiträge unterstellt wird. Stattdessen zeigen dieser Beitrag, dass moderne Wähler sich im Wesentlichen risikoneutral verhalten.

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1 Einleitung

Wissenschaftliche Modelle sind Abstraktionen. Sie sollen nicht wahr, sondern beim Erreichen von deskriptiven oder kausalen Schlussfolgerungen nützlich sein, und sie gründen regelmäßig auf einer Reihe notwendiger, teils restriktiver Annahmen. Die räumliche Theorie des Wählens unterstellt zum Beispiel, dass Wähler und Parteien fixe und eindeutig identifizierbare Positionen in *n*-dimensionalen ideologischen und/oder politischen Räumen einnehmen und dass die Relation von Wähler- und Parteipositionen das Wahlverhalten mindestens mitbestimmt.

In diesem Beitrag greife ich zwei dieser Grundannahmen auf: Die erste betrifft die Informiertheit der Wähler und, somit, die Sicherheit ihrer Entscheidung. Beinah alle Beiträge des Feldes unterstellen umstandslos, dass Wähler vorgegebene Dimensionen eines politischen Raums verstehen, auf jeder dieser Dimensionen bestimmte persönliche Idealpunkte einnehmen und dass sie die Positionen der wesentlichen Kandidaten und/oder Parteialternativen auf denselben Dimensionen genau kennen und exakt vermessen können. Eine ganze Reihe von Gründen spricht jedoch gegen diese idealtypischen Annahmen: Wähler sind unterschiedlich stark an politischen Fragen interessiert, und sie verfügen über unterschiedliche Ressourcen bei der Verarbeitung politischer Informationen. Politische Parteien haben oft ein strategisches Interesse daran, unklare (oder gar keine) Positionen zu bestimmten Sachfragen zu beziehen, sie sind manchmal intern gespalten und richten unterschiedliche programmatische Botschaften an unterschiedliche Segmente der Wählerschaft (vgl. Alvarez 1998; Bartels 1986; Brady und Ansolabehere 1989; Lupia 2016; Shepsle 1972; Tomz und Van Houweling 2009).

Die zweite Annahme betrifft die Wahrnehmung von geometrischer Nähe und Distanz in politischen Räumen und die Übersetzung dieser Maße in die jeweiligen Nutzenfunktionen einzelner Wähler. Die Auswahl einer geeigneten Distanzmetrik ist dabei kein triviales oder bloß technisches Detail, sondern sie hat weit reichende Konsequenzen für das theoretische Verständnis und die empirische Modellierung des Wahlverhaltens. In der Literatur sind vor allem Spezifikationen verbreitet, die à priori entweder quadratische und lineare Metriken unterstellen:

- Quadratische Verlustfunktionen setzen risikoaverse Wähler voraus. Sie sind mathematisch recht einfach handhabbar und bestimmen seit Langem den Status quo der „Neo-Downsianischen" Modelltradition (vgl. anstatt vieler weiterer Beiträge, Adams et al. 2005; Davis et al. 1970; Enelow and Hinich 1984; Feddersen 1992; Merrill und Grofman 1999; Schofield and Sened 2006).

- Lineare Verlustfunktionen unterstellen dagegen das risikoneutrale Verhalten der Wähler. Der ursprüngliche Beitrag von Downs (1957) benutzt lineare Distanzen, und sie werden weiterhin von einer sichtbaren Minderheit der Beiträge zur räumlichen Theorie des Wählens aus theoretischen oder aus empirischen Erwägungen bevorzugt (vgl. zum Beispiel Degan und Merlo 2009; Eguia 2012; Grynaviski und Corrigan 2006; Humphreys und Laver 2010; Kramer 1977; Patty et al. 2009; Wittman 1973, 1977).
- Neben diesen einfachen Basisoptionen haben theoretische und empirische Beiträge eine Reihe komplexerer Alternativen diskutiert und verwendet, etwa exponentielle (Carroll et al. 2013; Poole und Rosenthal 1985) oder nicht-euklidsche Verlustfunktionen (McKelvey 1976; Plott 1967). Schließlich verlassen einige weitere Modelle vollständig die Logik von Nähe und Distanz in euklidschen Räumen: Das Modell der Richtungswahl begreift politische Sachfragen als binär. Es begründet den eher affektiven und symbolischen Nutzen einer Parteialternative durch die richtungspolitische Übereinstimmung und die Intensität, mit der sie von Wählern und Parteien vertreten wird (Rabinowitz und Macdonald 1989).

Das Interesse an diesen Fragestellungen ist gleichermaßen aus methodischer wie aus inhaltlicher Perspektive wohlbegründet. Die beiden oben dargestellten Grundannahmen bestimmen (1) wie und in welchem Maße Unsicherheit über die Position bestimmter Parteien in ideologischen oder politischen Räumen entsteht und (2) auf welche Weise einzelne Wähler mit dem jeweiligen Maß an Unsicherheit umgehen. Theoretische wie statistische Modelle müssen diese Prämissen adäquat formulieren, sodass substanzielle Einsichten in den politischen Prozess, hier in die Determinanten des Wahlverhaltens, überhaupt erst möglich werden. Die hier nur kursorisch vorgestellten Modellspezifikationen gründen auf ganz unterschiedlichen Konzepten zum Umgang mit Risiko in Entscheidungssituationen. Diese Risikoorientierungen bestimmen wiederum wesentlich, wie Wähler mit begrenzten oder unsicheren Informationen umgehen, wie sie auf unbestimmte oder unklare Parteipositionen reagieren und wie sie Parteialternativen bewerten und letztlich auswählen.

Freilich können Wahlforscher nicht einfach à priori und axiomatisch unterstellen, welche konkreten Überlegungen Wähler anleiten, wenn sie räumliche Distanzen wahrnehmen und in bestimmte Nutzenfunktionen übersetzen. Und diese Frage kann auch nicht (allein) am Maßstab von mathematischer Schlichtheit und praktischer Verwendbarkeit entschieden werden. Vielmehr ist die Auswahl einer geeigneten Distanzmetrik eine letztlich empirische Frage, die durch die Analyse geeigneter Wahlstudien beantwortet und gerechtfertigt werden muss.

Beim gegenwärtigen Stand der Forschung sind nicht viele Autoren diesen Weg gegangen: Am Material amerikanischer Präsidentenwahlen haben zum Beispiel Berinsky und Lewis (2007), Grynaviski und Corrigan (2006) und Ye et al. (2011) gezeigt, dass lineare Distanzen den Datenstrukturen besser entsprechen und eine höhere Güte der Schätzung ermöglicht haben als quadratische Nutzenterme. Bei ähnlich angelegten Analysen deutscher Bundestagswahlen haben Thurner (2000) und Shikano und Behnke (2009) diese Befunde im Wesentlichen bestätigt. Schließlich hat Singh (2014) die bislang einzige ländervergleichende Studie vorgelegt, die am Material der ersten beiden Wellen des CSES-Projekts zeigt, dass lineare Nutzenterme, wieder gemessen an der Güte der Schätzung, bessere Erklärungen bieten für die Wahlbeteiligung und die Wiederwahl von Amtsinhabern.

In diesem Beitrag möchte ich über die bisherigen Fallstudien hinausgehen und gleichermaßen theoretische Überlegungen anstellen und empirische Befunde vollständig spezifizierter räumlicher Modelle präsentieren. Diese systematisch-vergleichende Analyse greift auf umfangreiche Daten der ersten drei Wellen des Projekts „The Comparative Study of Electoral Systems“ zurück (CSES; http://cses.org/). Substanziell konzentriere ich mich auf Wahlen zu nationalen Parlamenten, weil andere Urnengänge, etwa Präsidentenwahlen, oft eine ganz unterschiedliche Salienz aufweisen, meist viel deutlicher personalisiert und deshalb kaum sinnvoll vergleichbar sind. Die empirischen Befunde unterstreichen, über eine weite Bandbreite unterschiedlicher politischer Kulturen, Parteiensysteme und institutioneller Systeme hinweg, dass einzelne Wähler und ganze Elektorate sehr viel weniger risikoavers sind als bislang theoretisch angenommen und in den meisten Modellen empirisch spezifiziert wird. Über neunzig unterschiedliche Parlamentswahlen hinweg identifizieren die Ergebnisse nicht einen Fall, bei dem die Wähler derart risikoscheu sind wie es das „Neo-Downsianische“ Standardmodell quadratischer Verlustfunktionen unterstellt. Vielmehr tendieren die meisten Elektorate hin zu einem risikoneutralen Verhalten.

Um diese Argumentation vorzustellen und empirisch abzusichern, gehe ich in drei wesentlichen Schritten vor: Zunächst wiederhole ich einige konzeptionelle Schlüsselbegriffe und theoretische Grundlagen der räumlichen Theorie des Wählens. Dabei betone ich besonders die Vermessung räumlicher Nähe und Distanz und die theoretische Eigenschaften der Nutzenfunktionen (Abschn. 2). Im Anschluss stelle ich das empirische Datenmaterial vor, definiere und operationalisiere Schlüsselvariablen des räumlichen Modells und einige nicht-räumliche Kontrollvariablen. Die folgende empirische Analyse verwendet ausschließlich Daten der Umfragemodule, die in den ersten drei Wellen des Projekts „The Comparative Study of Electoral Systems“ abgefragt und erhoben wurden

(Abschn. 3). Im Anschluss stelle ich im Detail die empirischen Modelle der Wahlentscheidung vor und präsentiere vergleichende Befunde zur Rolle von Risikoorientierungen bei der Wahlentscheidung (Abschn. 4). Das Fazit fasst wesentliche Resultate zusammen, diskutiert ihre Implikationen für die theoretische und empirische Wahlforschung und skizziert Grundrisse eines zukünftigen Forschungsprogramms.

2 Grundannahmen des räumlichen Modells

Dieser Abschnitt beginnt mit einer kurzen Wiederholung einiger Grundlinien und zentraler Bausteine der räumlichen Theorie des Wählens. Wie bereits ausgeführt unterstelle ich, dass Wähler- und Parteipositionen in einem politischen Wettbewerbsraum sinnvoll verortet, dargestellt und aufeinander bezogen werden können. Um die theoretischen und statistischen Modelle nicht zu komplex werden zu lassen, beziehe ich mich jedoch in diesem Beitrag allein auf *eindimensionale* politische Räume.

Deterministische und probabilistische Versionen der räumlichen Theorie stimmen darin überein, dass ein Wähler v sich für denjenigen Kandidaten oder diejenige Parteialternative p entscheidet, von dem/der er sich den höchsten Nutzen $u(v,p)$ verspricht. Die rationale Nutzenkalkulation der Wähler enthält deshalb eine Verlustfunktion. Sie definiert, wie der „Nutzen“, den jeder Wähler i jeder Partei j zuschreibt, mit ihrer räumlichen Distanz im Koordinatensystem sinkt:

$$u(v,p) = f(v,p) = -\|v-p\| = -\left\|v_i - p_{i,j}\right\|. \tag{1}$$

Die Variablen v_i und $p_{i,j}$ benennen hier individual- und alternativenspezifischen Wähler- und Parteipositionen innerhalb des eindimensionalen politischen Raums. Zum Beispiel bezeichnet v_i den Idealpunkt von Wähler i, während $p_{i,j}$ die idiosynkratisch wahrgenommene Verortung von Partei j durch Wähler i angibt. $\|\bullet\|$ ist eine beliebige Distanzfunktion, die bestimmt, wie Entfernungen im politischen Raum in wähler- und parteispezifische Nutzenterme übertragen werden. Obgleich lineare und, mehr noch, quadrierte Distanzen am häufigsten zur Bestimmung von Entfernungen in politischen Wettbewerbsräumen verwendet werden, ist ein binärer Vergleich dieser Alternativen dennoch unnötig restriktiv. Vielmehr können lineare ($\beta = 1$) und quadrierte Abstandsmaße ($\beta = 2$) als Sonderfälle einer allgemeineren Spezifikation $L_\beta(v,p)$ verstanden werden (vgl. zum weiterführenden Konzept der Minkowski-Distanz Eguia 2013; Humphreys und Laver 2010; Laver und Hunt 1992; Thurner 2000; Ye et al. 2011):

$$u(v,p) = -L_\beta(v,p) = -\left|v_i - p_{i,j}\right|^\beta \tag{2}$$

Ich unterstelle zunächst, dass die Wahlentscheidung und die Parameter v_i und $p_{i,j}$ fixierte und gemessene Daten sind. Der Formparameter β wird dagegen durch ein statistisches Modell geschätzt; β gibt die Form der Nutzenfunktion an und bestimmt damit die Risikoorientierung der Wähler. Die Festlegung einer konkreten Distanzmetrik geht weit über den bloß technischen oder statistischen Teil des Arguments hinaus. Inhaltlich bestimmt die Form der jeweiligen Nutzenfunktion wie ein Wähler mit Risiko und Unsicherheit umgeht. Anstatt das Risikoprofil aller *(sic!)* Wähler vorab als quadratisch ($\beta = 2$) oder linear ($\beta = 1$) festzulegen, bestimme ich in diesem Beitrag den Koeffizienten empirisch: Schätzer von $\beta > 1$ zeigen generell riskikoaverse Elektorate an, $\beta \sim 1$ identifiziert risikoneutrale, und $\beta < 1$ bezeichnet risikoaffine Elektorate.

Beim eindimensionalen Parteienwettbewerb bleibt die Rangfolge der Alternativen unabhängig von der gewählten Distanzmatrix erhalten: Ordinale Nutzenfunktionen, die Reihung der Präferenzen und vorhergesagtes oder „korrektes“ Wahlverhalten werden durch die Spezifikation linearer oder quadratischer Verlustfunktionen nicht verändert. Die Auswahl der Distanzmetrik modifiziert jedoch kardinale Nutzenfunktionen: Bei linearen Verlustfunktionen fällt der Nutzen einer Partei j für Wähler i gleichmäßig mit der zunehmenden Distanz von Wähler und Parteialternative $\left|v_i - p_{i,j}\right|$. Bei quadratischen Verlustfunktionen fällt dagegen der Nutzen immer stärker ab, wenn sich das Differential von Wähler- und Parteiposition $\left|v_i - p_{i,j}\right|^2$ weiter erhöht. Lineare Distanzfunktionen reagieren deshalb vergleichsweise empfindlich auf kleinere Differenzen in der Nähe des Idealpunktes der Wähler, quadratische Verlustfunktionen fallen dagegen schroffer ab, wenn Wähler- und Parteipositionen ohnehin weit voneinander entfernt sind.

Abb. 1 charakterisiert und vergleicht risikoaverse (=konkave) und risikoaffine (=konvexe) Nutzenfunktionen in eindimensionalen politischen Wettbewerbsräumen. Diese exemplarische Darstellung wirft zusätzliches Licht auf die theoretischen Implikationen unterschiedlicher Verlustfunktionen und illustriert darüber hinaus die Bedeutung von Unsicherheit über die ideologischen oder politischen Positionen der Parteien. Innerhalb eines stilisierten, eindimensionalen politischen Raums konkurrieren zwei Parteien um Wählerstimmen: p_j benennt eine präzise Partei, die durch einen Punktschätzer abgebildet wird; $p_{j'}$ benennt eine unsichere Plattform, die nicht mit Bestimmtheit wahrgenommen, sondern durch eine Wahrscheinlichkeitsverteilung unscharf charakterisiert wird:

- Das linke Teilbild trägt mit einer durchgezogenen Linie die konventionelle Annahme risikoaverser Wähler und konkaver Nutzenfunktionen ab. Bei risikoaversen Nutzenfunktionen steigt der marginale Nutzenverlust mit zunehmender räumlicher Distanz stetig an, und bei gleicher Positionierung

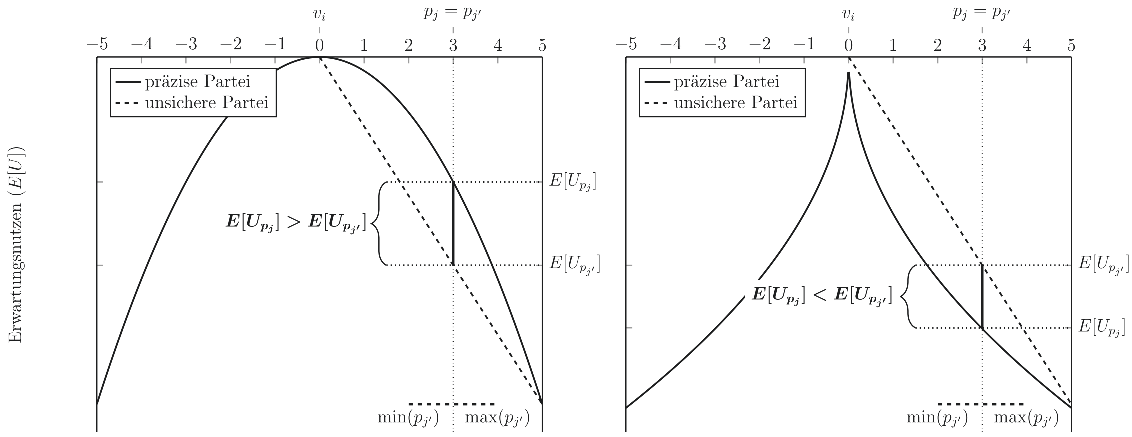

Abb. 1 Die Form der Nutzenfunktion und Risikoeinstellungen der Wähler

beider Parteien $(p_j = p_{j'})$ trägt die präzise Partei dem WÄhler stets einen höheren erwarteten Nutzen ein als die unsichere Partei $\left(E\left[U_{p_j}\right] > E\left[U_{p_{j'}}\right]\right)$.

- Das rechte Teilbild illustriert die Konsequenzen von Unsicherheit bei den (mindestens in Teilbereichen) konvexen Nutzenfunktionen risikoaffiner Wähler. Bei risikoaffinen Wählern nimmt der marginale Nutzenverlust hin zu den Rändern der Funktion stetig ab, und bei einer identischen (durchschnittlichen) Aufstellung beider Parteien $(p_j = p_{j'})$ erreicht die unsichere Plattform einen höheren erwarteten Nutzen als die präzise $\left(E\left[U_{p_{j'}}\right] > E\left[U_{p_j}\right]\right)$.

Generell haben risikoaverse Wähler konkave Nutzenfunktionen, und ein Wähler gilt immer dann als risikoavers wenn er den erwarteten Nutzen einer präzisen Parteialternative einer fairen Lotterie mit demselben Erwartungsnutzen vorzieht. Im Gegensatz haben risikoaffine Wähler mindestens teilweise konvexe Nutzenfunktionen und bevorzugen eine unsichere Lotterie gegenüber der fixen Auszahlung ihres Erwartungsnutzens.

3 Daten und Variablen

Dieser Beitrag schätzt angemessene Distanzmetriken empirisch und bestimmt die Form der Nutzenfunktionen in räumlichen Modellen der Wahlentscheidung in einer systematisch-vergleichenden Perspektive. Die vergleichende Wahlforschung verwendet für derartige Fragestellungen meist die parallel konstruierten Umfragemodule des Projekts „The Comparative Study of Electoral Systems“ (CSES; http://www.cses.org/). Der umfangreiche Datenbestand des CSES-Projekts hat sich als *die*

Datenquelle für systematische, ländervergleichende Analysen des Wahlverhaltens etabliert. Diese Untersuchung des Wahlverhaltens unter Risiko und Unsicherheit verwendet eine Kumulation der ersten drei Wellen des CSES-Projekts. Diese Datenbasis umfasst 129 einzelne Umfragemodule zu nationalen Wahlen in mehr als vierzig Vergleichsstaaten. Konkret benutze ich den „CSES Module 1–3 Harmonized Trend File", der am Wissenschaftszentrum Berlin von Heiko Giebler, Josephine Lichteblau, Antonia May, Reinhold Melcher, Aiko Wagner, und Bernhard Weßels zusammengestellt wurde. Daten und detaillierte Dokumentation sind über die CSES-Website verfügbar (http://cses.org/datacenter/trendfile/trendfile.htm; vgl. auch Anhang 6).

Die CSES-Module der unterschiedlichen Wellen enthalten alle Fragen und alle empirischen Informationen, die für die Schätzung einfacher räumlicher Modelle erforderlich sind. Die Questionnaires erheben Daten über die Wahlentscheidung, über die Idealpunkte der Wähler auf der Links-Rechts-Dimension ($v_i \in [0, 10]$) und über idiosynkratische Platzierungen der Parteialternativen auf identisch konstruierten Elfpunkteskalen $\left(p_{i,j} \in [0, 10]\right)$.

In diesem Beitrag verwende ich zur Messung von Parteipositionen durchgehend idiosynkratische Platzierungen, weil jeder einzelne Wähler die Bewertung von Parteialternativen und die Auswahl unter diesen Alternativen letztlich nur von denjenigen Informationen abhängig machen kann, die ihm unmittelbar selbst vorliegen, ganz gleich ob sie nun, nach welchem Maßstab auch immer, „richtig" oder „falsch" sind. Mittelwerte dieser Platzierungen, die Urteile politikwissenschaftlicher Experten oder die richtungspolitische Verortung von Manifestos werden dagegen durch Informationen begründet, die einzelne Wähler regelmäßig nicht haben oder auch gar nicht haben können, und diese Maße eignen sich deshalb nur wenig zur Erklärung der individuellen Wahlentscheidung. Mit ganz ähnlicher Begründung verzichte ich darauf, Parteiplatzierungen einzelner Wähler durch weitere Skalierungsverfahren zu „korrigieren" (vgl. etwa die Techniken bei Aldrich und McKelvey 1977; Hare et al. 2015; Lo et al. 2014) oder potenzielle Projektionseffekte zu korrigieren (vgl. Grand und Tiemann 2013; Krosnick 2002).

Neben diesen Schlüsselvariablen des räumlichen Modells erhebt das CSES-Projekt eine Reihe zusätzlicher, nicht-räumlicher Prädiktoren der Wahlentscheidung: die Art und Stärke einer langfristigen Parteiidentifikation, die Zufriedenheit mit dem jeweiligen demokratischen System, eine Batterie von drei Items, die politisches Wissen der Befragten erheben, und demografische Kontrollvariablen wie Alter, Geschlecht und Bildungsstand.

Der kumulierte Datenbestand des CSES-Trendfiles deckt den Zeitraum von 1996 bis 2011 ab. Ich verwende ausschließlich Wahlen zu nationalen Parlamenten (zu den Unterhäusern bei bikameralen Systemen), sodass die empirische Datenanalyse auf neunzig einzelnen Wahlsegmenten gründet und insgesamt 44 unterschiedliche, heterogene Vergleichsstaaten einschließt. Die Auswahl eines derart heterogenen Vergleichssets ermöglicht es, entlang der Linien eines „most different systems design“ zu prüfen, ob die Resultate einzelner Umfragesegmente deutlich durch jeweilige Kontexteffekte geformt sind oder ob die Analyse generelle Befunde erbringt, die über heterogene Kontextfaktoren hinweg „halten“ und verallgemeinerbar sind (vgl. bereits Przeworski und Teune 1970; Tiemann 2003).

Mit Blick auf die einzelnen Survey-Segmente musste ich einige sehr kleine Nachwahlbefragungen ausschließen, andere Segmente des CSES-Trendfiles konnte ich aufgrund von inkonsistenten oder fehlerhaften Daten nicht sinnvoll für die Analyse aufbereiten und verwenden. Zudem habe ich einige Umfragesegmente, die keine Angaben zur Parteiidentifikation abgefragt haben, von der empirischen Analyse ausgeschlossen. (Die für die vergleichende Analyse ausgewählten Segmente des CSES-Trendfiles sind in Anhang 6 aufgelistet und dokumentiert.).

Schließlich habe ich einige individuelle Beobachtungen nicht berücksichtigen können: Das betrifft zuerst Nichtwähler und die Unterstützer kleinerer Parteien, für die in den CSES-Questionnaires keine alternativenspezifischen Parteipositionen erhoben werden. Weiterhin nicht in die Analyse aufgenommen werden diejenigen Befragten, die keinen Idealpunkt auf der Links-Rechts-Dimension angegeben oder eine Mehrzahl der möglichen Parteialternativen, also mehr als fünfzig Prozent der Wettbewerber bei einer Parlamentswahl, nicht im politischen Wettbewerbsraum platziert haben. Umfassender spezifizierte, gesättigte Modellen, die auch Effekte nicht-räumlicher Variablen einschließen, konnten zudem Befragte nicht berücksichtigen, die keine Angaben zur Demokratiezufriedenheit, Alter, Geschlecht oder Bildungsstand gemacht haben.

4 Die Form der Nutzenfunktion

In diesem Abschnitt übersetze ich die Forschungsfrage und die theoretischen Modelle in ein stringentes Forschungsdesign. Ich stelle zunächst den Ansatz zur empirischen Datenanalyse vor, diskutiere und rechtfertige Grundentscheidungen bei der Auswahl und Spezifizierung des statistischen Modells und beschreibe die Operationalisierung und Messung wesentlicher Modellvariablen. Im folgenden Abschnitt stelle ich die empirischen Befunde vor und zeige, dass die über die unterschiedlichen, heterogenen Kontexte des kumulierten CSES-Datensatzes hinweg Wähler eher risikoneutrale als risikoaverse Positionen einnehmen, wenn sie Parteialternativen bewerten und ihre Stimme abgeben. Sodann erweitere ich sukzessive das Modell und berücksichtige die Effekte nicht-räumlicher Kontrollvariablen und eines empirisch bestimmten Unsicherheitsindikators.

4.1 Von theoretischen zu statistischen Modellen

Das statistische Modell bildet die Grundlinien des theoretischen Arguments genau ab: Wahlforschung sollte sich mit der Wahlentscheidung beschäftigen, und die Mehrheit an einschlägigen Beiträgen versucht zu erklären, warum einer (und nur einer) der Wahlvorschläge ausgewählt wird. Ein alternativer Ansatz versteht Parteiskalometer oder Angaben zur „Propensity to Vote" (PtV) als direkte Messung des kardinalen Nutzens (vgl. van der Eijk et al. 2006). Diese Perspektive ist somit nicht nur voraussetzungsvoller, sondern sie gründet auch auf recht unklar definiertem Material. Skalometer- oder PtV-Skalen innerhalb derselben Befragung unterscheiden sich oft deutlich, und ihre konkrete Formulierung greift anstatt kurzfristiger oder strategischer Wahlmotive oft längerfristige Einschätzungen auf. Die Analyse von Skalometerdaten findet meist mit linearen Modellen statt, und diese Ergebnisse richten sich nicht direkt auf die Wahlentscheidung, sondern sie sind kontaminiert durch die Bewertungen verschiedener Parteialternativen, die in der Präferenzordnung der Wähler weit unten stehen.

Ich analysiere in diesem Beitrag deshalb nicht allgemein das Gesamtspektrum aller Parteibewertungen und -präferenzen, sondern konzentriere mich spezifisch auf die jeweilige Entscheidung eines Wählers. Deshalb verwende ich Verfahren zur Modellierung diskreter Entscheidungen, um die räumlichen und nicht-räumlichen Determinanten des Wahlverhaltens zu bestimmen, die Form der Nutzenfunktionen zu schätzen und die Unsicherheit der Befunde zu erschließen (eine allgemeine Einführung bietet Train 2009). Bei konventionellen „Conditional Logit"-Modellen der Wahlentscheidung gilt für die Wahrscheinlichkeit, dass Wähler i die Parteialternative j auswählt:

$$Pr(v_i = j) = \frac{\exp\left[u(v,p)\right]}{\sum_{j=1}^{J} \exp\left[u(v,p)\right]} \tag{3}$$

Dieses statistische Zufallsnutzenmodell wird durch räumliche und nicht-räumliche Prädiktoren der Wahlentscheidung operationalisiert und durch α, β und $\boldsymbol{\lambda}$ und $\boldsymbol{\theta}$ parametrisiert. Deshalb ergibt sich der implizite Gesamtnutzen von Partei j für Wähler i mit:

$$u(v,p) = -\alpha\left|v_i - p_{i,j}\right|^{\beta} + \lambda_j c_j + \boldsymbol{\theta_{i,j} D_i} + \epsilon_{i,j} \tag{4}$$

In einem einfachen Grundmodell ist der implizite Gesamtnutzen die Summe eines räumlichen Anteils, der auf den geometrischen Distanzen von Wählern und

Parteien im politischen Raum gründet $\left(\left|v_i - p_{i,j}\right|\right)$, und einer parteispezifischen Konstante c_j, die nicht explizit modellierte und über alle Wähler(gruppen) hinweg konstante Eigenschaften einer Partei j abbildet. Im folgenden Schritt spezifiziere ich besser gesättigte Modelle und füge eine Matrix an nicht-räumlichen Einflüssen auf das Wahlverhalten ein. $\boldsymbol{D_i}$ enthält als Kontrollvariablen Angaben zur Parteiidentifikation, Demokratiezufriedenheit, Indikatoren, die den Grad an politischer Information einzelner Wähler aufgreifen, das Bildungsniveau, Alter und Geschlecht jedes Wählers.

Auf der Grundlage dieser Spezifikation und der empirischen Informationen werden die Parameter des vereinigten Modells geschätzt: $\hat{\alpha}$ gibt die Salienz der räumlichen Modellkomponente an und sollte somit stets numerisch positiv und substanziell bedeutsam sein. β modelliert die Form der Nutzenfunktionen und zeigt deshalb die Risikoorientierung des Elektorats an. An dieser Stelle ist es wesentlich, daran zu erinnern, dass beide empirische geschätzte Parameter, $\hat{\alpha}$ und $\hat{\beta}$, die Salienz und die Formparameter auf der Ebene des gesamten Elektorats und nicht auf der Ebene einzelner Wähler modellieren. Mit diesem Vorbehalt identifiziert $\hat{\beta} > 1$ risikoaverse Einstellungen, $\hat{\beta} < 1$ verweist auf risikoaffine Einstellungen, und $\hat{\beta} \sim 1$ kennzeichnet vorherrschende Risikoindifferenz.

Der gleichfalls empirische bestimmte Parametervektor $\boldsymbol{\lambda}$ fängt unmodellierte Eigenschaften der konkurrierenden Parteien durch alternativenspezifische Konstanten auf, und der Parametervektor $\boldsymbol{\theta}$ modelliert nicht-räumliche Effekte auf das Wahlverhalten. Die Hinzunahme dieser Parameter vermeidet die Missspezifikationen der empirischen Modelle; dennoch schreibe ich den bestimmten Konstanten λ_j keine umfassende inhaltliche Bedeutung zu und verstehe sie nicht umstandslos als „Valenz“ oder als Indikator für die Qualität einer Partei oder ihrer Kandidaten. Bei den nicht-räumlichen Prädiktoren der Wahlentscheidung versammelt die Koeffizientenmatrix $\boldsymbol{\theta_{i,j}}$ Effektparameter für individualspezifische Variablen, die im Modell allein vor dem Hintergrund einer ausgewÄhlten Referenzkategorie interpretierbar sind. Schließlich identifiziert $\varepsilon_{i,j}$ ein stochastisches Modell, und für die Verteilung der Fehlerterme wird eine Gumbel/Type-I-Extremwertverteilung unterstellt.

Bereits Grynaviski und Corrigan (2006, S. 394) haben auf „a dizzying combination of options“ bei der Modellierung von Wahlentscheidungen und der statistischen Abbildung der räumlichen Theorie des Wählens verwiesen. Dieser Beitrag konzentriert sich auf den Umgang der Wähler mit Risiko und Unsicherheit und kann zu einigen weiteren konzeptionellen Entscheidungen nur kursorisch Auskunft geben:

Erstens kann an dieser Stelle kein ausführlicher Beitrag zur Diskussion um die angemessene Bestimmung von Parteipositionen gleistet werden. Parteialternativen werden durch die idiosynkratische Wahrnehmung jedes einzelnen Wählers und nicht durch ihre Mittelwerte oder irgendwelche übrigen „objektiven" Positionsmaße platziert. Diese Strategie wird besonders durch das Argument gerechtfertigt, dass jeder einzelne Wähler seine ohnehin subjektive und nie vollständig „rationale" Wahlentscheidung nur an denjenigen Informationen festmachen kann, über die er selbst verfügt.

Zweitens kann ich an dieser Stelle nicht zur Diskussion über die Vorzüge und/oder Nachteile des „Conditional Logit"-Modells gegenüber alternativen, komplexeren Techniken wie dem generalisierten Extremwertmodell oder dem alternativenspezifischen Probit-Modell beitragen (vgl. Alvarez und Nagler 1998). Als besonderer Vorzug der komplexeren Modelle wird häufig die Aufgabe der „independence of irrelevant alternatives"-Annahme genannt. Eine Reihe von einschlägigen Beiträgen zeigt jedoch, dass die IIA-Annahme selten empirisch relevant oder besonders restriktiv ist und die alternativen Modelle, insbesondere das multinomiale Probit-Modell, regelmäßig an Identifikationsproblemen leiden, die ihre praktische Verwendbarkeit sehr limitieren (vgl. Adams et al. 2005; Dow und Endersby 2004; Merrill und Adams 2002).

Drittens beschränke ich meine empirische Applikation auf eindimensionale politische Räume. Ohne Frage ist Politik in der modernen Massendemokratie stetig komplexer geworden, und der politische Wettbewerb ist häufig in zwei- oder mehrdimensionalen politischen Räumen organisiert. Dennoch hat eine Vielzahl neuerer Beiträge gezeigt, dass eine übergreifende ideologische Dimension, meist charakterisiert als Gegensatz von „links" und „rechts" oder „konservativ" und „liberal", weiterhin wesentliche Einzeldimensionen des politischen Wettbewerbs einschließt und, insbesondere im Kontext zunehmender politischer Polarisierung, die Struktur des politischen Wettbewerbs angemessen abbildet (Gabel und Huber 2000; Jessee 2012; Munger und Munger 2015). Die Konzentration auf einfache, eindimensionale Räume hat aber auch konzeptionelle und theoretische Gründe. Neben vielen anderen hat zum Beispiel Eguia (2013) eindringlich illustriert, dass die Charakterisierung politischer Präferenzen durch euklidsche Verlustfunktionen in mehrdimensionalen Räumen oft problematisch und voraussetzungsvoll ist und insbesondere die unterschiedliche Salienz der räumlichen Dimensionen und die Frage nach ihrer Separierbarkeit eine Fülle weiterer Komplikationen in theoretische und statistische Modelle der Wahlentscheidung hineinträgt.

4.2 Empirische Befunde des reinen räumlichen Modells

Ich beginne die Vorstellung der empirischen Befunde mit einigen simplen Basismodellen. Die Daten sind dem kumulierten CSES-Datensatz entnommen, und ich schätze für jeden einzelnen der neunzig Wahlkontexte individuelle „Conditional Logit"-Modelle der Wahlentscheidung. Schlüsselvariable sind die räumlichen Distanzen von Wählern und Parteioptionen auf einer unidimensionalen Links-Rechts-Skala $\left(v_i; p_{i,j} \in [0, 10]\right)$.

Das einfache Basismodell mit den Effektparametern α und β kann gleichermaßen durch konventionelle Maximum-Likelihood-Verfahren oder mit bayesianischer MCMC-Simulation ausgewertet werden, und für einfache Modelle erreichen beide Verfahren tendenziell identische Resultate. Bei der folgenden Verallgemeinerung hin zu komplexeren Modellen sind die Bayes-Verfahren jedoch deutlich flexibler und bieten zudem verlässlichere Angaben der statistischen Unsicherheit. Ich verwende in diesem Beitrag durchgehend nicht-informative, diffuse Prioriverteilungen und simulieren die Posterioriverteilungen von α, β und den alternativenspezifischen Konstanten λ_j mit `JAGS`. Bei jedem der neunzig heterogenen Wahlkontexte benutze ich vier Markow-Ketten mit jeweils 100.000 Iterationen nach einem „burn in" von 50.000 Iterationen. Eine Reihe von Spezifikationstests zeigt, dass jedes einzelne dieser parallel konstruierten „Conditional Logit"-Modelle das Datenmaterial angemessen abbildet. Konventionelle Geweke- und Heidelberger-Welch-Tests verweisen auf die Stationarität der Posterioriverteilungen. Ich speichere für die Beschreibung der Verteilungen nur jede zehnte Iteration („thinning"), um mit der vorhandenen Autokorrelation der Markow-Ketten angemessen umzugehen. Schließlich simuliere ich die Modelle effizienter, indem ich Kernparameter wie α und β blocke, also aus einer multivariaten Normalverteilung ziehe (theoretische und praktische Aspekte der MCMC-Simulation werden genauer beschrieben bei Gelman et al. 2013; Jackman 2009).

Abb. 2 zeigt Zusammenfassungen der Posterioriverteilungen des Salienzparamters (α; x-Achse) und des Formparameters (β; y-Achse). Beide Teilbilder bieten im wesentlichen identische Informationen: Das obere summiert bayesianische Punktschätzer und Konfidenzintervalle; das untere ersetzt diese Informationen durch die konkrete Benennung einzelner Umfragesegmente, sodass die meisten Wahlkontexte identifizierbar werden.

Die horizontale Dimension von Abb. 2 bestätigt eindeutig die Aussagekraft der räumlichen Theorie der Wahlentscheidung über eine Vielzahl heterogener Wahlen und Kontexte des kumulierten CSES-Datensatzes hinweg. Bei jeder einzelnen Wahl ist die Nähe oder Distanz von Wählern und Parteien im politischen

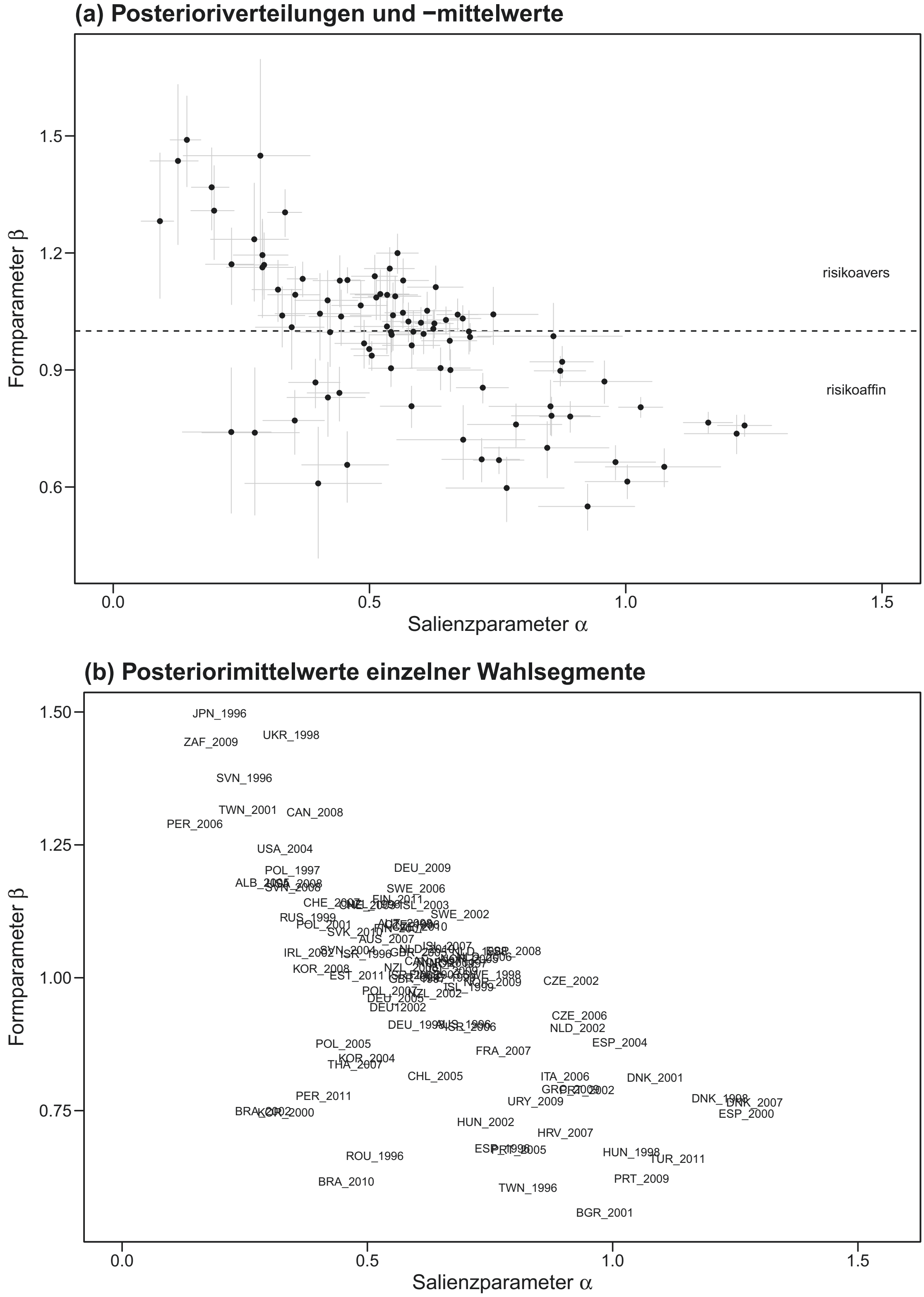

Abb. 2 Schätzer für Salienz- und Formparameter $\left(\hat{\alpha} \text{ und } \hat{\beta}\right)$

Raum eng und substanziell bedeutsam mit der Wahlentscheidung verbunden $\left(\hat{\alpha} > 1\right)$. Die simulierten Modellparameter zeigen eindeutig, dass die räumliche Theorie des Wählens über ganz unterschiedliche soziale, politische und institutionelle Kontexte hinweg wesentliche Einsichten in die Bewertung politischer Parteien und in die Wahlentscheidung eröffnet. Nicht nur in abstrakten theoretischen Modellen, sondern auch in konkreten empirischen und statistischen Analysen bevorzugen Wähler diejenigen Parteien, die ihren eigenen Idealpunkten möglichst genau entsprechen gegenüber denjenigen Plattformen, die sie weiter entfernt im politischen Spektrum verorten.

Während die räumliche Nähe auf der Links-Rechts-Dimension über alle Kontexte hinweg einen Effekt auf die Wahlentscheidung hat, unterscheiden sich die Stärke dieses Effekts und die Salienz des ideologischen Konflikts von links und rechts erheblich. Der Einfluss der räumlichen Nähe von Wählern und Parteialternativen nimmt ab, wenn politische Parteien und Parteiensysteme sich im stetigen Umbruch befinden (zum Beispiel in Albanien, Brasilien oder Peru). Die Wirkungen ideologischer Nähe auf der Links-Rechts-Dimension sind dagegen sehr viel ausgeprägter wenn der politische Wettbewerb hinreichend institutionalisiert, strukturiert oder auch polarisiert ist (etwa in Dänemark, Spanien oder Ungarn.) In diesem kurzen Beitrag kann ich mögliche Ursachen nur andeuten: Auf der einen Seite erzeugen die Stabilisierung und Institutionalisierung notwendige Funktionsbedingungen für die Entstehung programmatisch strukturierter Verbindungen zwischen Wählern und politischen Parteien. Zudem charakterisiert der Gegensatz von Links und Rechts den politischen Wettbewerb in verschiedenen Vergleichsstaaten unterschiedlich gut. Besonders im Kontext schroffer Polarisierung wird häufig die Dimensionalität des politischen Raums reduziert und die Struktur des politischen Wettbewerbs korrespondiert dann regelmäßig besser mit einem übergreifenden, eindimensionalen Links-Rechts-Schema.

Eine systematische Metaanalyse unterstreicht die Stabilität der empirisch bestimmten Parameter über wiederholte Wahlen innerhalb desselben Vergleichsstaats hinweg: Die Inspektion der Salienzparameter $\hat{\alpha}$ verweist auf sehr hohe Autokorrelation unter Beobachtungen aus denselben Vergleichsstaaten $\left(p\left[\hat{\alpha}\right] = 0.72\right)$. Deshalb liegt nahe, dass die Bedeutung ideologischer Nähe oder Distanz für die Wahlentscheidung auch eine systemische Variable ist, die von politischen Institutionen wie Parteiensystemen, Konfliktstrukturen oder institutionellen Regeln mindestens mitbestimmt wird.

Die generelle Aussagekraft des räumlichen Modells ist eine notwendige Bedingung für das Kernargument dieses Beitrags. Das grundlegende Interesse zielt ja darauf, die Form der Nutzenfunktionen innerhalb verschiedener Elektorate empirisch zu ermitteln und diese Befunde systematisch zu vergleichen, um so Aufschlüsse über die (durchschnittliche) Risikoorientierung der Wähler zu gewinnen. Schätzer für den Formparameter $\hat{\beta}$ sind in der vertikalen Dimension von Abb. 2 abgetragen. Diese Befunde, jeweils über neunzig einzelne, unabhängig voneinander geschätzte „Conditional Logit"-Modelle hinweg, illustrieren eindringlich, dass in der überwiegenden Mehrzahl der untersuchten Wahlen die Wähler eine risikoneutrale Position einnehmen ($\hat{\beta} \sim 1$). Nicht in einem einzigen Kontext dagegen entsprach der empirisch bestimmte Formparameter der strikten Annahme risikoaverser Wähler im „Neo-Downsianischen" Standardmodell, die bei der Verwendung quadratischer Verlustfunktionen unterstellt wird ($\hat{\beta} = 2$). Dennoch sind bei einer Reihe von Parlamentswahlen, zum Beispiel in Japan, Südafrika oder in der Ukraine, die Elektorate tendenziell risikoavers. In anderen Kontexten, zum Beispiel in Bulgarien, Rumänien, oder Ungarn, verläuft die Nutzenfunktion des Elektorats dagegen mindestens teilweise konvex und die Elektorate sind somit tendenziell risikoaffin.

Wie zuvor verweist eine Metaanalyse auf substanzielle Autokorrelation ($p\left[\hat{\beta}\right] = 0.48$) und unterstreicht deshalb auf die Stabilität dieser Befunde bei wiederholten Wahlen innerhalb derselben Vergleichsstaaten. Damit liegt nahe, dass auch die Risikoorientierung der Elektorate durch langfristige, stabile institutionelle Rahmenbedingungen mindestens mitgeprägt wird.

Diese Befunde hinterfragen einige wesentliche Prämissen, die der Spezifikation der räumlichen Theorie des Wählens regelmäßig zugrunde liegen. Über ein weites Spektrum unterschiedlicher Vergleichsstaaten und heterogener Kontexte hinweg sind Elektorate und Wähler nicht generell risikoavers, sondern die Mehrheit der Wähler verhält sich risikoneutral. Die Annahme konsistent risikoaverser Wähler, die direkt durch die Benutzung quadratischer Verlustfunktionen abgebildet wird, ist deshalb viel zu restriktiv, und sie wird durch die umfangreichen Daten, auf denen diese Analysen basieren, nirgendwo gedeckt.

Am Schluss dieses Abschnitts möchte ich auf einige potenzielle Probleme dieses Modells verweisen. Diese Bedenken betreffen besonders die mögliche Modellabhängigkeit der Resultate. Ganz unabhängig davon, ob bestimmte Formen der Nutzenfunktionen à priori definiert (wie bei Grynaviski and Corrigan, 2006 oder Singh 2014) oder empirisch geschätzt werden (wie in diesem Beitrag) beeinflusst die definierte oder geschätzte numerische Größe des Formparameters das Niveau des Salienzparameters. Die theoretische Spezifizierung der Nutzenfunktionen und die einfache Inspektion der empirischen Modellparameter in Abb. 2 verweisen nachdrücklich darauf, dass $\hat{\alpha}$ und $\hat{\beta}$ invers miteinander verbunden sind. Hohe Schätzer für die Salienz der ideologischen Dimension $\hat{\alpha}$ korrelieren recht deutlich mit geringeren Werten für den Formparameter $\hat{\beta}$.

Freilich sollte auf der Grundlage der hier benutzen Daten und mit den Bedenken über die mögliche Modellabhängigkeit dieser Befunde daraus nicht gefolgert werden, dass hohe Salienz und risikoaffine Wählerdispositionen einander kausal und theoretisch bedingen.

4.3 Berücksichtigung und Modellierung nicht-räumlicher Kontrollvariablen

Nur wenige empirische Analysen begründen die Wahlentscheidung *allein* mit der Bewertung ideologischer oder programmatischer Passfähigkeit eines Wählers mit den verschiedenen Parteialternativen. Stattdessen schlagen die meisten neueren Beiträge „vereinigte" Modelle vor, die simultan räumliche, nicht-räumliche und valenzbasierte Motive der Wahlentscheidung aufnehmen (Adams et al. 2005; Jessee 2012; Merrill und Grofman 1999; Schofield und Sened 2006). Das umfassende Datenmaterial des CSES-Projekts enthält eine Reihe von nicht-räumlichen Variablen, die die Wahlentscheidung mitbeeinflussen *und* die gleichermaßen mit der Nähe oder Distanz im räumlichen Modell verbunden sind, zum Beispiel die Art und Stärke der Parteiidentifikation, die Zufriedenheit mit dem demokratischen Prozess, den Grad an politische Information eines Befragten und grundlegende demografische Kategorien wie Alter, Geschlecht und Bildungsniveau.

Dieser Beitrag interessiert sich weniger für die direkten Wirkungen dieser Variablen auf das Wahlverhalten, sondern bestimmt, ob die Aufnahme dieser Kontrollvariablen die Befunde einfacher räumlicher Modelle erhält oder modifiziert. Die Prädiktormatrix $\boldsymbol{D_i}$ bildet die Effekte einer Reihe von nicht-räumlichen Einflüssen ab:

1. Die Parteiidentifikation, bestimmt als die Angabe eines langfristigen Näheverhältnisses zu einer konkreten politischen Partei, ist per se ein individualspezifisches Charakteristikum. Für die Aufnahme in das statistische Modell wird diese Variable in eine Reihe binärer Dummy-Variablen heruntergebrochen, die anzeigen, ob sich ein Wähler i mit einer bestimmten Parteialternative j identifiziert oder nicht.
2. Als eine weitere nicht-räumliche Kontrollvariable geht die Zufriedenheit mit dem politischen Prozess in die Modelle ein. Dieser Indikator wird auf einer einfachen Vierpunktskala erhoben und als ein Proxy für die Zufriedenheit mit der amtierenden Regierung verwendet. Konkrete Bewertungen der

Regierungsarbeit werden leider nur in der zweiten und dritten Welle der CSES-Umfragen erhoben.
3. Jede Analyse von Wahlentscheidung, die Risiko und Unsicherheit angemessen berücksichtigen möchte, benötigt Daten zur politischen Informiertheit einzelner Wähler. Die CSES-Questionnaires enthalten drei binäre Wissensfragen, die zu einer Skala politischer Informiertheit aufsummiert werden.
4. Schließlich gehen als demografische Kontrollvariablen das Alter, das Geschlecht und das Bildungsniveau der Befragten in die jeweiligen Modelle ein.

Abb. 3 vergleicht die Posterioriverteilungen für den Salienzparameter ($\hat{\alpha}$; linke Bildseite) und den Formparameter ($\hat{\beta}$; rechte Bildseite). Auf der x-Achse sind jeweils die Schätzer für einfache räumliche Modelle abgetragen, und die y-Achse kontrastiert die Parameter für reichhaltigere, vereinigte Modelle mit den spezifizierten Kontrollvariablen. Bereits eine oberflächliche Inspektion beider Modelle zeigt deutlich, dass die Hinzunahme nicht-räumlicher Kontrollvariablen die Parameter des Modells kaum modifiziert. Die zusätzliche Berücksichtigung signifikanter Einflüsse wie Parteiidentifikation, Demokratiezufriedenheit, politische Information und demografischen Basisvariablen verringert oder erhöht nicht die generelle Aussagekraft der räumlichen Modellkomponente $\left(\hat{\alpha}\right)$, sie hat keinen Einfluss auf die Entstehung und Veränderung konkaver oder konvexer Nutzenfunktionen und sie betrifft nicht das hiervon abgeleitete Risikoverhalten der Wähler $\left(\hat{\beta}\right)$.

4.4 Berücksichtigung und Modellierung unsicherer Parteipositionen

Ich komplettiere nun das statistische Modell und füge die Unsicherheit von Wählern über die politischen Standorte der politischen Parteien als messbare und gemessene Variable hinzu. Dabei sind valide Bestimmungen individueller Unsicherheit mit den durch weitgehend standardisierte Wahlforschungsprojekte wie CSES bereitgestellten Daten oft konzeptionell schwierig und innerhalb der Wahlforschung umstritten (vgl. die detaillierte Diskussion und Bewertung alternativer Unsicherheitsmaße bei Alvarez 1998).

Bei der Operationalisierung von individueller Unsicherheit über die Positionen einzelner Parteien dominieren im Wesentlichen zwei, teils verschiedene, teils komplementäre Perspektiven: Bartels (1986) argumentiert, dass im Grunde alle Wähler unsicher seien über die Positionen aller Parteien, dass sie sich jedoch teils

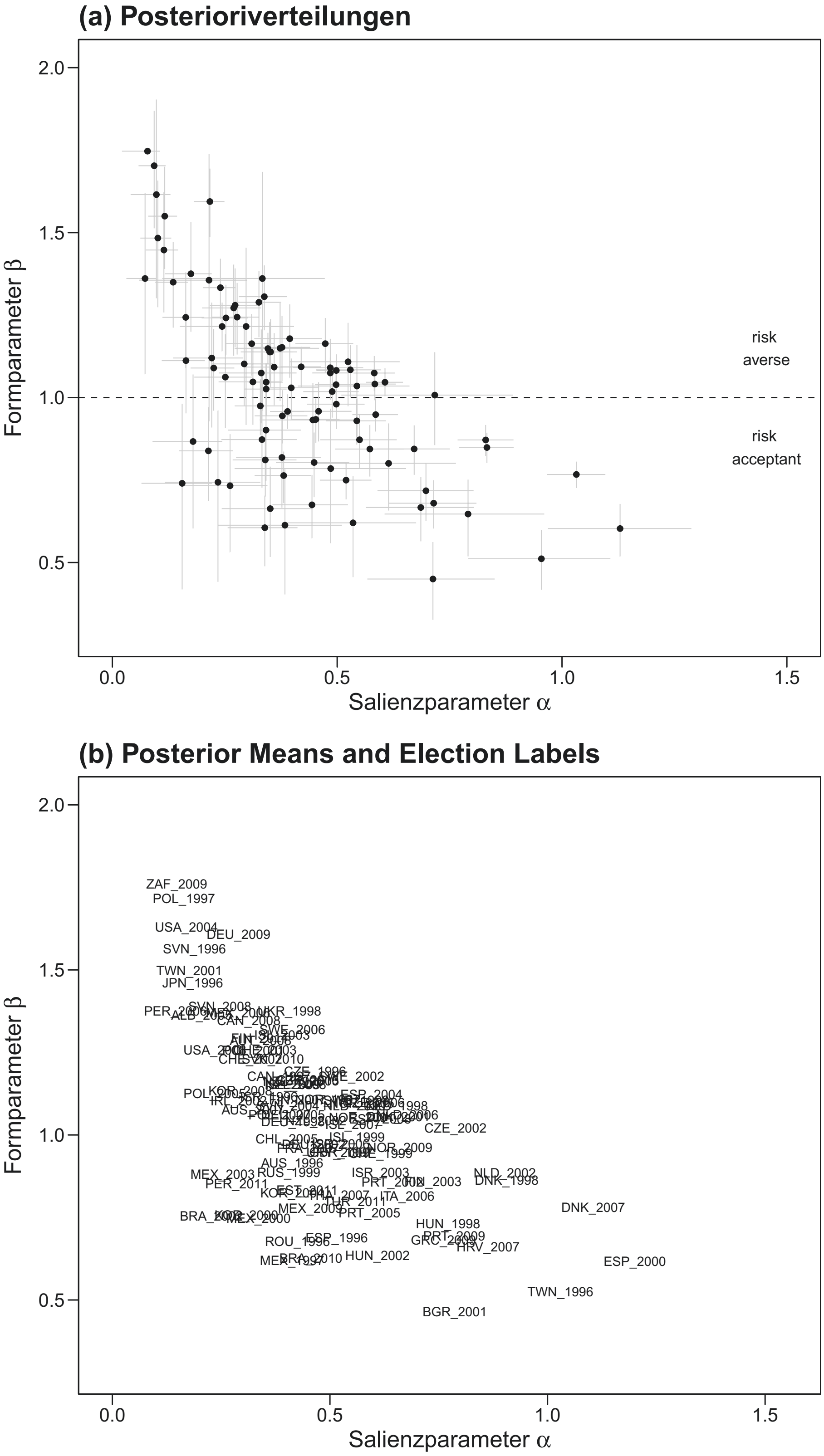

Abb. 3 Vergleich von einfachen und gesättigten Modellen $\left(\hat{\alpha} \text{ and } \hat{\beta}\right)$

erheblich im Grad dieser Unsicherheit unterscheiden. Wenn jedoch die Unsicherheit eines Wählers i über die Platzierung einer Partei j ein bestimmtes Niveau überschreite, werde dieser Befragte eine Verortung dieser Partei auf einer ideologischen oder politischen Skala ablehnen. Unsicherheit kann in diesem Sinne nicht direkt beobachtet werden, sondern Muster an Anwortverweigerungen werden als Maßstab für die Unsicherheit der Wähler benutzt. Im zweiten Schritt formuliert Bartels (1986) ein statistisches Modell, das fehlende Antworten durch eine Reihe von Erklärungsfaktoren abbildet und zum Beispiel Bildung, politische Informiertheit, politisches Interesse und die Verarbeitung von Kampagneninformationen und Parteiwerbung berücksichtigt. Die mit den empirischen Modellparametern vorhergesagte Wahrscheinlichkeit einer Antwortverweigerung wird dann als Maßstab und Instrument von Sicherheit oder Unsicherheit über die jeweilige Parteiposition bewertet.

Anstatt dieser inferenzbasierten Modellierungsstrategien unterstreicht Alvarez (1998) die Vorzüge direkt beobachteter Indikatoren. Er unterstellt, dass jede Platzierung einer Partei durch einen Wähler $p_{i,j}$ aus ihrem realen, „wahren" Standort p_j und aus einem wähler- und parteispezifischen Fehlerterm $\delta_{i,j}$ zusammengesetzt sei: $p_{i,j} = p_j + \delta_{i,j}$. Die Unsicherheit eines Wählers i bei der Verortung von Partei j wird deshalb durch die Differenz individualspezifischer und „wahrer" oder durchschnittlicher Parteipositionen bestimmt: $\delta_{i,j} = \left|p_{i,j} - p_j\right|$. Falls jedoch ein Befragter für eine bestimmte Partei keine Position angibt, definiere ich für den Indikator die maximal mögliche Distanz, also bei einer Elfpunkteskala gilt $\delta_{i,j} = 11$. Bei der Messung politischer Unsicherheit folge ich diesem Vorschlag von Alvarez (1998), weil die Qualität des von Bartels (1986) vorgeschlagenen Instruments in empirischen Applikationen oft nicht sehr hoch ist, sodass ich direkte gegenüber inferentiellen Indikatoren politischer Unsicherheit vorziehe.

Der Unsicherheitsindikator $\delta_{i,j}$ nimmt verschiedene Werte ein, die für jeden Wähler i und jede Partei j separat bestimmt werden. Konkrete Werte von $\delta_{i,j}$ reichen von 0, bei $p_{i,j} = \overline{p}_j$, bis hin zu 11, wenn $p_{i,j}$ und $\overline{p}_j$ an den gegenüber liegenden Extrempositionen der Skala positioniert sind oder wenn der Wähler i keine Platzierung von Partei j vornehmen kann oder möchte. Durch diese Regeln zum Umgang mit fehlenden Platzierungen sind die Werte von $\delta_{i,j}$ stets bimodal verteilt. Spezifische Werte des Indikators $\delta_{i,j}$ schwanken recht deutlich, und zwar sowohl innerhalb einzelner Segmente des CSES-Datensatzes und über diese Segmente hinweg. Unter den neunzig Umfragesegmenten in der Analyse waren sich zum Beispiel die Befragten in Tschechien 2002 am sichersten über die Standorte der Parteien $\left(\overline{\delta}_{i,j} = 1{,}26\right)$, in Italien 2006 waren sie dagegen besonders unsicher $\left(\overline{\delta}_{i,j} = 4{,}97\right)$.

Das statistische Modell korrespondiert genau mit dem theoretischen Konzept von Unsicherheit, das eingangs dargestellt wurde. Es gründet im Wesentlichen auf einer Modellbildungsstrategie, die Bartels (1986) entworfen, Berinsky and Lewis (2007) durch die explizite Modellierung verschiedener Risikoorientierungen weiterentwickelt und shikano and Behnke (2009) repliziert und für Mehrparteiensysteme verallgemeinert haben. Wie zuvor wird die Auswahl unter mehr als zwei Parteialternativen mit einem bayesianischen „Conditional Logit"-Modell abgebildet. Anstatt jedoch jede einzelne Partei durch einfache Punktschätzer im politischen Raum zu verorten, charakterisiere ich die Wahrnehmung unsicherer Parteien durch eine Wahrscheinlichkeitsverteilung über ihrer jeweiligen Position. Dabei wird bei jeder Iteration des MCMC-Prozesses ein neuer Wert für die Parteiposition aus dieser Verteilung gezogen und für die Schätzung der Modellparameter verwendet.

Im nächsten Schritt verbinde ich die beiden eben vorgestellten Ideen, die direkte Messung der Wählerunsicherheit und die Charakterisierung als unsicher wahrgenommener Parteipositionen durch Wahrscheinlichkeitsverteilungen. Die Grundannahme lautet, dass beide Größen proportional zueinander sind: Je höher die Unsicherheit eines Wählers i über die Position von Partei j, desto größer ist die Standardabweichung einer Normalverteilung über dieser Position: $\sigma\left(p_{i,j}\right) \propto \delta_{i,j}$. Bartels (1986, 717) unterstellt dabei, der Unsicherheitsindikator sei entworfen „to reflect variances of candidate perceptions up to an (unknown) positive scale factor". Genau deshalb müssen die empirisch bestimmten Werte von $\delta_{i,j}$ durch einen weiteren Modellparameter γ reskaliert werden, und das Produkt $\gamma\sigma\left[p_{i,j}\right]$ wird so als die tatsächliche empirische Variation der Parteipositionen interpretierbar:

$$\Theta_{i,j} = N\left(p_{i,j}, \sigma\left[p_{i,j}\right]\right) = N\left(p_{i,j}, \gamma\delta_{i,j}\right) \tag{5}$$

Schließlich ersetze ich die fixierten Parteipositionen $p_{i,j}$ durch die Wahrscheinlichkeitsverteilungen $\Theta_{i,j}$:

$$u\left(v, \Theta_{i,j}\right) = -\alpha\left|v_i - \Theta_{i,j}\right|^{\beta} + \lambda_j c_j + \varepsilon_{i,j};\, Pr(v_i = j) = \frac{exp\left[u\left(v, \Theta_{i,j}\right)\right]}{\sum_{j=1}^{J} exp\left[u\left(v, \Theta_{i,j}\right)\right]} \tag{6}$$

Bei diesem komplexeren Modell ist die technische Bestimmung der Modellparameter durch MCMC-Simulationen nicht einfach, weil die einzelnen Markow-Ketten häufig eine hohe Autokorrelation aufweisen und bei der Parameterschätzung nur langsam mischen. Der MCMC-Prozess simuliert nicht nur die einzelnen Modellparameter, sondern auch die Parteipositionen müssen jeweils aus Unsicherheitsverteilung $\Theta_{i,j} = N\left(p_{i,j}, \gamma\delta_{i,j}\right)$ gezogen werden. Deshalb habe ich für die

Simulation dieser komplexeren Modelle in jedem der neunzig Wahlsegmente vier Ketten und ein „burn in“ von 100.000 Iterationen vorgesehen und weitere 200.000 Iterationen für die Charakterisierung der Posterioriverteilungen gespeichert. (Ein Ausschnitt des für die MCMC-Simulationen benutzten `JAGS`-Codes ist in Anhang B dokumentiert.).

Abb. 4 fasst die wesentlichen Modellbefunde zusammen und charakterisiert die Posterioriverteilungen für die Salienz- ($\hat{\alpha}$), Form- ($\hat{\beta}$) und Unsicherheitsparameter ($\hat{\gamma}$). Die linke Spalte (a) charakterisiert zunächst den Salienzparameter $\hat{\alpha}$. Die geschätzten Parameter sind über neunzig heterogene Wahlumgebungen hinweg stets positiv, und diese Befunde belegen, dass, auch wenn die Unsicherheit von Parteiplatzierungen explizit modelliert wird, die räumliche Nähe auf der ideologischen Links-Rechts-Dimension weiterhin einen erheblichen Einfluss auf die Bewertung politischer Parteien ausübt und die Wahlentscheidung mitbestimmt. Über die einzelnen Kontexte hinweg variiert dieser Einfluss jedoch oft erheblich.

Diese Resultate bestärken einerseits die bereits vorgestellten und diskutierten Befunde einfacher räumlicher Modelle. Während die Verteilungen der Parameter $\hat{\alpha}$ und $\hat{\beta}$ über die einzelnen Wahlsegmente hinweg beinah unverändert bleiben, modifiziert die systematische Berücksichtigung unsicherer Parteipositionen die Befunde für einzelne Umfragesegmente des CSES-Datensatzes oft deutlich. Eine Konsequenz ist, dass die Modellabhängigkeit der Parameter im einfachen räumlichen Modell ($-\alpha \sim \beta$) in den komplexeren Modellen nicht reproduziert wird. Benutzt man die Wählerunsicherheit $\delta_{i,j}$, um die Verteilungen unsicher wahrgenommener Parteipositionen $\Theta_{i,j}$ zu charakterisieren, sind die Salienz- und Formparameter empirisch unabhängig voneinander.

Die mittlere Spalte (b) in Abb. 4 fasst die Posterioriverteilungen des Form- oder Risikoparameters $\hat{\beta}$ zusammen. Die systematische Berücksichtigung der Wählerunsicherheit ändert zunächst nur wenig an seiner empirische bestimmten Bandbreite über die heterogenen Wahlkontexte hinweg. Die Mittelwerte der Posterioriverteilungen liegen meist recht eng bei $\hat{\beta} = 1$, unterstreichen die empirische Angemessenheit linearer Distanzen und zeigen eine klare Tendenz hin zu risikoneutralen Haltungen der meisten Elektorate an. Dagegen werden einige Standardannahmen des „Neo-Downsianischen“ Modells, die unbedingte Bevorzugung quadratische, konkaver Nutzenfunktionen und die Unterstellung einer strikt risikoaversen Perspektive der Wähler und Elektorate, durch empirische Modellbefunde weiterhin beinah nirgendwo gestützt.

Schließlich stellt die rechte Spalte (c) in Abb. 4 die Posterioriverteilungen des Unsicherheitsparameters γ dar. Stets positive Schätzer und beinah durchgehend

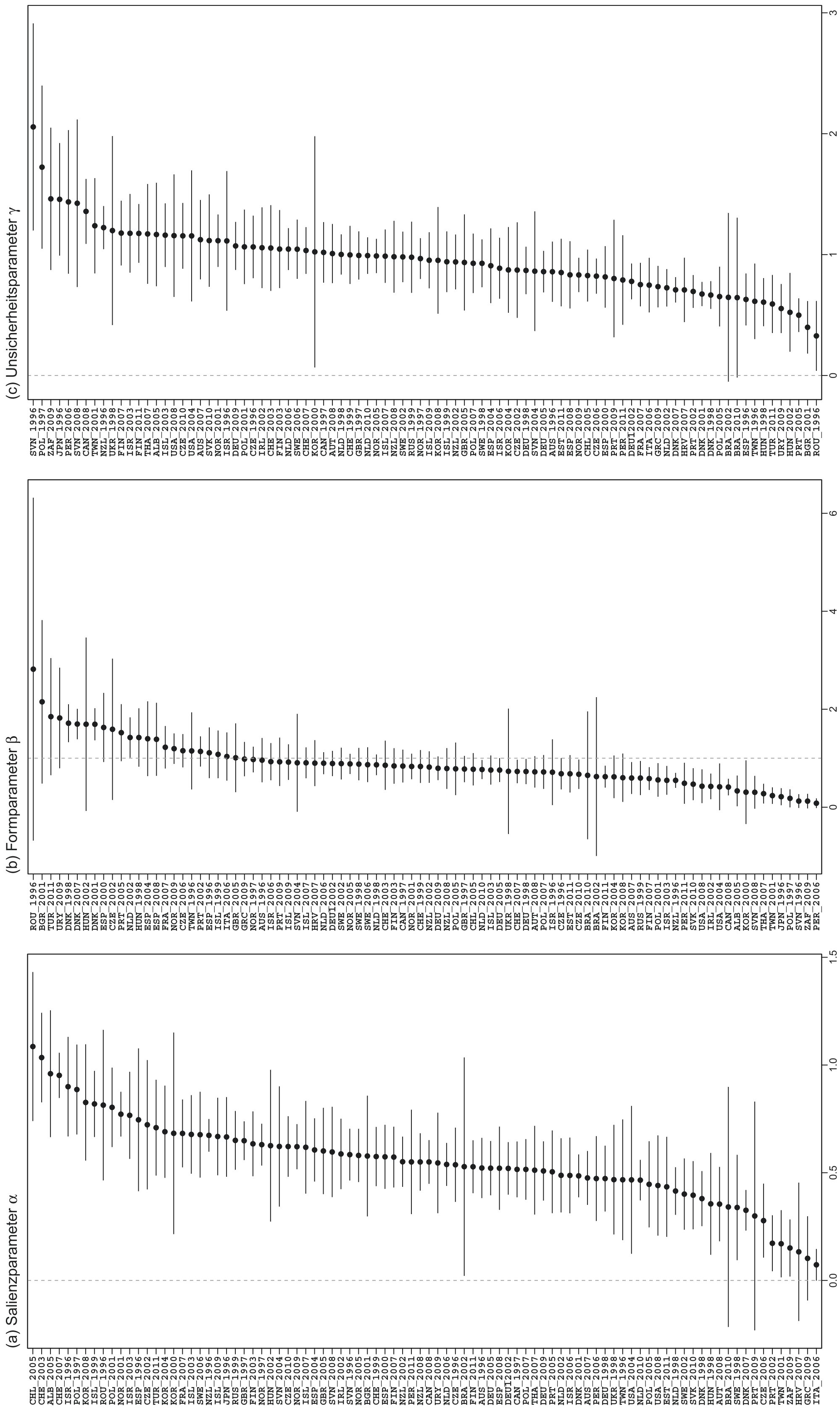

Abb. 4 Posterioriverteilungen der Salienz-, Form- und Unsicherheitsparameter $\left(\hat{\alpha},\ \hat{\beta},\ \text{and}\ \hat{\gamma}\right)$

positive Konfidenzintervalle des Unsicherheitsparameters verweisen auf den Einfluss unsicher wahrgenommener Parteipositionen auf die Wahlentscheidung. Höhere Parameterschätzer für $\hat{\gamma}$ korrespondieren mit weiteren Unsicherheitsverteilungen über der jeweiligen Parteiposition $p_{i,j}$ und verweisen auf die gesteigerte Bedeutung von Unsicherheit für die Bewertung von Parteialternativen und die Wahlentscheidung.

5 Fazit und Ausblick

Theoretische und empirische Beiträge zur räumlichen Theorie des Wählers haben oft einfach unterstellt, wie Wähler mit Risiko und Unsicherheit umgehen. Dieser Beitrag geht dagegen einen anderen Weg und bestimmt die Salienz räumlicher Nutzenterme und die Logik, mit der Wähler räumliche Distanzen in Nutzenterme verwandeln, in einer empirisch-vergleichenden Analyse über eine Fülle heterogener Kontexte hinweg. Diese Vogelperspektive auf die Risikoorientierung gesamter Elektorate erschließt mindestens zwei wichtige Einsichten:

1. Die standardisierte Annahme konkaver, quadratischer Nutzenfunktionen ist für die valide Modellspezifikation zu restriktiv. Die empirischen Resultate für die meisten der heterogenen Kontexte des CSES-Trendfiles widersprechen dieser Annahme ausdrücklich und verweisen stattdessen tendenziell auf die Angemessenheit und Nützlichkeit linearer Verlustfunktionen für die meisten Entscheidungssituationen.
2. Weil die Form der Verlustfunktionen bestimmt, wie Wähler geometrische Distanzen in die Nutzenterme des räumlichen Modells verwandeln, betrifft sie direkt die Risikoorientierung der Wähler. Die empirischen Resultate widersprechen der klassischen Annahme strikt konkaver und risikoaverser Präferenzen. Über heterogene Kontexte hinweg verfolgen Wähler stattdessen tendenziell risikoneutrale Haltungen, und in einer Minderheit der Wahlsegmente handeln Wähler sogar risikoaffin.

Die inhaltliche Bewertung dieser Befunde sollte stets mit großer Vorsicht erfolgen, und sie wird deshalb mit einigen Qualifikationen vorgenommen. Zunächst, das wurde bereits diskutiert, erschließen die hier vorgestellten Modelle die Risikoorientierung ganzer *Elektorate,* nicht einzelner Wähler. Nun kann aber nicht ohne Weiteres unterstellt werden, dass alle Befragten oder gar alle Wähler bei allen Parteien, die in einem Land zur Wahl stehen, dieselben Risikobewertungen anlegen. Diese Verlustfunktionen spezifisch für Gruppen von Wählern und für einzelne Parteien zu bestimmen, bleibt deshalb ein wichtiges Desiderat der Forschung.

Nicht nur die fehlende Möglichkeit, konkrete Risikoorientierungen einzelner Wähler zu bestimmen, zu modellieren und für empirische Analysen aufzugreifen, spricht für die mindestens komplementäre Hinzunahme experimenteller Studien. Die Verwendung einfacher Umfragedaten, auch wenn das hier aus einer empirisch gesättigten und strikt vergleichenden Perspektive geschieht, kann nur Aufschlüsse über empirische Assoziationen von Unsicherheit und Risikoorientierung bieten. In welche Richtung die Effekte genau weisen, was Eplanans und was Explanandum ist, bleibt jedoch nur teilweise bestimmt. Formale Modelle definieren häufig die Risikoorientierung der Wähler als eine strikt exogene Größe und postulieren, dass Wählerverhalten und Parteistrategie einfach darauf reagieren. Vice versa kann jedoch auch argumentiert werden, dass politische Unsicherheit, Informationsdefizite, widersprüchliche Informationen am Ausgangspunkt der Kausalketten stehen und ihrerseits Risikokalkulationen rationaler und nichtrationaler Wähler erst begründen.

Zudem ist nicht nur in der formalen Literatur weiterhin umstritten, ob die hier vorgestellten Befunde umstandslos als stringente Belege für risikoneutrales oder sogar risikoaffines Verhalten bewertet werden können. Bei der formalen Modellspezifikation verweist zum Beispiel Eguia (2013) eindringlich auf Probleme euklidscher Nutzenfunktionen in mehrdimensionalen politischen Räumen, die Gewichtung unterschiedlicher Wettbewerbsdimensionen und die analytische Separierbarkeit dieser Nutzenkomponenten. Schließlich wenden Shikano und Behnke (2009) ein, dass selbst belegbar konvexe Verlustfunktionen nicht unbedingt und ausschließlich auf auf risikoaffines Verhalten verweisen. Stattdessen könnten konvexe Verlustfunktionen auch für die mangelnde Bereitschaft sprechen, politische Positionen angemessen zu differenzieren und Güterabwägungen vorzunehmen. Sie wären damit eher ein empirischer Hinweis auf die Rigidität der Wähler als auf ihre Risikobereitschaft.

Schließlich kann der Effekt von Unsicherheit und Risikoorientierung in räumlichen Modellen nicht mit der exklusiven Perspektive auf einzelne Wähler und gesamte Elektorate untersucht werden. Eine wirklich komplette Analyse müsste auch die „Angebotsseite“ politischer Parteien angemessen berücksichtigen und bestimmen, wann und warum sie im politischen Wettbewerb unklare ideologische und/oder programmatische Positionen anbieten.

Danksagung Ich danke Susumu Shikano für Rat und Hilfe bei der Spezifikation der statistischen Modelle, und ich danke den beiden anonymen Gutachtern für die ihre sehr wertvollen Kommentare und Vorschläge. Der Beitrag entstand im Rahmen des von der Fritz-Thyssen-Stiftung geförderten PRojekts „Lost in Space? The Emptiness of the Center and Centrifugal Determinants of Vote Choice and Party Competition in EP Elections“ (Az. 10.17.1.039PO).

Appendix

Das Datenmaterial des CSES-Projekts

Umfragesegmente des kumulierten „CSES-Trendfile“

Die empirischen Analysen in diesem Beitrag stützen sich ausschließlich auf das reichhaltige Datenmaterial des CSES-Projekts. Ich verwende den „CSES Harmonized Trend File“, der länder- und wahlspezifischen Umfragesegmente von 1996 bis 2011 standardisiert und für die empirische Datenanalyse aufbereiten. Dieser integrierte Datensatz wurde am Wissenschaftszentrum Berlin von Heiko Giebler, Josephine Lichteblau, Antonia May, Reinhold Melcher, Aiko Wagner und Bernhard Weßels zusammengestellt. Daten und Dokumentation sind auf den Webseiten des Projekts verfügbar (http://cses.org/datacenter/trendfile/trendfile.htm; Stand: 15. Dezember 2015).

Diese Aufbereitung ersten drei Wellen des CSES-Projekts integriert und harmonisiert diejenigen Variablen, die wiederholt, also in mindestens zwei Wellen, abgefragt wurden. Insgesamt vereinigt der CSES-Trendfile Datenmaterial aus 129 einzelnen Umfragesegmenten, die im Kontext von Wahlen zu nationalen Parlamenten und zur nationalen Präsidentschaft abgehalten wurden. Um die Vergleichbarkeit der untersuchten Wahlbefragungen zu gewährleisten, habe ich allein Wahlen zu nationalen Parlamenten (bei bikameralen Systemen zu den jeweiligen Unterhäusern) ausgewählt, die stärker personalisierten Präsidentschaftswahlen jedoch ausgeschlossen. Auch einige weitere Segmente konnten ich nicht berücksichtigen: Teils fehlten einige Schlüsselvariablen der Modelle, teils waren die Daten fehlerhaft oder inkonsistent, und einige Umfragesegmente waren so klein, dass keine sinnvollen Inferenzen möglich waren. Mit diesen Kriterien kann die Analyse nur neunzig der ursprünglich 128 Wahlsegmente aus insgesamt 44 verschiedenen Staaten aufnehmen:

Albanien (2005), Australien (1996, 2007), Bulgarien (2001), Brasilien (2002, 2010), Chile (2005), Deutschland (1998, 2002, 2005, 2009), Dänemark (1998, 2001, 2008), Estland (2011), Finnland (2003, 2007, 2011), Frankreich (2007), Großbritannien (1997, 2005), Griechenland (2009), Ungarn (1998, 2002), Irland (2002), Island (1999, 2003, 2007, 2009), Israel (2003, 2006), Italien (2006), Japan (1996), Kanada (1997, 2007), Korea (2000, 2004, 2008), die Niederlande (1998, 2002, 2006, 2010), Norwegen (1997, 2001, 2005, 2009), Kroatien (2007), Neuseeland (1996, 2002, 2008), Österreich (2008), Peru (2006, 2011), Polen (1997, 2001, 2005, 2007), Portugal (2002, 2005), Rumänien (1996), Russland (1999), die Schweiz (1999, 2003, 2007), die Slowakei (2010), Slowenien (1996, 2004, 2008), Schweden (1998, 2002), Spanien (1996, 2000, 2004, 2008), Südafrika (2009), Thailand (2007), die Türkei (2011), Taiwan (1996, 2001), die Tschechische Republik (1996, 2002, 2006, 2010), die Ukraine (1998), Uruguay (2009) und die USA (2004, 2008).

Auswahl und Operationalisierung von Schlüssel- und Kontrollvariablen werden im Text knapp vorgestellt und begründet. Genauere Information zur Formulierung, Erhebung und Aufbereitung der einzelnen Indikatoren stehen kumuliert auf der Seite des CSES-Trendfiles und individuell auf den jeweiligen Seiten der drei benutzten CSES-Wellen bereit (http://cses.org/). Komplette Replikationsarchive stelle ich auf Anfrage gern zur Verfügung.

JAGS Code für die Schätzung des Unsicherheitsmodells

```
model{
for(i in 1:N_V){
    for(j in 1:N_P){
        mu[i,j]<- beta[1,j] - alpha[1] * (abs(lr_i[i]-lr_
        ij.dist[i,j])^alpha[2])
        emu[i,j]<- exp(mu[i,j])
        p[i,j]<- emu[i,j]/sum(emu[i,1:N_P])

        prec.lr[i,j]<- pow(unc_ij[i,j] * gamma,-2)
        lr_ij.dist[i,j]~dnorm(lr_ij[i,j], prec.lr[i,j])
    }
    vote[i]~dcat(p[i,1:N_P])
    }
# PRIORS;
# ALPHA;
  alpha[1:N_ALPHA]~dmnorm(a0, A0)
# BETA;
# identifying restriction;
    for(i in 1:N_BETA){
     beta[i,1]<- 0
    }
    for(i in 2:N_P){
      beta[1:N_BETA,i]~dmnorm(b0,B0)
      }
# GAMMA;
  gamma~dunif(0,3)
}
```

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Local Districts, National Contexts, and the Number of Parties

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To date, electoral systems are conceptualised as setting an 'upper bound' to, or defining a 'carrying capacity' for, the number of parties or lists, and their effect is assessed at the district level. This article adds to the empirical study of electoral systems by analysing a vast database of district-level electoral returns. The argument focuses on the demand and supply of viable electoral candidates, which are conditioned by the interplay of strategic entry (by the party rank and file) and strategic voting (by the electorate). Drawing on a database of almost 18,000 electoral districts taken from 15 West European countries, the empirical analysis yields a number of insights: most specifically, (1) district magnitude only becomes binding and effective when a higher social demand meets a lower carrying capacity of the electoral district; (2) the provision of upper tiers undermines the emergence of Duvergerian equilibria within the primary electoral districts.

This article provides a comparative analysis of the political consequences of electoral systems at the district level. It concentrates on the effects of (local) district magnitudes on the fragmentation of (local) electorates and explicitly models the context-dependency of electoral system effects. Relating electoral system features, predominantly the 'all-important factor' of district magnitude, to the number of candidates or lists has been labelled the 'core of the core' of electoral system research, and many have argued that it is time to close the book on these largely settled research questions (Shugart 2005: 31). Various studies have effectively compiled a vast body of empirical evidence and detected robust associations of district magnitude with fragmentation indicators at the national level (see Amorim Neto and Cox 1997; Brambor *et al.* 2006; Clark and Golder 2006; Filippov *et al.* 1999; Golder 2006; Lijphart 1994; Ordeshook and Shvetsova 1994; Taagepera and Shugart 1989, 1993).

This contribution aims at moving the perspective to the district level and testing the robust evidence gathered by national-level analyses with a rich and context-sensitive dataset that includes almost 18,000 electoral districts from 17

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West European countries. Although the principal findings of this article actually reinforce many results previously demonstrated with national-level electoral data, it is far too early to close the book on these important matters. The focus on the district level addresses the arena where the causal forces behind Duverger's law actually operate and exploits rich datasets that have only recently been made available for empirical analysis. The key analytical benefit is to explore contextual variation in the empirical links of district magnitude with the number of viable electoral contenders. This cogently relates interactions of the demand for (as given by the number of societal groups seeking representation) and the supply of political parties (as determined by the permissiveness of the electoral system). While there are currently only a few studies which introduce district-level electoral returns to a full-fledged comparative analysis (see Chhibber and Kollman 1998, 2004; Singer 2013; Singer and Stephenson, 2009), I am not aware of any previous attempt to systematically model contextual heterogeneity and thereby explore the consequences of district magnitude across various elections and countries.

Theoretically, this article builds upon two different theoretical strands which their respective authors have labelled 'generalisations' of Duverger's law: The 'generalized Duverger's law' by Taagepera and Shugart (1989, 1993) links district magnitude and the number of viable competitors by unconditional comparative statics, specific magnitudes produce a specific number of candidates or lists plus an error interval. The more recent 'direct generalization of Duverger's law' by Cox (1994, 1997) instead proposes conditional comparative statics: district magnitude merely provides a specific carrying capacity for candidates or lists and whether local district magnitudes may become binding and effective or not actually depends on the societal demands by groups seeking representation. The empirical results underscore both the significance of electoral systems for the structure of (local) electorates and their profound context-dependency: (1) district magnitude only becomes binding when a higher social demand meets a lower carrying capacity of the electoral district; (2) the magnitude of primary electoral districts only tends to be binding when the electoral structure provides no upper tiers or allocates only a very limited share of seats in the upper tiers.

The Two 'Laws' of Electoral Systems

This paper explores the empirical links of district magnitude and the number of electoral contenders both within and across heterogeneous electoral and national contexts. So as to clarify these pivotal issues of electoral systems research, this section builds upon and discusses two different generalisations of Duverger's law: the 'generalized Duverger's law' derived by Rein Taagepera and Shugart (1993); and the 'direct generalization of Duverger's law', the $M + 1$ rule proposed by Gary W. Cox (1994, 1997). The first one refers to an inductive generalisation at the macro level, and the second one spells out a theoretically valid model at the district level.

Comparative Statics of District Magnitude

Taagepera and Shugart (1989, 1993) aimed at unifying 'Duverger's law' and 'Duverger's Hypothesis'. Searching for a general expression that ties together both propositions, the authors turn towards an inductive inspection of empirical electoral returns. They reason that their measure of electoral fragmentation, being the effective number of electoral parties, is linked to district magnitude by something like

$$E[N|M] = 1.25 \cdot \lg M + 2.5$$

with N_v denoting the effective number of electoral parties and M denoting the average district magnitude at the national level. The equation allows for the prediction of fragmentation levels that will, as Taagepera and Shugart (1989) argue, usually fall within an interval of +/–1 unit from the proposed comparative statics. Note that the (common) logs capture the diminishing marginal effects of increasing district magnitudes.

Linking the effective number of electoral parties with logged district magnitudes has been one of the cornerstones of the empirical literature. Although diverse cases have been included, different decisions have been taken to measure key concepts and to include controls for any seminal analyses regarding the political consequences of electoral systems started out by linearly relating some party count with some logged or otherwise transformed indicator of district magnitude (Lijphart 1994; Taagepera and Shugart 1989, 1993). Subsequent studies regularly added interactions of institutional and societal structures and allege that more proportional electoral systems are empirically associated if (and only if) there is also a high level of societal diversity. This line of thinking is clearly more in line with the original emphasis made by Duverger (see Amorim Neto and Cox 1997; Brambor *et al.* 2006; Clark and Golder 2006; Filippov *et al.* 1999; Golder 2006; Ordeshook and Shvetsova 1994).

While this tradition has contributed to a cumulative research programme, there are still a number of ambiguities and conceptual problems. First, any linkage between the national number of parties and the national average features of the respective electoral system will be a rough, inductive generalisation. Secondly, Taagepera and Shugart (1989) also stick with the problematic idea of a multiplying effect of electoral rules: The more permissive an electoral system (i.e. the higher the district magnitude), the more viable parties will be in the electorate. This notion is established as a quasi-deterministic argument which does not take into account intervening factors such as social heterogeneity or the collective action problems partisans face when setting up their platforms (as spelled out in detail by Olson 1965).

Establishing an Upper Bound

For Duverger, the primary dependent variable is the number of political parties at the national level, but he was ready to assert, 'the true effect of the simple-majority [plurality] system is limited to local bi-partism ... the creation of a two-party system inside the individual constituency' (Duverger 1954: 223). Gary W. Cox has picked up the earlier arguments. His work utilises and combines both classic statements of formal theory and the inductive tradition of data analysis in comparative political science. He was able to strengthen the causal argument, broaden its applicability, and attach conditions to its functionality (Cox 1994, 1997, 1999).

Formal modelling suggests that in each single-member plurality district (M = 1) there will be only two vote-getting candidates in a game-theoretic equilibrium. Subsequently, Cox (1997: 99) extends his arguments to multi-member districts (M > 1). As a 'direct generalization of Duverger's law', he suggests that there may be no more than M + 1 viable candidates in each district of the magnitude M. This proposition, labelled the 'M + 1 rule', is the central building block of Cox's contribution to the analysis of electoral systems. The formal statement does not imply any unconditional comparative statics of electoral systems, but in contrast conceptualises institutional constraints as erecting an upper limit or a carrying capacity for the number of viable candidates or lists in a game-theoretic equilibrium.

Cox also provides a thorough discussion of the conditions and the limits of the M + 1 rule. Focusing on the analysis of real-world data, he posits that strategic voting seems 'to fade out rapidly for district magnitudes above 5' (Cox 1997: 100, 141). He continues to argue that larger magnitudes do not tend to be binding, because these districts imply much smaller and effectively almost indistinguishable vote gaps, which separate winners from losers and thus undermine strategic behaviour by voters and parties. These claims are supported by district-level returns from Columbian, Japanese, and Spanish elections. The notion that strategic entry and strategic voting happen to be ineffective in medium-sized and large proportional representation (PR) districts has indeed been contested. Empirical contributions have referred to evidence on strategic behaviour that also applies to larger districts. For instance, Gschwend (2007) and Gschwend and Stoiber (2012) urge for a closer look at the expectation formation processes and demonstrate that voters may effectively also react strategically when incentives are allegedly much weaker. Their analyses of district-level data from Portuguese and Finnish elections show that voters even continue to strategically desert trailing parties when district magnitudes are at about M = 20.

A Context-Sensitive Model of Electoral Systems

This contribution follows the trails of Duverger (1954) and Cox (1997) in conceptualising electoral systems as an institutional means to regulate the supply of candidates and lists by setting an 'upper bound' for or a carrying capacity of viable competitors within an electoral district. The sum of these features is key to understanding how electoral systems are dependent upon and embedded within their social and political context.

Providing the Demand for Candidates or Political Parties

Electoral system theories, however, usually do not address the demand side, meaning that the actual number of societal groups or interests seeking representation is usually considered to be exogenous and left unexplained. The sociological perspective has traced back the nature and number of candidates or political parties to the type and the number of substantial cleavages in a society (Grumm 1958; Lipset and Rokkan 1967; Lipson 1964; Rokkan 1970).

The sociological approach comes in pure variants which put forward an all-out criticism of institutional analysis: 'the method of voting remains a rather small consideration among the complex and infinitely diverse factors that, combined differently in each national society ... condition political life' (Lavau 1953: 46). Authors who suggest a more actor-centred perspective have contested these notions. Most notably, Olson (1965) has sharply criticised the apparent belief that social interests will somehow automatically organise in political parties to ignore imminent collective action problems which may, among other sources, arise from the institutional context. Therefore, most scholars to date subscribe to an interactive perspective that traces the number of parties back to a societal demand of groups seeking representation and an institutional supply by the electoral system.

Party Formation from Bottom Up and from Top Down

Note that the model proposed by Cox (1997) effectively implies a bottom-up logic of electoral competition and party formation. Strategic coordination happens in two successive stages within each analytically isolated electoral district and across the various electoral districts in a country. It therefore evolves in another, second-stage coordination game when the locally viable competitors decide to link (or not to link), to coordinate (or not to coordinate) across the various districts of an electoral system so as to form a nationalised political party (or to remain a regionalised political force). Cox (1997: 181–202) specifies national policy goals, the presidency, the premiership, and the existence of upper-tier seats in the electoral system as potential factors pressing towards cross-district linkage. Chhibber and Kollman (1998, 2004) add that the regionalisation or nationalisation of the British, Canadian, Indian, and US party systems depends on the competencies assigned to hierarchically ordered levels of government.

This perspective could, however, be contested by the possibility of a reverse top-down logic when parties have been previously established at the national level and subsequently field candidates or lists in local electoral districts. Once political parties are established as organised, national-level political institutions, local lists and candidates may frequently be launched top-down by political elites who, by branding a candidate, make him/her eligible. Any organised and ambitious political party will tend to field candidates in any possible electoral district, regardless of the actual chances of winning a seat, in order to showcase itself as a serious contender for national-level political office. In an early reply to Duverger, Colin Leys has tried to redirect the focus towards the macro level and argued that strategic voting 'occurs in favor not of the two parties which are in the lead locally, but *in favor of the two parties which have the largest number of seats in Parliament, regardless of their local strength*' (Leys 1959: 149; emphasis original). Leys emphasises that voters may care more about party competition at the national than at the local level, because any vote that does not contribute to strengthening a party that could realistically enter government may be considered wasted. Furthermore, voters might also know much more about national than about local party competition and thus use the national race as a proxy to reason about potential local circumstances. If this argument can be sustained, the presence of top-down logics may seriously undermine the independence of strategic coordination in each isolated electoral district and thereby limit the applicability of the $M + 1$ rule.

The balancing of pressures from above and from below is certainly crucial for the decision-making by voters and party elites and affects the emergence of strategic voting and strategic entry. If the top-down argument can be theoretically defended and empirically sustained, the effectiveness of strategic entry may be seriously undermined and the ballots presented to the voters in small and large districts may effectively look the same. In turn, this implies that the Duvergerian logic would be exclusively (or at least predominantly) driven by strategic voting. But this also implies that voters are able to sort out strategic incentives at the local and at the national level when they cast their vote. The argument by Duverger (1954) and Cox (1997) assumes that voters are aware of local district magnitudes, the thresholds set for the representation of small parties, the presence and effects of upper tiers, and the potential outcome of the local electoral race. In contrast, the argument by Leys (1959) implies that voters potentially know a lot more about and react to the political programmes and candidates, the potential outcome of the elections, and options for coalition-building at the national level. Sure enough, voting for a locally hopeless candidate does not make any 'rational' sense at either the local or the national level, but the voters' perception might of course be a different one.

These arguments leave us with two competing perspectives on the goals and the informational resources that condition strategic coordination: The bottom-up hypothesis implies that voters and party elites are informed about and respond to local electoral standings and incentives within their respective district. In contrast, the top-down perspective presumes that voters care very little and know very little about the institutional features of and likely electoral results within each local district and instead use proxies taken from the national level.

Single-Tier and Multi-Tier Districting

The district structure also comes with strong implications for strategic coordination and determines whether district magnitudes may be empirically binding or not. According to Cox (1997: 48), any electoral district that cannot be partitioned into smaller districts in which seats are allocated are 'primary' electoral districts, while those which can be partitioned into a number of primary districts are called 'secondary' districts.

In single-tier systems, when only primary districts are present, district magnitude as specified by the $M + 1$ rule may be binding. In the model proposed by Duverger (1954) and Cox (1997), parliamentary seats can only be won and lost within each of the individual districts, and both voters and party elites need to react to the institutional incentives established by district magnitude (and other electoral system features). Whether these district magnitudes are actually binding depends on the district structure. When there are only primary electoral districts, voting for a locally hopeless candidate or list does not make any instrumental sense, district magnitude may be binding and induce the emergence of Duvergerian equilibria.

In contrast, if the primary districts are grouped into secondary districts, the impact of the $M + 1$ rule might easily be watered down since voters and political elites might support lists or candidates which are out of the running in the primary districts in order to potentially win a seat in one of the secondary, upper-tier districts. As a consequence, votes that were unsuccessful in the primary districts could still be converted to seats in the secondary electoral districts if both tiers are connected. The applicability of the $M + 1$ rule is seriously undermined when the individual districts cannot be modelled as a series of independent coordination games. In single-tier electoral systems local district magnitudes may be binding given that the aforementioned contextual conditions are also met. More complex two- or multi-tier electoral systems effectively render district magnitude less binding and potentially irrelevant for the number of viable competitors in local districts.

Testing the Consequences of District Magnitude

This section aims at deriving a statistical model that is suitable for a systematic test of the key hypotheses. In general terms, I conceptualise the emergence of a party system to be driven by demand and by supply: (1) the demand for political parties is driven by the multiplicity of societal factions seeking representation in the political system (e.g. ethnic or religious groups); (2) the supply of political parties is not only affected by these groups' organisational capabilities, but also limited by the institutional rules of the game, most significantly by the electoral system. Given that the aforementioned contextual and informational conditions which facilitate the emergence of Duvergerian equilibria at $M + 1$ are met, I explain the interaction of societal demand and institutional supply of political parties by three exemplary scenarios:

(1) In the first scenario, the societal demand for parties, as proxied by the number of societal groups seeking representation, is comparatively low or at least consistently lower than the electoral district's carrying capacity. As a consequence, district magnitude will not be binding and there should be no association between district magnitude and the number of viable parties or lists (or at least no association that is generated by electoral system).
(2) The second scenario implies that the number of candidates or lists which emerge from societal conflicts is consistently higher than the carrying capacity of $M + 1$ candidates or lists established by the district magnitude. From this, it follows that district magnitude will be binding and there should be a co-variation of magnitude and the number of candidates or lists in Duvergerian equilibrium.
(3) The third scenario specifically applies to electoral systems, which employ a wide range of different district magnitudes. Smaller districts, which have a carrying capacity below the number of groups seeking representation, may be limiting the proliferation of candidates and lists, while larger district magnitudes that have a carrying capacity higher than the number of groups will often not be binding at all. This also implies that N_v does not linearly increase with M, but that the empirical association of these core variables levels off for larger districts.

While strategic voting and strategic entry still operate even in larger PR districts (Gschwend 2007; Gschwend and Stoiber 2012), there are a number of reasons to assume that the marginal effect of district magnitude decreases at higher levels. It is much more difficult to gather precise information about which candidates or lists are competitive and which are trailing when districts become larger, although voters may of course also act upon less precise guesses about the standing of electoral races. Beyond the key features of the electoral system, the formal requirements which regulate the inclusion on the ballot and the organisational prerequisites for gaining sufficient visibility and, in turn attracting success, clearly increase when district magnitudes and local electorates become larger. Therefore, in the subsequent analysis, I will adhere to the idea of modelling district magnitude on a logged scale.[1]

As indicated before, the $M + 1$ rule does not posit unconditional comparative statics of district magnitude. The second and third scenarios, however, propose a conditional and context-sensitive association of these key variables. M and N will be related whenever the number of electoral contenders reaches or exceeds the carrying capacity of the electoral district and the specified contextual conditions for strategic coordination on $M + 1$ viable candidates or lists are fulfilled. An upper bound is therefore either established by the institutional carrying capacity of the electoral system, as defined by the $M + 1$ rule, or set by the ethnic or religious fragmentation of a society. This implies that these two equations may be treated as exchangeable:

$$\mathrm{E}[N|M;F] = \min(M+1,F) = \beta_0 + \beta_1 \mathrm{lg}M + \beta_2 F + \beta_3 \mathrm{lg}M \cdot F$$

This implies that the number of viable contenders in a local district (N) is less than or equal to the minimum of the carrying capacity of the district (as defined by $M + 1$) and the degree of societal fragmentation (F). The interactive relationship of electoral systems and societal heterogeneity implies that $\beta_1 = \beta_2 = 0$ and $\beta_3 > 0$.

Concerning the institutional context, the provision of upper tiers conditions the mechanisms of strategic voting and strategic entry. When upper and lower tiers are linked by some sort of remainder, seat, or vote transfer, upper tiers might actually weaken the incentives set forth by the electoral system, inflate its carrying capacity, and render district magnitude less binding. In connected multi-tier systems, any votes that did not elect a candidate or list in the primary electoral districts could still be successful in the secondary ones, so that it may often be instrumentally rational in the short term to cast a vote for a locally hopeless candidate. Upper tiers may thus undermine the incentives produced by the wasted resources and wasted vote arguments and thus weaken strategic entry and strategic voting:

$$\mathrm{E}[N|M;F] = \beta_0 + \beta(\mathrm{lg}M + F + \mathrm{upper})$$

$$= \beta_0 + \beta_1 \mathrm{lg}M + \beta_2 F + \beta_3 \mathrm{upper} + \beta_4 \mathrm{lg}M \cdot \mathrm{F} + \beta_5 \mathrm{lg}M \cdot \mathrm{upper}$$

Since the provision of upper tiers tends to undermine the logics of strategic entry and strategic voting, I generally expect a higher number of candidates or lists in equilibrium when there is a significant share of upper-tier seats (this implies $\beta_3 > 0$). In addition, when seats are not exclusively allocated within the primary electoral districts, but also within secondary districts, I do not expect any covariance of local district magnitudes and the number of locally viable candidates or lists (i.e. $\beta_5 < 0$).

In order to fully address empirical variation in the rich dataset at hand, we estimate hierarchical linear models, which include random intercepts and random slopes at the election level (j). We assume that the random effects and the error term are independent and identically distributed (i.i.d.) and drawn from a multivariate normal distribution with $\zeta_{0j} \sim N(0,\psi_0); \zeta_{1j} \sim N(0,\psi_1)$ and $\varepsilon_{ij} \sim N(0,\theta)$:

$$\begin{aligned}\mathrm{E}[N|M;F] = {} & (\beta_0 + \zeta_{0j}) + (\beta_1 + \zeta_{1j})\mathrm{lg}M + \beta_2 F + \beta_3 \mathrm{upper} + \beta_4 \mathrm{lg}M \cdot F \\ & + \beta_5 \, \mathrm{lg}\, M \cdot \mathrm{upper} + \varepsilon_{ij}\end{aligned}$$

Data and Measurement

This section turns from theoretical concerns towards the presentation of the empirical data at hand. In the first step, the data and data sources are briefly reviewed, and in the second step basic decisions regarding the operationalisation of the key indicators at the district level and at the national level are introduced and defended.

District-Level Data Sources

For a long period of time, there has often been an asymmetry of electoral system theories, which focused on the district level, and the availability of reliable empirical information that was regularly only accessible for the national level. In past years there have been considerable achievements in the collection and publication of district-level electoral returns. One of the most exhaustive datasets was prepared by Daniele Caramani (2000), who compiled systematic and standardised district-level general election results, where available, for 18 West European countries from as early as 1830 until the end of the twentieth century.

In this study, I draw heavily on this rich database, but I settle on the interval from 1945 to 1998 in order to obtain parallel, synchronous observations taken from as many countries as possible. In this period, detailed district-level data is available for 169 elections to the respective national parliaments. Thus, the dataset comprises 17,248 single observations (i.e. electoral districts). This number includes 14,152 single-member districts in Germany, Italy, and the United Kingdom and 3,096 multi-member districts in Austria, Belgium, Finland, Iceland, Ireland, Italy, Luxembourg, the Netherlands, Norway, Portugal, Spain, Sweden, and Switzerland. In addition, the dataset has, where available, been carefully augmented by information on the magnitude of electoral districts, several other electoral system features and national-level data regarding social heterogeneity or basic features of the political systems taken from a variety of sources. (More extensive information on data and data sources, the elections covered, and the variation of district magnitude are provided in the Online Appendixes A, B, and C.)

Key Indicators in a Hierarchical Context

Previous studies have settled on a consensus regarding the measurement of the effects, which are assessed at the district level (the key variables effective number of electoral parties and effective magnitude) and at the national level (the conditioning variables social heterogeneity and the provision of upper tiers):

The effective number of electoral parties (N). In any analysis of Duverger's law the primary dependent variable is the number of viable contenders in an election. Analyses at the national level usually apply the effective number of parties suggested by Laakso and Taagepera (1979). Students of district-level electoral returns, however, did not unanimously settle on a specific measure of (local) party system fragmentation. Some contributions have concentrated on counts of the raw numbers of candidates or party lists. Other approaches have somewhat narrowed that down and have addressed a more or less defined number of 'viable' contenders within an electoral district (see Cox 1994, 1997). Others have also applied the effective number of parties as a weighted count of the number of parties at the district level.

Certainly, the choice of an appropriate measure depends on the research question. Focusing on strategic entry, counts of raw or 'viable' parties would clearly be appropriate. Focusing on strategic voting, I prefer the weighted party count, as given by the effective number of parties, also for the analysis of district-level data. The basic rationale provided by Cox refers to the desertion of trailing candidates and the strategic switching towards more competitive candidates in an electoral district, meaning the further weakening of weak and the further strengthening of strong competitors for the public vote. I believe that an unweighted count of candidates or parties addresses only one side, namely the desertion of trailing platforms. In contrast, weighted vote counts address directions, vote gains and vote losses due to strategic coordination. The effective number of electoral parties is given by $N = [\sum v_i^2]^{-1}$.

Local effective magnitude (M_{eff}). Since Douglas W. Rae (1967) referred to district magnitude as the 'all-important factor', this quantity has become the key concept for measuring the impact of alternative electoral systems. In the equilibrium of strategic coordination – that is, strategic entry by political elites and strategic voting by the electorate – larger district magnitudes tend to be permissive towards the proliferation of candidates, while lower magnitudes, above all single-member districts, tend to limit the number of viable contenders. Taagepera and Shugart (1989: 117, 126–41, 266–9, 274–7) have modified this concept so as to introduce the 'effective magnitude' combining information on district magnitudes and electoral thresholds. The authors propose the expressions $T = 50\%/M$ and $M = 50\%/T$ to roughly express formal thresholds in district magnitudes equivalents and vice versa.

Upper tiers (upper). The presence and proportion of upper-tier seats are crucial features of any electoral system and determine whether local district magnitudes can be considered strictly binding or not. If there is only one tier, parliamentary seats can only be achieved within the respective electoral districts of magnitude M, and local district magnitudes may, given the aforementioned contextual conditions, be binding. When there are additional tiers so that, for instance, unsuccessful votes within a specific district are used in a second, regional or national stage of seat allocation, local district magnitudes cannot be binding in a strict sense. Voting for a locally hopeless candidate or list may still be instrumentally rational in the short term when these votes potentially help the specific party to gain parliamentary seats at the upper levels of the electoral system. The independent variable upper therefore captures the share of the seats that is awarded in the upper tiers of the electoral system, roughly ranging from 0 (single-tier electoral systems, for instance in Finland, Portugal, or Spain) up to about 0.5 for more complex systems like the German mixed-member proportional system (Shugart and Wattenberg 2001).

Social fractionalisation and heterogeneity (F_{eth}; F_{rel}). In addition to the key electoral system variables, the main arguments of this contribution require the careful consideration of the effects of social heterogeneity and the interactions of institutional and societal structures. In any sociological argument, party system fragmentation is supposed to increase with the number of political cleavages or issue dimensions (see Lipset and Rokkan 1967). According to this view, longstanding multipartyism is determined by the existence of multiple, stable, and politicised lines of division within a society, whereas two-party systems are to be explained by fewer lines of division in politics and by the relative mildness of these issues.

Although ethnic heterogeneity features prominently in the existing literature, one has of course to admit that these measures provide a very poor proxy for capturing social and political conflict/diversity as outlined by Lipset and Rokkan (1967). Given that the heterogeneity of West European societies is defined by more than just the ethnic dimension, I have opted for more detailed data on ethnic and religious diversity which has been published by Alesina *et al.* (2003). So as to measure empirical fragmentation levels, the authors utilised a fragmentation index which is computed as one minus the Herfindal index of the respective group shares: $F = 1 - [\sum g_i^2]$. Substantively, this index captures the probability that two randomly selected individuals are members of different ethnic, linguistic, or religious groups. An in-depth discussion of more adequate and advanced empirical measurement strategies for assessing the societal inputs to the political process is provided by Stoll (2013: chapter 3).

The Empirics of Supply and Demand

There are numerous methods and perspectives for measuring the political consequences of electoral systems. Some contributions have turned to the microfoundations of strategic voting and tried to determine the share of voters that responded strategically to the incentives of the electoral systems (Alvarez and Nagler 2000; Blais and Nadeau 1996; Gschwend 2007; Gschwend and Stoiber 2012). Other contributions have looked for patterns of strategic desertion of trailing candidates at the district level and, for instance, focused on the vote ratio of the second and first losers (the so-called 'SF-Ratio') in any primary electoral district (Cox 1997). Instead of focusing on the whole battery of potential indicators, this contribution adopts a straightforward perspective on the number of candidates or lists competing in a local district as the substantively most meaningful dependent variable.

Merging Supply with Demand

Table 1 presents three multi-level models which consider that electoral districts i are nested in the contexts of elections j and their more or less stable social and institutional contexts. Random coefficients and random slopes at the election level capture contextual heterogeneity in the association between local effective magnitudes and the number of viable candidates or lists. Evident contextual variation is, in turn, accounted for by the higher-level covariates: ethnic and religious heterogeneity and the presence of upper tiers.

I begin the presentation of the results with some simple bivariate evidence which does not straightforwardly relate to the $M + 1$ rule, but rather adopts the inductive perspective embodied in the 'generalized Duverger's law': Model 1 in Table 1 shows empirically that, as hypothesised, more permissive electoral districts (i.e. larger effective magnitudes) are closely associated with higher fragmentation levels of the local electorates, while lower district magnitudes are regularly coupled with a lower number of locally viable candidates or lists. The estimates for intercept and slope closely resemble the inductive rule of thumb provided by the 'generalized Duverger's law' (Taagepera and Shugart 1989) and also correspond with previous empirical findings with national-level data (see Amorim Neto and Cox 1997; Clark and Golder 2006; Golder 2006; Lijphart 1994; Ordeshook and Shvetsova 1994). Plugging in the unconditional empirical estimates from Model 1 yields something like $N = 1.21 \cdot \lg M_{\text{eff}} + 2.49$.

TABLE 1
EFFECTIVE MAGNITUDE AND THE NUMBER OF PARTIES

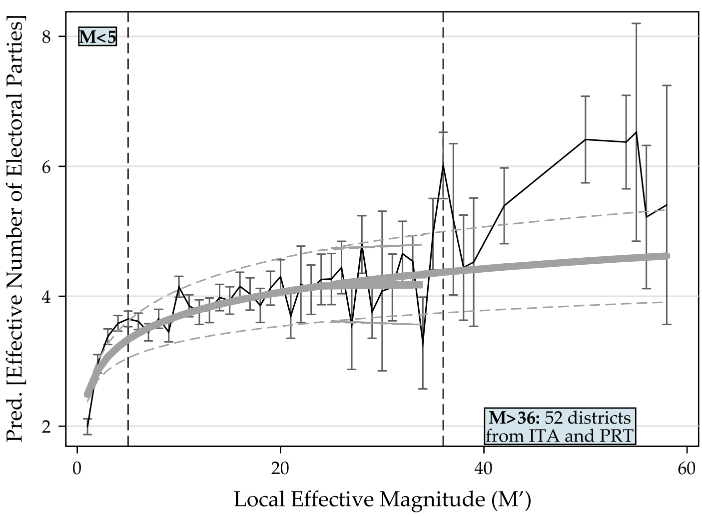
	(1)		(2)		(3)	
1: Electoral system						
lg M_{eff}	1.21***	(0.09)	0.44*	(0.19)	0.71***	(0.18)
upper					1.23***	(0.31)
2: Social diversity						
– ethnic: F_{eth}			−0.73**	(0.27)	−0.74**	(0.27)
– religious: F_{rel}			−0.65**	(0.24)	−0.81***	(0.25)
3: Interaction terms						
lg M_{eff} * F_{eth}			1.23**	(0.40)	1.62***	(0.39)
lg M_{eff} * F_{rel}			1.30**	(0.42)	0.92*	(0.41)
lg M_{eff} * upper					−4.20***	(0.94)
Constant	2.49***	(0.05)	2.93***	(0.12)	2.85***	(0.12)
4: Random effects						
$\sqrt{\varphi_0}$	0.89	(0.012)	0.75*	(0.11)	0.64**	(0.09)
$\sqrt{\varphi_1}$	0.35***	(0.05)	0.31***	(0.05)	0.29***	(0.04)
$\sqrt{\theta}$	0.22***	(0.01)	0.22***	(0.01)	0.22***	(0.01)
N	17,248		17,248		17,248	
log. Likelihood	−11787.42		−11773.50		−11757.87	

Notes: Dependent variable = *N*.
Hierarchical linear models; standard errors in parentheses.
*$p < 0.05$.
**$p < 0.01$.
***$p < 0.001$.

Figure 1 summarises these findings graphically and provides some additional checks on the association of the number of candidates or parties and its specific functional form: the thick line indicates the logarithmic mapping of effective magnitude upon the effective number of electoral parties, the thin dashed lines show the respective confidence intervals (all printed in grey). The diagram displays another model so as to facilitate specification checks: instead of treating logged effective magnitude as a continuous predictor, I break district magnitude down into a series of dummy variables ($1 \leq M \leq 58$). The vertical error bars capture the effect of district magnitude on the effective number of parties and the related confidence intervals (all printed in black). Both specifications produce almost identical results as long as district magnitudes are small. When district magnitudes become very large (e.g. from $M > 35$), the predictions derived from the series of dummy variables are increasingly unstable and somewhat off the scale. Note that these predictions are only based on about 50 districts, which are all taken from Italy (up until 1992) and Portugal (until 1991).

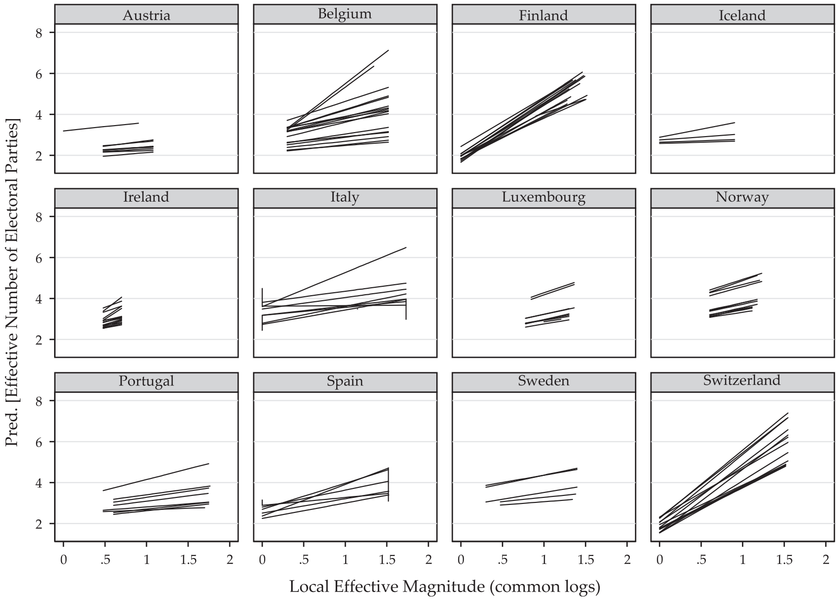
The random effects part of the model also reveals considerable contextual heterogeneity regarding both the baseline levels of electoral fragmentation (the random intercepts) and the marginal effect of logged district magnitude on the number of candidates or lists.[2] Figure 2 explores the context dependency of

FIGURE 1
DISTRICT MAGNITUDE AND THE NUMBER OF PARTIES



Notes: This figure is based on Model 1 in Table 1. The thick curve illustrates the impact of effective magnitude M_{eff} on the effective number of electoral parties N_v, and the thin dashed lines represent the 95 per cent confidence bands. Moreover, the thin solid line and the vertical bars capture the coefficients for the categorical representation of district magnitude and their respective confidence intervals. (The large nationwide districts employed in the PR elections to the Dutch 'Tweede Kamer' [N = 100 in 1952 and N = 150 since 1956] have been omitted to make the figure more readable.)

FIGURE 2
ELECTION- AND COUNTRY-WISE VARIANCE



Notes: The dashed lines represent distinct elections, while the bold solid lines refer to the theoretical baseline given by the 'generalized Duverger's law' (Taagepera and Shugart 1989). Germany and the UK have been excluded because both these countries employ, as assessed by this dataset, exclusively single-member districts and thus exhibit no empirical variation. Denmark and Greece have also been excluded due to some missing data and interruptions in the time series.

electoral systems in more detail. The effect of local district magnitudes on the fragmentation of local electorates, as indicated by the random slope on logged effective magnitude, significantly varies by election and by country. Each of the lines in the various panels of the trellis captures the empirical Bayes predictions for the marginal effects of district magnitude. Flat regression profiles refer to the absence of electoral system effects and the number of candidates or lists is about the same in small and in large districts. In contrast, steep profiles indicate that strong marginal effects and district magnitudes may potentially be binding. In this case, small district magnitudes are linked with low, and large magnitudes are linked with high numbers of electoral parties.

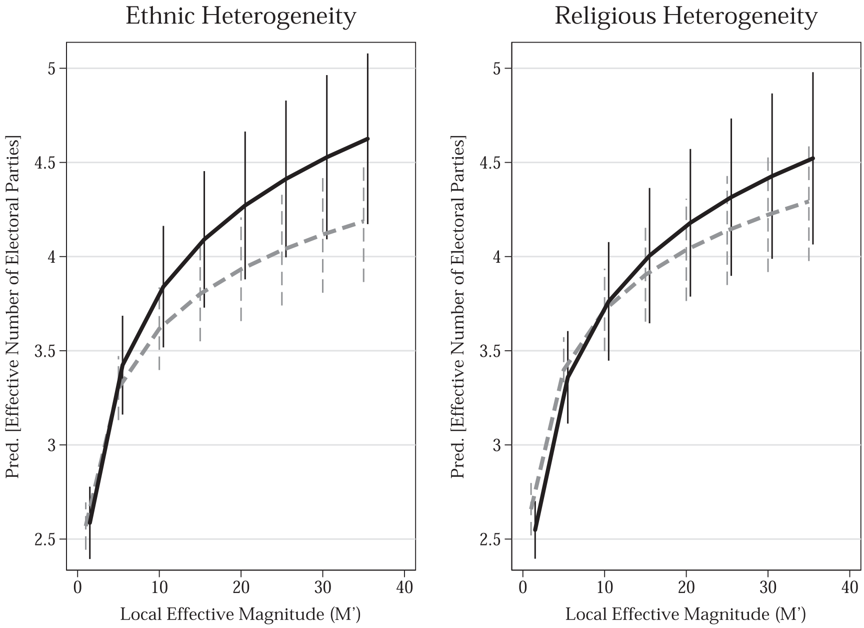
This allows us to shed some light on the political consequences of effective magnitude: In the first group of countries, there is a particularly steep slope indicating a significant marginal effect of local district magnitudes on electoral fragmentation. This is clearly visible in countries that are comparatively more heterogeneous in ethnic and religious terms and/or have implemented single-tier PR systems. Countries like Finland, Ireland, and Switzerland and, somewhat less clearly, Portugal and Spain fall into that category. The second cluster of countries also unambiguously reveals significant marginal effects, but these are somewhat weaker or unstable across repeated elections: Belgium, Italy, and Luxembourg. In a final cluster of countries, there is only limited evidence for the marginal effects of district magnitude. Almost flat predictions for the election-specific regression profiles indicate that institutional incentives are obviously not effective and/or binding. This refers to countries which are either socially homogeneous and/or have implemented more complicated PR electoral systems with upper tiers and vote or seat transfers such as Austria, Iceland, Norway, or Sweden.

So far, we have presented some inductive generalisations and unconditional comparative statics which follow the traits of the 'generalized Duverger's law' (Taagepera and Shugart 1989, 1993), but which do not fully address the original argument made by Duverger and do not pick up the logics which lie at the heart of the $M + 1$ rule. When evaluating whether district magnitude becomes binding, the analytical scope cannot be limited to the supply side given by the electoral system, but also needs to consider the demand side given by the number of societal groups, which organise into political parties and strive for representation in parliament. In empirical terms, the extensive database on ethnic, linguistic, and religious fragmentation, compiled by Alesina *et al.* (2003), provides the principal indicators for social heterogeneity. Because the measures of ethnic and linguistic diversity are highly collinear, only the ethnic and religious dimensions are available for the empirical analysis.

Model 2 in Table 1 provides very conclusive evidence in favour of the joint supply and demand model. Note that interactive regression models usually (need to) include both the constitutive (here: $\lg M_{\text{eff}}$ with F_{eth} or F_{rel}, respectively) and their multiplicative terms ($\lg M_{\text{eff}} * F_{\text{eth}}$ and $\lg M_{\text{eff}} * F_{\text{rel}}$, respectively).[3] Introducing interactions of effective magnitude and the two diversity indicators renders the constitutive terms of effective magnitude statistically barely significant and substantively smaller. Notwithstanding, there are significant and substantively meaningful interactions of effective magnitude with both ethnic and religious fractionalisation. If there is a high demand for parliamentary representation by various ethnic or religious groups in a society and if local electoral districts have a sufficient carrying capacity, the upper bound may become binding.

Figure 3 displays the predictive margins of these interactive effects. Regarding ethnic (the left-hand panel) and religious diversity (the right-hand panel), the marginal effects of district magnitude are substantively stronger in heterogeneous than in homogeneous countries. By and large, the marginal effect of ethnic heterogeneity seems to be somewhat more meaningful and robust than the consequences of religious diversity. These arguments conform to the analytical concept of the $M + 1$ rule in a more straightforward manner. Effective magnitude becomes binding when there is a high number of groups forming parties and contesting the elections; in contrast, the marginal effects of

FIGURE 3
PREDICTIVE MEANS OF LEVELS OF SOCIAL HETEROGENEITY



Notes: The solid black lines and confidence intervals refer to countries with higher than average diversity, the dashed grey lines and confidence intervals refer to more homogeneous countries with lower than average diversity.

district magnitude are much weaker in more homogeneous countries where the proliferation of candidates or parties will often be below the carrying capacity established by the electoral rules from the outset.

That said, there are also some limitations to the preceding analyses: we could only obtain comparable indicators of social diversity at the national level, but we lack any similar data for the district level of the 17 countries in the analysis. This implies an emphasis on what we have called a top-down perspective on strategic entry. The downside, of course, is that we lack any data on controls for district-level variables other than their magnitude, which may also affect the number of candidates or lists. For instance, smaller districts are often (but not always) employed in the countryside, while larger districts often (but not always) tend to be set up in larger cities. One could perhaps speculate that larger communities tend to employ larger districts and are at the same time characterised by more complex and diverse coalitions of societal interests. This might easily run counter the institutional argument and, lacking data on the social composition of the primary districts, this alternative scenario is difficult to refute.

When I utilise national averages as proxies for the diversity of individual districts, I introduce measurement bias to a key independent variable, because the 'real' social diversity scores will be different from their average. Therefore, there is a risk of introducing bias to our estimates of these key effects (see King *et al.* 1994: 163–8). In order evaluate its magnitude (and thus to check the robustness of the empirical evidence), I have conducted a series of statistical simulations with varying degrees of unsystematic intra-election variation. More specifically, I have added i.i.d. distributed random terms to the indicators of ethnic and religious diversity (F_{eth} or F_{rel}), and, for each level of variation, I have repeated the simulation and estimation procedures 1,000 times. Note that these counterfactuals are even biased against the key hypotheses since one may usually assume that larger, urban districts tend to be more heterogeneous than small, rural ones.

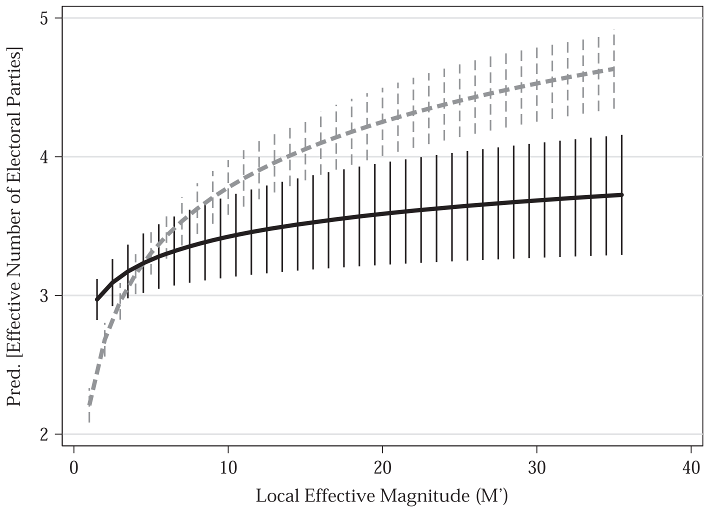
Below the line, the evidence from these simulation studies suggests that the core findings on the interactive and conditioning effects of social diversity are very robust. Both concerning ethnic and religious diversity, the interactive effects of electoral and social structures are robustly maintained even if I allow for significant levels of within-context heterogeneity. (For the details of simulation and estimation please refer to Online Appendix D, especially Figures D-1 and D-2 summarising the simulations.)

The Effects of Upper Tiers

As hypothesised, the provision of upper tiers may render local district magnitudes less binding and the $M + 1$ rule less applicable or invalid. In single-tier electoral systems, where seats are exclusively allocated within the primary electoral districts, votes for unsuccessful candidates or lists are wasted. When two- or multi-tier systems allow for the transfer of unused votes or seats towards the regional or national level, this is not necessarily the case. Model 3 in Table 1 adds the share of seats which are awarded in upper tiers to the interactive specification. The empirical findings indicate a strongly significant negative interaction of local effective magnitudes with the provision of the upper tier at the election level. Effectively, the larger the share of seats assigned to secondary or higher districts, the smaller the marginal effect of district magnitude on the number of viable electoral contenders.

Figure 4 illustrates the predictive margins associated with this model. Controlling for social heterogeneity, the number of candidates or lists visibly increases with district magnitudes if there are only primary electoral districts, and local district magnitudes thus tend to be binding. In contrast, the effects of district magnitude almost completely diminish when we focus on those electoral systems that do employ secondary (or even tertiary) districts. In the presence of upper tiers, the local number of candidates or parties in small and in large districts is almost the same.

FIGURE 4
PREDICTIVE MEANS OVER SINGLE- AND MULTI-TIER PR



Notes: The dashed grey line and confidence bands refer to single-tier electoral systems without any secondary districts, the solid black line and confidence bands indicate more complex two- or multi-tier electoral systems that emply upper tiers for seat allocation.

Summary and Conclusion

Any generalisation of a well-established hypothesis, most notably the proposition of a 'Law', suggests the very impression of scientific progress. Theoretical studies have, during recent decades, made significant progress in terms of the clarity of the argument, the formalisation and sophistication of propositions. However, empirical analyses still lag way behind these theoretical achievements. While the strategy of gross national-level generalisation seems to be exhausted and does not provide any new insight, sophisticated research based on district-level data, for instance the growing literature on mixed-member electoral systems, is too often confined to single-case studies yielding heavily context-dependent results and almost no scope for systematic generalisation.

The research design adopted in this contribution has picked two attempts to restate Duverger's original propositions as the conceptual point of departure: (1) the inductive generalisation implied by the 'generalized Duverger's law' (Taagepera and Shugart 1989, 1993) and (2) the causal argument stated by the 'direct generalization of Duverger's law' (Cox 1997). The first perspective implies unconditional, the second conditional comparative statics of district magnitude.

The multi-level framework employed in the empirical analysis allows for systematic assessments of heterogeneous contexts and the fruitful procession of more extensive and detailed empirical datasets. The findings presented here are threefold. First, the analysis of district-level electoral returns refers to the adequacy of a multi-level perspective, and the significant causal force and empirical validity of the key causal mechanism embodied in the $M + 1$ rule. Both within and across the contexts of individual elections the number of viable contenders in a district co-varies with its magnitude. The political consequences of district magnitude cannot, however, be modelled by unconditional comparative statics, but are dependent upon the provision of candidates or political parties by social (for instance ethnic or religious) division lines.

The second finding regards the fundamental context-dependency of electoral coordination. In some countries, district magnitude and local fragmentation are closely connected, for instance in Finland or Switzerland, while in other polities the empirical association is much weaker or even absent. The constraining effects of (low) district magnitudes are thus moderated by contextual factors at the election or country levels. Individual electoral districts, taken from the same contexts, tend to closely resemble each other. Significant intraclass correlations refer to potential top-down effects of strategic entry that, in turn, interact with the bottom-up logics of strategic voting.

Thirdly, I have proposed some causal mechanisms that might either encourage or prevent the strategic moves by voters and by partisans that lie at the heart of Duverger's psychological effect. The marginal effect of (logged) district magnitude becomes significantly smaller when upper tiers of the electoral systems render local district magnitudes less binding.

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Notes

1. The utilisation of logged scales was pioneered by Taagepera and Shugart (1989, 1993) and it has been common practice in a number of analyses which focused on national-level electoral returns (see Amorim Neto and Cox 1997; Benoit 2002; Brambor *et al.* 2006, Clark and Golder 2006; Filippov *et al.* 1999; Golder 2006; Ordeshook and Shvetsova 1994). Logged scales have also been utilised for the study of district-level electoral returns (see Singer and Stephenson 2009). Other contributions even go beyond that and suggest an inverse quadratic relationship of district magnitude and the number of viable candidates or lists.
2. A Lagrange multiplier test confirms these results. The random slope model fits the data considerably better than the more restrictive random intercept specification ($\chi 2 = 1174.79^{***}$) so that any model that fails to account for this variant of contextually induced heterogeneity has to be considered severely misspecified.
3. The application and interpretation of interactive regression models in political science are discussed in depth by Kam and Franzese (2007).

Notes on Contributor

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Projection effects and specification bias in spatial models of European Parliament elections

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Abstract

Substantial empirical evidence suggests that voters cast their ballot not only by considering the different policy positions of parties or candidates, but also appear to pull candidates/parties they prefer closer to their own ideal position ('assimilation') while pushing candidates/parties they dislike, farther away ('contrast'). These effects are called 'projection effects'. We illustrate that voters' perceptions of policy positions of candidates/parties are contaminated by non-spatial considerations. Building on data from the EES series, we empirically demonstrate that projection effects are substantively meaningful and statistically significant in elections to the European Parliament. We moreover distinguish between unsystematic projection bias that only depends on the closeness to a specific candidate or party and systematic projection bias that is also affected by party-, voter-, and context-specific determinants.

Keywords

Comparative politics, European elections, projection effects, spatial modelling

Introduction

The spatial model of voting, as pioneered by Downs (1957), is often considered the workhorse of modern electoral studies, and belongs now to the standard toolbox of applied empirical researchers. Any spatial voting model, in principle, claims two things: (i) there exists a meaningful relationship between the ideal points of voters

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and the programmatic positions of political parties in a one- or multidimensional political space, and (ii) these relations matter for party evaluation and vote choice.

The validity of empirical research therefore depends on the ability of voters to meaningfully locate parties within a competitive political space. However, eminent contributions to political psychology have shown that there are limits to this. Particularly, concerns are raised over whether voter-specific party placements may be biased due to so-called projection effects. This implies that individuals may locate parties they like, for whatever reason, closer to their personal ideal point ('assimilation'), while they push parties they oppose, for whatever reason, further away ('contrast').[1]

Empirical analyses have compiled valid and robust evidence for the presence and consequences of projection effects (Judd et al., 1983; Miller et al., 1986; Granberg et al., 1988; Krosnick, 1990; Rahn et al., 1994). There are still some significant gaps. Most studies have focused on US elections and have placed a strong emphasis on the personal vote of presidential or congressional candidates. Further, political scientists often limited their methodological perspectives to binary party competition, concentrating most of the time therefore on the distinction between one preferred and one disliked party (notable exceptions are provided by Granberg and Holmberg, 1986; Granberg and Brown, 1992; Merrill et al., 2001; Markowski and Tucker, 2007). European voters are, in contrast, regularly faced with more fragmented party systems, and with much more complicated patterns of inter-party competition and cooperation that may well be expected to have an impact on their preference structures. Instead of binary preference structures, they will more often support various parties to varying degrees.

In this contribution, we fill this crucial research gap and analyse projection effects within a systematic and comparative framework. We compare the magnitude and the determinants of projection effects in elections to the European Parliament (EP). More specifically, we focus on the most recent wave of the European Election Studies (EES) series, the PIREDEU study of 2009. To our knowledge, we provide the first systematic comparative assessment of projection effects and utilize the most comprehensive dataset to date. The PIREDEU study also offers the opportunity to compare projection effects along a highly salient national left–right dimension and a supranational European integration dimension of low salience.

The presence of projection effects may affect both actual vote choice and scientific models of vote choice. Regardless of whether voting behaviour is motivated by spatial proximity, direction, or by discounting, distorted party placements influence electoral behaviour and lead to wrong inferences about electoral choice. Empirical evidence for projection effects thus calls into question the principal findings of the spatial voting framework regarding national and European elections. This is particularly true when projection effects are heterogeneous and context-dependent, and when the presence and the extent of assimilation and contrast differ across alternative dimensions of political contestation. Further, voter- and party-specific determinants may introduce systematic distortion and thus bias studies of EU issue voting. The presence of projection effects is therefore not confined to socio-psychological labs or insider discussions in niche journals. Rather, the presence of assimilation and contrast potentially modify vote choice and bias scientific analyses of EP elections.

Our article addresses these issues in three consecutive steps. We first present our principal argument, develop a formal definition, and propose a measurement model for the differences between subjective and 'objective' party placements. The empirical analysis provides descriptive evidence of the presence and the magnitude of projection bias. In order to check and cross-validate the consistency and robustness of our findings, we compare projection effects on a high salience dimension, the left–right scale, and a low salience dimension, the European integration scale. In a third step, we present and justify the distinction between unsystematic and systematic projection effects and review evidence on the causal determinants of projection effects at the voter, party, and system levels. The final section concludes with some remarks on assimilation and contrast and a word of caution to applied researchers.

Defining projection effects

Explanations for projection effects build on consistency perspectives positing that individuals try to overcome and work against cognitive dissonance; predominant are dissonance theory (Festinger, 1957), balance theory (Heider, 1958), and congruity theory (Osgood and Tannenbaum, 1955). Individuals experience cognitive inconsistency if two cognitions do not correspond, and the perceived discomfort increases with the level of substantive importance attributed to these cognitions. The dualism of assimilation and contrast is thus an established building block in the classical literatures on social psychology and cognition theory.

These classical approaches have been adopted by electoral studies and used for studies of voter perceptions of party ideologies and positions.[2] Applied to voter-specific party placements, psychological consistency may be achieved by moving either the voter's personal ideal point or by manipulating the spatial positions of the political parties. The existing literature often builds on rather ambiguous assumptions, on the one hand, regarding the difference between persuasion and projection effects, and on the other hand, concerning the relation of assimilation and contrast. Thus, we start by tailoring these key terms from cognition theory to electoral studies:

1. The voter may adjust or change her individual spatial position due to efforts taken by political parties to persuade potential voters of specific party policies ('persuasion').
2. The voter may review and modify the spatial positions taken by alternative parties in order to bring consistency back in that:
   a. she might pull positions of political parties she likes for non-policy reasons closer to her unchanged spatial position ('assimilation');
   b. she might push parties she opposes for non-policy reasons further away from her unchanged ideal point ('contrast').

The inherent causal mechanisms are established with reference to classical contributions to political psychology: following Heider's 1958 balance theory, the individual need for establishing cognitive consistency first depends on the existence of 'unit relations' with the respective candidate or party, in other words, the stronger the attachment to a specific candidate or party, the higher the individual pressure to eradicate cognitive dissonance. Second, dissonance theory also stresses the role of choices in strengthening the need for projection effects. For instance, casting a vote for a specific candidate or party increases the chance of assimilation effects to eliminate any possible post-decisional cognitive dissonance and vice versa. Third, not only electoral preferences and choices, but also the importance attributed to specific issues, regulate how sensitive an individual is to cognitive inconsistencies. The salience of an issue dimension therefore determines the extent of projection effects.

This contribution focuses on the extent and determinants of projection effects rather than the electoral consequences of party cues and persuasion effects. Assimilation and contrast label causal effects which cannot be directly observed (cf. Krosnick, 1990, 2002). Political scientists are thus not in a position to unambiguously establish whether a voter has modified her individual position due to successful persuasion, or whether she has altered the spatial positions she assigns to the party alternatives as a result of assimilation and/or contrast (cf. Conover and Feldman, 1982; Tomz and van Houweling, 2008).

Krosnick (2002) points out that cross-sectional data alone is not particularly helpful for the straightforward isolation and differentiation between persuasion and projection effects.[3] Dynamic panel studies potentially facilitate the isolation of causal factors due to the temporal sequencing of events, but panel data is usually not available and comes with problems of its own. In multi-party elections, applied researchers may validate their findings by exploring contextual data. Recall that policy-based persuasion effects will result in a movement of the voter position towards the policies represented by her favoured candidate or party; changes in the spatial distances towards other parties are merely a by-product of this. In contrast, the parallel dynamics of assimilation and contrast involve the movement of *all* party placements and thus produce a much more structured pattern.

Furthermore, issue salience may help to separate persuasion and projection. As indicated above, higher issue salience is expected to increase projection effects. Regarding persuasion, the voters' policy preferences are stable and tend to resist change when they attribute high importance and salience to a certain dimension of political conflict. In turn, voter positions are fragile, unstable, prone to change, and may be significantly altered by persuasion when attributed issue salience diminishes (Krosnick, 2002: 124).

Given that projection effects, assimilation and contrast, imply that non-spatial motives systematically bias voter-specific party placements, our measurement strategy proceeds in two consecutive steps:

1. Differences of subjective and some kind of 'objective' party placements constitute a necessary, but not a sufficient, condition for the presence of

projection effects. Therefore we compute the deviation of voter-specific party placements from some baseline measure of 'true' positions.
2. A sufficient condition for the definition of projection bias exists if voters deflate (inflate) spatial distances for candidates or parties they like (oppose). To measure whether this condition is met, we relate the calculated placement differences to some measure of party attachment and evaluate the marginal effects as a sufficient condition indicating projection effects.

'Objective' party positions

A crucial point for the measurement of projection effects is whether we need some yardstick of 'objective' party positions and, if so, how such a specific baseline value may be theoretically justified and empirically assessed. Of course, the very notions of 'objective' or 'true' party positions are conceptually dubious, and guessing 'where a party really stands' intimates somewhat esoteric assumptions and goals. In fact, operationalizations and measurement strategies suggested by political scientists are often ambiguous. Established measures of party ideology frequently start out with different assumptions, are often designed for different purposes, and differ empirically.[4]

In empirical terms, the studies by Granberg and Holmberg (1986), Granberg et al. (1988), Merrill et al. (2001), and Adams et al. (2005) reach similar conclusions: there is ample, solid, and robust evidence for the existence of projection effects. Its two components, assimilation and contrast, appear to be of approximately similar significance in US presidential and congressional races. Beyond the US as the last remaining example of classical bi-partism, the assimilation of preferred parties appears to have a much more meaningful effect than the contrast of disliked ones. This has often been called the 'asymmetry' of projection effects which is defined by the alleged dominance of assimilation over contrast effects (Granberg and Holmberg, 1986; Granberg et al., 1988; Merrill et al., 2001; Krosnick, 2002).

Approaches to measuring projection effects frequently focus on the departure of subjective, voter-specific party placements from some baseline measure of 'objective' or 'true' party positions. The most common choice is to proxy unified party positions by the arithmetic mean of voter-specific party placements. This is based on the hope that the differences of subjective and 'objective' party positions and projection bias cancel each other out (Macdonald et al., 1991, 1995, 1998, 2001; Rabinowitz and Macdonald, 1989, 2007).

Alternative measurement strategies focus on external data sources such as roll calls, expert judgements, or manifesto data as indicators of 'true' party positions. External data may help to address problems such as endogeneity, but often come with problems of their own. Scalability problems complicate the use of any data source as an indicator of 'true' party positions. Differences of voter-specific party placements and external data, such as expert judgements, manifestos, or roll-call data, may not unambiguously relate to misplacement and projection, but are also affected by conceptual and empirical inconsistencies among the data sources.

While we subscribe to most of the criticism put forward against unified and 'true' party positions, we nevertheless think that a straightforward quantification of projection bias needs to rely on some proxy for 'objective' party positions. We continue to use mean party placements over all respondents, because this measure still provides the least problematic yardstick for 'true' party positions. Notably, mean placements help to minimize scalability problems and avoid the loss of observations due to incomplete matching with other data sources.[5] However, we have replicated the analyses with alternative proxies of 'true' party positions such as expert judgements. While these results obtained with different indicators do not differ systematically, the combination of diverse measures and datasets adds a substantive amount of random noise. The web appendix documents these robustness checks.

Placement differences and projection bias

A necessary condition for the assertion of projection effects is a difference between spatial distances based on subjective and 'objective' party placements. Considering that our unit of analysis is not the location voter i assigns to party j, but rather the distance between the voter's self-placement and the assigned spatial party position, our unit of analysis may be written analogously to the difference of two city-block spatial distances or utilities:

1. Subjective placement differences U_s are given by the negative distance of the voter's self-placements (v_i) and subjective, voter-specific party placements p_{ij} ($U_s = -|v_i - p_{ij}|$).
2. 'Objective' placement mismatch U_o is given by the negative distance of the voter's self-placements and 'objective', unified party positions p_j ($U_o = -|v_i - p_j|$).
3. The difference between subjective and 'objective' placements ΔP is, in turn, given by

$$\Delta P = U_s - U_o = -|v_i - p_{ij}| - \left[-|v_i - p_j|\right]$$

Note that we label the difference of voter-specific and mean party locations (p_{ij} and p_j) relative to the voter v_i placement differences with ΔP. Actual values of ΔP yield some first insight into potential projection effects and differentiate between three alternative scenarios:[6]

1. $\Delta P = 0$: if ΔP equals 0, the voter's assessment of a political party equals the 'true' position and projection bias is absent by definition;
2. $\Delta P > 0$: if ΔP is positive, subjective party placements yield a higher utility than mean party positions and this *may* indicate projection bias (assimilation);

3. $\Delta P < 0$: if ΔP is negative, utilities based on 'objective' placements exceed subjective ones and this *may* also point to projection bias (contrast).

There are many potential reasons why voter-specific party placements may differ from 'objective' positions of political parties. Projection bias, be it assimilation or contrast, is only one possible explanation. The formal identification of projection bias requires one to relate the placement differences ΔP to some measure of party attachment and to compute the marginal effect of party identification on placement differences. Let us therefore assume that we obtain an indicator A for the attachment of the voter to any of the parties contesting in an election. Lower (higher) values of A indicate that she opposes (likes) this specific party. If there is projection bias, placement differences ΔP and party attachment A should be positively related: $\beta_1 > 0$: $\Delta P = \beta_1 A_{ij} + \beta_0$. The corresponding marginal effect is given by $\partial \Delta P / \partial A_{ij} = \beta_1$. Projection effects are only present when $\beta_1 > 0$. The intersection of the upward-sloping predicted effects with the zero line ($\Delta P = 0;\ U_s = U_o$) by definition separates contrast and assimilation.

Describing projection effects

We begin the empirical part with a justification of our case selection and briefly introduce as well as review the data at hand. The subsequent section turns to descriptive empirical findings about placement differences and projection bias regarding the two dimensions that define our political space, the left–right and the European integration scales.

Data and data sources

In this article, we present a secondary analysis of the EES data. The rich information gathered by the PIREDEU questionnaires includes data on self- and party-placements on the left–right and European integration dimension, a fine-grained indicator of party attachment, and some contextual variables which will be used to explore the causal determinants of projection effects.[7]

The PIREDEU questionnaires explore voter self-placements on a left–right and a European integration scale ($v_i \in [0, 10]$). The voters are required to place each party contesting in the EP elections on parallel programmatic scales ($p_{ij} \in [0, 10]$). This yields a total of four different scales which provide sufficient information to relate voter and party positions and to compute the principal dependent variables. Placement differences are given by the subtraction of 'objective' from subjective spatial distances which are computed separately for each dimension of political contestation: $\Delta P = -|v_i - p_{ij}| - [-|v_i - \bar{p}_{ij}|]$.

Turning to the principal independent variable, the EES include a fine-grained indicator for attachment to any of the platforms competing in the EP elections. Respondents are asked to indicate the probability with which they would 'ever' vote for a certain party on a scale ranging from 0 ['not at all probable'] to 10 ['very probable'] yielding the 'propensity to vote' indicator (PtV). The empirical distribution of the PtV is, in statistical terms, strongly skewed to the right. This implies that more than 40 percent of those who are asked whether they would consider casting a vote for a specific political party promptly indicate that they are 'not at all' willing to do this (with PtV at or close to zero). While the remaining categories are about evenly distributed, the distribution has two additional, but less sharp-cut, peaks that indicate neutrality (with PtV at about five) and extremely favourable evaluations of the respective party (with PtV at or close to ten).

At this point, we need to specify a brief caveat and justify our use of PtV. Previous research has often been criticized for evaluating PtV instead of actual vote choice and thus confusing political preferences with political behaviour. While this survey question provides a lot of empirical variation, it is not always obvious what the probability to *ever* vote for a certain party actually means. This is even more complicated regarding the EES data given that empirical tests have shown that stated PtV is consistent with vote choice in national rather than EP elections (cf. Van der Brug et al., 2007). On these grounds, we believe that it is fundamentally flawed to use PtV, for example, as a proxy for stated vote choice. In contrast, we believe that it is justified to use PtV as a self-contained variable which is not modelled as a proxy for vote choice. For all practical purposes, the PtV item captures long-term attachments to political parties, provides an indicator of preferences rather than political behaviour, and is essentially merely another feeling thermometer which offers a fine-grained perspective on party attachment.

In this analysis, building on PtV as a more fine-grained measure of party attachment, A provides additional information and goes beyond preceding studies that persistently employed some kind of binary distinction between supporters and non-supporters of a political party. This kind of gross measure may well be sufficient for the analysis of two-party competition in US presidential and congressional elections. In multi-party systems, we believe this is an all too large simplification of the more complex and graduated preference structures. When there are multiple political parties competing, as in any member state of the European Union (EU), lumping a variety of different voters who possibly hold very different partisan and political preferences into the broad and diffuse category of 'non-supporters' does not make much sense, potentially introduces bias to our measurement, and is likely to decrease the empirical significance of contrast effects.

In order to facilitate the statistical analysis, the complex data matrix given by the PIREDEU dataset is transformed into a stacked, alternative-specific layout. Altogether, the dataset includes 27,069 individual voters, more than 155,000 party placements on the left–right dimension, and almost 130,000 party placements regarding European integration. The PIREDEU data provides sufficient empirical information to explore projection effects within the 27 national segments and for a total of 198 individual parties competing in the national segments of the 2009 elections to the EP.

Measures of placement differences

We begin the empirical part with some descriptive evidence on the average mismatch of subjective and 'objective' party placements on either political dimension. Note that the label 'mismatch' refers to absolute values of placement differences ΔP which algebraically boil down to the absolute mismatch of voter-specific and mean party placements: $|\Delta P| = |U_s| - |U_o| = |p_{ij} - p_j|$. We provide a brief descriptive summary of the distances between voter-specific placements and mean positions on both dimensions. The average mismatch on the left–right dimension is lower than on the European integration dimension. In the PIREDEU dataset, voter-specific party positions are on average about 1.64 scale points off the mark on the left–right dimension; the average mismatch amounts to 1.92 points on the European integration scale. The empirical distributions on both dimensions are positively skewed; in other words, the empirical values are concentrated on the left-hand side of the distribution, and the right-hand tail is considerably longer. More than fifty percent of the misplacements lie below 1.26 for the left–right and below 1.56 for the European integration scale.

The higher precision of left–right placements corresponds to the significantly higher salience of the encompassing left–right dimension. Political parties tend to send clear signals on their overall political orientation and voters are sufficiently informed about these positions. In contrast, party positions towards European integration are considerably less salient, not communicated explicitly, and voters often hold very little knowledge about EU politics. In sum, most voters are able to make sense of spatial party competition and to agree to a very significant extent on placements which closely resemble their 'true' positions which are proxied by the mean placements over all respondents. There is certainly also some visible disagreement concerning party placements, and we have to analyse these first findings in more detail to ascertain whether these differences occur due to projection effects or other factors such as limited cognitive resources and low salience of European integration matters.

Presence and magnitude of projection effects

As indicated before, we model the effects of party attachment A on placement differences ΔP to capture the sufficient conditions for projection effects. Given that voters are nested in party alternatives which are, in turn, nested in national contexts we apply hierarchical linear models with random intercepts at the party and voter levels and an additional random coefficient on party attachment at the party level.

There are, in principle, two alternative ways to estimate the descriptive model: PtV, our indicator of party attachment, is captured by an eleven-point scale and could either be treated as (1) a categorical or as (2) an interval variable. Building on this model, Table 1 provides a brief overview of projection effects on either dimension of the political space. We converted the fixed part of the hierarchical model to

Table 1. Describing projection effects; PtV as a categorical predictor

	Left–right		European integration	
Party attachment [PtV_{ijk}]	$\Delta\hat{U}_{lr}$	σ_{lr}	$\Delta\hat{U}_{ei}$	σ_{ei}
0	−1.2144	(0.0465)	−0.5459	(0.0429)
1	−0.7352	(0.0527)	−0.3125	(0.0550)
2	−0.5063	(0.0514)	−0.1692	(0.0522)
3	−0.2732	(0.0524)	−0.0503	(0.0536)
4	−0.1570	(0.0541)	0.0547	(0.0560)
5	0.0635	(0.0535)	0.1700	(0.0540)
6	0.1376	(0.0585)	0.2108	(0.0619)
7	0.1826	(0.0605)	0.3394	(0.0646)
8	0.2669	(0.0624)	0.3463	(0.0669)
9	0.3691	(0.0695)	0.3638	(0.0776)
10	0.6642	(0.0674)	0.4245	(0.0733)
N	128,145		105,813	

Notes: The dependent variables capture the placement differences on the left–right (ΔP_{lr}) and on the European integration dimension (ΔP_{ei}).

average predicted effects across all country and party contexts. Beginning with left–right, a voter tends to pull parties she strongly prefers by 0.66 scale points ($\sigma_{lr} = 0.067$) toward her personal position (assimilation; PtV = 10). On the other hand, she tends to push parties she strongly opposes by more than 1.21 scale points ($\sigma_{lr} = 0.047$) away (contrast; PtV = 0). Turning to the second policy dimension, European integration, projection effects are considerably weaker. Voters tend to pull strongly favoured parties by about 0.42 scale points ($\sigma_{ei} = 0.073$) toward their personal preference (assimilation) and push disfavoured ones about 0.55 points ($\sigma_{ei} = 0.043$) away (contrast).

The empirical estimates confirm highly significant effects of PtV on placement mismatch and, thus, meaningful projection effects regarding both left–right and European integration. The results obtained by the categorical model also corroborate the asymmetry hypothesis and provide a stronger contrast than assimilation effects on both political dimensions. Apparently, the magnitude of projection effects on the left–right is higher than on the European integration dimension and this reinforces our presumption that individuals feel a higher need to compensate cognitive inconsistencies when political issues are more salient.

The second option, namely treating PtV as a continuous variable, is more convenient regarding the ease of statistical estimation and the inclusion of interaction terms. Admittedly, this strategy involves assumptions about the linearity of parameters which are not supported by theory and may not be tenable. Recall that the previous model indicated substantive non-linearities of projection effects on both

Table 2. Describing projection effects; PtV as a continuous predictor

	Left–right		European integration	
	ΔP_{lr}		ΔP_{ei}	
Fixed effects				
Party attachment [ln PtV$_{ijk}$]	0.7076***	(0.0169)	0.3958***	(0.0204)
Constant	−1.2230***	(0.0483)	−0.5459***	(0.0469)
–Random effects–				
1: Country level				
RI [$\sqrt{\psi_k}$]	0.1922***	(0.0466)	0.2064***	(0.0384)
2: Party level				
RS [$\sqrt{\psi_{j,1}}$]	0.2178***	(0.0134)	0.2558***	(0.0169)
RI [$\sqrt{\psi_{j,0}}$]	0.4109***	(0.0240)	0.2940***	(0.0214)
3: Voter level				
Residual [$\sqrt{\theta}$]	1.9647***	(0.0044)	2.4755***	(0.0054)
N	128,145		105,813	
log. lik.	−269224.2756		−246455.6659	

Standard errors in parentheses.
*$p < 0.05$, **$p < 0.01$, ***$p < 0.001$.
Notes: The dependent variables capture placement differences on the left–right (ΔP_{lr}) and European integration dimensions (ΔP_{ei}). We report the results of hierarchical linear models with random effects at the country and party levels. RI indicates random intercepts, RS indicates random slopes and both give the standard deviation of the random effects at the respective level. The individual variance components are assumed independent, and all covariances are set to zero.

dimensions that constitute our political space. Because contrast effects appear to be much stronger than assimilation, we have decided to use the natural logs of PtV as our key independent variable.[8] Table 2 presents two hierarchical models that also substantiate the presence of projection effects on the left–right and the European integration dimension. The empirical findings reveal a positive, statistically significant, and substantively meaningful effect of logged PtV on placement differences regarding both dimensions.

Figure 1 illustrates these descriptive inferences graphically and provides a comparison to previous findings. The subpanels for left–right and for European integration plot the predicted effects from both models along with their respective confidence intervals. The thin continuous line and the grey-shaded confidence interval refer to the categorical predictors (cf. Table 1), the thick dashed line displays the predictions from the continuous model, and the thin dashed lines mark the associated confidence interval. The fact that both predictions are almost congruent and identical corroborates our confidence in the descriptive inferences and in the specification of *logged* PtV as the principal independent variable. Our confidence in

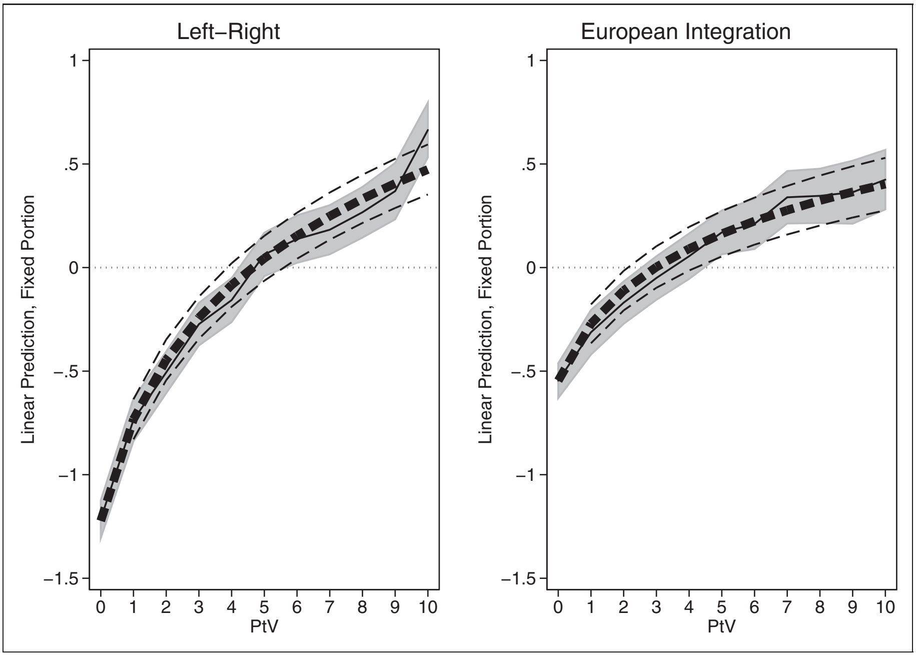

Figure 1. Comparing projection bias in the European political space.
Notes: The subpanels reproduce the mean trajectories for projection bias on the left–right (left-hand side) and on the European integration dimension (right-hand side). There are two statistical expressions: (1) the thin solid line and the grey-shaded area specify the predicted values and their 95 percent confidence interval from the categorical specification in Table 1; (2) the thick dashed line and the thin dashed lines represent the predicted values and confidence intervals from the continuous specification in Table 2. The dashed horizontal line ($\Delta p = 0$) separates contrast (below) from assimilation effects (above).

these descriptive inferences is further reinforced by a series of additional robustness checks which we cannot fully document here due to space constraints.[9]

The descriptive inferences reveal three principal insights into the nature of projection effects:

1. Projection bias affects party placements both in a statistically significant and a substantively meaningful manner.
2. Projection bias is more effective on the left–right than on the European integration dimension. This finding confirms the idea of classical contributions to social psychology that voters try to work around perceived inconsistencies the more the specific issue matters to them, and most voters perceive the encompassing left–right dimension as much more salient than European integration. In contrast, if we measured the persuasion of voters by advertised party politics, we would be likely to find higher levels on the by far less salient European integration dimension.

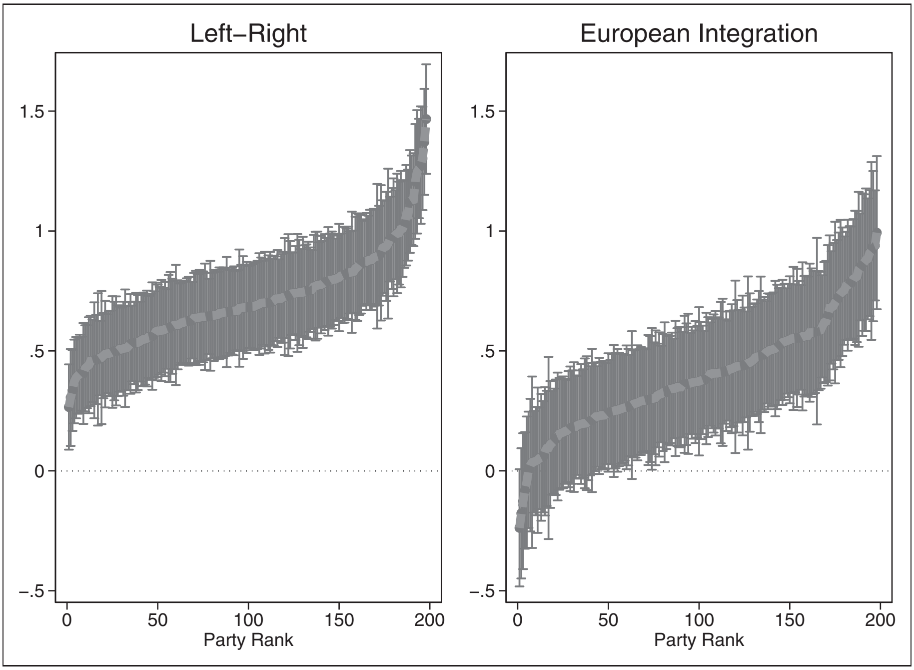

Figure 2. Party-specific random slopes on the left–right and European integration dimension in European Parliament elections.
Notes: The subpanels reproduce the election- and party-specific slopes on PtV and their respective 95 percent confidence intervals.

3. The novel operationalization of party attachment with a straightforward focus on the PtV explicitly refers, in stark contrast to the findings of previous studies, to the dominance of contrast over assimilation effects.

To probe more deeply into the context dependency of our findings, we now examine the random effects part of our models. The descriptive inferences in Table 2 stipulate random slopes on logged PtV which allow projection effects to vary across heterogeneous national and party contexts. Likelihood-ratio tests confirm that there is considerable empirical heterogeneity and indicate that random coefficient models yield a far better fit than simpler random intercept models.[10]

Figure 2 presents a more detailed assessment of causal heterogeneity at the party level. Note that the total projection effect is given by the estimated coefficients β plus the respective party-specific predictions for the random slopes ζ_j. While empirical evidence for projection effects on the left–right scale is robust and stable, the specific magnitude of assimilation and contrast varies considerably. Among the 198 political parties included, the predicted slopes on logged PtV range from a low of 0.26 up to 1.39. The empirical findings regarding European integration are similar, but effects are considerably weaker. For as many as 191 of the 198 political parties we estimate an overall positive slope which indicates the presence of projection effects. The individual predictions of party-specific coefficients vary between -0.41 and 1.02.

Explaining projection effects

While we have presented reliable and robust descriptive evidence for the presence of projection effects, we still lack empirical evidence on whether assimilation and contrast occur independently of different contexts or whether these phenomena are actually caused by specific contexts. In this section, we first introduce a distinction between unsystematic and systematic projection effects and discuss potential explanatory variables. Subsequently, we demonstrate that assimilation and contrast are unsystematic rather than systematic phenomena and are not linked to key contextual determinants at the voter, party, or national level in any robust fashion.

Unsystematic and systematic projection effects

Regarding the causal inferences, we distinguish between unsystematic and systematic projection effects. We have, roughly speaking, less to worry about when projection effects are unsystematic. Without the important role of contextual and systematic determinants, the presence and the magnitude of assimilation and contrast effects are exclusively defined by the evaluation of a specific party alternative and an i.i.d. random error. Systematic projection effects in contrast warrant additional concerns. When assimilation and/or contrast effects are systematically conditioned by contextual effects at the voter, party, or country level, our voter-specific party placements are not only systematically distorted, but also much more difficult to correct or rescale.

Because the descriptive inferences have shown the empirical asymmetry of projection effects and indicated that contrast is more effective than assimilation, we have decided to split our empirical sample and to analyse the determinants of assimilation and contrast in separate samples. Effectively, we focus on one dataset of cases that fulfills the necessary condition for assimilation ($\Delta P = U_s - U_o > 0$) and another that meets the necessary condition for contrast effects ($\Delta P < 0$). Cases of 'leapfrogging' have again been removed from the analysis. To yield comparable results and facilitate the interpretation of our findings, we use absolute values of our dependent variables. Therefore, they indicate placement mismatch $|\Delta P|$ so that lower values refer to potentially weak, while higher values indicate potentially strong, levels of assimilation or contrast.

To build our models, we regress the absolute values of placement mismatch $|\Delta P|$ on a set of explanatory variables and on the interactions of these independent variables with logged PtV: $|\Delta P| = \beta_0 + \beta_1 \mathrm{PtV}_{ijk} + \beta_2 \boldsymbol{X} + \beta_3 \mathrm{PtV}_{ijk} \boldsymbol{X}$. This is done separately for assimilation and contrast and for the left–right and the European integration dimension. Note that an association of absolute placement mismatch with a contextual variable does not sufficiently define projection effects, but rather indicates whether a contextual variable facilitates or exacerbates the correct placement of a political party.

To assess the effects of contextual variables on assimilation and contrast in a more straightforward fashion, we interact contextual characteristics $\boldsymbol{X}$ with the key indicator of party attachment PtV_{ijk}. The marginal effects of $\boldsymbol{X}$ indicate the impact of context variables on assimilation and contrast. We consider projection effects to be systematic if (and only if) the interaction of PtV with an explanatory variable is statistically significant and substantively meaningful. In contrast, projection effects may be considered unsystematic and independent of the electoral contexts when there are no statistically *and* substantively noteworthy interactions. Ultimately, we add random intercepts at the party ($\zeta_{0,j}$) and country ($\zeta_{0,k}$) levels and also include a random slope on logged PtV at the party level ($\zeta_{1,j}$) in order to control for the effects of unmodelled contextual variables:

$$\begin{aligned}\Delta \mathrm{P} &= \beta_{0,jk} + \beta_{1,jk} \ln \mathrm{PtV}_{ijk} + \beta_2 \ln \mathrm{PtV}_{ijk} \cdot \mathrm{X}_i + \beta_3 \ln \mathrm{PtV}_{ijk} \cdot \mathrm{X}_j + \beta_4 \ln \mathrm{PtV}_{ijk} \mathrm{X}_k + \varepsilon_{ijk} \\ &= [\beta_0 + \zeta_{0,j} + \zeta_{0,k}] + [\beta_1 + \zeta_{1,j}] \ln \mathrm{PtV}_{ij} + \beta_2 \ln \mathrm{PtV}_{ijk} \cdot \mathrm{X}_i + \beta_3 \ln \mathrm{PtV}_{ijk} \cdot \mathrm{X}_j \\ &\quad + \beta_4 \ln \mathrm{PtV}_{ijk} \mathrm{X}_k + \varepsilon_{ijk} \end{aligned} \tag{1}$$

Having defined systematic and unsystematic projection effects, we will now review and operationalize a number of potential causal effects at the individual, the party, and the national level and present testable hypotheses regarding (1) their direct effects on the overall magnitude of placement differences and (2) their interaction with logged PtV which indicates their modifying impact on, respectively, assimilation and contrast.

Voter sophistication. We start with voter-specific causes of projection bias and discuss whether some voters are more likely to misplace parties and/or more likely to project their own positions upon preferred parties than others. The voters' varying levels of political information and sophistication should correlate with the capability to process political information and to locate parties within a political space. Preceding research has demonstrated that educated, informed, and sophisticated voters are more likely to base their electoral choice upon the evaluation of policy positions.

Each of the subsequent hypotheses comes in two parts and includes suppositions about the effects of contextual variables on the absolute level of mismatch $|\Delta P|$ (the main effects) and about the modifying impact of the variable on, respectively, assimilation or contrast (the interactive effects):

(i) Regarding the absolute level of mismatch, our voter sophistication hypothesis posits that sophisticated voters are, by and large, more likely to place political parties close to their 'true' position.
(ii) With reference to the impact of voter sophistication on both assimilation and contrast, we posit that sophisticated voters ascribe a higher salience to issues,

tend to rely more frequently on issue voting, and are thus more likely to shift party positions either away from or towards their own policy position.

The PIREDEU survey modules allow us to construct an index of political sophistication, since the questionnaires include a total of seven items on the factual political knowledge of both domestic and European politics. Our indicator of political knowledge simply adds up the number of correctly answered knowledge questions, therefore ranging from zero to seven (every question answered correctly).

Voter extremism. Individual ideological or programmatic orientations may affect mismatch, assimilation, and contrast. We assume that extremist voters tend to interpret and understand ideological and/or programmatic scales differently from centrist voters:

(i) Party placements of extremist voters will therefore be more 'off the mark' than those of moderate voters with regard to the absolute mismatch of subjective and objective party placements $|\Delta P|$.
(ii) We further assume that these differences in the interpretations of spatial dimensions also affect the specifics of projection bias and increase the magnitude of assimilation and contrast effects.

In terms of left and right, we consider voters extremist when they lean either to the very left or to the very right, and we operationalize left–right extremism with the squared values of the centered left–right placements ($[v_i^{\mathrm{lr}}]^2$). In terms of European integration, we look at voter extremism via squared voter self-placements on the European integration scale ($[v_i^{\mathrm{ei}}]^2$).

Party extremism. Possible determinants of projection effects are not restricted to the voter level, but are also considered with regard to the issue positions of political parties and basic features of party competition. Usually, when parties move towards the political extremes, their position is clearly visible, their extremist ideology is discussed in public discourses, and their positions may be easily discerned from those of the mainstream parties:

(i) Voters will therefore have less problems agreeing on the positions of extremist than on those of mainstream parties, and
(ii) Party extremism will increase the magnitude of projection effects.

For instance, voters who like an extremist party tend to rank it not as extreme as voters do who do not like or even strongly dislike it. We assess party extremism with the squared values of the average, centered party placements by all respondents on the left–right ($[p_j^{\mathrm{lr}}]^2$) and on the European integration dimension ($[p_j^{\mathrm{ei}}]^2$).

The ambiguity of party positions. We continue with features of political parties and inspect the clarity of programmatic party positions. Quite often voters simply lack sufficient information to place parties on issue scales, because these parties deliberately present ambiguous policy positions. A broad range of contributions have shown that parties that blur their issue positions may actually be better off (cf. Shepsle, 1972; Tomz and van Houweling, 2009; Rovny, 2012). When parties present clear, explicit, and unified positions, it will be harder for voters to ignore these prevalent signals and then to project their own preferences upon party platforms:

(i) With regard to the main effects of party unity and clarity, we hypothesize that clear and unified party positions tend to reduce the average mismatch of subjective and 'objective' party placements.
(ii) We also posit that the ambiguity of programmatic party positions will increase and that the unity and clarity of party positions will diminish the extent of assimilation and contrast. Programmatic party ambiguity is measured by the standard deviation, σ_j, of party placements on the left–right and European integration scales. Low scores of σ_j indicate unified party positions, and higher values indicate ambiguous party positions.

Incumbency. Projection bias may be causally related to incumbency. Parties in national government have a high degree of visibility in the media and in public discourse. Their track record in office usually provides sufficient information to locate incumbent parties on a number of issue dimensions. In contrast, opposition parties often demand and/or propose substantive political change that is often only vaguely defined. We thus expect that voters generally have more difficulties in accurately placing opposition than government parties and that opposition parties are more likely to be evaluated with high levels of assimilation and contrast, because they lack a visible track record and are regularly associated with decisive, but often insufficiently defined, policy change. In the empirical analysis, incumbency is assessed by the dummy variable G_j. Data on the incumbency status on the election date was compiled from the ParlGov database and merged with the PIREDEU dataset (cf. Döring, 2013).

Fragmentation of party systems. We now shift our focus from individual parties to systemic features of party competition. In terms of overall placement mismatch, it will be easier for voters to process information and to correctly place political parties on competitive issue dimensions when there is only a limited level of party system fragmentation. However, focusing more explicitly on projection effects, a lower number of parties opens up more room for manoeuvre and for shifting individual parties within the political space. The concentration of party systems also allows for the emergence of more clear-cut, manifest patterns of party evaluation. We thus hypothesize that a lower number of parties tends to allow for, on average, more accurate party placements, but also for more intense projection bias.

We measure the fragmentation of national electorates in EP elections by the effective number of parties (Laakso and Taagepera, 1979). Building on the vote shares gained by political parties, the index is given by $N_v = [\sum_{j=1}^{n} v_j^2]^{-1}$. Empirical data on the individual parties' vote shares that is required to compute fragmentation scores was also taken from the ParlGov dataset.

Polarization of party systems. Ultimately, we broaden our perspective to also cover features of party system polarization. Expectations about potential effects of polarization on assimilation and contrast may be considered from various angles:

(i) Significant levels of polarization among political parties indicate the politicization of an issue and the provision of crystallized and discernible party positions. We argue that clear programmatic divisions allow voters to clearly capture individual party positions and this helps to reduce the average mismatch of subjective and 'objective' placements.
(ii) In contrast, if polarization also induces bitter conflicts, voters may develop more uncompromising evaluations of political parties which, in turn, propel and reinforce projection effects. While polarization therefore reduces the average mismatch of subjective and 'objective' party placement, we expect that, at the same time, it increases the propensity to assimilate liked and to contrast disliked candidates.

We use the Dalton index to measure party system polarization (Dalton, 2008: 906). Generally speaking, the index builds on the average deviation of individual party positions from the mean party position, $\bar{p}$, and these differences are weighted by the respective parties' vote shares v_j:

$$D = \left[\sum_{j=1}^{n} v_j (p_j - \bar{p})^2\right]^{0.5}.$$

Given that we evaluate seven independent causal determinants on the mismatch of subjective and objective party placements and on assimilation and contrast, our theoretical argument is complex and involves both the main effects of the explanatory variables and their interaction with party attachment. For convenience, Table 3 wraps up the hypotheses at the voter, party, and national levels.

Empirical determinants of assimilation and contrast

Table 4 presents the empirical findings and summarizes our causal inferences on the determinants of assimilation and contrast. We have analysed the absolute values of mismatch for potential cases of assimilation ($\Delta P > 0$; Models 1 and 3) and contrast ($\Delta P < 0$; Models 2 and 4) separately. We also ran separate models for the left–right and the European integration dimension. Recall that evidence for assimilation and contrast is present whenever the difference of subjective and mean party placements systematically co-varies with the voters' propensity to cast a vote for a certain party.

Table 3. A summary of the causal hypotheses

Variable	(1) Mismatch $\|\Delta P\|$	(2a) Assimilation	(2b) Contrast
–Voter level–			
Sophistication	decrease	increase	increase
Extremism	Increase	increase	increase
–Party level–			
Extremism	decrease	increase	increase
Ambiguity	decrease	decrease	decrease
Incumbency	decrease	decrease	decrease
–Party system level–			
Fragmentation	increase	increase	increase
Polarization	decrease	increase	increase

Beginning with the individual level, voter sophistication tends to reduce placement differences so that informed voters regularly place parties much closer to their 'true' positions. Across the board, voters who hold extremist preferences or support extremist ideologies tend to place parties less accurately than centrist, mainstream voters, because their understanding of issue dimensions and programmatic competition obviously 'differs'.

Turning towards the party level, we find a clear association of programmatic party ambiguity with the average levels of misplacement. Parties which offer unified and clear programmatic positions are significantly less likely to be misplaced by the voters in any of the four models. In contrast, the findings with regard to the effects of party extremism are mixed: a majority of our models indicates that party extremism reduces placement error (Models 2–4), but there also appears to be conflicting evidence (Model 1).

Moving up levels, we find, contrary to our hypotheses, almost no evidence for the empirical consequences of party system features such as fragmentation and polarization. Across our four models there is no significant or substantively meaningful association of placement differences with the fragmentation of electoral party systems and no consistent and meaningful association with our polarization indicator either.

These results shed some light on the determinants of party misplacement. Yet, as indicated, assimilation and contrast are causal effects and cannot be directly observed. Differences among subjective and 'objective' party placements per se do not sufficiently identify projection effects. Instead, causal determinants are captured by conditions that modify the effect of party attachment on utility differences. With regards to the interactions of (logged) PtV with contextual variables from the three hierarchical layers of the model, the empirical evidence is much more ambiguous. There are only a few indications of systematic projection effects. The empirical estimates are either statistically insignificant, substantively meaningless, or differ

Table 4. Causal inferences on projection effects

	Left–right				European integration			
	Assimilation; $\Delta P > 0$ (1)		Contrast; $\Delta P < 0$ (2)		Assimilation; $\Delta P > 0$ (3)		Contrast; $\Delta P < 0$ (4)	
–Main effects–								
Party attachment [ln PtV_{ij}]	0.6927***	(0.1250)	−0.4080**	(0.1270)	−0.1235	(0.1479)	−0.1961	(0.1667)
Pol. sophistication [S_i]	−0.0601***	(0.0059)	−0.0511***	(0.0042)	−0.0652***	(0.0060)	−0.0187**	(0.0059)
Voter extremism [v_i^2]	0.0575***	(0.0010)	0.0304***	(0.0007)	0.0926***	(0.0010)	0.0250***	(0.0010)
Party extremism [p_j^2]	0.0752***	(0.0059)	−0.0694***	(0.0042)	−0.0178*	(0.0082)	−0.0754***	(0.0076)
Party unity [δ_j]	0.6399***	(0.0495)	0.7228***	(0.0357)	0.4265***	(0.0610)	0.9174***	(0.0538)
Incumbency [G_j]	0.0892*	(0.0453)	−0.0760*	(0.0336)	0.1155**	(0.0377)	0.0790*	(0.0369)
Fragmentation [N_v]	0.0026	(0.0170)	−0.0080	(0.0143)	0.0023	(0.0153)	−0.0030	(0.0120)
Polarization [D]	0.0113	(0.0319)	−0.0554*	(0.0264)	−0.0066	(0.0344)	0.0647*	(0.0284)
Constant	−0.7175**	(0.2362)	0.8773***	(0.1859)	−0.0760	(0.2690)	−0.2860	(0.2349)
–Interactive effects–								
ln $PtV_{ij}.S_i$	−0.0044	(0.0035)	−0.0071*	(0.0034)	0.0203***	(0.0038)	−0.0027	(0.0043)
ln $PtV_{ij} \cdot v_i^2$	0.0007	(0.0006)	−0.0013*	(0.0007)	−0.0072***	(0.0006)	−0.0002	(0.0007)
ln $PtV_{ij} \cdot p_j^2$	−0.0268***	(0.0036)	0.0362***	(0.0037)	0.0347***	(0.0056)	0.0623***	(0.0062)
ln $PtV_{ij} \cdot \delta_j$	−0.1589***	(0.0287)	−0.0619*	(0.0288)	0.0223	(0.0340)	−0.0887*	(0.0387)
ln $PtV_{ij} \cdot G_j$	−0.0692*	(0.0270)	0.0554*	(0.0275)	−0.0811***	(0.0237)	−0.0436	(0.0268)
ln $PtV_{ij} \cdot N_v$	−0.0116	(0.0080)	0.0028	(0.0081)	−0.0027	(0.0074)	0.0298***	(0.0081)
ln $PtV_{ij} \cdot D$	−0.0302	(0.0161)	0.0321	(0.0165)	0.0052	(0.0174)	−0.0325	(0.0196)

(continued)

Table 4. Continued

	Left–right				European integration			
	Assimilation; $\Delta P > 0$ (1)		Contrast; $\Delta P < 0$ (2)		Assimilation; $\Delta P > 0$ (3)		Contrast; $\Delta P < 0$ (4)	
–Random effects–								
1: Country level								
RI [$\sqrt{\psi_k}$]	0.0891***	(0.0300)	0.0875***	(0.0221)	0.0848***	(0.0247)	0.0427***	(0.0262)
2: Party level								
RS [$\sqrt{\psi_{j,1}}$]	0.1329***	(0.0104)	0.1385***	(0.0105)	0.0981***	(0.0103)	0.1112***	(0.0114)
RI [$\sqrt{\psi_{j,0}}$]	0.2142***	(0.0185)	0.1568***	(0.0130)	0.1502***	(0.0166)	0.1473***	(0.0153)
3: Voter level								
Residual [$\sqrt{\theta}$]	1.1061***	(0.0038)	1.2363***	(0.0036)	1.2451***	(0.0043)	1.4001***	(0.0047)
N	43397		58912		42250		44627	
log. lik.	−66258.1384		−96372.4656		−69421.4484		−78531.9015	

Standard errors in parentheses.

$^{*}p < 0.05$, $^{**}p < 0.01$, $^{***}p < 0.001$.

Notes: RI indicates random intercepts, RS indicates random slopes and both give the standard deviation of the random effects at the respective level. The individual variance components are assumed independent, and all covariances are set to zero.

between our models for assimilation and contrast or with regard to the two dimensions of political contestation.

In sum, the causal analysis has identified some empirical determinants of the absolute mismatch of subjective and 'objective' party placements. Particularly, low levels of voter sophistication, extremist voter ideologies, programmatic ambiguity, and discordance within political parties contribute to placement differences and errors. However, our analysis failed to point out any specific interaction between these (and other) context effects on the association of placement mismatch and the PtV. Projection effects thus appear to occur rather independently of socio-political and institutional contexts. Thus, assimilation and contrast are unsystematic rather than systematic effects.

Findings, remedies, and perspectives

The theoretical arguments and empirical findings in this article are important and consequential for a number of reasons: established hypotheses from political psychology are effectively corroborated by our analysis of party placements in the most recent wave of the EES. Party placements are systematically distorted by the individuals' PtV for a certain party, and these projection effects are much stronger on the salient left–right scale than on the less salient European integration scale.

Building on what is, to our knowledge, the most comprehensive and exhaustive comparative dataset on projection effects, we were also able to assess the robustness of our findings across the currently 27 member states of the EU, 198 political parties or lists competing in the 2009 EP elections, and two specific dimensions of political competition. While the well-established patterns of assimilation and contrast play out on both dimensions, our novel empirical operationalization, which builds on the fine-grained PtV as the key independent variable, reveals that contrast appears to outweigh assimilation effects in multiparty electoral competition.

Our analysis ultimately aimed to explore the causal determinants of assimilation and contrast. We find much more evidence for the lack of, than for the significance of, causal and contextual effects. By and large projection effects appear to be unsystematic and context-independent rather than systematic and context-dependent. In subsequent research, our descriptive and causal inferences may be utilized to develop rescaling techniques which correct for the distortion of idiosyncratic party placements by assimilation and contrast.

In the realm of practical politics, projection effects may easily affect an individual's vote choice and facilitate a spillover from non-policy to policy utilities. Whenever voters like a specific candidate or party due to non-policy considerations, these convictions may help to artificially fabricate more favourable policy utilities as well. As a result, the overall utility function which consists of policy and non-policy utility terms (cf. Adams et al., 2005) may favour that candidate or party even more, increase the PtV for that party, and help to change the voter's choice.

While projection effects are an interesting phenomenon per se, there are also significant consequences for models of electoral choice. Supporters of the directional voting camp have repeatedly argued that projection effects imply a contamination of proximity utilities with non-spatial considerations so that the significance of proximity voting may be overestimated and the impact of alternative models of spatial voting and of non-policy motives underestimated (Macdonald et al., 1991, 1998, 2001). The bias implied by projection effects potentially leads to biased inferences on the determinants of electoral choice.

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Notes

1. Rabinowitz and Macdonald refer to this as 'rationalization'. However, we will stick with the term 'projection effects', composed of assimilation and contrast, throughout this text. This terminology has previously been utilized by Merrill et al. (2001) and Adams et al. (2005) and it dates back to the earlier works in cognitive psychology cited above (cf. Heider, 1958; Sherif and Hovland, 1961; Wedell et al., 2007).
2. Regarding the limited perspective of this article, we cannot present a thorough review of the political psychology literature. Instead, Jon Krosnick (2002) provides an excellent overview of classical contributions to cognition theory and their application to the study of voter perceptions and projection effects.
3. While cross-sectional analyses often corroborate the existence of projection effects, the empirical evidence is also consistent with a whole gamut of alternative explanations, for example '(1) perspective effects, (2) policy-based evaluation and persuasion, (3) variation in candidate's issue statements (Rovny, 2012), and (4) the agreement principle' (Krosnick, 2002, 121–122).
4. There is an extensive debate concerning the merits and setbacks of various measures such as the analysis of party manifestos, computerized text analysis, expert surveys on party locations, or roll-call data. Due to space constraints, we cannot adequately summarize these controversies. A good starting point for this literature includes Marks (2007) and other articles in a special issue of *Electoral Studies* 1/2007.
5. Of course, no operationalization of 'objective' party positions fully captures the underlying latent concept. However, we like to check that separately for each measure and feel a little uneasy by estimating latent variables from divergent data sources. It is very difficult to understand the position scores produced by such a procedure, to assess options and limitations of these data, and to meaningfully apply these scores to real-world political analysis. We simply do not know enough about the properties of position scores that are extracted from a variety of different issues. Such a task would need to be addressed in additional publications.
6. At this point we hasten to add one additional caveat: there are a number of cases where the voter does not only pull a party towards her own position, but also reverses the ordering of party and voter positions ('leapfrogging'). Although the results do not significantly differ, we decided to remove all cases from our sample where the party position did not remain on the same 'side' either on the left–right and/or the European integration dimension. The extent of 'leapfrogging' amounts to roughly 16 percent of all respondents on the left–right dimension and approximately 23 percent on the European integration dimension.

7. Further details about the aims, the study components such as the voter and candidate surveys or the media and manifesto studies, and the methodological foundations of the PIREDEU project can be accessed on the internet: http://www.piredeu.eu.
8. Given that the original PtV scale ranges from zero to ten and that the logarithm of zero is not defined, we have shifted the scale by one point. Thus, our logged scale may take on values from ln 1 to ln 11.
9. For example, we substituted a number of alternative external proxies as yardsticks for 'true' party positions such as manifesto data and alternative expert surveys. These results are available via the web appendix.
10. Random intercept models (not shown here) are rejected in favour of random slope models. The statistically significant test statistics are $\chi^2 = 658.81$ for left–right and $\chi^2 = 406.28$ for the European integration dimension.

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