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Social cleavages and political dealignment in contemporary Chile, 1995–2009

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Abstract

There is abundant research on how social cleavages shape political preferences in developed countries with uninterrupted democracies, but we know less about this topic for middle income countries with recently restored democracies. In this analysis of the Chilean case, we examine with Latinobarometer survey data from 1995 to 2009 the evolution of social cleavages as shapers of political preferences (measured with a left–right self-placement scale). We find a general process of dealignment across time, indicated by the decreasing association between political preferences on the one hand, and class, religion and regime preferences on the other. We tentatively link dealignment at the mass level to the strategies pursued by political parties operating in a political and economic context that encourages ideological moderation and convergence to the centre. These strategies weaken the differentiated signals needed for sustaining an aligned citizenry.

Keywords

dealignment, Chile, social cleavages, vote choice/preference

Introduction

In 1989, Chileans elected their national political authorities by popular vote, ending the 17-year-long military dictatorship of General Augusto Pinochet and honouring the country's pedigree as one of the most robust democracies in Latin America. As Chilean democracy consolidated in the following years, scholars began addressing pressing questions. To what extent did social cleavages shape the political preferences of Chileans in the new democratic context? How did cleavages evolve as democracy consolidated and socio-economic modernization ensued? What might explain the observed changes?

Past research provided two answers to these questions. One was developed by Valenzuela and Scully (Scully, 1992; Valenzuela and Scully, 1997; Valenzuela et al., 2007). Inspired by Lipset and Rokkan's (1967) sociological model of party systems, they argued that political preferences in post-authoritarian Chile – including voting choices and ideological positions in a left–right scale – were essentially shaped by traditional religious and class cleavages. These cleavages were not new – they had structured political conflict in Chile since the mid-19th century (religion) and early 20th century (class).

Scholars such as Tironi and Agüero (Tironi and Agüero, 1999; Tironi et al., 2001) and Torcal and Mainwaring (2003), however, claimed that a new 'regime preferences' division between those who supported the Pinochet regime and those who opposed it had displaced traditional cleavages like class or religion. This division, which was epitomized by the 1988 plebiscite (in which Pinochet was voted out and the path for re-democratization opened), shaped the new political landscape and had an enduring impact on electoral preferences. Those supporting Pinochet favoured the centre-right coalition (*Alianza por Chile*) and those opposing him favoured the centre-left (*Concertación de Partidos por la Democracia*). Several studies based on cross-sectional survey data supported this claim (Agüero

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et al., 1998; Alvarez and Katz, 2009; Bonilla et al., 2011; Ortega Frei, 2003; Tironi et al., 2001).

In this article we examine the evolution of social cleavages in shaping Chileans' political preferences (measured as self-identification in a left–right political scale) between 1995 and 2009. We find that, for the entire period, both sociological and regime preference variables shape locations on the left–right scale. However, we also find a systematic reduction in the size of the marginal effects of multiple cleavage variables, namely education, religious denomination and regime preferences of respondents. We relate this trend to a general process of ideological convergence among Chilean parties, for which we provide some exploratory evidence. This convergence process can be understood as a progressive political learning whereby political parties adapted to the institutional and economic environment inherited from the Pinochet regime.

We contribute to the debate in four ways. First, by exploring a case of relatively recent democratic consolidation, we expand social cleavage research, which mostly focuses on cases with uninterrupted democratic rule since World War II (i.e. Western Europe and North America), but barely so in middle-income nations with recent democratic transitions. This allows us to consider a new division – that between supporters and opponents of a previous authoritarian regime, which for obvious reasons has not been explored in uninterrupted democracies. We also look at the more traditional class and religion cleavages.

Second, we test the widespread hypothesis about a generalized decline of social cleavages (Dalton, 2002; Inglehart, 1990) by looking at cleavage strength year after year – therefore not assuming linear patterns of evolution. This is consistent with an issue evolution perspective (Carmines and Stimson, 1989), which claims that the strength of cleavages varies in different directions according to the issues opportunistically activated by political elites.

Third, by using yearly data for a 14-year period (1995–2009) we present a truly longitudinal study of the evolution of Chilean cleavages. This is an improvement over past studies about Chile, which typically use cross-sectional surveys and therefore cannot assess whether the strength of cleavages increases or decreases across time (two partial exceptions are Torcal and Mainwaring [2003] and Raymond and Feltch [2012]). We use a single dataset (the Latinobarometer survey) containing comparable questionnaires and the same model specification across time.

Finally, we take into account the fact that not all Chileans express a preference on the dependent variable of our analysis (the left–right ideological self-placement scale). Consequently, we employ a Heckman selection model (Heckman, 1979), which allows us to simultaneously estimate individuals' propensity to express any ideological preference as well as their position on the scale.

We first review the literature on social cleavages and political agency and then describe the particularities of the Chilean political system. We then present our data, methods and results. In the discussion section we try to make sense of the results, considering party strategies and the institutional arrangements inherited from the Pinochet regime. We conclude and present pending research tasks.

Social cleavages and political agency

One of the most enduring debates in political sociology and political science revolves around social cleavages, i.e. the extent to which structural traits like gender, class, religion or ethnicity shape electoral choices and political ideologies. The basic idea is that people from different social groups systematically develop political preferences and make electoral choices that seem to advance their group interests, values and identities.

While few would discuss that the association between social categories and preferences exists under certain circumstances, a more pressing question is why cleavage strength varies across time or place. One answer comes from a sociological approach, which suggests that variations arise from changes in social-structural factors, such as inequality between and within groups, socio-economic modernization, value change and changes in group size (Manza and Brooks, 1999; Inglehart, 1977; Lipset and Rokkan, 1967). But because social structures typically change slowly, this approach alone cannot explain changes in cleavage strength that take place over short time periods – such as those we diagnose below for Chile. Thus, we emphasize a more dynamic 'political agency' approach, one which focuses on how political actors react to the political, economic and social setting in which they are embedded (Chhibber and Torcal, 1997; Evans and Tilley, 2012; Przeworski and Sprague, 1986).

According to this approach the associations between social categories and political preferences result from the choices of political actors embedded in particular contexts. Parties and politicians develop strategies for gaining the support of certain social groups. Groups respond to these appeals and increase their support towards the party, yet at the same time other groups feel alienated and therefore support disproportionately another party. This creates or deepens certain political conflicts and identities – though not necessarily intentionally (Hetherington, 2001; Mainwaring et al., 2013; Posner, 2004; Torcal and Mainwaring, 2003). Conversely, conflicts may be deactivated when parties develop catch-all strategies or when they move to the centre of the relevant axis of competition, because citizens may stop perceiving substantial differences in political supply (Enyedi, 2005; Kriesi, 1998).

One key aspect for determining party strategies is the institutional context. For instance, electoral rules may encourage parties to seek the median voter (Cox, 1997; Downs, 1957), and this may require downplaying some conflicts and identities while activating latent others.

The impact of institutional arrangements may not be immediate but require time, as actors learn how to operate within certain rules. Political learning helps in explaining why cleavage strength may change across time due to institutional factors, even if rules do not change. We suggest that this is the case regarding the impact of the binomial system in post-transitional Chile. Before presenting the data on Chile we review research showing how the political agency approach accounts for changes in cleavage strength regarding the three cleavages we consider in our analysis (class, religion and regime preferences).

Class and political preferences

Classic post-war electoral studies have shown that in most industrial societies the working class tended to support leftist parties while the middle and upper classes supported liberal and conservative parties (G. Evans, 1999; J. Evans, 2004; Knutsen, 2007; Lazarsfeld et al., 1948; Lipset, 1963; Manza et al., 1995). Social-structural explanations of variations in the class cleavage include the embourgeoisement thesis, the decline of labour unions, increasing occupational mobility (Dalton, 2002: 153; Evans, 2004: 56; Manza et al., 2005: 215) and cultural change (Inglehart, 1990).

Recent studies have emphasized the political agency model. Evans and Tilley (2012) found that the decline of class voting in Britain resulted from the Labour Party's move to the political centre in the 1990s and 2000s rather than from an increase in class heterogeneity. Chhibber and Torcal (1997) argue that the resurgence of class voting in Spain during the 1990s should be attributed to the policies of the governing PSOE. With the raising of unemployment benefits and increasing social expenditures, these policies increased workers' support for the PSOE and alienated the upper classes. For contemporary Latin America, Mainwaring et al. (2013) have found that class voting in Latin America is higher in countries with a strong and viable leftist candidate. This is because these candidates emphasize themes such as land reform, redistribution and social justice, which polarize the electorate along class lines.

Religion and political preferences

Religious identities are powerful in shaping political preferences (Manza and Brooks, 1999). Individuals with certain religious identities may perceive that a given political party furthers their interests, values or beliefs to a greater extent than others, thus supporting them disproportionately. These alignments may vary across countries and regions, and some scholars focus on social-structural factors in explaining them (Dalton, 2002: 161; Esmer and Pattersson, 2007: 499; Manza and Wright, 2003).

Other scholars consider the strategies and choices of parties and candidates. For instance, conservative or confessional parties may choose to emphasize moral issues with strong religious overtones when they feel threatened by liberal governments (Kalyvas, 1998), therefore mobilizing religious and alienating secular supporters (Mohseni and Wilcox, 2008: 211). As a reaction to religious embattlements, self-confident irreligious parties may become more openly secular, as the American Democrats did in the 1980s to face the arousal of the Republican-aligned Christian Right (Mohseni and Wilcox, 2008: 211).

Parties and politicians may also make strategic choices that weaken religious cleavages. European Christian Democratic parties originally mobilized voters on the basis of religious identities, but as they became catch-all centre or centre-right parties – as in Italy and Germany – they softened religious issues and attempted to attract non-Christian groups and younger voters (Manza and Wright, 2003: 299; Mohseni and Wilcox, 2008: 218).

Attitudes toward the authoritarian regime: A new divide?

Finally, we focus on the authoritarian–democratic division. This is absent in cleavage studies of consolidated democracies simply because they do not have a recent authoritarian past. Authoritarian regimes may leave a powerful legacy in their societies – a legacy that colours citizens' views of most political issues once democracy is restored. Specifically, we argue that citizens' political preferences may be shaped by their positions toward the previous authoritarian regime (be it in favour or against it). If they favoured the authoritarian regime and the latter positioned itself as rightist, they should see themselves as rightist and favour rightist parties – and vice-versa. The political agency approach suggests that parties will activate or downplay this division as a means of obtaining votes or other kinds of political advantage (e.g. internal cohesion). For instance, parties that adhere to the regime will activate the division when the population holds a positive image about it, yet will try to deactivate it when the regime loses legitimacy (Kitschelt et al., 2010: ch. 8; Moreno, 1999).

The Chilean case

Within Latin America, Chile is an interesting case because its social cleavages are supposed to be comparatively strong (Dix, 1989; Mainwaring and Scully, 1995), partially as a result of a party system similar to multiparty continental systems. Since re-democratization in 1990, the Chilean party system has revolved around two multiparty coalitions: the centre-left *Concertación por la Democracia,* composed of the *Partido Socialista* (PS), *Partido Por la Democracia* (PPD), *Democracia Cristiana* (DC) and *Partido Radical Socialdemócrata* (PRSD) and the centre-right *Alianza por el Cambio,* composed of *Renovación Nacional* (RN) and the *Unión Demócrata Independiente*

(UDI). The *Concertación* governed between 1990 and 2010, when it was defeated by the Alianza – currently in power. The peculiar binomial electoral system – with only two members being elected in each district – granted a similar number of legislators to both coalitions.

Historically, the Chilean class cleavage stemmed from the early development of a strong labour movement tied to leftist parties – a tie that was reinforced during the Socialist government of Salvador Allende (1970–1973) – and the coalescence of the industrial and land-owning classes around rightist parties for protecting their privileges. While some studies suggest that class should continue shaping political positions in the post-authoritarian period (Valenzuela et al., 2007), others claim the opposite (Torcal and Mainwaring, 2003). Yet the issue has not been settled because we lack a comprehensive exploration of the evolution of such cleavages across time.

The religious cleavage shaped the origins of Chile's first party system in the 1850s. The conflict stemmed from divergences within the political class regarding the influence of the Church on state and society (Scully, 1992; Valenzuela, 1995). But there is less agreement about the *current* strength of religious cleavages or their evolution across time. Some believe it is very influential (Valenzuela and Scully, 1997; Valenzuela et al., 2007), others claim the opposite (Torcal and Mainwaring, 2003), and a recent study suggests an increasing salience of religion (Raymond and Feltch, 2012).

The regime division pits those who favour the authoritarian regime of Pinochet against those who prefer democracy. Prior studies have found a strong association between attitudes toward Pinochet's regime and political preferences, with Pinochet supporters favouring rightist parties and self-identifications and opponents favouring the left (Tironi and Agüero, 1999; Tironi et al., 2001; Torcal and Mainwaring, 2003). These associations stem from the heavy consequences of the regime for Chileans. The regime was enduring (it lasted 17 years) and highly repressive (thousands of its opponents were tortured or assassinated). Moreover, it engaged in multiple market reforms that reduced the regulatory role of the state, privatized social services such as health, education and pensions, and increased the flexibility of the labour market. Additionally, Pinochet sent clear clues that his regime was a rightist one, e.g. he presented himself as saving the country from the Marxist left.

Data and methods

We use Chilean survey data from the Latinobarometer project between 1995 and 2009, which employs a probability sample of voting age citizens conducted every year with the exception of 1999. For reasons detailed at length in the online supplement, this is the best dataset for a longitudinal analysis of political preferences in Chile. We also provide in the supplement a brief description of the methodological details of the Latinobarometer surveys.

Dependent variable

We measure the political preferences of Chileans using respondents' self-placement on a scale ranging from 0 (left) to 10 (right). This scale is widely used in political behaviour research because it is shorthand to people's orientation 'toward a society's political leaders, ideologies and parties' (Mair, 2007: 207). Prior research shows that Chileans consistently order their political parties along the scale and that it represents a meaningful construct for a majority of the population (Fontaine, 1995; Harbers et al., 2012; Kitschelt et al., 2010: ch. 5). Moreover, the political scale has been previously used in research on social cleavages and is strongly correlated with voting choices not only internationally (Barone et al., 2007; Norris and Inglehart, 2004: 448; Mair, 2007: 218 f.), but also in Chile (Tironi et al., 2001; Torcal and Mainwaring 2003).[1] Measures of party preference or vote recount might capture electoral preferences more directly, but have high levels of non-response that hinder multivariate analysis (Morales, 2010). Also, voting behaviour has limitations for the study of social cleavages and political preferences (Barone et al., 2007). Because electoral choices are strongly influenced by contingencies of political supply (Alvarez and Nagler, 2000; Cox, 1997), many individuals vote for parties that do not reflect their true preferences. In this respect, the left–right scale presumably better reflects long-term ideological orientations.

Still, the left–right scale in Chile has the problem that a sizeable and increasing proportion of the population (around 30 percent in recent years) refuses to locate on it. Figure 1 shows the proportion of the population that mentions any position on the ideological scale between 1995 and 2009. After 1997, when this proportion peaked at 84 percent (according to the weighted data), it decreased to a record low of 63 percent in 2003. Thereafter the proportion of identifiers has remained around 70 percent. This proportion is too high to ignore and, as we show below, there are important differences between the population that mentions any position on the scale and the population that does not. This illustrates a classical selection problem where the observed sample is not a random subset of the entire sample (Achen, 1986). To address this problem we employ a Heckman selection model (Heckman, 1979) and simultaneously estimate the propensity of individuals to express any ideological preference as well as their position on the scale.

Independent variables

We measure the democracy–authoritarian divide with the following question: 'Which of the following statements do you agree most with? (a) Democracy is preferable to any other kind of government; (b) In certain situations, an

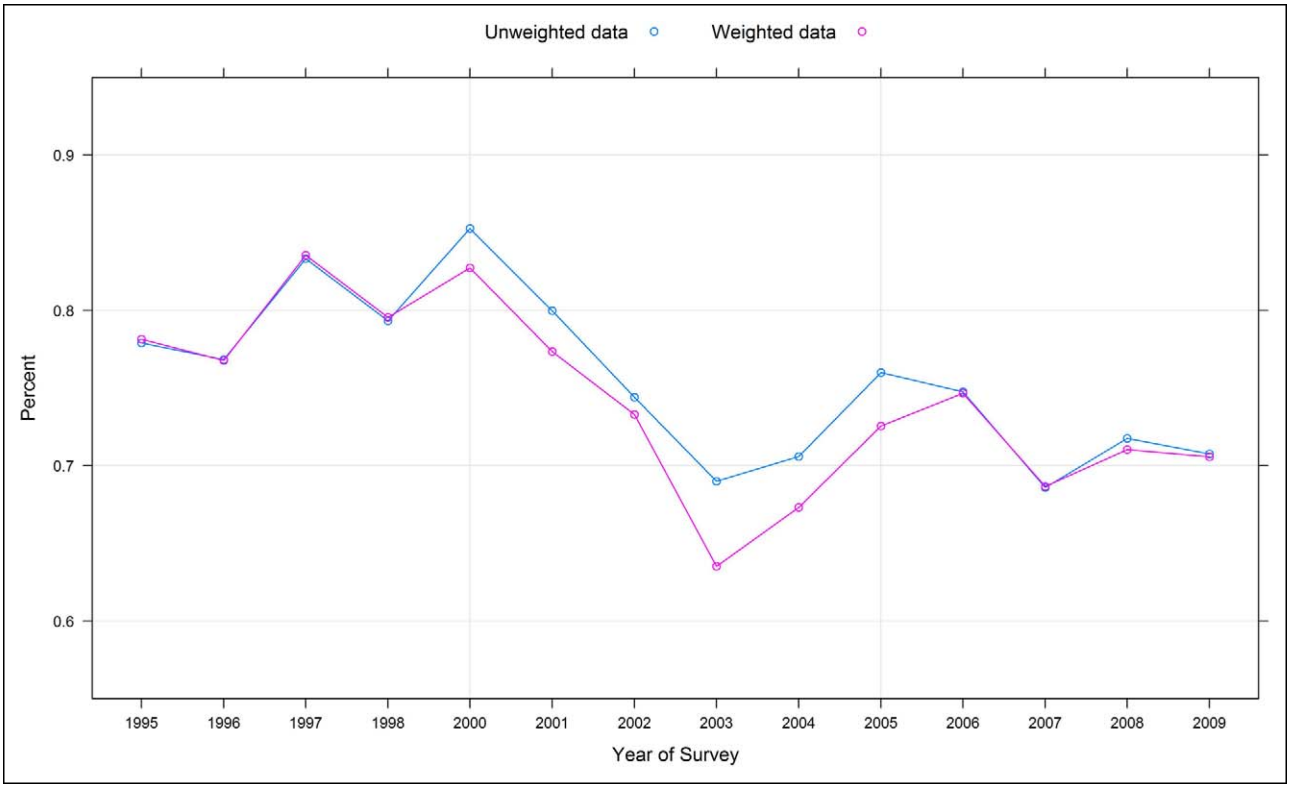

Figure 1. Proportion of the population that identifies with the left–right scale.

authoritarian government can be preferable; or (c) For people like me, it does not matter whether we have a democratic or a non-democratic regime'. Hunneus and Maldonado (2003) and Valenzuela et al. (2007) have convincingly argued that in Chile answers to this question reflect people's attitudes (in favour or against) the Pinochet regime rather than abstract appraisals about regime types. Therefore, those who prefer democracy favour the democratic regime inaugurated in 1990, while those who prefer authoritarianism favour the previous authoritarian regime. For our regression model we create two dummy variables that indicate the authoritarian (alternative b) and indifference choice (alternative c), with the latter also including the 'don't know' responses. Alternative (a) (full democrats) is the reference category.

We measure the class cleavage with two indicators of an individual's socio-economic position: the level of education (with eight categories ranging from illiterate to complete university degree, and treated as a continuous predictor) and a household goods index, which is an additional index that counts the number of goods each survey respondent reports possessing or has in his/her household.[2] We would have preferred a measure of household income, but this is not available in the Latinobarometer surveys. Nonetheless, this index is highly correlated with household income (in surveys of the *Centro de Estudios Públicos*, correlations range between 0.65 to 0.7 during different years; see section 4 online supplement for details).

The religious cleavage is captured through a religious denomination question. We introduce this variable in the statistical models with three dummies (Catholics, Evangelicals and a residual 'others' category, with people with no religion as the reference category). All estimates are calculated controlling for respondent's gender (dummy for male) and age (which we divided into five age-group dummies).

Lastly, to avoid identification problems in our Heckman regression model, we include a four-point interest in politics variable as an exclusive predictor of the selection equation. This variable is assumed to affect a respondent's propensity to locate on any position on the left–right scale but not the position they prefer. Exploratory analyses confirmed that the correlation between interest in politics and respondent's left–right position is significant (0.12), but also much smaller than that between interest in politics and respondent's propensity to locate on the scale (0.32).[3]

Results

To evaluate the effect of cleavages on ideological preferences we conduct two sets of analyses. First, we model respondent's propensity to mention a left–right position and their preferred position using the entire Latinobarometer pooled dataset. Given some missing variables during

Table 1. Heckman selection model for left–right ideological scale (pooled data).

	Selection Eq.	Outcome Eq.
Intercept	–0.748***	2.139***
	(0.081)	(0.17)
Male	0.132***	–0.162***
	(0.027)	(0.047)
26–35 years	0.15***	0.151**
	(0.041)	(0.07)
36–45 years	0.131***	0.222***
	(0.041)	(0.072)
46–55 years	0.204***	0.253***
	(0.046)	(0.079)
56–65 years	0.238***	0.354***
	(0.051)	(0.087)
66 years or more	0.136***	0.438***
	(0.052)	(0.094)
Education	0.08***	0.123***
	(0.01)	(0.018)
Household goods index	0.022***	0.101***
	(0.008)	(0.015)
Catholic	0.154***	0.765***
	(0.041)	(0.071)
Evangelical	0.129**	0.702***
	(0.053)	(0.096)
Other religion	0.037	0.331***
	(0.064)	(0.111)
Don't care about gov. type / Dk	–0.23***	0.621***
	(0.03)	(0.061)
Authoritarian gov. can be preferable	0.108***	1.926***
	(0.039)	(0.062)
Interest in politics	0.54***	
	(0.018)	
Inverse Mills ratio		1.299***
		(0.134)
Rho		0.566
Sigma		2.297
N obs / N censored	12900 / 3047	

***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.
Note: Model includes year effects for each survey.

a few applications, this accounts for a total of 11 annual surveys conducted between 1995 and 2009.[4] In this model we incorporate all the independent variables mentioned in the last section, as well as year-specific dummy variables for absorbing the impact of specific events and circumstances that took place during application of the surveys. The parameters from this model indicate the (partial) association of each independent variable and the left–right scale for the entire period between 1995 and 2009. Second, we apply this same model specification, with the exception of the survey-year dummy variables, to each survey separately. Then we calculate the 'marginal effects' for each of the socio-economic, religious and regime preference variables, and plot the estimates in order to capture the evolution of the association between the different cleavage variables and respondents' ideological preferences.

Results from the pooled model are given in Table 1. Given space constraints we cannot review the results from the selection equation, but we notice that with the exception of the 'others' religion dummy variable, all coefficients are statistically significant at a 99 percent level of confidence or higher. This reinforces the importance of accounting for the dependency between mentioning a position and location on the ideological scale.

The outcome equation also contains many significant estimates. With the exception of the gender dummy variable, all coefficients are positive, which indicates an increase towards a more right-wing position. Older respondents locate themselves more to the right than younger ones, and the coefficients grow monotonically as the groups grow older. Respondents with more formal education and household goods also locate to the right of those with less of each type of resources. Catholics, Evangelicals and those identified with another religion are more rightist than irreligious respondents. The estimates for Evangelicals and Catholics are particularly stark, leading a positive change equivalent to around three-quarters of a point on the scale.

Lastly, the two dummy variables reflecting regime preference also show a positive and significant association with the left–right scale. For instance, those who mentioned that an authoritarian government can be preferable are almost two points more rightist (in the 1–10 scale) than those who always prefer democracy.

In sum, several social and political divisions have significant effects on Chileans' ideological preferences over the entire 1995–2009 period. We claim that these results provide simultaneous support for both the classical sociological notion that emphasizes the role of traditional cleavages, particularly social class and religious denomination, and for the approach that emphasizes the relevance of the division between supporters and opponents of the military dictatorship. Moreover, the Chilean case is consistent with the issue evolution perspective (Carmines and Stimson, 1989), which is that alternative social and political divisions do not necessarily substitute, but can complement each other (see also Raymond and Feltch, 2012).

While the pooled model provides valuable information, it also hides important levels of heterogeneity in the predictive strength of the cleavage divisions across time. Thus, we estimated separately for each survey the same model specification given in Table 1 (though excluding the survey dummy variables), and calculated the 'marginal effects' of each cleavage variable. We show the results in Figures 2, 3 and 4. Each figure plots the marginal effect of the specified independent variables along with their 95 percent confidence intervals (calculated via non-parametric bootstrap) and adds a local fit curve that makes the temporal patterns more interpretable.[5] We provide the full details of all estimated models, plus a comparison with OLS estimates, in the online supplement.

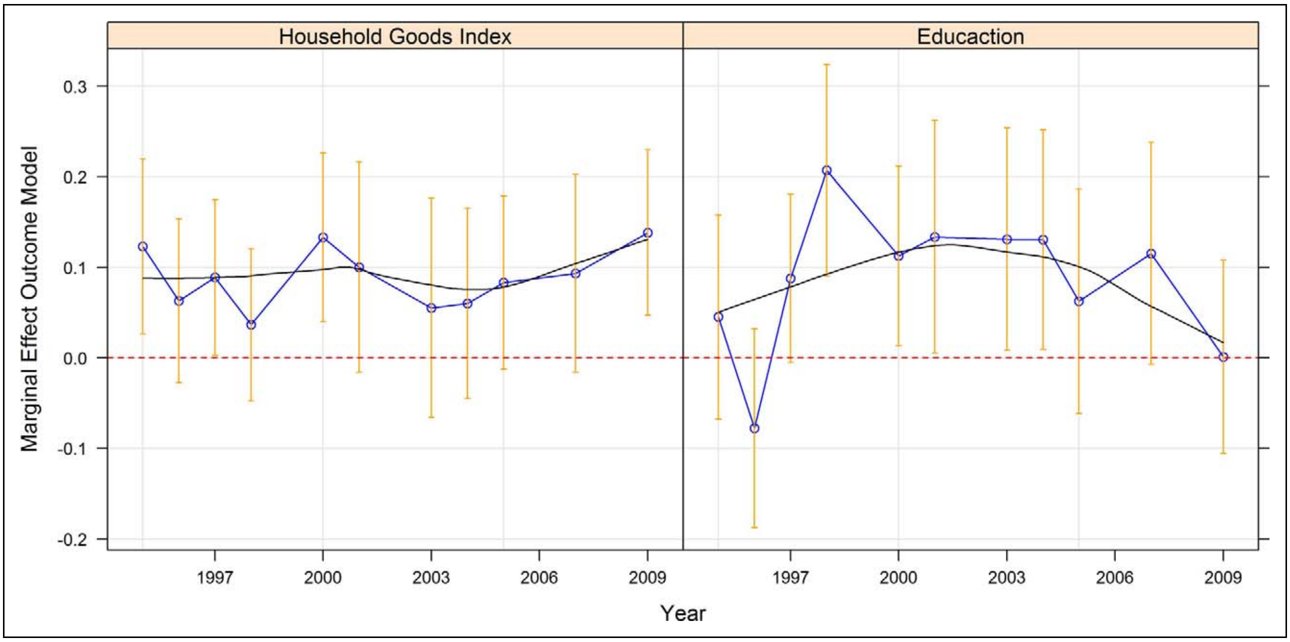

Figure 2. Evolution of socio-economic marginal effects applied to each available year of Latinobarometer data.

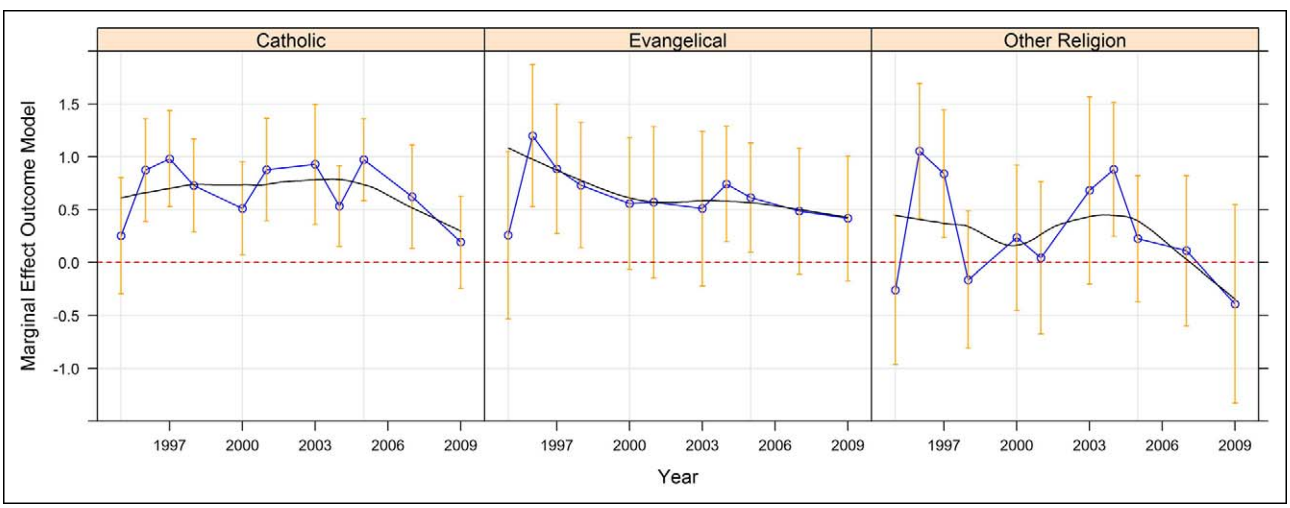

Figure 3. Evolution of religious group marginal effects applied to each available year of Latinobarometer data.

The evolution of the socio-economic indicators is shown in Figure 2 showing that during the last few years a simultaneous modest increase in the magnitude of the marginal effect of the household goods index and a sharp decrease in the marginal effect of education. The decrease of this last variable has been relatively systematic since 1998 up to the most recent survey, but sharpens after the 2004 survey, though the 2007 survey registers a small recovery. Note that in the last survey the marginal effect of education is non-significant and the point estimate is very close to zero. This sharp decline is preceded by four years of a relatively stable and significant estimate (2000 to 2004), which in turn is preceded by a more unstable period registering a very sharp increase in the association between education and self-location on the political scale. In contrast to this movement, the marginal effect of the household good index registers a weak but stable decrease during the first eight years of available data (and significant only on some occasions). After this the trend reverses, and from 2004 onwards the marginal effect becomes larger year after year. Its marginal effect becomes significant in the 2009 survey.

Figure 3 shows the results for the religious denomination variables. We can see again a reduction in the marginal effect of a cleavage variable. Being Catholic is consistently associated across the period with a more right-wing position on the left–right scale, but this divergence has been

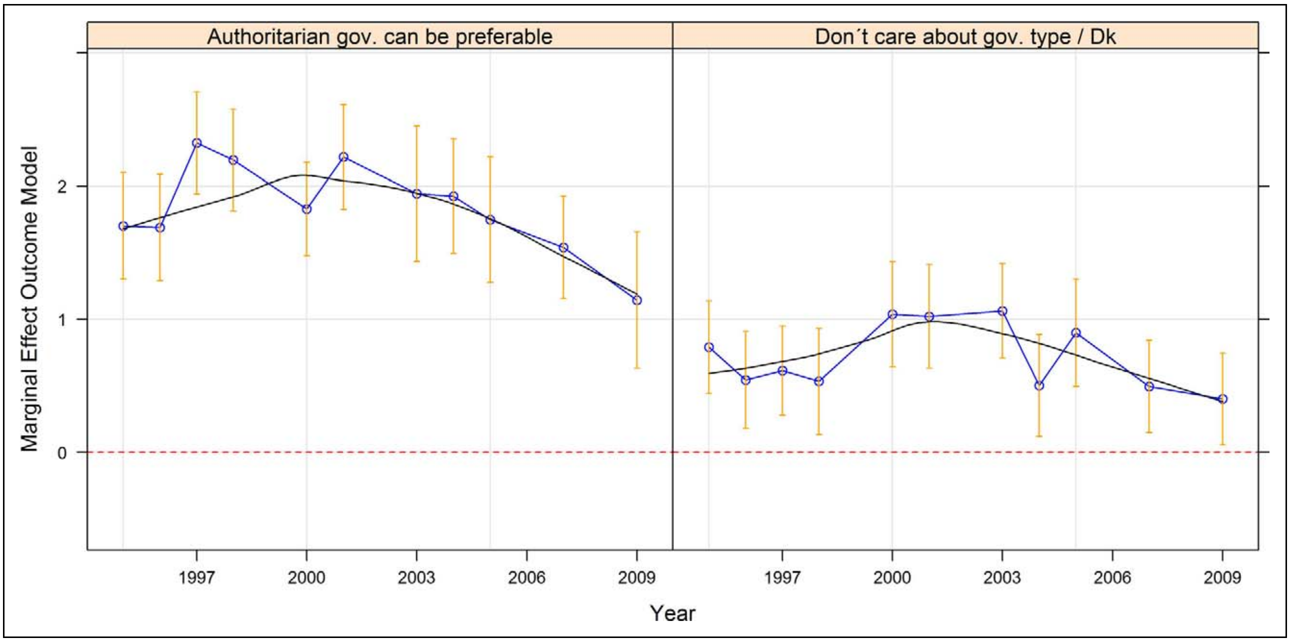

Figure 4. Evolution of attitudes towards regime marginal effects applied to each available year of Latinobarometer data.

decreasing monotonically since 2005. In fact, the marginal effect during the 2009 survey is non-significant, something that only occurs one more time in 1995. Among Evangelicals, one can also observe a continued decrease in the magnitude of the positive marginal effect for the period studied, though most of the changes occur between the years 1996 and 2001. Thereafter, identifying with an Evangelical denomination is associated with a positive, relatively stable and sometimes significant marginal effect. Results among the residual ‘others’ religion category also show a reduction in the size of the coefficients, and even becomes negative during the 2009 survey.

Finally, the marginal effects of respondents’ attitudes to democracy appear in Figure 4. We can see in both cases a spectacular and constant decrease (which is monotonic in the case of the authoritarian response) in the size of the marginal effects starting around the period 2000–2001. Although the ‘authoritarian government can be preferable’ option remains highly significant during all the available years, the size of the marginal effect in the 2009 survey is just about half of its size in the 2001 survey. The ‘don’t care about government’ option remains positive, but is only marginally significant during the last survey. Compared to the estimates around the years 2000–2003, the marginal effect of the 2009 survey is less than half the size.

In sum, with the only exception of the household goods index showing a modest increase in its marginal effect during the last surveys, there is a systematic reduction in the size of the marginal effects of multiple cleavage variables, namely education, religious denomination and regime preferences of respondents. While it is too early to make definitive claims, it seems that during the period we cover Chilean society experienced a generalized dealignment of the social and political basis of ideological preferences.

Discussion

How can we make sense of these trends? A thorough answer is beyond the limits of this article. However, we suggest – and provide some admittedly non-conclusive evidence – that the observed cleavage decline can be partially traced to a process of ideological convergence and moderation that has taken place among Chilean parties since redemocratization on issues related to class, religion and political regime. This convergence at the elite level weakened the differentiated signals needed to sustain strong cleavages among the masses (for a similar argument, see Evans and Tilley, 2012). We argue that this process, in turn, relates to the political and economic legacy of the Pinochet regime and how political actors reacted to it. Specifically, and consistent with a political agency approach, we claim that several elements of the legacy of the Pinochet regime, which we identify below, encouraged Chilean political parties, through a learning process, to adopt centrist political positions and strategies.

Evidence of ideological convergence among Chilean political parties comes from the Political Elites in Latin America (PELA) parliamentary survey (visit http://americo.usal.es/oir/elites/index.htm). This project surveyed Chilean deputies on four occasions during the period under study (1993, 1998, 2002 and 2006). In tapping the socio-economic cleavage we consider a question about the

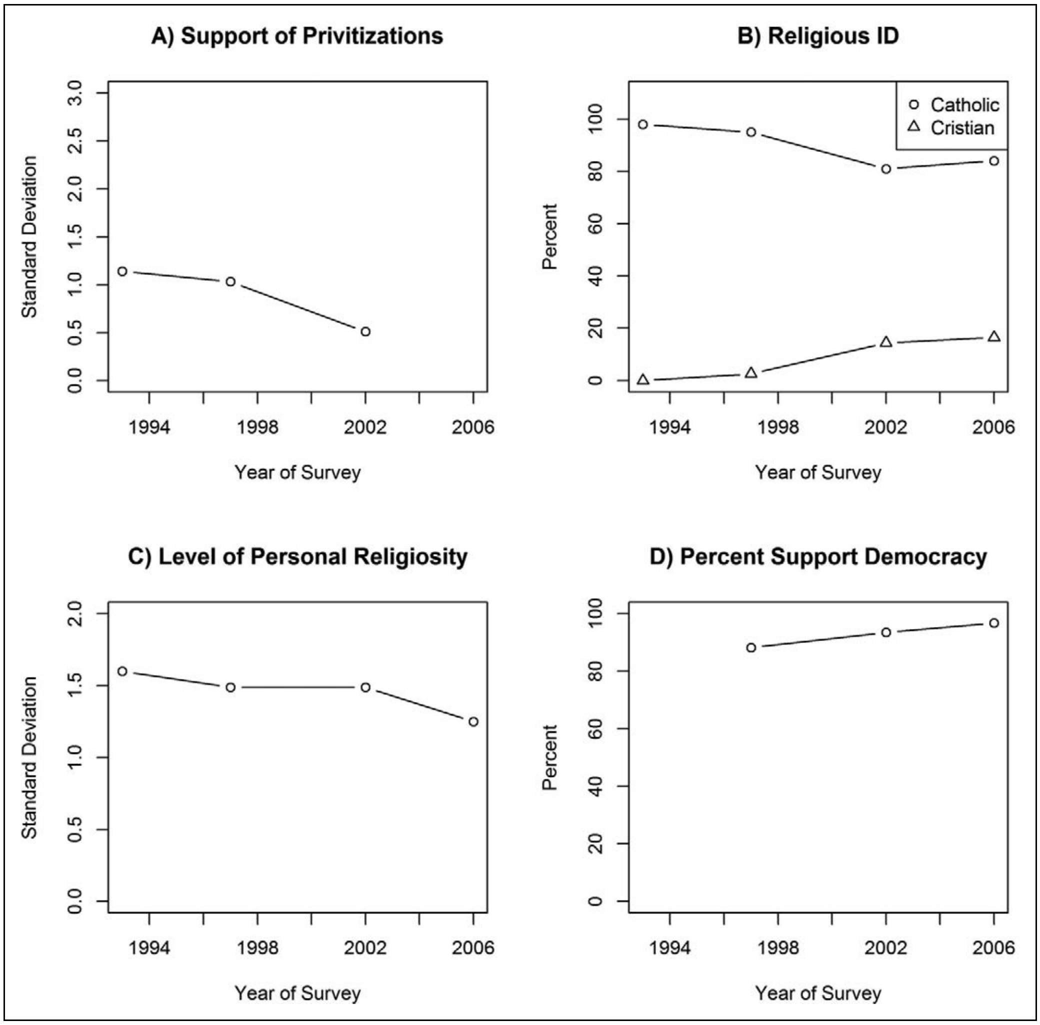

Figure 5. Evolution of congressmen attitudes to cleavage related issues.

socio-economic model that should prevail in society. Each interviewed congressman was asked to locate their opinion on a 5-point scale on which 1 equals 'would privatize all public services' while 5 was 'would not privatize any public service'. The results for this variable are shown on plot A of Figure 5, which indicates for each wave with available data the standard deviation of the average position of each party on this question. As indicated, there is a sizeable reduction in the size of the standard deviation of about 50 percent between 1994 and 2002 (the question was not asked in the 2006 survey). This, we claim, indicates an increasing level of homogeneity, or ideological convergence, in the views of congressmen on this matter. We found increasing moderation particularly among leftist legislators. The mean scores of PS legislators changed from 4,42 in 1993 to 2,66 in 2002, indicating a movement towards privatization. For PPD legislators the trend was similar: 4,36 in 1993, 2,58 in 1998 and 2,93 in 2002.

The religious attitudes of legislators also show convergence patterns. While a majority of deputies identify themselves as Catholics across all PELA waves, the corresponding percentage decreased from 98 percent in 1993 to 84 percent in 2006. Conversely, in the 2000s there was a sharp increase in those identifying themselves as 'Christians' – from 0 percent in 1993 to 17 percent in 2006. This trend is most visible among Christian Democrats and to a lesser extent among RN legislators and might indicate a reluctance of congressmen to identify with the institutional structure of the Catholic Church while still maintaining a diffuse link with religion. There is also evidence of convergence on religiosity. According to the PELA survey, the standard deviation of party means in a question about personal religiosity (where 1 = minimum and 10 = maximum) decreased steadily across time, indicating lower religious polarization, as shown on plot C in Figure 5.

Lastly, changes in attitudes toward democracy among legislators are also consistent with the decline of the regime divide at the mass level. The percentage of Chilean legislators agreeing with the sentence 'Democracy is better than any other form of government' increased across time from 84 percent in 1997 to 96 percent in 2006 (question not asked in 1993). This change resulted from an increasing valuation of democracy by RN and UDI legislators – all *Concertación* legislators chose the democratic option in the three survey waves. As legislators achieved near consensus regarding democracy, there was little room for activating the political regime divide; elite politics was less capable of 'feeding' this divide at the mass level.

The Pinochet legacy and political learning

How can we explain this apparent process of ideological convergence and moderation among political elites? No doubt many factors matter, including social-structural trends. Increasing secularization and religious pluralism, as well as rising levels of economic wealth, have perhaps undermined the politicization of identities along religious and socio-economic lines. However, we believe there is a more immediate factor; namely, the political and economic legacy of the Pinochet regime, and, particularly, the set of incentives around which political competition was structured. Importantly, the adaptation to this new setting did not happen overnight but through a process of political learning across the 1990s and 2000s. Much of the timing of the results shown above is relatively consistent with this perspective.

From the political point of view, the most relevant feature of the inherited institutional setting refers to the binomial electoral system. This system – in which only two representatives are elected in each district – makes it difficult for a single party to gain representation by running alone. Thus, parties have incentives for building enduring electoral pacts sufficiently attractive at the local level to obtain at least one representative. This results in large coalitions composed of several parties which exclude extremist forces and therefore reduce polarization in the electoral supply (Alemán, 2009; Siavelis, 2002). Moreover, due to the electoral thresholds provided by the system, in most cases each coalition obtains one representative per district (rarely one of them obtains both), which grants predictability of results and limits real competition (Alemán, 2009). Arguably, these factors hinder the clear and differentiated party signals that promote strong cleavages (Evans and Tilley, 2012).

A second relevant institutional element refers to the 'authoritarian enclaves' (Garretón, 2003) of which the 'designated senators' were particularly emblematic. Once a president's period was over, and Pinochet being the first one, he had the constitutional power to assign a number of senators to the upper chamber. During the 1990s this granted veto power to the political right over many areas of legislation, forcing the *Concertación* governments to abandon deep social reforms with no chance of being approved in Congress (Navia, 2009; Roberts, 2011). Facing a moderate, non-revolutionary left restricted by 'authoritarian enclaves', the right was much more receptive to playing by the rules of democracy.

Though anecdotal, several empirical patterns in the behaviour of parties can be connected with the incentives derived from this general institutional setting, and from the binomial system in particular. First, internal religious heterogeneity within both large coalitions conspired against any attempt to politicize religious identities. The *Concertación* encompasses the religious and (in moral issues) conservative Christian Democracy along with the secular and liberal PS, PPD and PRSD. The *Alianza* includes the mildly liberal RN and the very conservative UDI. While both coalitions resist some internal diversity, any party that disproportionately favours some religious identities to the detriment of others may create strains within its coalition and indirectly favour the rival coalition (Luna, 2008). Moreover, because parties may need time to learn this logic, religious cleavages may be initially strong and decline after a certain time. This partially explains why the leftist parties of the *Concertación* decided to moderate their originally innovative bills on paternity, divorce and abortion – all issues with strong religious overtones – and frame them in ways that emphasized 'family values', therefore being palatable to the Christian Democrats (Haas and Blofield, 2005: 47). Along the same line, Alemán and Saiegh (2007) argue that *Concertación* leaders strategically removed certain moral issues from the legislative agenda in order to avoid confrontation among coalition parties.

The UDI provides a second example of how outreach party strategies weakened social cleavages. Faced with a large centre-left coalition and an internal coalition partner (RN) also seeking the right-wing electorate, UDI carried out a successful strategy in capturing the support of the popular sectors. This involved developing personal contacts and grassroots mobilization in poor communities, distributing particularistic benefits and publicly downplaying the elite character of UDI's core constituencies while highlighting its 'popular' side (Luna, 2010). Joaquín Lavín, a former UDI mayor and party leader who almost wins the 2000 presidential election, emphasized the need for better social protection for the poor and the unemployed, and implemented high-impact targeted measures – such as building a beach for the popular classes that remained in Santiago during the summer. This neopopulist style, which was widely replicated in UDI's municipal governments, marked a sharp contrast with the traditional upper-class bent of the Chilean right and promoted rightist political identifications among the popular classes. Perhaps not coincidentally, the class cleavage (measured with education) declined dramatically after Lavín's arousal as the leader of the right (Figure 2).

Lastly, in the face of a series of incidents during the 1990s and early 2000s, a consensus emerged among political parties regarding the political legacy of Pinochet's regime itself. The political right had good reason to remain attached to the Pinochet regime when democratization ensued. Different from other authoritarian experiences in the region – Argentina in sharp contrast – when Chilean democracy was restored Pinochet was supported by wide sectors of the population and his institutional legacy was protected by the 'authoritarian enclaves'. However, several processes soon motivated part of the Chilean right to distance itself from Pinochet's legacy and endorse democracy. First, as it became clear that the *Concertación* governments did not destabilize social order through grassroots mobilization, and that they successfully promoted economic growth and decreased poverty within the socio-economic model inherited from Pinochet, there were few reasons for resenting democracy (Roberts, 2011). Second, several official reports commissioned by the *Concertación* governments (the last one being the Valech report published in 2004) confirmed, beyond doubt, massive human rights violations by the military regime. Third, high impact media scandals such as the international arrest of Pinochet in London in 1998 and the 'Riggs affair', which revealed that Pinochet had committed millionaire fraud with public funds, inflicted severe damage on Pinochet's reputation among the citizenry. These factors encouraged right-wing parties to develop a strategy that emphasized distance between them and the authoritarian regime to avoid losing their more moderate supporters. Once again, this was best incarnated by UDI leader Joaquín Lavín, who praised Pinochet's economic model but energetically deplored the abuses of the military regime and predicated his absolute support for democratic rule (Luna, 2010: 346; Navia, 2009). The emergence of this newer and more moderate right relaxed the links between rightist self-positioning and support for the authoritarian regime, opening the way for a decline in the cleavage as observed in Figure 4.

The economic legacy of the Pinochet regime, in which multiple areas of public spending – such as education, pensions and health insurance – were partly transferred to the private system, also encouraged ideological convergence among political actors. Although Pinochet's economic policies were controversial by the time the first *Concertación* government took office, the country had experienced an average economic growth rate of 6 percent during the last six years of the dictatorship, and in the following decade this continued. During these years the *Concertación* administrations certainly expanded social programmes (such as *Chile Solidario*, the health reform AUGE and a social security reform), but they also accepted private property, promoted growth and savings, attracted domestic and foreign investment, and promoted international trade. Also, the *Concertación* governments did not replace the labour code or the privatized educational, pension and health systems. Moreover, the first *Concertación* government with a socialist president, Ricardo Lagos, eased regulations on private companies, increased private participation in mining and infrastructure projects and signed free trade agreements with the US and the European Union.

These examples clearly indicate that the Chilean left, particularly the Socialist Party, abandoned their historical preferences towards radical socio-economic reform and accepted the market-centred model of society imposed by Pinochet (Siavelis, 2002), though, of course, with important corrections. This process of political learning ultimately undermined the fear of the upper classes of the left and weakened the link between the left and the popular classes. Moreover, this helps us in understanding why, in a national survey carried out in late 2005, almost half of those respondents self-identified with the right or centre-right approved Lagos's performance (Navia, 2009). In fact, as with Lavín, it was during the Lagos administration that the class cleavage declined most.

Conclusions

Past research on social cleavages and politics in Chile revolved around the debate between the role of social categories (e.g. class and religion) versus that of political divisions (such as attitudes in favour of or against the Pinochet regime). Yet because this research was based on cross-sectional data for one or at most two years (Raymond and Feltch, 2012 for a notable exception), we do not know whether the effect of cleavages changed across time. Using yearly surveys for the 1995–2009 period, we found a generalized process of dealignment. We show that the effect of the division between those favouring and those opposing democracy dwindled too. Interestingly, our household goods index goes against this pattern and becomes more important as a predictor of political preferences from 2004 onwards.

We provide some preliminary evidence indicating that cleavage decline is paralleled by a process of ideological convergence and moderation among Chilean parties that weakened the differentiated signals needed to sustain strong cleavages among the masses. Also, we argue that this process can be understood as a progressive political learning dynamic whereby political parties adapted to the institutional and economic environment inherited from the Pinochet regime.

Of course, more research must be done before we can reach firmer conclusions. For instance, a comprehensive assessment of the class cleavage would require more refined measures, particularly empirical operationalizations of class position based on occupational categories such as those used in the social stratification literature (for examples, see Nieuwbeerta (1995), Manza and Brooks (1999) and Evans and Tilley (2012)). Also, the religious factor merits more research that includes indicators of public and private religiosity that cut across religious denominations.

Appendix Article: 'Social cleavages and political dealignment in contemporary Chile, 1995–2009'

In the following document we provide further methodological and statistical information related to the empirical analysis presented in the above article. Section 1 gives sample information of the survey data we employ and discusses the reasons for not using other well-known data sources. Section 2 provides detailed empirical estimates of the Heckman selection model for each survey separately. Section 3 gives empirical estimates showing the strong association between respondents' positions on the left–right scale and their electoral preferences. And, finally, section 4 provides some empirical results validating our household goods index as a reasonable socio-economic measure.

1. Latinobarometer survey data

According to documentation of the Latinobarometrer Corporation, the Chilean surveys between 1995 and 1998 employed probability with quota surveys that included the male and female population aged 18 years and older living in the 29 cities over 40,000 inhabitants between regions I to X region of the country (and thereby excluded regions XI and XII). This equates to 70 percent coverage of the adult population of the country. Respondents within households were chosen using age and gender quotas, though households and census districts were chosen randomly. The surveys conducted between 2000 and 2004 have the same coverage as the earlier ones, but employed multi-stage probability samples including respondent selection within households. Thereafter, Latinobarometer surveys employed multi-stage probability samples and covered the entire adult population of the country. The data and abundant methodological information can be downloaded at www.latinobarometro.org.

It is worth mentioning that the analysis we carry out cannot be done using the well-known Centro de Estudios Públicos (CEP) surveys, which have employed multi-stage probability samples since 1994. Several reasons justify this claim. First, the CEP failed to regularly include in their surveys a measure capable of capturing the regime preference dimension, and therefore cannot be used as a measure of the evolution of the influence of this fundamental political cleavage. According to our search there is one question that was asked on several occasions. This was: 'Considering both the good and bad things of the governments I'm going to name, what grade from 1 to 7, where 1 is bad and 7 is excellent, you would put the government of Augusto Pinochet?' The question was applied during 7 years between 1994 and 2003. Not only is the time span shorter, but during the surveys where this measure was included some other key social cleavage variables were excluded. For example, between 1994 and 1999 not a single survey included both the Pinochet government evaluation and a religious affiliation question. Considering this kind of data limitation, we could potentially reproduce the type of analysis done in the article (with simultaneous controls for all relevant cleavage variables) for the years 1999, 2000, 2001 and 2003. This short time period clearly represents an unsatisfactory option.

In second place, the CEP surveys did not ask the left–right self-identification 10-point scale up to the year 2004,[6] and even after that the question wording has one or two significant changes that undermine the temporal comparability of the data. The CEP surveys do include, since 1994 up to this date, an ideological preference question asking respondents to mention the ideological position that best describes themselves (with nominal response categories: right, centre right, centre, centre left and left). Unfortunately this last question has a much higher non-response rate than the more abstract left–right 10-point scale. Indeed, CEP surveys from 2004 to 2009 included both questions in the same questionnaires, which allowed comparisons between both measures. For this entire period (and using un-weighted data) the response rate of the nominal questions was, on average, 18 percentage points lower. Consequently, we favoured using the left–right scale because it maximizes the number of survey respondents that provide substantive information.

2. Heckman model selection and OLS results

We begin by providing the full results from each of the Heckman equations employed to calculate Figures 2, 3 and 4 of the article, and, for comparison, include the OLS estimates of the outcome equation. The results are given in Tables A, B, C and D.

It is interesting to note that OLS estimates either understate or overstate the magnitude of the associations between some of the independent variables and the left–right scale. Probably the most dramatic case refers to the coefficients of education. In this case, the OLS estimates are on several occasions much smaller than the respective coefficients of the outcome equation of the Heckman model. Consider the surveys from 1995, 2004, 2005, 2007 and 2009. In these cases the coefficients of the Heckman model at least double the size of the OLS estimate. These results, however, should be of no surprise since education is a strong positive predictor of whether each respondent mentioned a position on the left–right scale. The OLS estimates of the household goods index are also smaller than the Heckman estimates in several surveys, the most dramatic cases being observed in the surveys of 1995, 2000 and 2009.

The opposite can be observed among estimates of the indifference option of the regime preference question ('For people like me, it does not matter whether we have a democratic or a non-democratic regime.'). In this case, the OLS

Table A. Heckman selection model for left–right ideological scale in 1995–1997.

	1995			1996			1997		
	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS
Intercept	–0.958***	1.862***	3.286***	–1.161***	3.206***	3.637***	–0.29	2.159***	2.721***
	(0.257)	(0.557)	(0.426)	(0.27)	(0.519)	(0.431)	(0.287)	(0.527)	(0.372)
Male	0.147	–0.035	–0.186	0.104	–0.382**	–0.44***	0.109	–0.382***	–0.429***
	(0.09)	(0.166)	(0.152)	(0.093)	(0.151)	(0.147)	(0.096)	(0.134)	(0.127)
26–35 years	0.087	0.294	0.269	–0.167	0.084	0.106	–0.043	0.07	0.094
	(0.127)	(0.221)	(0.206)	(0.131)	(0.204)	(0.204)	(0.13)	(0.183)	(0.177)
36–45 years	–0.098	0.24	0.366	–0.322**	0.165	0.201	0.15	0.162	0.132
	(0.137)	(0.25)	(0.232)	(0.138)	(0.228)	(0.224)	(0.145)	(0.195)	(0.187)
46–55 years	0.143	0.47*	0.424*	0.06	0.379	0.368	0.02	–0.071	–0.068
	(0.156)	(0.275)	(0.256)	(0.163)	(0.244)	(0.243)	(0.165)	(0.228)	(0.222)
56–65 years	0.191	0.354	0.292	–0.108	–0.127	–0.095	0.267	0.534**	0.438*
	(0.177)	(0.323)	(0.302)	(0.18)	(0.291)	(0.291)	(0.191)	(0.251)	(0.236)
66 years or more	–0.464**	0.592	1.047***	–0.093	0.311	0.335	0.113	0.413	0.385
	(0.19)	(0.421)	(0.388)	(0.189)	(0.308)	(0.307)	(0.202)	(0.283)	(0.275)
Education	0.1***	0.119*	0.016	0.103***	–0.053	–0.075	0.067*	0.108**	0.074
	(0.033)	(0.065)	(0.056)	(0.031)	(0.057)	(0.053)	(0.038)	(0.054)	(0.048)
Household goods index	0.036	0.149***	0.103**	0.046*	0.074	0.055	0.055**	0.105***	0.085**
	(0.027)	(0.052)	(0.048)	(0.026)	(0.046)	(0.046)	(0.027)	(0.04)	(0.036)
Evangelical	–0.229	–0.426	–0.383	0.08	1.075***	1.048***	0.257	0.918***	0.855***
	(0.212)	(0.397)	(0.374)	(0.224)	(0.36)	(0.36)	(0.239)	(0.325)	(0.314)
Catholic	0.003	0.262	0.321	0.139	1.232***	1.233***	0.199	0.945***	0.915***
	(0.202)	(0.396)	(0.375)	(0.202)	(0.343)	(0.343)	(0.212)	(0.302)	(0.294)
Other religion	0.142	0.359	0.259	0.194	0.921***	0.88***	0.113	1.016***	1.005***
	(0.161)	(0.291)	(0.272)	(0.159)	(0.259)	(0.258)	(0.154)	(0.216)	(0.211)
Don't care about gov. type / Dk	–0.097	0.72***	0.93***	–0.117	0.516***	0.561***	–0.301***	0.526***	0.691***
	(0.1)	(0.2)	(0.182)	(0.102)	(0.19)	(0.183)	(0.11)	(0.197)	(0.161)
Authoritarian gov. can be preferable	0.19	1.842***	1.797***	0.327**	1.765***	1.721***	–0.182	2.268***	2.357***
	(0.125)	(0.215)	(0.2)	(0.135)	(0.19)	(0.186)	(0.132)	(0.192)	(0.178)
Interest in politics	0.603***			0.708***			0.332***		
	(0.065)			(0.074)			(0.061)		
Inverse Mills ratio		1.997***			0.567			1.067	
		(0.452)			(0.389)			(0.679)	
Sigma		2.563	2.311		2.172	2.166		2.009	1.927
Rho		0.779			0.261			0.531	
N obs / N censored	964/271			890/251			957/182		

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

Table B. Heckman selection model for left–right ideological scale in 1998–2001.

	1998			2000			2001		
	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS
Intercept	–0.595**	1.248**	2.299***	–0.265	2.054***	2.663***	–0.062	2.551***	2.935***
	(0.262)	(0.519)	(0.38)	(0.314)	(0.528)	(0.428)	(0.296)	(0.493)	(0.454)
Male	0.128	0.075	0.009	0.234**	–0.145	–0.263*	0.051	–0.286*	–0.338**
	(0.091)	(0.156)	(0.141)	(0.105)	(0.155)	(0.142)	(0.099)	(0.162)	(0.159)
26–35 years	0.357***	0.726***	0.449**	0.198	0.128	0.083	–0.096	–0.476*	–0.464*
	(0.125)	(0.229)	(0.201)	(0.164)	(0.224)	(0.217)	(0.149)	(0.251)	(0.248)
36–45 years	0.355***	0.606**	0.302	0.048	0.353	0.355	0.234	0.094	0.014
	(0.137)	(0.248)	(0.217)	(0.16)	(0.225)	(0.218)	(0.154)	(0.25)	(0.245)
46–55 years	0.155	0.73***	0.591**	–0.019	0.054	0.092	0.115	–0.094	–0.14
	(0.15)	(0.275)	(0.254)	(0.17)	(0.247)	(0.241)	(0.156)	(0.254)	(0.25)

(continued)

Table B. (continued)

	1998			2000			2001		
	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS
56–65 years	0.409**	0.764**	0.4	0.074	0.238	0.265	0.4**	0.174	0.058
	(0.184)	(0.316)	(0.275)	(0.19)	(0.274)	(0.267)	(0.199)	(0.31)	(0.302)
66 years or more	0.471**	1.165***	0.831***	–0.173	0.404	0.542**	–0.017	0.026	0.023
	(0.196)	(0.332)	(0.292)	(0.185)	(0.287)	(0.275)	(0.19)	(0.326)	(0.324)
Education	0.045	0.233***	0.182***	0.026	0.122**	0.102**	0.028	0.141**	0.113*
	(0.039)	(0.067)	(0.06)	(0.035)	(0.051)	(0.048)	(0.039)	(0.063)	(0.062)
Household goods index	0.037	0.058	0.043	0.043	0.148***	0.115**	–0.043	0.089	0.089
	(0.026)	(0.048)	(0.044)	(0.034)	(0.051)	(0.048)	(0.034)	(0.058)	(0.058)
Evangelical	–0.101	–0.22	–0.072	0.353	0.358	0.224	–0.127	0.01	0.054
	(0.2)	(0.349)	(0.32)	(0.279)	(0.368)	(0.353)	(0.238)	(0.386)	(0.382)
Catholic	0.008	0.737**	0.979***	0.04	0.572*	0.584*	0.173	0.619*	0.553*
	(0.175)	(0.322)	(0.295)	(0.21)	(0.324)	(0.316)	(0.22)	(0.335)	(0.33)
Other religion	0.192	0.839***	0.826***	0.117	0.553**	0.508**	–0.081	0.857***	0.893***
	(0.134)	(0.225)	(0.206)	(0.171)	(0.252)	(0.245)	(0.155)	(0.239)	(0.235)
Don't care about gov. type / Dk	–0.413***	0.296	0.819***	–0.378***	0.908***	1.152***	–0.178	0.972***	1.111***
	(0.1)	(0.229)	(0.168)	(0.111)	(0.218)	(0.182)	(0.11)	(0.203)	(0.192)
Authoritarian gov. can be preferable	–0.017	2.185***	2.175***	0.159	1.882***	1.853***	0.2	2.272***	2.263***
	(0.136)	(0.211)	(0.19)	(0.148)	(0.187)	(0.181)	(0.128)	(0.195)	(0.193)
Interest in politics	0.44***			0.459***			0.641***		
	(0.059)			(0.064)			(0.07)		
Inverse Mills ratio		2.056***			1.285**			0.903**	
		(0.556)			(0.599)			(0.452)	
Sigma		2.417	2.123		2.288	2.201		2.323	2.288
Rho		0.85			0.561			0.389	
N obs / N censored	931/243			1006/170			875/202		

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

estimate tends to be larger than the coefficient of the outcome equation of the Heckman model. In some cases (e.g. in 1998 and 2005) the differences are quite large.

Lastly, a second piece of information justifying the use of a selection model, instead of simply using OLS, can be found in the parameter estimate of the inverse mills ratio. As shown in the tables, the estimate of the inverse mills ratio is significant at a 0.05 or lower in 8 out of the 11 surveys, and is significant at a 0.10 level in 9 out of the 11 surveys.

3. Predictive capacity of the left–right scale

In the article we employ the left–right self-identification scale as our measure of political preferences, and not the more commonly employed vote recall or vote intention question. In the article, we provide a theoretical justification for this decision, as well as some bibliographical references of previous work that employed the same measure. Nonetheless, in this current appendix we offer some further empirical results that show a very strong association between respondents' positions on the left–right scale and their declared electoral preferences. Through this analysis we seek to remove any possible doubt about the validity of our chosen dependent variable.

In addition to the left–right self-location, the Latinobarometer survey asked respondents since the year 2001 to state for which party they would vote if there were general elections next Sunday. We decided not to use this variable not only because it covers a narrower time period, but also because it suffers from a very large proportion of non-response. Indeed, if we add the proportion of responses 'Don't know', 'Does not vote' and 'No response', this comes to a total of 51 percent for the period covered between the years 2002 and 2009 (but excluding 2003). Despite this problem it is still worth exploring whether people's responses to the left–right scale and vote intention are associated or not with the segment of the population that declares both a position on the left–right scale and an electoral preference. Table E provides the results of binary logit models predicting vote intention for one of the two main political coalitions (*Alianza* or *Concertación*).[7] Table F gives more disaggregated estimates predicting vote preference for either of the *Alianza* parties (UDI or RN), the Christian Democrats (which are part of the *Concertación*, but stand at the political centre), the *Concertación* centre-

Table C. Heckman selection model for left–right ideological scale in 2003–2005.

	2003			2004			2005		
	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS
Intercept	–1.806***	1.945***	2.718***	–1.453***	1.289**	2.947***	–0.609**	2.385***	3.401***
	(0.251)	(0.566)	(0.461)	(0.257)	(0.588)	(0.424)	(0.261)	(0.523)	(0.431)
Male	–0.002	–0.384**	–0.428***	0.35***	–0.028	–0.318**	0.126	–0.104	–0.214
	(0.086)	(0.167)	(0.165)	(0.087)	(0.17)	(0.147)	(0.088)	(0.152)	(0.142)
26–35 years	0.125	0.43	0.406	0.17	0.288	0.257	0.278*	0.074	–0.112
	(0.134)	(0.27)	(0.269)	(0.135)	(0.253)	(0.238)	(0.142)	(0.25)	(0.232)
36–45 years	0.084	0.403	0.371	0.109	0.545**	0.459*	0.211	–0.12	–0.272
	(0.136)	(0.274)	(0.273)	(0.136)	(0.251)	(0.235)	(0.141)	(0.252)	(0.237)
46–55 years	0.367**	0.587**	0.477*	0.366**	0.641**	0.414*	0.171	–0.292	–0.412*
	(0.147)	(0.284)	(0.279)	(0.148)	(0.269)	(0.247)	(0.149)	(0.262)	(0.247)
56–65 years	0.177	0.371	0.325	0.315**	0.488*	0.303	0.258	0.02	–0.12
	(0.155)	(0.315)	(0.314)	(0.158)	(0.29)	(0.269)	(0.158)	(0.28)	(0.264)
66 years or more	0.463***	0.782**	0.626*	0.311*	0.762**	0.692**	0.195	0.253	0.137
	(0.165)	(0.33)	(0.322)	(0.166)	(0.321)	(0.303)	(0.168)	(0.304)	(0.289)
Education	0.138***	0.186***	0.125*	0.173***	0.267***	0.108*	0.078**	0.108*	0.027
	(0.033)	(0.07)	(0.064)	(0.035)	(0.072)	(0.059)	(0.037)	(0.065)	(0.058)
Household goods index	0.03	0.067	0.06	–0.031	0.036	0.045	–0.037	0.062	0.067
	(0.029)	(0.061)	(0.061)	(0.031)	(0.059)	(0.055)	(0.03)	(0.051)	(0.049)
Evangelical	0.131	0.736*	0.674*	0	0.883**	0.88**	0.075	0.269	0.266
	(0.198)	(0.396)	(0.394)	(0.21)	(0.383)	(0.359)	(0.209)	(0.352)	(0.333)
Catholic	0.123	0.56	0.545	–0.007	0.738**	0.867***	0.012	0.621**	0.642**
	(0.164)	(0.346)	(0.346)	(0.164)	(0.323)	(0.306)	(0.157)	(0.283)	(0.27)
Other religion	0.245*	1.026***	0.961***	0.131	0.638***	0.571***	0.097	1.03***	0.984***
	(0.129)	(0.259)	(0.257)	(0.128)	(0.224)	(0.208)	(0.127)	(0.213)	(0.201)
Don't care about gov. type / Dk	–0.067	1.036***	1.182***	–0.068	0.449**	0.757***	–0.337***	0.703***	1.083***
	(0.094)	(0.206)	(0.197)	(0.099)	(0.208)	(0.187)	(0.098)	(0.208)	(0.176)
Authoritarian gov. can be preferable	0.249*	2.041***	2.003***	0.218*	2.096***	2.026***	0.188	1.857***	1.83***
	(0.128)	(0.225)	(0.224)	(0.132)	(0.226)	(0.209)	(0.152)	(0.241)	(0.228)
Interest in politics	0.717***			0.558***			0.582***		
	(0.064)			(0.056)			(0.061)		
Inverse Mills ratio		0.834**			1.799***			1.62***	
		(0.348)			(0.394)			(0.41)	
Sigma		2.38	2.347		2.344	2.091		2.298	2.106
Rho		0.35			0.767			0.705	
N obs / N censored	828/369			845/349			911/287		

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

left parties (PS, PPD or PRSD), or for the far left coalition *Juntos Podemos* (dominated by the Communist party). We estimate coefficients separately for each year for which we have data available, as well as a pooled model that includes all the observations together.[8]

We first consider the binary logit model predicting coalition vote. Several important results emerge here. First, for each year the left–right scale has a highly significant and substantially large coefficient. According to the pooled model, a unit change on the left–right scale is associated with a 171 percent increase in the odds of the Concertación being favoured. Similarly, the McFadden Pseudo R square and the percentage of correct predictions from the model indicate an overall strong fit. Considering the pooled model again, using the left–right scale alone we can correctly predict 86 percent of cases. This is a 43 percent increase in the predictive capacity of the model with the left–right scale compared to a null model that only correctly predicts 60 percent of cases.[9] Second, there is a remarkable level of stability in the size of the coefficients across each wave. While somewhat smaller in the 2002 survey, the left–right coefficient still involves a very strong effect over electoral preferences. Indeed, for this year a unit change on the left–right scale is associated with a 97 percent increase in the odds of a *Concertación* party being favoured. Therefore, the individual's position on the left–right self-identification scale is not only highly predictive of their vote intention, but this statistical association between variables tends to be relatively constant across time.

Table D. Heckman selection model for left–right ideological scale in 2007–2009.

	2007			2009		
	Selection Eq.	Outcome Eq.	OLS	Selection Eq.	Outcome Eq.	OLS
Intercept	–1.184***	1.388**	3.074***	–1.196***	3.023***	3.562***
	(0.231)	(0.681)	(0.428)	(0.256)	(0.528)	(0.435)
Male	0.17**	0.165	–0.001	0.127	–0.143	–0.202
	(0.083)	(0.166)	(0.15)	(0.086)	(0.153)	(0.148)
26–35 years	0.342**	0.169	–0.073	0.381***	–0.08	–0.175
	(0.138)	(0.278)	(0.253)	(0.137)	(0.265)	(0.259)
36–45 years	0.205	0.115	–0.026	0.436***	–0.054	–0.168
	(0.13)	(0.265)	(0.247)	(0.134)	(0.261)	(0.252)
46–55 years	0.308**	0.154	–0.114	0.453***	0.096	–0.028
	(0.145)	(0.289)	(0.264)	(0.148)	(0.283)	(0.273)
56–65 years	0.26*	0.283	0.104	0.522***	0.737**	0.574**
	(0.151)	(0.304)	(0.282)	(0.168)	(0.306)	(0.291)
66 years or more	0.198	–0.009	–0.129	0.278*	0.286	0.205
	(0.165)	(0.339)	(0.317)	(0.163)	(0.311)	(0.305)
Education	0.067**	0.161**	0.076	0.047	0.014	–0.002
	(0.032)	(0.066)	(0.058)	(0.033)	(0.056)	(0.054)
Household goods index	0.031	0.114*	0.075	0.025	0.145***	0.128**
	(0.029)	(0.059)	(0.054)	(0.028)	(0.052)	(0.051)
Evangelical	0.372*	0.364	0.062	–0.272	–0.464	–0.347
	(0.213)	(0.426)	(0.392)	(0.208)	(0.419)	(0.409)
Catholic	0.299**	0.688**	0.461*	0.395**	0.523*	0.458
	(0.143)	(0.301)	(0.276)	(0.164)	(0.283)	(0.279)
Other religion	0.36***	0.866***	0.637***	0.107	0.221	0.222
	(0.116)	(0.249)	(0.223)	(0.12)	(0.214)	(0.213)
Don't care about gov. type / Dk	–0.218**	0.349*	0.637***	–0.322***	0.317	0.487***
	(0.091)	(0.201)	(0.175)	(0.091)	(0.199)	(0.175)
Authoritarian gov. can be preferable	–0.039	1.513***	1.55***	–0.043	1.133***	1.155***
	(0.109)	(0.203)	(0.189)	(0.14)	(0.238)	(0.235)
Interest in politics	0.439***			0.591***		
	(0.052)			(0.057)		
Inverse Mills ratio		1.616***			0.68*	
		(0.469)			(0.372)	
Sigma		2.342	2.110		2.13	2.101
Rho		0.69			0.319	
N obs / N censored	828/369			845/349		

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

We also estimated for each wave the association between the left–right scale and a more disaggregated version of individuals' vote preferences. Despite the different measurement strategies, the results (shown in Table F) indicate very similar results to those just discussed. The coefficients of the left–right scale are highly significant and large for each year, and their magnitude remains relatively stable across time, though once again with the partial exception of the 2002 survey. The results from this table also show a high degree of consistency between the magnitude of the coefficients and the ideological location of each set of parties. Indeed, the coefficients for the left–right scale predicting votes for the centre-left parties (PPD/PS/PRSD) are more negative than those associated with votes for the centrist Christian Democrats; the coefficients associated with the *Juntos Podemos* coalition are, in turn, even more negative than the coefficients of the centre-left parties. In other words, a unit increase in the left–right scale (with higher values indicating a more right-wing position) implies a larger reduction in the probability of voting for the *Juntos Podemos* coalition than for the centre-left parties, and an even larger reduction than voting for the Christian Democrats. These varying magnitudes perfectly reflect the ideological alignment of the Chilean political parties.

Lastly, the fit statistics of the models in Table F are lower than the ones observed in the simpler models predicting vote preference for either of the main political coalitions. Model assessment analysis indicated that the reduction of fit is related to a certain inability of the model to clearly differentiate respondents that vote for the Christian Democrats and those who prefer the centre-left parties. This, of course, should not be surprising given that both sets of parties are

Table E. Binary Logit for Coalition Vote Intention.

	2002	2004	2005	2006	2007	2008	2009	All years
Intercept	4.358	7.455	6.86	6.319	5.668	6.063	6.485	5.881
	(10.366)	(10.918)	(11.596)	(10.805)	(10.251)	(9.419)	(10.135)	(–28.319)
Left–right scale	–0.677	–1.231	–1.179	–1.058	–0.939	–1.172	–1.134	–0.999
	(9.862)	(10.162)	(10.638)	(9.959)	(9.939)	(9.33)	(9.658)	(26.631)
Log likelihood	–175.866	–141.524	–198.338	–165.22	–149.815	–173.705	–152.292	–1191.495
Pseudo R^2	0.323	0.512	0.451	0.411	0.394	0.386	0.424	0.404
Correct predictions	0.843	0.881	0.877	0.883	0.823	0.807	0.86	0.859
N cases	395	455	567	446	373	410	401	3047

Notes: 1) Numbers in parentheses are *t* statistics; 2) Model predicts vote for Concertación parties.

Table F. Multinomial logit for vote intention.

	2002	2004	2005	2006	2007	2008	2009	All Years
DC								
Intercept	2.608	5.64	4.402	4.663	3.786	3.338	5.229	3.975
	(5.803)	(8.008)	(6.882)	(7.67)	(6.405)	(5.176)	(8.129)	(18.389)
Left–right scale	–0.49	–1.034	–0.943	–0.887	–0.719	–0.801	–0.997	–0.788
	(6.876)	(8.339)	(7.803)	(8.07)	(7.259)	(6.516)	(8.477)	(20.427)
PPD/PS/PRSD								
Intercept	4.673	8.038	7.112	6.585	6.452	6.172	6.415	6.205
	(9.889)	(10.837)	(11.494)	(10.481)	(10.094)	(9.445)	(9.508)	(27.589)
Left–right scale	–0.877	–1.468	–1.297	–1.234	–1.239	–1.312	–1.293	–1.185
	(10.258)	(10.675)	(10.979)	(10.382)	(10.314)	(9.924)	(9.936)	(27.705)
Juntos Podemos								
intercept	3.575	7.595	6.284	5.802	6.053	5.377	6.442	5.526
	(6.015)	(8.81)	(8.602)	(7.929)	(7.434)	(6.91)	(8.4)	(20.392)
Left–right scale	–1.215	–2.198	–1.889	–1.69	–1.923	–1.689	–1.95	–1.712
	(7.822)	(9.542)	(10.182)	(9.462)	(8.219)	(8.672)	(9.727)	(24.117)
Log likelihood	–393.573	–379.903	–465.838	–430.127	–332.648	–374.7	–386.623	–2831.913
Pseudo R^2	0.223	0.323	0.290	0.247	0.284	0.261	0.269	0.262
Correct predictions	0.602	0.666	0.715	0.627	0.639	0.664	0.601	0.637
N cases	415	473	590	469	388	432	424	3191

Notes: 1) Numbers in parentheses are *t* statistics; 2) Coalition for the Change parties as reference category.

Table G. OLS for household goods index.

	1996	2001	2008
Intercept	1.646***	1.429***	1.132***
	(0.066)	(0.097)	(0.123)
Household income	0.459***	0.425***	0.421***
	(0.013)	(0.013)	(0.015)
N cases	1368	1280	1091
R	0.70	0.66	0.65
R^2	0.49	0.44	0.42

members of the same coalition and share similar ideological positions on many issues. Despite these difficulties, the left–right scale still obtains a good fit. The McFadden Pseudo R^2 is never smaller tan 0.20, and the percentage of correctly predicted cases is always above 60. The pooled model, with 64 percent of correctly predicted cases, implies an 85 percent increase in predictive accuracy compared to a null model correctly predicting only 34 percent of cases.[10]

4. Validating the household goods index

In our article we employ the household index as a socio-economic measure of the survey respondent. It is calculated as an additive index that counts the number of goods each survey respondent reports possessing in his/her household. It includes the following goods: television, refrigerator, computer, washing machine, landline phone, car and also hot water. This index has a Cronbach alpha of 0.72.

To validate this measure we compared it with a household income variable using three different surveys from the Centro de Estudios Públicos applied in 1996, 2001 and 2008 (using CEP surveys 32, 41 and 58). The household income variable is an ordinal scale with 14 income ranges which vary across surveys.

Using these data it is possible to recreate the exact same household goods index measure. As shown in Table G, according to the CEP data both variables – household income and the household goods index – are highly correlated, with Pearson correlations ranging between 0.65 and 0.7.

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Notes

1. In section 3 of the online supplement we provide empirical results that show a very strong association between respondents' positions on the left–right scale and their declared electoral preferences.
2. The goods included in the index are television, refrigerator, computer, washing machine, landline phone, car and, additionally, hot water. This index has a Cronbach alpha of 0.72.
3. Adding interest in politics into the outcome equation does not produce any meaningful change in the estimates we present below.
4. To be exact, we have data for all the years in the 1995–2009 period except 1999, 2002, 2006 and 2008.
5. Given that in our specification the cleavage variables appear in both the selection and outcome equations, the marginal effect of each covariate does not correspond to its coefficient in the outcome equations (as in linear regression). Instead, each independent variable has a direct effect captured through its respective parameter in the outcome equation, and an indirect effect captured through its estimated parameter in the selection equation. Formally, the marginal for each independent variable k corresponds to:

$$\frac{\partial E[y_i|Z_i^* > 0]}{\partial x_{ik}} = \beta_k - \gamma_k(\rho\sigma_\varepsilon)\delta_i$$

where β_k and γ_k are the parameters from the outcome and selection equations, respectively, ρ is the correlation between the errors of the outcome and selection equations, and $\delta_i = \lambda_i^2 + w_i\gamma\lambda_i$ where λ_i is the inverse Mills ratio and $w_i\gamma$ is the linear predictor of the probit selection equation. See Green (2003: 783) for the full derivation. The marginal effects shown in Figures 2 to 4 are calculated for an 'average' survey respondent who is male, between 36 and 45 years of age, Catholic, chooses the 'democracy is always preferable' option and has a sample average value on education, the household goods index and interest in politics. The confidence intervals of each figure are normal-theory intervals estimated using 1,000 samples from non-parametric bootstrapping simulation.
6. There is one survey in 1995 that includes a 10-point left–right self-identification scale, but its question wording is completely different from the surveys of 2004 and after.
7. Recall that both political coalitions date back to the first presidential election held in 1989. Up to this day they have remained stable and their party membership has remained almost untouched. While the main parties included in each coalition have not changed at all, some small parties have dropped over the years. Work on congress roll-call votes shows that members of the two main coalitions tend to vote as ideological blocs (Aleman and Saiegh, 2007).
8. This analysis cannot be done for the years 2000, 2001 and 2003 given that the codes for the political parties are not available from the Latinobarometer website. We do not calculate the respective model for the surveys before the year 2000 given that the vote intention question was worded differently, and therefore is not comparable to the question used later.
9. In Table E we consider a prediction correct if the predicted probability of voting for a Concertación party is higher than 0.5 and the respondent declared that she would vote for the Concertación. Also, we consider a prediction correct if the predicted probability of voting for the Concertación is lower than 0.5 and the respondent declared that she would vote for a Coalition for a Change party.
10. In Table F we consider a prediction correct if the party mentioned by the respondent is also the party with the highest predicted probability of being voted for by the respondent.

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