IMT Institute for Advanced Studies, Lucca

Lucca, Italy

Essays in Applied Econometrics

PhD Program in Economics

XXVIII Cycle

By

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2016

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2016

Contents

List of Figures

List of Tables

Acknowledgements

Chapter 2 is the reproduction of the working paper "Procedures vs. Incentives: The Case of the University Promotion System in Italy", *IZA Discussion Paper 9386*, (2015). The material first originated from my master thesis supervised by Enrico Rettore[1] and Lorenzo Rocco[2] whose advices I greatly acknowledge.

Chapter 3 comes from a join work with Enrico Rettore. I acknowledge enlightening discussion with Paolo Rossi and Luciano Modica. I am also indebted to the Italian Minister of Education and the Italian Data Protection Commissioner (Garante della privacy) for making available the relevant data. Results are preliminary and currently under revision.

Chapter 4 is the reproduction of the article "Voting Behavior, Coalitions and Government Strength through a Complex Network Analysis" *PlosOne*, (2014). I greatly benefited from contributions of my coauthors: Gabriele Pompa[3], Michelangelo Puliga[4], Gianni Riotta[5] and Alessandro Chessa[6]. The authors acknowledge support from the Progetto di Interesse CNR "CRISISLAB", and the FET project SIMPOL 610704. All authors acknowledge enlightening discussions and support from people in LINKALAB. The funders had no role in study

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Working Papers

1. Dal Maso, C., Rettore E., Rocco L., "Procedures vs. Incentives: The Case of the University Promotion System in Italy", *IZA Discussion Paper 9386*, (2015)

Abstract

Intense political debate surrounded the Italian university system in the last ten years. The latter was widely perceived as auto-referential and inefficient so that each government proposed its own solutions and reforms. These targeted the number of classes offered, the way professors are recruited as well as parameters and amount of public funding toward universities. The present dissertation investigates how interventions affected Italian professors. The first work look at recruitment through the lens of game theory and represent the selection process as a bargain between members of the evaluation committee over the possible candidates. We bring the model to the data and evaluate how the reform of selection rules affected some of the parameters of the model. The second essay, focus on the funding system and in particular on the norms that limit personnel expenses of universities. We exploit the variation of budget constraints in time and across institutions in a quasi-experimental framework and estimate the effect of staff turn over limits on the probability of voluntary leaving academia. Finally, the last work looks directly into the political process and studies the formation and break of coalitions in the Italian parliament through concepts of complex network literature.

Chapter 2

A common observation is that individuals strive to neutralize the effect of procedural rules designed to drive choices away from their private optimum. An example of this phenomenon is offered by the reaction of Italian academia to two

reforms that modified the procedures of recruitment and promotion, by introducing random selection of the examiners not appointed by the recruiting school and reducing from two to one the number of candidates to be qualified. We model the negotiation occurring within evaluation committees and test the decision rule implied by the theoretical model on the sample composed of all selections to associate and full professorship initiated by the Italian schools of economics between 2004 and 2011. Particularly, we investigate whether these reforms decreased the relative weight of the examiner appointed by the recruiting school on committee's decision. Empirical results suggest that both reforms had little if no effect on examiners' weights.

Chapter 3

The number of professors employed in Italian universities dropped by 15% between 2008 and 2013. This resulted from two opposite trends: a reduction in new hires and an increment in professors leaving academia. While the first trend can be easily explained by government limits on personnel expenses the second one deserves further explanations. In this paper we investigate whether budget constrains on personnel expenses backfired by pushing Italian professor to quit their positions for private companies or foreign institutions. We exploit the selectivity of the new regulation along with its time variation in a quasi-experimental framework. Results indicate an increase in voluntary leaves among institutions with more severe staff limitations. Surprisingly the latter show lower voluntary leaves on average in both pre and post reform periods. We provide some explanations for these results and point out the limits of our methodologies for future investigations.

Chapter 4

We analyze the network of relations between parliament members according to their voting behavior. In particular, we examine the emergent community structure with respect to political coalitions and government alliances. We rely on tools developed in the Complex Network literature to explore the core of these communities and use their topological features to develop new metrics for party polarization, internal coalition cohesiveness and government strength. As a case study, we focus on the Chamber of Deputies of the Italian Parliament, for which we are able to characterize the heterogeneity of the ruling coalition as well as parties specific contributions to the stability of the government over time. We find sharp contrast in the political debate which surprisingly does not imply a relevant structure based on establised parties. We take a closer look to changes in the community structure after parties split up and their effect on the position of single deputies within communities. Finally, we introduce a way to track the stability of the government coalition over time that is able to discern the contribution of each member along with the impact of its possible defection. While our case study relies on the Italian parliament, whose relevance has come into the international spotlight in the present economic downturn, the methods developed here are entirely general and can therefore be applied to a multitude of other scenarios.

Chapter 1

Introduction

Education plays a fundamental role for the economic well being of individuals and society. Micro evidence suggests education is critical in modern labour markets with schooling being positively related to higher wages and lower unemployment (Cohn and Addison, 1998; Card, 1999). On a macro level, education seems to benefit societies through positive externalities such as technological progress (Lucas, 1988), reduction in crime rates (Lochner and Moretti, 2004) or more active voters (Wantchekon et al., 2015). This explains the high share of education in states expenditure (OECD, 2014) as well as the intense political debate that surrounds the education system. Part of this debate is mirrored in the empirical literature where features of the educational system, their evolution and economic impact is evaluated over time. Brunello and Checchi (2007) for example investigate the impact of early tracking of schools on human capital accumulation; Battistin and Schizzerotto (2012) and De Paola and Scoppa (2014) analyse the effect of grade retention on student achievements and Checchi (1999) and Bagues (2012) analyse the determinants of promotions to academic chairs. The first two works in the present dissertation relate to this stream of literature and in particular to the program evaluation of educational reforms. We exploit two exogenous shifts in regulation to assess their effects on the selection of professors in Italian academia. We look at two types of selection: the positive selection re-

lated to the recruitment process and the negative one related to outflows of Italian researchers.

In the first one, we look at the main features of the recruitment system to associate and full professors through the lens of game theory. We build up a simple bargain model and empirically estimate its parameters. Unlike Checchi et al. (2014) our players are commissioners who negotiate over candidates. Once the rules underpinning the selection of commissioners change according to an exogenous shift in regulation we are able to partially identify their bargaining power and possibly conclude on whether the political goals of the reform were achieved or not. Although our focus differs from civic capital, our findings are similar in spirit to Putnam et al. (1994) and Durante et al. (2011) in that the regulatory framework may have unexpected outcomes depending on the incentives in place. In our case, we show how repeated interaction and the payoffs at stake play against the ultimate objectives of policy makers as we are able to structurally estimate some of the parameters of the theoretical model.

In the second work, we move from inflows to outflows of researchers occurred in conjunction with the economic downturn. In the aftermath of the economic crisis many OECD member countries saw fiscal deficits soaring due to stimulus packages, reduction in revenues and in some cases banking assistance packages (OECD, 2011). Fiscal deficits rose to 7.9% of GDP and many countries, faced with interest rate hikes on sovereign bonds, implemented fiscal consolidation plans. Italy adjusted fiscal imbalances mainly through tax increases and reduction of retirement benefits but, nevertheless, the adjustment impacted also the education sector. Indeed, on the expenditure side many norms targeted public employees and professors in particular, with limits to the amount of staff turn over in public universities. The latter were divided in two types depending on a threshold on the incidence of personnel expenses: universities below the limit could hire only up to 50% of the savings from retired personnel while universities above that were completely prevented from hiring new professors. Not surprisingly, the norm induced a reduction in new hires and promotions that brought the number of professors

employed in Italian universities down by 15% between 2008 and 2013. Besides the block on entries, the decline resulted also from an increasing number of professors leaving academia. In this paper we investigate whether budget constrains on personnel expenses backfired by inducing Italian professor to quit their positions for private companies or foreign institutions. We argue that a prolonged alt in promotions may have affected careers outlooks and may have reduced expected earnings of Italian professors. Differentials in labour incomes are indeed regarded as the most important economic determinants of migration (Sjaastad, 1962, Borjas, 1991). Besides expected earnings, whether individuals decide to quit their positions depends, among others, on the size of outside opportunities. With this respect, Italian researchers were not the only one dealing with reduced funding. Similar concerns rose, for example, in the US during the negotiation between democrats and republicans over debt-ceiling deals (Gulledge, 2011). Yet in Italy these budget constraints were planned on multiple years (in fact they are still in place) and they were relatively easy to predict resulting possibly in greater impact on expectations. Empirical work on the link between economic crisis, public universities funding and potential outflows of researchers is surprisingly scarce. To fill this gap, we exploit the selectivity of the new regulation along with its time variation in a quasi-experimental framework. Preliminary results indicate an increase in voluntary leaves among institutions with more severe staff limitations. Surprisingly the latter show lower voluntary leaves on average in both pre and post reform periods. We provide some explanations for these results and point out the limits of our methodologies for future investigations.

Finally, the last essay looks directly into the political process and studies the formation and break of coalitions in the Italian parliament through concepts of complex network literature. This approach is well established for the US Congress with a network representation being given to committees and subcommittees that share the same members (Porter et al., 2005), to members of the Congress who co-sponsor bills (Fowler, 2006; Tam Cho and Fowler, 2010) or those who place the same roll-call votes (Waugh et al., 2009a). Similarly to Zhang et al. (2008) and Waugh

et al. (2009b) we make use of the network science concept of modularity in order to reconstruct the community structure of the parliament while being agnostic about political ties. We introduce a novel method to characterize the position of each deputy in the community of reference, based on its contribution to the modularity score, proposing a more intuitive interpretation compared to that based on the spectral decomposition developed in Waugh et al. (2009b) and in Porter et al. (2005). In this way we are able to characterize the heterogeneity of the ruling coalition as well as parties specific contributions to the stability of the government over time. We find sharp contrast in the political debate which surprisingly does not imply a relevant structure based on establised parties. We take a closer look to changes in the community structure after parties split up and their effect on the position of single deputies within communities by exploiting the events leading to the formation of the first Renzi government. While our case study relies on the Italian parliament, whose relevance has come into the international spotlight in the present economic downturn, the methods developed here are entirely general and can therefore be applied to a multitude of other scenarios.

Chapter 2

Procedures vs. incentives: the case of the university promotion system in Italy

2.1 Introduction

When a formal procedure is introduced to constrain individual behavior, a strategic reaction should be expected. Often, people struggle to complying with the procedure without being forced to change their optimal choice, in so doing bypassing the spirit and the purpose of the norm.

An example of this phenomenon is provided by two reforms of recruitment and promotion of professors in Italian academia. The first reform, enacted in 2008, introduced the principle of randomization in the composition of the local evaluation committees. Before, one member of the committee was appointed by the school calling for the position and four members were professors elected by all peers nationwide. After the reform the four external examiners were randomly drawn from a list of peers. The second reform, introduced in 2005, suspended in 2008 and soon restored, reduced from two to one the number of candidates to be qualified in each local selection (called *concorso*). In both cases the declared purpose was that of enhancing the independence of the external

examiners, reducing collusion within each committee and by that way making recruitment and promotion more meritocratic.

In this paper we evaluate the effect of such reforms. We first propose a model of candidates' qualification from which we derive the committee's optimal decision rule. The latter has a simple empirical counterpart that we estimate on all *concorsi* held in Italy between 2004 and 2011 and called by the schools of economics. Particularly, we estimate the relative weight of the internal examiner on committee decisions.

Our results suggest that both reforms did not significantly altered the preeminent role of the internal examiner on recruitment and promotion decisions, supporting the claim that Italian academia was able to neutralize the effect of the new procedures and continue its business as usual.

This result is not surprising given that the reforms modified only procedural aspects of recruitment but did not alter the fundamental incentives of academia and academics. Indeed, both reforms were introduced in a context of low-powered incentives, where funds were equally distributed among universities, professors salary depended on seniority only, while high quality teaching or strong scientific productivity played no role. Professors directly managed their institutions and their power and rents depended on colleagues' consensus. Hence, rather than being considered a resource, the most productive candidates in the *concorsi* were perceived as a threat, as they could have altered the *status quo* in the schools. Quite the opposite, the preferred candidates were those offering enough guarantees regarding their willingness to preserve and support the *status quo*.

Our analysis is related with Perotti (2002) and Durante et al. (2011). They look at the selections held in Italy immediately after the decentralization of recruitment occurred in 1998. Perotti (2002) concludes that decentralization did not achieve a higher degree of meritocracy compared to the previous system. Being a candidate affiliated or well connected to the recruiting school had a dominant influence on the probability of being qualified and appointed, especially when the scientific productivity of the evaluation committee was below the median. Interestingly, Perotti (2002) analyzes the determinants of the number of votes that external

examiners received and finds no relation with their scientific record. Durante et al. (2011) study familism/nepotism in academic promotions by looking at the pattern of family names in the recruiting university and among the candidates. They find a strong association between civic capital measured at the province level and nepotism and argue that while decentralization increased the possibility to behave opportunistically everywhere, misconduct increased mainly in the areas poor of civic capital. A recent contribution by De Paola et al. (2011) exploits the random composition of the evaluation committees (the same reform we analyze in this paper) to establish whether promotions are affected by gender bias. They follow the same approach adopted by Bagues and Esteve-Volart (2010) that investigate gender bias in national public exams to access four main corps of the Spanish Judiciary over 1987-2005, Zinovyeva and Bagues (2011) that estimate gender bias in academic promotions in Spain, and in Bagues (2012) that study the role of connections compared to productivity. Also Combes et al. (2008), by using the French data of the *agrgation nationale* between 1984 and 2003, assess the relative importance of scientific production compared to professional links between candidates and evaluators.

Our paper is organized as follows. Section 2 illustrates some features of the institutional background, Section 3 discusses the dataset and Section 4 introduces the model. Section 5 presents the empirical implementation. Results are reported in Section 6. Possible alternative explanations of our findings are discussed in Section 7. Conclusions and technical appendices follow.

2.2 Institutional Background

Italian *concorsi* have been widely criticized for not selecting the best candidates, reflecting the scarce attention of academia to meritocracy. The features of the public university funding prevailing between 1999 and 2010 help understanding what incentives universities, schools and academics faced in that period. There were neither a nationwide evaluation of university productivity or a firmly established link between funding

and performance. Before 2009 universities were assigned government funds - the so called *Fondo di Finanziamento Ordinario* (FFO) - on a completely egalitarian basis, regardless of their performance in research and teaching. Salaries increased each other year for everyone by the same proportion and career progression depended very little on publications and productivity (Perotti, 2002). Professors were subject to very few obligations regarding teaching and research. Overall, university provided decent rents to academics, in terms of a prestigious social position, a good salary, the possibility of carrying on external professional activities.[1]

There is some consensus - even if based only on anecdotal evidence - that Italian academy was quite successful in preserving and maximizing its rent, an achievement generally incompatible with a meritocratic recruitment.[2] The decentralized promotion system introduced in 1998, by which a school with a vacancy was allowed to call a *concorso*, offered schools the possibility of recruiting the candidates fitting better with their needs and, by enlarging the opportunities of collusion, favored the purpose of maintaining the *status quo*. Academics affiliated to different schools jointly determined the outcome of several *concorsi*, by mutually agreeing on the principle that each school had the indisputable right of deciding who was to be hired, regardless of any consideration of candidate's scientific merits. As discussed by Perotti (2002) also examiners with good scientific records that participate to evaluation committees often accepted this principle, because they knew that any deviation from the collusive equilibrium could have been punished in the future *concorsi* that their own schools would have opened.

[1]Nevertheless, the absence of a link between productivity and remuneration or career progression does not imply that the generality of Italian professors were inactive, nor that they did not care about their publication record. It means only that good research was not promoted by the university system and that it ultimately rested on the intrinsic motivation and good will of individuals and schools.The situation might improve in the next future. In 2011 the national agency of evaluation (ANVUR) became fully operational and the first nationwide assessment of research production (VQR 2004-2010) terminated in 2014. Since 2009 a small share of the FFO has been linked to some productivity indicators (this share was 7 percent of total funding in 2009 and increased up to 15 percent in 2013).

[2]Indeed, recruiting highly productive scholars, more interested in scientific achievements than on rents, would likely alter the *status quo*.

Collusion was also favored by the rules governing *concorsi*. Candidates were evaluated by a committee made up of five professors: one member of the committee was appointed by the recruiting school (internal examiner) and the other four came from different universities (external examiners). The evaluation committee selected a fixed number of candidates to be qualified (called *idonei*) and the recruiting school had the option, but not the obligation, of appointing one *idoneo* of its choice. The appointed *idoneo* was the winner of the *concorso*. The remaining *idonei* could be appointed by any other university, including those the external examiners came from, within two or three years. Between 1999 and 2008 external examiners were elected by the whole national body of full professors in the same scientific field, a procedure that allowed the election of external examiners to be "managed" in order to obtain committees accommodating the will of the recruiting school **?**). Between 1999 and 2001 three candidates in each *concorso* could be qualified. Between 2002 and 2005, this number decreased to two (the *double qualification system*). Having more than one qualification left room for extensive "horse-trading" between internal and external examiners and between schools. The internal examiner could more easily obtain the qualification of his preferred candidate by promising of voting for the preferred candidate of an external examiner for the second *idoneità*.

This system underwent two reforms. The first regarded the composition of evaluation committees. Since 2007, external examiners were randomly selected from a (long enough) list of professors. The purpose of the reform was that of making more difficult to manipulate the composition of the evaluation committee and of increasing the independence of the external examiners, by this way reducing the weight of the internal examiner and the recruiting school on the final decision. The second reform regarded the number of *idonei* in each *concorso.* Since 2005 only one candidate could qualify for each position (the *single qualification system*). This decision lasted little and two qualifications were re-introduced in 2007. Eventually, following public condemnation, the single qualification system was re-established in 2008. Figure 1 depicts this dynamic.[3]

[3]Since 2012 a new centralized recruitment system has been introduced and the first pro-

Figure 1: Single and double qualifications overtime

These reforms modified two formal rules of the *concorsi* with the purpose of inducing committees to select better deserving candidates even if in contrast with the desires of the recruiting school. In fact, these reforms contradicted the logic of the decentralized system and recognized that schools had little incentive to hire and promote on a meritocratic basis. However, hardly are downstream interventions effective if not accompanied with more radical reforms able to modify fundamental incentives. Indeed, if schools agreed - even implicitly - on the principle that each of them is free to decide who is to be hired, then neither the way external examiners are selected, nor the number of *idonei* would matter for the final outcome of a *concorso*.

2.3 Theoretical model

We model the selection procedure as a negotiation between internal and external examiners. The model presented in this section refers to the case of double qualification. The single qualification setting may be interpreted as a special case and will be discussed in Appendix B.

Suppose that a continuum of aspirant professors apply to a *concorso*. Define the quality of each candidate as the set of his relevant characteristics and let it be summarized by a uni-dimensional measure, denoted h, with $h \geqslant 0$. Each candidate has a particular value of h.

Evaluation committees receive and examine candidates curricula and eventually decide who are the *idonei*. To simplify the analysis we assume that the committee is composed by one internal examiner and only one

motions and recruitment based on the new system have taken place in 2014.

external examiner.[4] We also assume that the internal examiner shares the same preferences of the recruiting school, so that any agency problem can be set aside.

Internal and external examiners negotiate over which applicants should qualify. Next the internal examiner, on behalf of the recruiting school, decides whether to appoint and what candidate to appoint.[5]

Each party conceives an ideal candidate and weights each applicant according to the distance from such ideal. Each party's ideal is a candidate that perfectly reflects party's characteristics, i.e. with a h score equal to party's h. By this characterization we capture the preference for preserving the *status quo* that was prevalent in the Italian university in those years.[6]

We formalize the negotiation between internal and external examiners and the subsequent decision of the internal examiner as a modified "collective model" (see Chiappori, 1988)[7] with a guarantee of minimum payoff to the internal examiner. Thus, committee preferences are described by a weighted average of its members payoffs. In this welfare function the "Pareto weights" reflect the relative bargaining power of the two parties or their relative influence on the committee decision. The guarantee of minimum payoff accounts in a simple way for the option that the internal examiner has to refuse of appointing one of the *idonei* if they are both too far from his ideal. Upon refusal he receives a minimum

[4]The latter assumption implies that the preferences of the external examiners can be summarized by those of a representative agent.

[5]This formalization is consistent with empirical data: typically committees seek an agreement among all examiners or at least of four out of five. Both anecdotal and empirical evidence exclude that the final decision in a *concorso* could be obtained by means of majority voting or other procedures that do not involve communication, coordination, mediation and compromise among examiners. Partly the search for an agreement is sponsored by the rules governing the recruitment procedures (e.g. most evaluation criteria are stated by the law and additional criteria need to be approved by the committee and bind all examiners evaluation), partly unanimity serves to discourage future plaints by the non qualified candidates and partly unanimity is the result of collusion and the mutual agreement on the principle of "right-to-decide" of each recruiting school.

[6]A school with an open vacancy that calls for selection procedure wants to preserve and reproduce its average quality, a strategy which guarantees the stability of the existing internal equilibria.

[7]The collective model is commonly used to represent the choice of a household composed of several members with different preferences.

payoff equal to the option value, exogenously fixed, of organizing another selection in the future and recruiting someone else possibly closer to his preferences. No such guarantee protects the external examiner. The inclusion of this device in the model serves to account for the relevant proportion of *concorsi* that end up with no appointment between 2004 and 2011, as we shall illustrate in the next section.

Formally, examiners utility function is defined as

$$U_i = -\alpha(h_i - h_1)^2 - (1-\alpha)(h_i - h_2)^2 \tag{2.1}$$

$$U_e = -\beta(h_e - h_1)^2 - (1-\beta)(h_e - h_2)^2$$

where subscript i and e identifies internal and external examiners respectively, U is the utility function, h_1 and h_2 are the qualities of the two qualified candidates, h_i and h_e are the qualities of the examiners and, finally, α and β are preference parameters over candidates of the internal and external examiners, respectively. We assume $\alpha > \frac{1}{2}$ so that *idoneo* 1 will weight more than *idoneo* 2 in the internal examiner utility function. Thus, at equilibrium *idoneo* 1 will be closer than *idoneo* 2 to the internal examiner ideal, and, in case of recruitment, he will be the chosen candidate (i.e. the winner).

Once two *idonei* have been identified, the internal examiner will appoint *idoneo* 1 if the utility U_i he derives from the selection procedure exceeds his minimum payoff $\bar{U}_i$. Otherwise, no one will be appointed and the *concorso* concludes with a *failure*. From the external examiner's viewpoint failure has pros and cons. On the pros side, in a *failure*, the external examiner obtains two *idonei* relatively close to his bliss point. On the cons side, although the two *idonei* may be appointed by other schools within two or three years, appointment is not guaranteed and it will certainly take time. Moreover the external examiner will be held responsible for the waste of resources suffered by the recruiting school and he will be at risk of being sanctioned in future selections procedures. We assume that the cost of a failure to the external examiner is exogenous and equal to δ.

Formally, the problem that internal and external examiners face is

that of maximizing

$$W = \begin{cases} \gamma U_i + (1-\gamma)U_c & \text{if } U_i \geq -\bar{U}_i \\ \gamma \bar{U}_i + (1-\gamma)(U_c - \delta) & \text{if } U_i < -\bar{U}_i \end{cases} \tag{2.2}$$

with respect to h_1 and h_2, where γ is the "Pareto weight" of the internal examiner. The first line corresponds to the *appointment regime,* where *idoneo* 1 will be eventually recruited, and the second to the *failure regime,* where two candidates are qualified but no one will be recruited.

Being the problem discontinuous, the two regimes have to be analyzed separately. Formally each regime involves a constrained maximization in its own. In Appendix A we show that, if δ is small enough, the solutions of problem (2.2) is either the unconstrained solution of the appointment regime or the unconstrained solution of the failure regime, depending on the values of the parameters[8].

Precisely, the optimum of (2.2), which represents the decision rule of the committee, is

$$\begin{cases} h_1^* = \dfrac{\alpha\gamma}{K} h_i + \dfrac{\beta(1-\gamma)}{K} h_e \\ \\ h_2^* = \dfrac{(1-\alpha)\gamma}{1-K} h_i + \dfrac{(1-\beta)(1-\gamma)}{1-K} h_e \end{cases} \qquad if\ (h_i - h_e)^2 \leqslant \Delta$$

and

$$\begin{cases} h_1^* = h_e \\ h_2^* = h_e \end{cases} \quad if\ (h_i - h_e)^2 > \Delta$$

where $K = \alpha\gamma + \beta(1-\gamma)$ and Δ is a complex function of the parameters $(\alpha, \beta, \gamma, \delta, \bar{U}_i)$ defined in Appendix A.

Indeed, if the distance between the internal and external examiner's bliss points is not too large, the solution of the negotiation will satisfy the internal examiner and the appointment of *idoneo* 1 will conclude the procedure. In this case both *idonei* have characteristics located between

[8] In other words the optimum of the problem never occurs at the discontinuity frontier of the objective function W.

the bliss points of the two examiners and they will be closer to the internal examiner bliss point the higher his weight (for any fixed preference parameters α and β). Instead, when such distance exceeds a well defined threshold, the outcome of the *concorso* will be a failure, no candidate will be appointed, but the two *idonei* h score will coincide with the bliss point of the external examiner.

Quite similarly, in the single-qualification case we have

$$h_1^* = \gamma h_i + (1-\gamma)h_e \quad if\ (h_i - h_e)^2 \leqslant \Theta \tag{2.3}$$

and

$$h_1^* = h_e \quad if\ (h_i - h_e)^2 > \Theta \tag{2.4}$$

where Θ is a function of the parameters $(\gamma, \delta, \bar{U}_i)$. We refer to Appendix B for a formal derivation of the single-qualification case.

2.4 Data

Our data include all *concorsi* to associate and full professorship called by Italian schools of economics (*facolt di economia*) and concluded between 2004 and 2011. We have discarded all competitions in non-core scientific areas (foreign languages, informatics, sociology) and focused on economics, management, law and quantitative disciplines (stats and maths). In Table 1 we report the number of *concorsi* by year of the call, distinguishing between *concorsi* with elected committees (*Pre-reform*) and with randomly selected committees (*Post-reform*) and between *concorsi* with one or two qualifications.[9]

Quite surprisingly, more than 10 percent of all *concorsi* ended up without winners, i.e. the recruiting school eventually decided not to appoint any qualified candidate. We refer to these cases as *failure*s (see Table 2). We have investigated what justification was provided by the recruiting schools in case of failure: in most cases the reason for non appointing

[9]Note that in 2007 there are no calls. This happened because the introduction of new rules forced the Minister to postpone all outstanding competitions. The procedures for the *concorsi* called for in 2008 where suspended for almost two years and they ended up in 2010 and 2011.

Table 1: *Concorsi* by year of the call

	Pre-reform		Post-reform		
	Single	Double	Single	Double	Total
2004	.	126	.	.	126
2005	24	115	.	.	139
2006	46	.	.	.	46
2007	.	.	.	.	.
2008	.	.	1	148	149
2009	.	.	12	20	32
Total	70	241	13	168	492

was a major misalignment between *idonei*'s characteristics and the characteristics of the position the school posted in the call.[10]

Table 2: Failures by type of *concorso*

	Pre-reform	Post-reform	Total
Double: Appointment	219	143	362
Double: Failure	22	25	47
Single: Appointment	56	10	66
Single: Failure	14	3	17
Total	311	181	492

Data include biographical information of all examiners and qualified candidates, the names of the *idonei* and of the winner (if any) for each competition, the dates of the call and of the end of the procedure. We do not have information on candidates who apply for the position but eventually did not qualify.

Perhaps the most critical choice in this study is how to measure candidates' and examiners' quality. To be fully consistent with the model we should consider an index that combines scientific productivity, teaching quality and all other characteristics that might be relevant in a selection procedure. Being data on all these dimensions unavailable - if not unob-

[10]There are two cases where faculty motivated the decision with budget problems or a major reorganization of teachings underway. We have excluded these cases from our dataset.

servable altogether - we focus on scientific productivity assuming that it is an unbiased predictor of candidates' and examiners' overall quality.

Scientific productivity is the only dimension that is measurable for all subjects involved in our analysis in an objective manner. We adopt the *h index* proposed by Hirsch, an index which takes into account both quantity and impact of each scholar scientific production. For instance, an h-index equal to *n* indicates that a researcher has published at least *n* papers cited by at least *n* other publications belonging to the same dataset. By construction, the index is robust to cases of scholars with few but highly cited papers as well as of authors with lot of papers with only a modest impact on the scientific community.

We have collected the h-index of all qualified candidates, examiners and the entire faculty of the recruiting school, from the "Publish or Perish" software (Harzing, Harzing) which exploits the Google Scholar database of publications and citations. Eventually, we have collected the h-index of candidates, examiners and faculty members for almost 500 *concorsi* between 2004 and 2011 making up a database of about 36000 entries.[11] The choice of measuring scientific productivity over the Google Scholar data, which includes not only published articles and books but also gray publication such as working papers and reports, depends on the fact that we need information for quite different scientific areas (e.g. law v.s. economics) with quite different publication traditions. Furthermore, Google Scholar is the most comprehensive search engine for social sciences (Harzing and der Wai, 2008).

We use as an unbiased predictor of h_1 and h_2, candidates qualities, their own h-index as of the date of the call, h_e by the average h-index of the four external examiners in each evaluation committee and h_i by the average h-index of the faculty of the recruiting school in the same field of the position to be awarded. This way we make operational our assumption that the internal examiner acts on behalf of his school and completely shares his school's preferences.

In Tables 3 and 4 we report summary statistics of the h-index of candi-

[11] The h-index is based exclusively on the publications appeared on Google Scholar before the date of call of each *concorso*

dates and examiners distinguishing between double and single qualification, by outcome of the *concorso* (failure or appointment), and by method of committee selection (elected, randomized)[12].

[12]The sample of *concorsi* that we consider coincides with that used in the empirical analysis. Compared to the full sample of *concorsi* reported in Table 1, 27 *concorsi* are missing for lack of some relevant information.

Table 3: Summary statistics - Double qualification

	Elected Committees				Randomized Committees				Pooled Sample			
	Mean	Median	IQR	Obs.	Mean	Median	IQR	Obs.	Mean	Median	IQR	Obs.
Failure												
h index candidate	3.6	1.5	7.0	34	4.3	2.0	5.0	38	4.0	2.0	6.5	72
h index internal examiner H_i	2.2	1.5	1.8	34	4.0	4.3	4.1	38	3.1	2.2	3.7	72
h index external examiner H_e	3.9	1.3	4.3	34	4.2	2.8	6.0	38	4.0	1.6	4.6	72
$\lvert H_i - H_e \rvert$	3.1	1.0	3.8	34	2.4	1.3	2.8	38	2.7	1.2	3.4	72
Appointment - Idoneo 1												
h index candidate	2.3	1.0	3.5	216	3.6	2.0	6.0	136	2.8	2.0	4.0	352
h index internal examiner H_i	2.4	1.8	3.1	216	3.7	2.7	5.0	136	2.9	2.3	3.4	352
h index external examiner H_e	2.5	1.5	3.0	216	3.1	2.0	3.5	136	2.7	1.8	3.5	352
$\lvert H_i - H_e \rvert$	1.9	1.3	2.2	216	2.6	1.9	3.2	136	2.2	1.5	2.4	352
Appointment - Idoneo 2												
h index candidate	2.4	1.0	3.0	216	3.8	3.0	6.0	136	2.9	2.0	5.0	352
h index internal examiner H_i	2.4	1.8	3.1	216	3.7	2.7	5.0	136	2.9	2.3	3.4	352
h index external examiner H_e	2.5	1.5	3.0	216	3.1	2.0	3.5	136	2.7	1.8	3.5	352
$\lvert H_i - H_e \rvert$	1.9	1.3	2.2	216	2.6	1.9	3.2	136	2.2	1.5	2.4	352
Total												
h index candidate	2.4	1.0	4.0	466	3.8	2.0	6.0	310	3.0	2.0	5.0	776
h index internal examiner H_i	2.4	1.7	3.0	466	3.7	2.7	4.8	310	2.9	2.3	3.6	776
h index external examiner H_e	2.6	1.5	3.0	466	3.2	2.0	3.5	310	2.8	1.8	3.5	776
$\lvert H_i - H_e \rvert$	1.9	1.3	2.2	466	2.6	1.8	3.2	310	2.2	1.4	2.5	776
Observations	466				310				776			

Note: IQR stands for interquartile range.

Table 4: Summary Statistics - Single qualification

	Elected Committees				Randomized Committees				Pooled Sample			
	Mean	Median	IQR	Obs.	Mean	Median	IQR	Obs.	Mean	Median	IQR	Obs.
Failure												
h index candidate	2.4	1.5	2.0	10	9.5	9.5	7.0	2	3.6	2.0	5.5	12
h index internal examiner H_i	2.3	2.3	2.5	10	7.1	7.1	0.8	2	3.1	2.6	3.2	12
h index external examiner H_e	1.5	1.1	2.0	10	17.3	17.3	16.5	2	4.1	1.9	2.6	12
$\lvert H_i - H_e \rvert$	1.0	1.0	1.0	10	10.1	10.1	17.3	2	2.6	1.1	1.5	12
Appointment												
h index candidate	2.4	1.0	3.0	55	3.2	2.5	4.0	10	2.5	1.0	3.0	65
h index internal examiner H_i	2.3	2.1	2.3	55	5.3	6.2	5.3	10	2.8	2.3	2.3	65
h index external examiner H_e	3.0	2.0	2.5	55	2.2	1.8	1.5	10	2.8	2.0	2.3	65
$\lvert H_i - H_e \rvert$	1.8	0.9	1.3	55	3.4	2.5	4.9	10	2.0	0.9	1.3	65
Total												
h index candidate	2.4	1.0	3.0	65	4.3	3.0	6.0	12	2.7	1.0	4.0	77
h index internal examiner H_i	2.3	2.1	2.3	65	5.6	6.9	5.0	12	2.8	2.3	2.5	77
h index external examiner H_e	2.7	2.0	2.5	65	4.7	2.1	3.3	12	3.0	2.0	2.5	77
$\lvert H_i - H_e \rvert$	1.6	0.9	1.2	65	4.5	2.6	4.9	12	2.1	0.9	1.3	77
Observations	65				12				77			

Note: IQR stands for interquartile range.

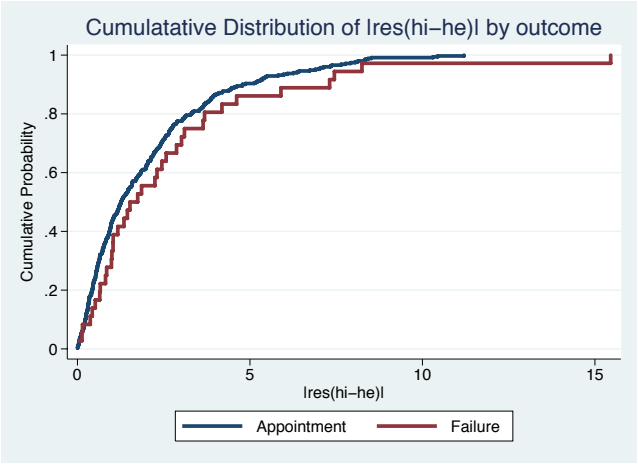
Descriptives based on the "Pooled Sample" (right-hand panel) immediately reveal that the *concorsi* ending up with a failure are those where the difference between the internal and the external examiners' h-index is relatively large, consistently with the predictions of our model. Figure 2 depicts the cumulative distribution of the absolute value of $(h_i - h_e)$ by type of outcome (after controlling for a full set of dummies for scientific field, year of the *concorso*, geographical area of the school, level of the *concorso*). In accordance with the implications of the theoretical model the (absolute) difference in quality between internal and external examiners is larger when a failure occurs.

Another interesting evidence emerging from the double-qualification case is that the median h-index of both *idonei* is in between h_i and h_e in case of appointment and close to the external examiner bliss point in case of failure. Qualitatively similar results holds for *concorsi* with a single qualification, although in the appointment case median h-index of the unique *idoneo* is outside the interval $[h_i, h_e]$.

Comparing the pre-reform (elected committees) and the post-reform (randomized committees) samples, we note that the h-index of post-reform candidates and examiners are moderately larger and more dispersed. Unexpectedly however, not only the variance between-*concorsi* of external examiners' h-index increased little after the reform (as the IQR values of Tables 3 and 4 indicate), but also the variance within-*concorsi* remained remarkably stable, suggesting that, overall, randomization had a relatively small impact on external examiners, at least as regards their h-index.[13] Actually, the more pronounced differences regarded internal examiners, i.e. the the members chosen by the recruiting schools, who were not directly influenced by the reform.

[13]The average standard deviation rose from 2.56 to 2,86 in the double qualification case and from 2.56 to 3.30 in the single qualification case.

Figure 2: Distribution of $(H_i - H_e)$ net of observable characteristics of the *concorso*



2.5 Empirical Implementation.

After rewriting the decision rule of the committee as

$$h_1^* - h_e = \begin{cases} \frac{\alpha\gamma}{K}(h_i - h_e) & \text{if} \quad (h_i - h_e)^2 \leq \Delta \\ 0 & \text{if} \quad (h_i - h_e)^2 > \Delta \end{cases} \tag{2.5}$$

$$h_2^* - h_e = \begin{cases} \frac{(1-\alpha)\gamma}{1-K}(h_i - h_e) & \text{if} \quad (h_i - h_e)^2 \leq \Delta \\ 0 & \text{if} \quad (h - h_e)^2 > \Delta \end{cases} \tag{2.6}$$

the *key implication* of the model becomes apparent: the distance between internal and external examiners $(h_i - h_e)$ determines both the *discrete outcome* (agreement vs. failure) and the *continuous* outcome (the quality for the qualified candidates).

To take to the data the equation for the decision rule we set

$$h_i - h_e = (H_i - H_e) + X\theta + \varepsilon \tag{2.7}$$

$$h_1^* - h_e = (H_1 - H_e) + X\theta + \varepsilon \tag{2.8}$$

$$h_2^* - h_e = (H_2 - H_e) + X\theta + \varepsilon \tag{2.9}$$

where H is for Hirsch h-index. The component $X\theta + \varepsilon$ (X observable and ε unobservable) is a *concorso*-specific prediction error of the unobserved quality of candidates and examiners. Since this component is *concorso*-specific we impose the restriction that θ is common across equations (9)-(11). Furthermore, ε is a random variable with zero-mean and standard deviation σ accounting for unobservable characteristics of each *concorso* that we assume to be orthogonal to $(H_i - H_e)$, $(H_1 - H_e)$ and $(H_2 - H_e)$ conditional to X.[14]

Replacing equations (2.7)-(2.9) into equations (2.5) and (2.6) yields the empirical counterparts of the optimal decision rules in the case of agreement:

$$H_1 - H_e = \frac{\alpha\gamma}{K}(H_i - H_e) + \left(\frac{\alpha\gamma}{K} - 1\right) X\theta + \left(\frac{\alpha\gamma}{K} - 1\right) \varepsilon \tag{2.10}$$

[14]Note that rewriting the model in terms of deviations from h_e implicitly imposes the constraint that the coefficients of h_i and h_e sum up to 1.

and

$$H_2 - H_e = \frac{(1-\alpha)\gamma}{1-K}(H_i - H_e) + \left(\frac{(1-\alpha)\gamma}{1-K} - 1\right) X\theta + \\ + \left(\frac{(1-\alpha)\gamma}{1-K} - 1\right)\varepsilon \tag{2.11}$$

Agreement occurs when $-\sqrt{\Delta} \leqslant (H_i - H_e) + X\theta + \varepsilon \leqslant \sqrt{\Delta}$ and failure when $(H_i - H_e) + X\theta + \varepsilon < -\sqrt{\Delta}$ or $(H_i - H_e) + X\theta + \varepsilon > \sqrt{\Delta}$.

Observable concorso characteristics include a dummy variable equal to one if the *concorso* is to appoint a full professor and equal to zero if it is to appoint an associate professor; dummies for the scientific field of the competition; dummies of geographic area (to capture the characteristics of the schools calling the selection); and year dummies. The inclusion of scientific field and geographic dummies captures fundamental differences among scientific fields in their propensity to publish and among schools as regards the criteria they adopt to appoint new scholars. To reduce the dimensionality of the regression problem we replace X in our equations with a *bivariate* Generalized Propensity Score including the estimated conditional expectation, $GPS1(X)$, and conditional variance, $GPS2(X)$, of $(H_i - H_e)$ on the control variables in X (see Imbens, 2000). The key mathematical property of the GPS is that if conditioning on the whole set of control variables warrants the identification of the causal parameter of interest, then conditioning on the GPS based on the very same set of control variables warrants identification as well.

The effect of the reform is captured by including an interaction between $(H_i - H_e)$ and a dummy *reform* which takes one for *concorsi* with a randomized committee.[15] To simplify estimation, we impose from the outset the restriction that the threshold Δ relevant for the outcome failure/agreement is *not* affected by the reform. At first glance this might seem inconsistent with the theoretical model, but in appendix C we suggest this is indeed the case at least approximately.

[15]The crucial identifying restriction we are imposing here is that conditional on the control variables the Pareto weight would have been stable overtime in the absence of reforms.

Hence, the equations we estimate in case of agreement are:

$$\begin{aligned} H_1 - H_e = {} & \pi_{11}(H_i - H_e) + \pi_{12}(H_i - H_e) \times reform + \\ & + (\pi_{11} - 1)\left(\theta_0 + \theta_1 GPS1(X) + \theta_2 GPS2(X)\right) + \\ & + (\pi_{11} - 1)\,\varepsilon \end{aligned} \tag{2.12}$$

and

$$\begin{aligned} H_2 - H_e = {} & \pi_{21}(H_i - H_e) + \pi_{22}(H_i - H_e) \times reform + \\ & + (\pi_{21} - 1)\left(\theta_0 + \theta_1 GPS1(X) + \theta_2 GPS2(X)\right) \\ & + (\pi_{21} - 1)\,\varepsilon \end{aligned} \tag{2.13}$$

where $\frac{\alpha\gamma}{K} = \pi_{11}$ and $\frac{(1-\alpha)\gamma}{1-K} = \pi_{21}$ in the *concorsi* with elected committees (*reform*=0) and $\frac{\alpha\gamma}{K} = \pi_{11} + \pi_{12}$ and $\frac{(1-\alpha)\gamma}{1-K} = \pi_{21} + \pi_{22}$ in *concorsi* with randomized committees (*reform*=1).

The key parameter is the Pareto weight of the internal examiner, γ, before and after the reform that introduced randomization. Our strategy identifies only $\frac{\alpha\gamma}{K}$ and $\frac{(1-\alpha)\gamma}{1-K}$ both before and after the reform. Note however that if $\frac{\alpha\gamma}{K} > \frac{(1-\alpha)\gamma}{1-K}$, a condition directly testable against the data, then $\alpha > K$. This fact bears two important implications:

1) $\pi_{21} < \gamma < \pi_{11}$ (and $\pi_{21} + \pi_{22} < \gamma < \pi_{11} + \pi_{12}$ after the reform) i.e. the estimated parameters π_{21} (resp. $\pi_{21} + \pi_{22}$) and π_{11} (resp. $\pi_{11} + \pi_{12}$) are lower and upper bounds, respectively, for the parameter of interest γ, since $\frac{1-\alpha}{1-K}\gamma < \gamma < \frac{\alpha}{K}\gamma$.

2) $\alpha > \beta$, i.e. the internal examiner weights more *idoneo* 1 compared to the external examiner.

Although without further assumptions on preference parameters α and β we can achieve only partial identification of γ, this will prove to be enough to draw conclusions.

The case of single qualification is alike. The empirical equation to estimate and the condition for agreement vs failure can be easily derived

from equations (2.3) and (2.4)

$$H_1 - H_e = \pi_{s1}(H_i - H_e) + \pi_{s2}(H_i - H_e) \times reform + \\ + (\pi_{s1} - 1)(\theta_0 + \theta_1 GPS1(X) + \theta_2 GPS2(X)) + \\ + (\pi_{s1} - 1)\varepsilon \tag{2.14}$$

where $\gamma = \pi_{s1}$ in the *concorsi* with elected committees and $\gamma = \pi_{s1} + \pi_{s2}$ in the *concorsi* with randomized committees. Note that in the single-qualification case exact identification of γ is obtained.

From now on consider only *idoneo* 1 - the cases for *idoneo* 2 and that of single-qualification being very similar.

2.5.1 Estimation procedure

To estimate the parameters of the model we proceed in two steps. First we estimate a modified Probit model for the probability to observe a *concorso* ending with an agreement. The standard Probit model is modified to take into account that a failure might occur either because

$$h_i - h_e = \theta_0 + (H_i - H_e) + \theta_1 GPS1(X) + \theta_2 GPS2(X) + \varepsilon < -\sqrt{\Delta}$$

or because

$$h_i - h_e = \theta_0 + (H_i - H_e) + \theta_1 GPS1(X) + \theta_2 GPS2(X) + \varepsilon > \sqrt{\Delta}$$

The resulting likelihood function is:

$$L^t = \prod_{\{failure\}} \left[\Phi\left(-\frac{1}{\sigma}(\Delta)^{1/2} - \frac{1}{\sigma}(H_i - H_e) \right) + \Phi\left(-\frac{1}{\sigma}(\Delta)^{1/2} + \frac{1}{\sigma}(H_i - H_e) \right) \right] \times \\ \times \prod_{\{agreement\}} \left[\Phi\left(\frac{1}{\sigma}(\Delta)^{1/2} - \frac{1}{\sigma}(H_i - H_e) \right) - \Phi\left(-\frac{1}{\sigma}(\Delta)^{1/2} - \frac{1}{\sigma}(H_i - H_e) \right) \right] \tag{2.15}$$

(for simplicity we omit from equation (17) the two components of the GPS and the constant). Maximization of (2.15) yields an estimate for $\frac{1}{\sigma}$ and for $\frac{1}{\sigma}(\Delta)^{1/2}$.

Second we estimate equations (2.12) and (2.13) on the sub-sample of *concorsi* ended with an agreement by means of a truncated regression estimator. The adjustment to take into account truncation is derived in Appendix D.[16]

Second step standard errors must be corrected to account for the use of estimated values for $\frac{1}{\sigma}$ and for $\frac{1}{\sigma}(\Delta)^{1/2}$ that enter in the adjustment for truncation. We implement the correction by running a Montecarlo simulation. From the distribution of ε, assumed to be normal with zero-mean and standard deviation equal to the value of σ estimated at the baseline, we draw 500 random vectors, denoted ε_j for $j = 1...500$. For each j we compute $I_j^* = \hat{\theta}_0 + (H_i - H_e) + \hat{\theta}_1 GPS1(X) + \hat{\theta}_2 GPS2(X) + \varepsilon_j$ and we generate a pseudo-outcome variable Y_j which takes 0 if $\left(I_j^*\right)^2 \leqslant \hat{\Delta}$ (predicted case of agreement) and 1 if $\left(I_j^*\right)^2 > \hat{\Delta}$ (predicted case of failure), where $\hat{\theta}_0$,$\hat{\theta}_1$,$\hat{\theta}_2$ and $\hat{\Delta}$ are baseline estimates. Next, for any Y_j we re-run our two-stage procedure where the dependent variables of the second-stage truncated regression (for both *idoneo* 1 and *idoneo* 2) are generated combining baseline predicted values with ε_j. Both dependent and independent variables include the correction for truncation which is derived from first-stage estimates for each j. Thus, we produce 500 replications of all parameter estimates that we use to derive parameter standard errors.

As a benchmark we also report standard OLS estimates of equations (2.12) and (2.13) on the subsample of *concorsi* where agreement was achieved, neglecting the problem of truncation. This is because in this specific instance the bias induced by the upper truncation and the symmetric bias induced by the lower truncation might cancel out.

2.6 Empirical results

Results of first and second stage are reported in Table 5 and 6, for the case of double and single qualification respectively.[17] As for the second

[16]Note that the coefficients of $GPS1(X)$ and $GPS2(X)$ in equations in (2.12) and (2.13) are non linear functions of the coefficients of $(H_i - H_e)$ and of the coefficients of $GPS1(X)$ and $GPS2(X)$ in the agreement equation, something we can test against the data.

[17]Note that in the first stage we do not impose the restriction that the censoring thresholds are symmetric around zero, since the intercept of the equation is left unconstrained.

stage, we report estimates of π_{11} , π_{12} and π_{21}, π_{22} respectively for *idoneo* 1 and *idoneo* 2, which are the bounds of γ, i.e. the Pareto weight of the internal examiner before and after the reform. For the case of single qualification we report the corresponding parameters π_{s1} and π_{s2}. There is a neat evidence that adjusting (two-step estimator) and not adjusting (OLS estimator) for truncation plays a minor role: point estimates are similar, although, as expected, OLS estimates are more precise. In the following we comment on the OLS.

Under the hypothesis that 1) evaluation committees aim to reach unanimous decisions to discourage subsequent appeals, and 2) external examiners perfectly coordinate (as assumed in the model), the Pareto weight of the internal examiner should be 0.2 (reflecting the fact that four external faced a single internal examiner). Let us take this value as a benchmark against which contrasting our results.

In the double-qualification case with elected committees (*reform* = 0) the bounds we find depict a narrow range for the Pareto weight of the internal examiner, going from 0.527 to 0.605. In the single qualification case, where exact identification is achieved, γ is slightly smaller (0.455), but still points to a preeminent role of the internal examiner as compared to the benchmark, even when reaching a decision is made more difficult by the impossibility of trading candidates within a given *concorso*. These estimates support the hypothesis that schools tend to replicate the *status quo* by recruiting candidates with a scientific standing similar to that of incumbent faculty.

In the double qualification case, the introduction of randomization in the composition of the committees (*reform* = 1) possibly widened the range of Pareto weights, although estimates are rather imprecise. We cannot thus exclude that γ decreased, but even in this case the internal examiner would have kept a substantial role. In the single qualification case, the post-reform Pareto weight of the internal examiner is not significantly different from its pre-reform level, thought the point estimate is negative.[18]

[18]Note also that there is no clear evidence against the hypothesis that the coefficients of $GPS1(X)$ and $GPS2(X)$ are indeed the product of $(\pi_{c1} - 1)$, for $c = 1, 2, s$ and θ_1 and

Table 5: Double qualification

Parameter	First Stage	Two Stage Procedure Second Stage Idoneo 2	Two Stage Procedure Second Stage Idoneo 1 (Winner)	Single Stage Procedure (OLS) Idoneo 2	Single Stage Procedure (OLS) Idoneo 1 (Winner)
$\frac{\Delta_H}{\sigma}$	1.345***				
	(0.391)				
$\frac{\Delta_L}{\sigma}$	-2.677***				
	(0.413)				
$\frac{1}{\sigma}$	0.074**				
	(0.036)				
$\frac{\theta_1}{\sigma}$	-0.211				
	(0.142)				
$\frac{\theta_2}{\sigma}$	-0.079***				
	(0.022)				
$\pi_{\cdot 1}$		0.535***	0.595***	0.527***	0.605***
		(0.179)	(0.152)	(0.092)	(0.080)
$\pi_{\cdot 2}$		-0.071	0.142	-0.174	0.027
		(0.169)	(0.153)	(0.122)	(0.121)
$(\pi_{\cdot 1}\text{-}1)\theta_1$		-0.206	0.100	-0.296*	0.073
		(0.394)	(0.305)	(0.161)	(0.162)
$(\pi_{\cdot 1}\text{-}1)\theta_2$		0.091**	0.034	0.030	-0.003
		(0.045)	(0.037)	(0.025)	(0.023)
σ	13.519**				
	(6.563)				
Δ_H	18.190*				
	(10.479)				
Δ_L	-36.192**				
	(17.646)				
θ_1	-2.846	0.442	-0.246		
	(1.902)	(0.920)	(0.733)		
test θ_1 [p-val]		[0.063]	[0.128]		
θ_2	-1.067*	-0.194*	-0.083		
	(0.604)	(0.093)	(0.733)		
test θ_2 [p-val]		[0.109]	[0.081]		

Note: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.
The upper panel reports estimates as obtained from modified probit (column 1), truncated regressions (columns 2 and 3), and OLS on the sub-sample of *concorsi* ended with an appointment (column 3). The lower panel reports the implied structural parameters. Standard errors of columns 1 to 3 are obtained by parametric bootstrapping the two stage procedure.

Table 6: Single qualification

	Two-Stage Procedure		Single-Stage Procedure (OLS)
Parameter	First Stage	Second Stage - Winner	Winner
$\frac{\Delta_H}{\sigma}$	0.950**		
	(0.395)		
$\frac{\Delta_L}{\sigma}$	-3.780**		
	(1.675)		
$\frac{1}{\sigma}$	0.178		
	(0.109)		
$\frac{\theta_1}{\sigma}$	-0.427*		
	(0.230)		
$\frac{\theta_2}{\sigma}$	-0.084**		
	(0.036)		
π_{s1}		0.458***	0.455***
		(0.162)	(0.136)
π_{s2}		-0.165	-0.151
		(0.248)	(0.302)
$(\pi_{s1}-1)\theta_1$		0.039	-0.088
		(0.231)	(0.203)
$(\pi_{s1}-1)\theta_2$		-0.025	-0.059
		(0.031)	(0.044)
σ	5.614		
	(3.437)		
Δ_H	5.332		
	(3.600)		
Δ_L	-21.221**		
	(10.597)		
θ_1	-2.396*	-0.072	
	(1.226)	(0.417)	
test θ_1 [p-val]		[0.064]	
θ_2	-0.470*	0.047	
	(0.274)	(0.059)	
test θ_2 [p-val]		[0.058]	

Note: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.
The upper panel reports estimates as obtained from modified probit (column 1), truncated regressions (column 2), and OLS on the sub-sample of *concorsi* ended with an appointment (column 3). The lower panel reports the implied structural parameters. Standard errors of columns 1 and 2 are obtained by parametric bootstrapping the two stage procedure.

2.7 Discussion

From a repeated game perspective, academia played an equilibrium that granted each university the right of deciding the candidate to be promoted or recruited, at its sole discretion. External examiners accepted rather passively the choice of the recruiting school, fearing retaliation in the *concorsi* called by their own school in future. Given schools' preference for the *status quo*, at equilibrium each school was allowed to recruit a candidate of quality comparable to that of the incumbent faculty. Our analysis indicates that government attempts to alter this equilibrium by introducing new procedural rule failed as academia devised successful strategies to bypass the new formal constraints.

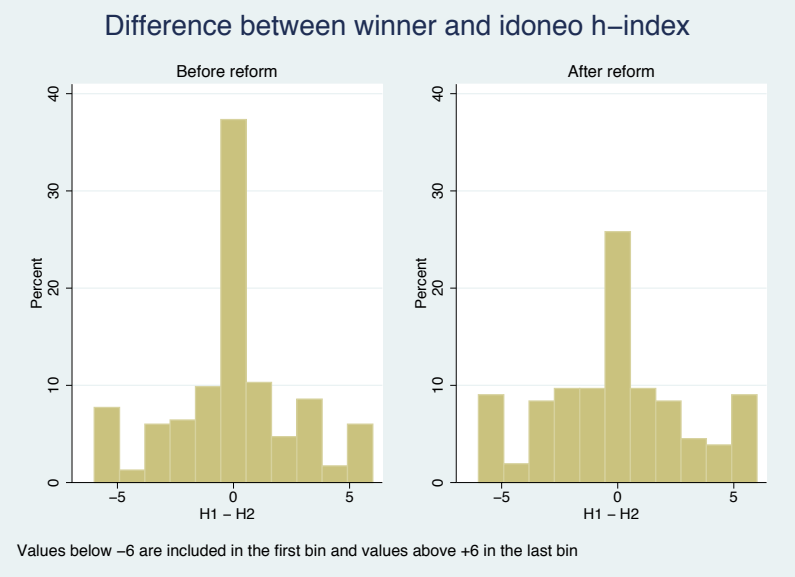
In this section we discuss three potential concerns regarding our results and interpretation.

First, a hiring pattern observationally equivalent to that obtained by schools recruiting scholars with a scientific quality similar to that of the incumbent faculty could be the result of a very different process. Suppose that candidates self-sort in the *concorsi* called by the schools closest to them in terms of scientific quality. If so, even if schools preferred to appoint the candidate with the highest standing to improve their quality, self-sorting would constraint their choice and schools would be forced to hire candidates that *de facto* replicate the *status quo*. If this were the case, however, in *concorsi* with double qualification, where we observe two applicants, we should systematically have that the winner's quality is higher than the second *idoneo* quality. In fact, only in one third of the cases H_1 exceeds H_2 as shown in Figure (3), where we report the distribution of $(H_1 - H_2)$ both before and after the reform.

Second, the lack of effect of the reforms that we find might be due to the countervailing influence of other policies. The only significant change occurred in the period under consideration was the rule enacted in 2009 that imposed of replacing at most 20 percent of all retirements in academia, a provision that significantly reduced the number of callable positions. However, this policy should have reinforced rather than off-

θ_2 respectively (see the bottom panels of Tables 5 and 6).

Figure 3: Distribution of $H_1 - H_2$



set the effect of the reform we analyze in this paper, given that it should have strengthen the competition among universities, being fewer positions available. Accordingly, the Pareto weight of the internal examiner should have decreased in contrast with the evidence we found.

Third, throughout the analysis we have maintained that all the characteristics of the examiners relevant for their preferences over candidates can be summarized by a unique index of scientific productivity, the h-index as computed on the Google Scholar dataset of publications. Even leaving aside the possible bias of the h-index as a measure of scientific productivity, there might be other individual characteristics, that influence the choice of the examiners, such as affinity in the field of research between candidates and examiners, a scientific potential of candidates not yet captured by the h-index, the role played by the affiliation of the candidates and the role played by academic and extra-academic ties. Nonetheless, summary statistics fit well with model implications and make us confident that the possible specification errors we commit are not such that of invalidating our main conclusions.

2.8 Conclusions

We have analyzed the effect of two reforms of the recruitment system in Italian universities whose purpose was that of increasing meritocracy by making external examiners more independent and by reducing the opportunities of collusion within committees. We have argued that the purpose of these reforms contrasted with the fundamental incentive of large proportions of schools and incumbent scholars of maintaining the *status quo* and continuing to benefit of sizable rents.

To study the effect of the reforms, we have derived a theoretical model that represents the process of negotiation within evaluation committees in a way that accounts for the evidence that a non negligible share of *concorsi* ended up without appointment (failures). The theoretical model implies a simple decision rule that turns out to be a linear function of internal and external examiners qualities and allows to (partially) identify the Pareto weight of the internal examiner. The empirical counterpart of

the decision rule has been estimated on the data from all *concorsi* called by the schools of economics in Italy and concluded between 2004 and 2011.

According to our estimations, neither the introduction of randomization of external examiners nor the reduction in the number of qualifications from two to one did significantly change the relative weight of the internal examiner on the evaluation committee.

The implication of our analysis is that procedural rules are little effective in modifying agents behaviors when they contrast with agents fundamental incentives. Changing procedural rules risks to load further administrative burden and costs without any substantial benefit. In the case of the recruitment system in Italian academia, reforms aiming to reward schools' scientific standing would be much more effective in making the adoption of meritocracy in schools' own interest.

Appendix A

In this appendix we characterize the solution of the negotiation problem between internal and external examiners.

We distinguish between two regimes, the appointment and the failure regime. In both cases we can have either an internal or a corner solution. Internal and corner solutions are linked to the structural parameters of the problem by different functional forms. As we are able to observe only what is the prevailing regime and not whether the solution is internal or corner, we need to determine under what conditions there is a one-to-one relation between observed regime and type of solution.

Assume $h_i > h_e$. The symmetric case is analogous.

Consider first the agreement regime.

The internal solution is

$$h_1^* = \frac{\alpha\gamma}{K} h_i + \frac{\beta(1-\gamma)}{K} h_e$$

$$h_2^* = \frac{(1-\alpha)\gamma}{1-K} h_i + \frac{(1-\beta)(1-\gamma)}{1-K} h_e$$

with

$$K = \gamma\alpha + (1-\gamma)\beta$$

The internal examiner utility at that solution is

$$U_i = -(h_i - h_c)^2\Gamma_i$$

the external examiner utility is

$$U_c = -(h_i - h_c)^2\Gamma_c$$

and the value of the problem is

$$W^A = -(h_i - h_c)^2\left[\gamma\Gamma_i + (1-\gamma)\Gamma_e\right]$$

where

$$\Gamma_i = (1-\gamma)^2\left[\frac{\alpha\beta^2}{K^2} + \frac{(1-\alpha)(1-\beta)^2}{(1-K)^2}\right]$$

$$\Gamma_e = \gamma^2\left[\frac{\alpha^2\beta}{K^2} + \frac{(1-\alpha)^2(1-\beta)}{(1-K)^2}\right]$$

and

$$P = \left[\gamma\Gamma_i + (1-\gamma)\Gamma_e\right] = \gamma(1-\gamma)\left[\frac{\alpha\beta}{K} + \frac{(1-\alpha)(1-\beta)}{1-K}\right]$$

It is easy to check that $0 < \Gamma_i < 1$, $0 < \Gamma_e < 1$ and $0 < P < \gamma$.

The condition which separates the internal from the corner solution is $-(h_i - h_c)^2\Gamma_i \geqslant -\overline{U}_i$ i.e. we observe an internal solution if

$$(h_i - h_c)^2 \leqslant \frac{\overline{U}_i}{\Gamma_i}$$

and a corner solution otherwise.

The corner solution satisfies the first order conditions of the Lagrangian

$$L = W - \lambda(-\bar{U}_i - U_i)$$

with $\lambda \geqslant 0$.

At the corner solution, the marginal rate of substitution between h_1 and h_2 needs to be equalized between the internal and the external examiners and, furthermore, the corner solution needs to satisfy the constraint

$U_i(h_1, h_2) = -\overline{U}_i$. These two conditions are enough to derive the constrained maximizer (h_1^c, h_2^c). The superscript c indicates the constrained solution. Moreover exploiting the positivity of the Lagrangian multiplier we obtain

$$h_1^c > \frac{\alpha\gamma}{K} h_i + \frac{\beta(1-\gamma)}{K} h_e$$

and

$$h_2^c > \frac{(1-\alpha)\gamma}{1-K} h_i + \frac{(1-\beta)(1-\gamma)}{1-K} h_e$$

Thus, for the parameter combinations such that the constraint is binding, the constrained solution is certainly preferable to the internal examiner compared to what he would have obtained if the constraint were not protecting himself. Indeed, the internal examiner, obtains $U_i^{Ac} = -\overline{U}_i$ regardless of his relative bargaining power and the values of $h_i$and h_e. Correspondingly, the utility of the external examiner, denoted U_e^{Ac} is certainly lower compared to what he could obtain were the constraint not active. The value of the problem is $W^{Ac} = \gamma(-\overline{U}_i) + (1-\gamma)U_e^{Ac}$ and note that it is certainly smaller than the value of the unconstrained problem for the same parameters profile.

Next, consider the failure regime.

In this case (h_1, h_2) are such that $U_i \leqslant -\overline{U}_i$. The internal examiner is protected and obtains $-\overline{U}_i$, but no qualified candidate will be appointed. This implies that the external examiner can freely decide about the candidates, however at the additional cost δ that represents the utility loss due to the fact that no one of his preferred candidates will be immediately appointed.

In this regime the internal solution corresponds to the bliss point of the external examiners, i.e.

$$h_1 = h_2 = h_e$$

At the internal solution $U_i < -\overline{U}_i$ (strictly smaller) and particularly $U_i(h_e, h_e) = -(h_i - h_e)^2$. Therefore, the condition for an internal solution can be expressed as

$$(h_i - h_e)^2 > \overline{U}_i$$

The corresponding value of the problem is

$$W^F = \gamma(-\overline{U}_i) + (1-\gamma)(-\delta)$$

For parameters configuration such that the constraint is binding, i.e. when $(h_i - h_e)^2 \leqslant \overline{U}_i$, the solution will correspond to the point on the constraint $U_i = -\overline{U}_i$ that is closer to the external examiner bliss point. Note that since internal examiner's preferences do not play any role in this choice, the constrained solutions of the failure and the appointment regimes will not coincide in general. The value of the problem is

$$W^{Fc} = \gamma(-\overline{U}_i) + (1-\gamma)(U_e^{Fc} - \delta)$$

where U_e^{Fc} is external examiner's payoff at the constraint solution. The value of W^{Fc} will be certainly smaller than the value of the unconstrained problem for the same parameters profile.

Finally, comparing the two regimes, we have:

- For $\overline{U}_i < (h_i - h_e)^2 < \frac{\overline{U}_i}{\Gamma_i}$, it is $W^A > W^F$ if

$$(h_i - h_e)^2 < \Delta = \frac{\gamma\overline{U}_i + (1-\gamma)\delta}{P} \tag{2.16}$$

  The quantity Δ belongs to the interval $\left[\overline{U}_i, \frac{\overline{U}_i}{\Gamma_i}\right]$ if

$$\overline{U}_i \geqslant \frac{\Gamma_i}{\Gamma_e}\delta \tag{2.17}$$

  and otherwise it is larger than $\frac{\overline{U}_i}{\Gamma_i}$. In all cases either the appointment or the failure internal solution is optimal.

- for $(h_i - h_e)^2 > \frac{\overline{U}_i}{\Gamma_i}$ we compare the appointment regime's constrained solution with the failure regime's internal solution. If condition (2.17) holds, the latter certainly dominates. Indeed, for all $(h_i - h_e)^2 > \Delta$ we have that $W^F > W^A$and thus it must also be that $W^F > W^{Ac}$.

- Similarly, for $(h_i - h_e)^2 < \overline{U}_i$, the solution of the problem is the agreement region's internal solution. Indeed for all $(h_i - h_e)^2 < \Delta$ we have that $W^A > W^F$ and thus is must also be that $W^A > W^{Fc}$.

Summing up, if condition (2.17) holds, or, in words, if the cost of failure is sufficiently small, either the internal solution of the appointment regime or the internal solution of the failure regime prevails. Otherwise it is possible to observe, for some $(h_i - h_e)^2 > \frac{\overline{U}_i}{\Gamma_i}$ the appointment regime's constrained solution. Intuitively, when the cost of the failure is large, it cannot be compensated by the additional utility enjoyed by the external examiner in the failure regime, corresponding to the freedom of choosing his preferred candidates.

Appendix B - one qualification

When only one candidate is to be qualified, the model simplifies as follows.

$$U_i = -(h_i - h_1)^2$$

$$U_e = -(h_e - h_1)^2$$

and

$$W = \begin{cases} \gamma U_i + (1-\gamma)U_e & U_i \geqslant -\bar{U}_i \\ \gamma(-\bar{U}_i) + (1-\gamma)\left(U_e - \delta\right) & U_i < -\bar{U}_i \end{cases}$$

Also in this case we treat separately the appointment and the failure regime and we assume $h_i > h_e$.

In the appointment regime, which occurs when $(h_i - h_1)^2 \leqslant \bar{U}_i$, the internal solution is

$$h_1 = \gamma h_i + (1-\gamma)h_e$$

and

$$U_i = -(1-\gamma)^2(h_i - h_e)^2$$

$$U_e = -\gamma^2(h_i - h_e)^2$$

$$W^A = -\gamma(1-\gamma)(h_i - h_e)^2$$

This regime prevails if

$$(h_i - h_e)^2 < \frac{\bar{U}_i}{(1-\gamma)^2}$$

Otherwise, for $(h_i - h_e)^2 \geqslant \frac{\bar{U}_i}{(1-\gamma)^2}$ the constrained solution occurs, which satisfies the Lagrangian $L = \gamma U_i + (1-\gamma)U_e - \lambda(-\bar{U}_i - U_i)$. At the constrained solution h_1^c satisfies $(h_i - h_1^c)^2 = \bar{U}_i$ and it is larger than it would have been in the unconstrained problem. This guarantees the internal examiner's utility to remain at $U_i = -\bar{U}_i$ at the expense of the external examiner utility denoted again U_e^c. The value of the problem is $W^{Ac} = \gamma(\text{-}\bar{U}_i) + (1-\gamma)U_e^c$.

The failure regime occurs when $(h_i - h_1)^2 \geqslant \bar{U}_i$. The internal solution corresponds to the bliss point of the external examiner

$$h_1 = h_c$$

The value of the problem is

$$W^F = \gamma(-\bar{U}_i) + (1-\gamma)(-\delta)$$

At the bliss point the constraint is slack and equals to $(h_i - h_e)^2 > \bar{U}_i$.

When $(h_i - h_e)^2 \leqslant \bar{U}_i$ the internal solution is not viable. At the constrained solution the external examiner does not obtain his bliss point. The selected candidate h_1^c is such that $(h_i - h_1^c)^2 = \bar{U}_i$ and the external examiner's utility is U_e^c.[19] The value of the problem is

$$W^{Fc} = \gamma(-\bar{U}_i) + (1-\gamma)(U_e^c - \delta)$$

Comparing the two regimes:
For $\bar{U}_i < (h_i - h_e)^2 < \frac{\bar{U}_i}{(1-\gamma)^2}$ we have $W^A > W^F$ when

$$(h_i - h_e)^2 < \Theta = \frac{\gamma\bar{U}_i + (1-\gamma)\delta}{\gamma(1-\gamma)} \tag{2.18}$$

[19]Compared to Appendix A, external examiner's utility at the constrained solution is the same in both regimes. This fact occurs because no substitution among candidates is possible in the problem with only one qualified candidate.

The quantity Θ is internal to $\left[\bar{U}_i, \frac{\bar{U}_i}{(1-\gamma)^2}\right]$ for

$$\bar{U}_i > \frac{(1-\gamma)^2}{\gamma^2}\delta \tag{2.19}$$

When condition (2.19) holds, either the internal solution of the appointment regime or the internal solution of the failure regime solve the maximization problem. Indeed, for $(h_i - h_e)^2 > \frac{\bar{U}_i}{(1-\gamma)^2}$, we have that $W^{Ac} < W^F$(since W^{Ac} would be dominated by W^A if the constrained were not active) and for $(h_i - h_e)^2 < \bar{U}_i$, we have that $W^A > W^{Fc}$ (since W^{Fc} would be dominated by W^F if the constrained were not active).

Finally, note that the case with a single qualification can be derived from the case with two qualifications by setting $\alpha = \beta = 1$. Indeed, in this case, $\Gamma_i = (1-\gamma)^2$, $\Gamma_e = \gamma^2$ and $P = \gamma(1-\gamma)$.

Appendix C - elasticity of the threshold.

The small size of the samples at hand has forced us to pool together *concorsi* with elected and randomized committees. As a result, the empirical model estimated over the pooled sample assumes that the threshold separating agreement from failure is the same in both types of procedures. Since the threshold depends on γ, this restriction is questionable, unless the threshold is little sensitive to variations of γ.

In this appendix we compute the elasticity of $\sqrt{\Delta}$, (see equation 2.16), with respect to γ, denoted $\varepsilon_{\sqrt{\Delta}\gamma}$ Such elasticity is a rather complicate function of γ and of all other parameters. To simplify its expression, consider the following reparametrization (up to a small loss of generality). Let be $\alpha = \frac{1}{2} + \frac{m}{2}$, $\beta = \frac{1}{2} - \frac{m}{2}$ so that $m = \alpha - \beta$.

A linearization of $\varepsilon_{\sqrt{\Delta}\gamma}$ around $\gamma = \frac{1}{2}$, the midpoint of the domain of γ, yields

$$\begin{aligned} \varepsilon_{\sqrt{\Delta}\gamma}(\gamma) \simeq \quad & \left(\tfrac{1}{2} - \tfrac{\delta}{\bar{U}_i+\delta}\right) + \\ & +2\left[(1-m^2) + 2\tfrac{\delta}{\bar{U}_i+\delta}\left(\tfrac{1}{2} - \tfrac{\delta}{\bar{U}_i+\delta}\right)\right]\left(\gamma - \tfrac{1}{2}\right) \end{aligned} \tag{2.20}$$

Since at $\gamma = \frac{1}{2}$, it is $\overline{U}_i > \delta$ (by condition 2.17), both the intercept and the slope coefficient are positive.

In particular, at $\gamma = \frac{1}{2}$, the elasticity simplifies to

$$\varepsilon_{\sqrt{\Delta}\gamma}\left(\gamma = \frac{1}{2}\right) = \frac{1}{2} - \frac{\delta}{(\overline{U}_i + \delta)}$$

and it is bounded above by $\frac{1}{2}$.

Moreover, the elasticity is monotonically decreasing in δ and largest when $\delta = 0$. In this case, the least favorable, expression (2.20) becomes

$$\varepsilon_{\sqrt{\Delta}\gamma}(\gamma) \simeq \frac{1}{2} + 2(1 - m^2)\left(\gamma - \frac{1}{2}\right)$$

Hence, at $\gamma = \frac{3}{4}$ it is $\varepsilon_{\sqrt{\Delta}\gamma}(\frac{3}{4}) = 1 - \frac{1}{2}m^2$ and at $\gamma = \frac{1}{4}$ it is $\varepsilon_{\sqrt{\Delta}\gamma}(\frac{1}{4}) = \frac{1}{2}m^2$. We conclude that for a large range of values of γ, $\sqrt{\Delta}$ is inelastic.

The elasticity of $\sqrt{\Theta}$ - the corresponding threshold in the single qualification case (see equation 2.18) - behaves similarly.

Appendix D - truncation adjustment.

In this section we derive the adjustment to account for truncation in equations (2.12) and (2.13). Define:

$$\begin{aligned} I^* = h_i - h_e &= \theta_0 + (H_i - H_e) + \theta_1 GPS1(X) + \theta_2 GPS2(X) + \varepsilon \\ &= \Omega + \varepsilon \end{aligned}$$

Its expectation conditional on agreement is

$$E\left(I^* | -\sqrt{\Delta} \leqslant I^* \leqslant \sqrt{\Delta}\right) = \Omega + \Xi$$

where

$$\begin{aligned} &\Xi = E\left(\varepsilon | -\sqrt{\Delta} \leqslant I^* \leqslant \sqrt{\Delta}\right) = \\ &= \sigma\left[\phi\left(\frac{-\sqrt{\Delta}}{\sigma} - \frac{\Omega}{\sigma}\right) - \phi\left(\frac{\sqrt{\Delta}}{\sigma} - \frac{\Omega}{\sigma}\right)\right] / \left[\Phi\left(\frac{-\sqrt{\Delta}}{\sigma} - \frac{\Omega}{\sigma}\right) - \Phi\left(\frac{\sqrt{\Delta}}{\sigma} - \frac{\Omega}{\sigma}\right)\right] \end{aligned}$$

is the conditional expectation of $\varepsilon \sim N(0, \sigma^2)$, ϕ and Φ are the density and the cumulative distribution of a standard normal respectively.

In the pre-reform period, equation (2.12) is

$$H_1 - H_e = \pi_{11}(H_i - H_e) + \\ + (\pi_{11} - 1)(\theta_0 + \theta_1 GPS1(X) + \theta_2 GPS2(X)) + (\pi_{11} - 1)\varepsilon$$

and its expectation conditional on agreement is

$$E\left(H_1 - H_e | -\sqrt{\Delta} \leqslant I^* \leqslant \sqrt{\Delta}\right) = \pi_{11}(H_i - H_e) + \\ + (\pi_{11} - 1)(\theta_0 + \theta_1 GPS1(X) + \theta_2 GPS2(X)) + (\pi_{11} - 1)\Xi \qquad (2.21)$$

Parameters of equation (2.21) can be estimated by means of the following adjusted regression

$$\widetilde{(H_1 - H_e)} = \pi_{10} + \pi_{11}\widetilde{(H_i - H_e)} + \pi_{13} GPS1(X) + \pi_{14} GPS2(X) + \xi \quad (2.22)$$

where

$$\begin{aligned} \widetilde{(H_1 - H_e)} &= (H_1 - H_e) + \Xi \\ \widetilde{(H_i - H_e)} &= (H_i - H_e) + \Xi \end{aligned} \qquad (2.23)$$

and ξ is a IID error term. Similarly we proceed after the reform, and for equation (2.13) and equation (2.14).

Chapter 3

Budget cuts and brain drain: the case of Italian universities

> It appears evidently from experience that a man is of all sorts of luggage the most difficult to be transported
>
> A. Smith, 1776

3.1 Introduction

Italy is one of the OECD countries with the highest rate of skilled labour emigration. In 2000, almost half million of high skill Italians resided abroad (Docquier and Marfouk, 2004). Such outflows coupled with the inability to attract skilled foreigners led many authors to point at an "Italian brain drain" (Becker et al., 2004; Biondo et al., 2012). High skill migration is a particular economic concern because of the high human capital transfers involved and its impact on productivity (Hornung, 2014) and innovation (Chellaraj et al., 2008; Hunt and Gauthier-Loiselle, 2010). Universities are generally regarded as fundamental for innovation, more

so in Italy where over 30% of research and development personnel works at universities (Istat, 2014). In this paper we focus specifically on these subjects and investigate whether the deterioration of careers prospects induced by budget constraints prompted researchers to leave Italian academia. Indeed, severe personnel expenses cap were selectively imposed on Italian universities in 2009. The new regulation prevented institutions with staff expenses above a certain threshold from hiring and promoting professors. Not surprisingly, the norm induced a reduction in new hires and promotions that brought the number of professors employed in Italian universities down by 15% between 2008 and 2013. Besides the block on entries, the decline resulted also from an increasing number of professors leaving academia. While aging and changes in the recruitment process may explain part of the outflows, both its timing and magnitude deserve further analysis. In this paper we investigate whether budget constrains on personnel expenses backfired by inducing Italian professor to quit their positions for private companies or foreign institutions. We argue that a prolonged alt in promotions may have affected careers outlooks and may have reduced expected earnings of Italian professors. Differentials in labour incomes are indeed regarded as the most important economic determinants of migration (Sjaastad, 1962, Borjas, 1991). Besides expected earnings, whether individuals decide to quit their positions depends, among others, on the size of outside opportunities. With this respect, Italian researchers were not the only one dealing with reduced funding. Similar concerns rose, for example, in the US during the negotiation between democrats and republicans over debt-ceiling deals (Gulledge, 2011). Yet in Italy these budget constraints were planned on multiple years (in fact they are still in place) and they were relatively easy to predict resulting possibly in greater impact on expectations. Empirical work on the link between economic crisis, public universities funding and potential outflows of researchers is surprisingly scarce. To fill this gap, we exploit the selectivity of the new regulation along with its time variation in a quasi-experimental framework. Preliminary results indicate an increase in voluntary leaves among institutions with more severe staff limitations. Surprisingly the latter show lower voluntary leaves on

average in both pre and post reform periods. We provide some explanations for these results and point out the limits of our methodologies for future investigations. The rest of the paper is organized as follows: section 3.2 presents the institutional setting and the dataset, section 3.3 summarizes the main trends in Italian universities in terms of promotions, new hires and voluntary leaves, section 3.4 presents the econometric models and their preliminary results, section 3.5 briefly discuss their common findings and limits along with alternative explanations. Finally section 3.6 concludes.

3.2 Institutional setting and dataset

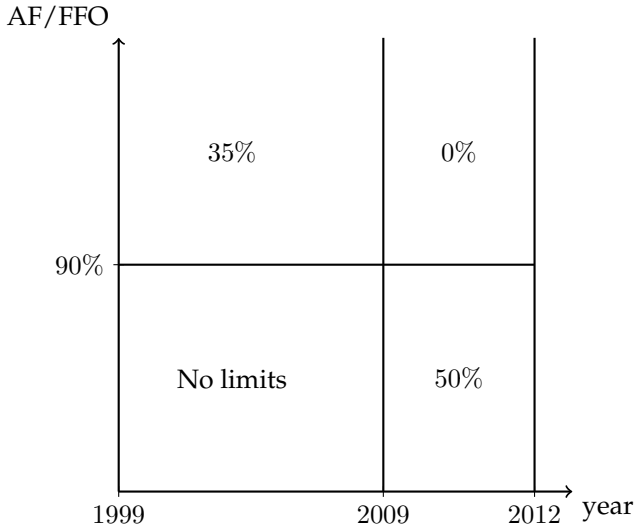
Italian universities fund their activities mostly through transfers from the central government. The larger among these goes under the name of "ordinary fund" (FFO[1]) possibly due to the fact that it is determined on historical basis. The high share of resources spent on personnel is among the reasons for the survival of such funding scheme. Indeed, cutting this fund in favor of, for example, mechanism tied to research performance, would expose many university to potential deficits given the rigidity of personnel expenses. An attempt to constrain staff expenses was fist set in 1999 when the Italian legislator set the limit of staff turn-over to 35% of the savings from retirements for those universities with personnel expenses exceeding 90% of the FFO[2]. Despite that, personnel expenses kept increasing inasmuch the legislator made the prescription more severe in 2009. From then until 2012, institutions exceeding the 90% prescription are prevented from hiring (0% turn-over) while those below the threshold are subject to a 50% limit. Finally, a more complex index to determine turn-over regimes was set in 2012[3]. Figure 4 summarizes this setting. The graph shows how maximum percentage staff turn-over varies with the ratio of personnel expenses over the ordinary fund (AF/FFO) on the y-axis and over time on the x-axis.

[1] From the Italian *fondo di finanziamento ordinario*

[2] Law 449/1997 art. 1 co. 4

[3] This index accounts for, among others, student enrollment fees and debt. See D.lgs 49/2012 art. 5 and 7.

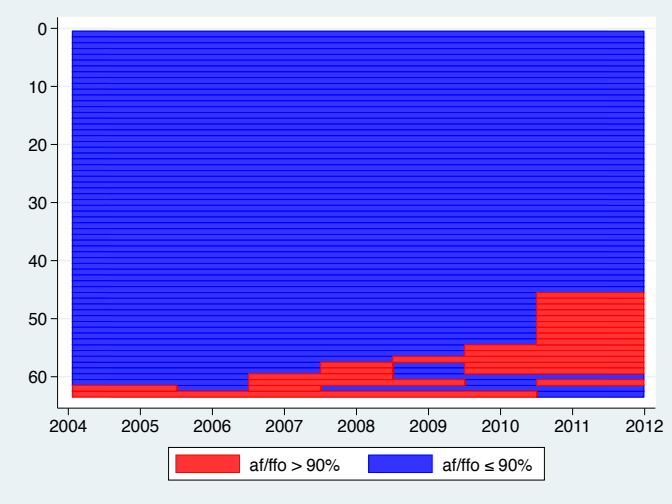
Figure 4: Maximum staff turn over



This paper wants to assess the impact of tighter job market conditions on the movements of researchers in Italian universities. For this reason we compare two legislative periods: i) *pre-reform* from 1999 to 2008 and ii) *post-reform* from 2009 to 2012. The data we need are essentially of two types. The first is financial data on university and in particular on the ratio between staff expenses and ordinary fund essential for determining which regime applies. This data is available for the period 2004 - 2011 and was provided directly from the Minister of Education. Interestingly, the minister had no data for the period 1999 - 2003, although personnel caps were already in place. This possibly points to low or even no enforcement of turn-over regulations in early periods. With respect to the period covered in the data, the majority of institutions are below the limit. The number of those crossing it gradually increases over time while a minority of institutions repeatedly move back and forth as shown in figure 5.

The second type of data regards the career of Italian researchers. With this respect, we track inflows, outflows and movements across Italian

Figure 5: Institutions exceeding staff limits



universities for each professor in the period 1996 - 2013 thanks to micro data also made available by the Minister of Education. Crucially, any outflow is motivated so that we can distinguish retirements and resignations from other personal reasons to leave. Table 7 summarizes the main trends relating professors movements. Note that the overall number of professors reaches a peak in 2008 and then decreases drastically due to a combined reduction in entries and increase in exits. Movements across universities are negligible throughout the period while also showing a contraction starting in 2009. When combined together, these sources of information provide us with a complete picture only from 2005 to 2011[4]. Nevertheless, we use the longer panel on professors in many descriptives and in part of the econometric analysis whenever the assumptions needed are plausible.

[4]Regulation looks at previous year personnel expenses so that, with this respect, 2004 is lost.

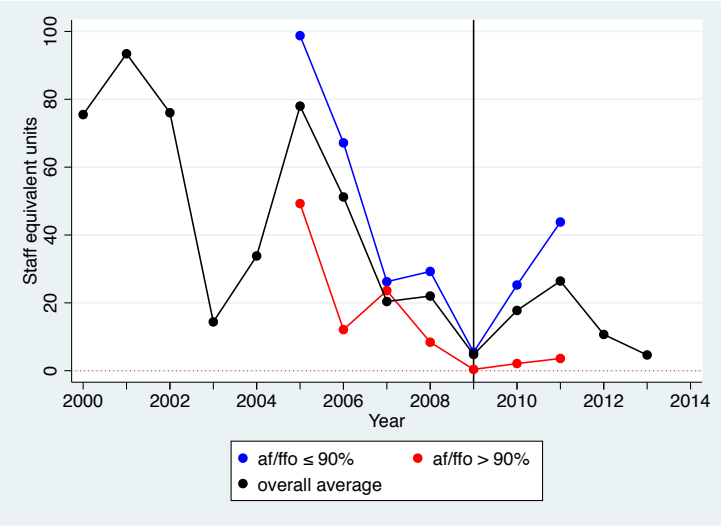
Table 7: Dataset overview

Year	Present	Inflows	Moves	Outflows	of which:	voluntary leave	retirement	other exits
1996	49469	1205	514	1177		259	731	187
1997	49187	787	584	1070		488	400	182
1998	49956	1546	2150	770		189	391	190
1999	50767	1765	758	947		302	440	205
2000	51988	2251	630	1022		352	473	197
2001	54852	3918	1007	1061		366	530	165
2002	57516	3753	683	1081		313	573	195
2003	56458	221	325	1253		463	600	190
2004	57363	2009	434	1118		407	540	171
2005	60240	4015	604	1141		375	595	171
2006	61986	3109	422	1381		534	678	169
2007	61908	1531	273	1599		630	796	173
2008	62768	2669	384	1822		573	1064	185
2009	60858	492	185	2378		831	1396	151
2010	57735	1194	115	4325		1340	2850	135
2011	56451	1454	244	2739		1005	1599	135
2012	54925	579	174	2101		704	1275	122
2013	53530	10	102	1546		396	1047	103

3.3 Descriptives

A drastic increase in outflows is visible in table 7 in 2009. The trend possibly starts in 2005 and reaches its maximum in 2010 when outflows are roughly four times the average in 2000-2005. Nevertheless, these exits are highly heterogeneous and refer to retirements, voluntary leaves, disciplinary lay-offs etc. The first two account for the majority of exits and show a similar pattern over time. While both exit decisions may be affected by the new regulation, we concentrate particularly on voluntary leaves in the sections that follow. Indeed, this is the most intuitive channel through which individuals may want to leave academia as a reaction to a worsening of careers outlook. With this respect, figure 6 reports increments in staff expenses due to promotions and new hires. The implications of the picture are twofolds. First, individuals face tighter labour market conditions with few advancements in career possible in the overall market. Second, institutions with higher imbalances (red lines) seem to contract increments in expenses more than the blue ones with no or few (after 2009) constraints. Despite the figures are generally in line with the regulatory setting in place, the latter still does not explain why insti-

Figure 6: Increments in personnel costs: promotions and new hires



Note: Staff equivalent units (y-axis) are based on the following equivalences: 1 full professor, 0.7 associate professor, 0.5 assistant professor. E.g. a promotion from associate to full professor would cost 0.3 equivalent units. Red and blue lines are limited due to lack of data on universities expenditures.

tutions below the limit should reduce promotions and new hires before 2009, when they are completely unconstrained. This may point to some anticipation effect or other confounders that we consider in the next sections.

3.3.1 Individual level evidence

We ask whether a worse job market outlook induced by the new regulation triggered extensive outflows from Italian academia. Whether this is the case depends on various factors such as information, rationality and outside options. Complete information allows individuals to predict worse career outlook and to form correct expectations with respect to future earnings. On the other hand, rationality and the existence

of alternatives abroad and outside academia determine how individuals react to these expectations. The higher the outside option the more likely researchers will leave their current positions. In the present setting, widespread political debate over the new regulation coupled with a simple and highly predictable prescription seems to be an ideal setting for individuals to form accurate expectations. Therefore, individuals in institutions above the personnel expenses limit should have worse expectation *ceteris paribus*. On the other hand, the presence of appealing outside options is more problematic. Indeed, the economic downturn reduced both chances of getting academic positions abroad (e.g. Carulli, 2013) and possibilities of being hired in the private sector. Further, the extent of outside options should clearly vary along scientific fields with "hard sciences" being generally less affected by the crisis (Oreopoulos et al., 2012) and more international disciplines such as medicine and engineering showing smaller costs of migration. Before moving to a more rigorous analysis we give a sense of the impact of such variability on the probability to voluntary leave in table 8. We report probit average partial effects separately for 1) pre and 2) post reform of the following equation:

$$y_i = \alpha + \beta\, age_i + \boldsymbol{\gamma}\, \mathbf{area_i} + \boldsymbol{\delta}\, \mathbf{seniority_i} + \boldsymbol{\eta}\, \mathbf{university_i} + \varepsilon_i$$

where y_i is a binary variable equal to 1 if individual i voluntary leaves university, $\mathbf{area_i}$ is a vector of dummies for scientific areas[5], $\mathbf{seniority_i}$ controls for individual i being assistant, associate or full professor and $\mathbf{university_i}$ is a vector of university controls. For simplicity we neglect the time dimension of data here. Time varying factors such as age and seniority are set equal to their levels in i) the year when the individual left or in ii) some random year for those who have never left.

We set as reference categories those with lower propensity to leave and indicate them with a dot (.) in table 8. These categories are civil engineering and architecture among scientific disciplines and full professor among seniority levels. As anticipated hard sciences show an higher propensity to leave along with medicine and economics. As for seniority we see two opposite effects: a positive one for age and a negative

[5]as defined by National University Council (CUN)

Table 8: Probability of voluntary leave

	(1)	(2)
	APE pre	APE post
age	0.00625***	0.00838***
	(0.000)	(0.000)
math	0.0308***	0.0355***
	(0.005)	(0.006)
fisics	0.0152***	0.0126*
	(0.005)	(0.005)
chemistry	0.0270***	0.0200***
	(0.005)	(0.005)
geology	0.0313***	0.0215**
	(0.006)	(0.007)
biology	0.0283***	0.0258***
	(0.004)	(0.005)
medicine	0.0395***	0.0302***
	(0.004)	(0.004)
agriculture and veterinary	0.0325***	0.0299***
	(0.005)	(0.006)
civil engineering, architecture	.	.
	(.)	(.)
industrial engineering	0.0306***	0.0173***
	(0.005)	(0.005)
litterature	0.0122***	0.0258***
	(0.004)	(0.004)
history, filosofy and psico	0.00946*	0.0228***
	(0.004)	(0.005)
law	0.00635	0.00832
	(0.004)	(0.005)
econ and statistics	0.0227***	0.0242***
	(0.005)	(0.005)
political sciences	0.00406	0.00254
	(0.005)	(0.006)
assistant professor	0.0869***	0.0759***
	(0.003)	(0.003)
associate professor	0.0362***	0.0365***
	(0.002)	(0.002)
full professor	.	.
	(.)	(.)
male	-0.0118***	-0.0160***
	(0.002)	(0.002)
Observations	65744	64470

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

one for tenure. The sign for associate and full professors are negative as expected since these individuals should have lower incentives to quit thanks to relatively established careers. Despite that, the sign of age is puzzling here since we expect youngsters to be more likely to leave due, for example, to weaker familiar ties. On the contrary, the sign of the coefficient is positive and its marginal effect is non negligible in magnitude. Finally, little changes occurs between pre and post reform.

3.3.2 University level evidence

We now move to investigate university variation. In particular, we want to see whether being subject to tighter personnel regulations results in higher dropouts. Figure 7 shows voluntary leave rates for university below and above the 90% limit separately for assistant, associate and full professors.

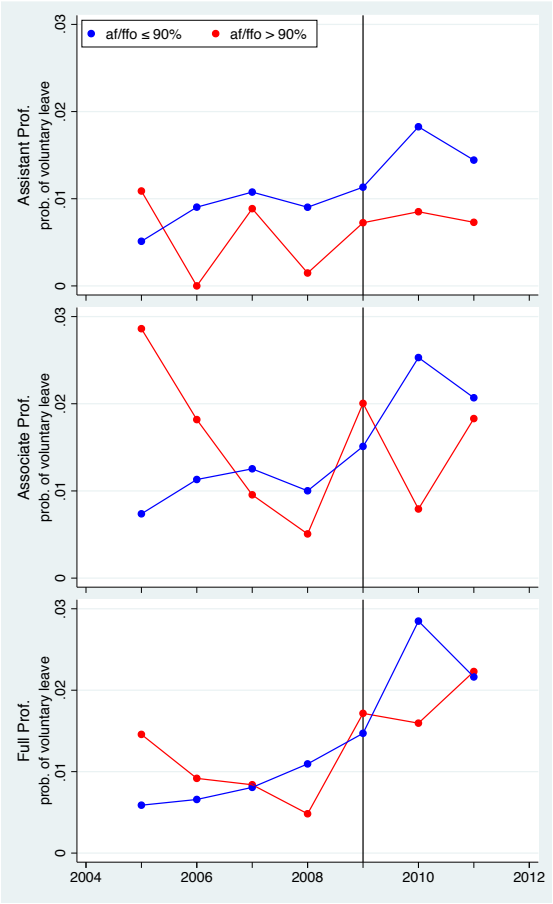
In all three plots the blue patter shows hikes in 2010: exactlty one year after the new regulation was set in place. On the other hand, institutions with higher personnel expenses (red line) present an unstable pattern with higher variation over time and a clear increasing trend only for full professors. Besides variation, these institutions also show lower exits compared to their reference group. This puzzling result maybe due to low cardinality / high sample variability of the group or to institutions switching from one group to the other one or more times as we reported in figure 5. While little can be done against the first potential problem the second one may be easily addressed by restricting the sample. These and other possible solutions are investigated in the next section.

3.4 Methods

We use terminology and notation developed in experimental analysis (Rubin, 1974) to formalize the relation between new regulation and voluntary leaves. Define the treatment for unit i at time t as:

$$D_{it} = 1(af/ffo_{it-1} > 90\%)$$

Figure 7: Incidence of voluntary leaves by seniority



Note: Weighted averages.

That is unit i is treated if personnel expenses in the previous year exceeded the regulatory limit[6]. Also define a binary variable $P_t = 1(t \geq 2009)$ to describe the introduction of tighter rules in 2009. Further assume for simplicity a linear relation of the type:

$$y_{it} = \beta_0 + \beta_1 D_{it} + \beta_2 (D_{it} \times P_t) + \boldsymbol{\beta}_3 \mathbf{C}_{it} + \varepsilon_{it} \tag{3.1}$$

where $\mathbf{C}_{it}$ is a matrix of controls including the vectors for seniority, university, scientific fields and year controls. Equation 3.1 is know as a difference in difference estimator (e.g. Card, 1992) because the interaction parameter β_2 compares the change in the time dimension (before minus after) in the outcome variable in the treatment group with the same change for the control group. In the present setting this amount to compare the propensity to leave before and after the introduction of new regulation (2009) for university below and above the threshold of personnel expenses. In the absence of confounders, the coefficient β_2 identifies the *causal effect* of the new regulation on the propensity to leave the system.
Specialized readers may notice a slight variation to the standard framework in equation 3.1. Indeed, we are including university controls in $\mathbf{C}_{it}$ that would be collinear with D_{it} were not for the fact that some universities happened to switch group (i.e. move below and above the threshold) over time (see figure 5). The possible implications of this peculiarity will be investigated in more details in section 3.4.1.
Before putting equation 3.1 into operation, we need to define the basic statistical unit i. This is non trivial since the determinants on the right hand side of equation 3.1 vary at different levels of aggregation. Indeed, treatment status D_{it} varies at university level while seniority and scientific field vary at individual level. Further, it would be tempting to use micro data at individual level but unfortunately we would derive misleading standard errors for the β_2 coefficient we are most interested

[6] An alternative specification with a continuous variable af/ffo_{it-1} was also investigated yielding similar results. Note that the latter is quite an appealing definition of treatment. Indeed, professors may update their career expectations progressively as their university approaches the threshold rather than abruptly once it is overcome. This specification yield similar results so that we present the simpler and more common binary specification of treatment.

in. Indeed, if we take individual as reference unit, the data are unlikely to be independent across observations. The propensity to leave tent to be correlated among professors from the same university because they share common characteristics and are exposed to the same environment. Moulton (1986) shows that by assuming a simple additive structure with an individual component and a group component the resulting standard errors increase sharply as the intra-class correlation and the group size increases. While more elaborate solutions exist[7], the most simple one is to use groups as fundamental statistical units in place of individuals. This amounts to set i as a generic university in equation 3.1 and to compute shares of individual varying factors such as seniority, scientific fields and voluntary leaves.

Estimates of equation 3.1 at university level are reported in table 9. The baseline in column 1 combines all levels of seniority while the remaining columns stratify on that. The interaction coefficient is positive as expected indicating that the new rules incremented voluntary leaves in universities with personnel imbalances of $.008$ more than those below the regulatory limit. Also the magnitude appears relevant given that that voluntary leaves are usually $1-2\%$ per year as reported in table 7 so the coefficient translates in an increase of $50-100\%$. Once we stratify by levels of seniority in the remaining columns, the significance seems to be driven mainly by assistant professors (column 2). Despite such increase, it is surprising that treated university experience lower average voluntary leaves before and even after the reform. To see this note the coefficient for D_{it} is always negative and almost always greater in magnitude than that for the interaction term. Leaving aside the effect of the reform, we need to figure out why professors in institutions with budget imbalances systematically show a lower propensity to voluntary leave their institutions. If we believe our model correct, then a tentative explanation maybe that these institutions are most likely of lower quality compared to those with sound budgets. Then their professors face worse outside opportunities and therefore are more strongly tied to their cur-

[7]See for example Angrist and Pischke (2008, pp. 233-235).

Table 9: Difference in difference

	(1) baseline	(2) assistant p.	(3) associate p.	(4) full p.
D_{it}	-0.0084**	-0.0131**	-0.0035	-0.0077
	(0.0032)	(0.0047)	(0.0051)	(0.0052)
$D_{it} \times P_t$	0.0080*	0.0110*	0.0035	0.0099
	(0.0036)	(0.0053)	(0.0057)	(0.0058)
β_0	-0.0215	-0.2043**	-0.2245	-0.1501
	(0.1327)	(0.0738)	(0.1145)	(0.1516)
% scientific field	Yes	Yes	Yes	Yes
% seniority	Yes	No	No	No
university controls	Yes	Yes	Yes	Yes
year controls	Yes	Yes	Yes	Yes
N	455	446	447	455
R^2	0.633	0.427	0.526	0.542

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

rent position [8]. Then despite the reform affects their career perspective more on average than colleagues in other institutions as measured by the interaction coefficient, the effect on expectations does not overcome the lower outside options these individual face in both pre and post reform periods.
Nevertheless, on the econometric side, there may be good reasons to delve further in equation 3.1. First, akin to survival analysis the decision to leave is clearly time dependent in that individuals must trivially have not left before. This time dependence introduces a dynamic structure in our model. Secondly, treated and untreated may be ill defined here. Indeed, each university may have imbalances one year and recover the year after. Ideally we should define a plan of treatments (Robins et al., 1999) and compare each plan with a slight variation of it. Finally, we may believe not the unconfoundness assumption and specifically point to each university having its own time dependent vicissitudes, or in technical terms its own trend, before and after the reform. These concerns will drive our analysis in the subsequent sections.

3.4.1 Structural break

In the present setting the treatment status is not homogeneously distributed within the two groups. Indeed, institutions generally experiencing different length of budget imbalances as reported in figure 5. With this respect, the difference in difference model of equation 3.1 neglects the history of treatment up to time t so that institutions with a long history of imbalances such as "L'Orientale di Napoli" are assumed to experience the same treatment effect as relatively sound institutions with just temporary excesses. Ideally, we would have many observations for each "treatment plan" and compare plans with a similar sequence of budget imbalances (Robins et al., 1999). Unfortunately, this is not possible in the present setting given the small number of observations and the marked differences in exposure to treatment. Following this line of reasoning, we may doubt the existence of suitable comparison groups, focus only

[8] This is similar to results found in managerial literature on declining organizations (Bedeian and Armenakis, 1998).

on one of them and exploit only the time variation in voluntary leaves. This model is known as structural break as it parametrizes a "jump" in a functional relation at a specific point in time. Identification depends on the absence of confounders other than the reform, that may be responsible for such "jump". Operatively, we pick up the most numerous group, namely the never treated, consisting of 42 university. We then extend our time span as much as we can in order to obtain reliable estimates of time trends. This comes at the cost of assuming that institutions never treated in the period 2005-2011 were also not treated in the period 1999-2004 for which financial data is missing[9]. This allow us to extend the time frame to the period 1996-2013 and possibly to obtain better estimates.

The plot of figure 8 gives us some graphical insights on how to parametrize the "jump" in the functional relation.

Note the pattern in figure 8 appears to increase in the years immediately following the reform and to decrease thereafter with a final level possibly below the average one in the pre-reform period. To allow such shift we need at least a second order polynomial function of the type:

$$y_{it} = \beta_0 + \beta_1 time\,from\,reform_t + \beta_2 time\,from\,reform_t^2 + \boldsymbol{\beta}_3 \mathbf{C}_{it} + \varepsilon_{it} \tag{3.2}$$

where $time\,from\,reform_t = (t - 2008)^+$ are the number of years after reform. Estimates are reported in table 10.

The positive and significant coefficient for $time\,from\,reform$ indicates an increase in dropouts after the introduction of the new personnel regulation. This trend subsequently fades out and becomes negative by 2011 as indicated by $time\,to\,reform^2$.

These results are useful in quantifying the magnitude of the increase but require further investigation. A more robust analysis would enrich equation 3.2 with parameters for other potential breaks as in Dincecco (2009). In this way, the presence of unexpected significant breaks along the series would cast doubt on the results of table 10 and indicate that our

[9] Actually the requirement is less stringent. We just need the treatment to be homogeneous within the group so that years of imbalance are possible as long as they are common to all universities in the group.

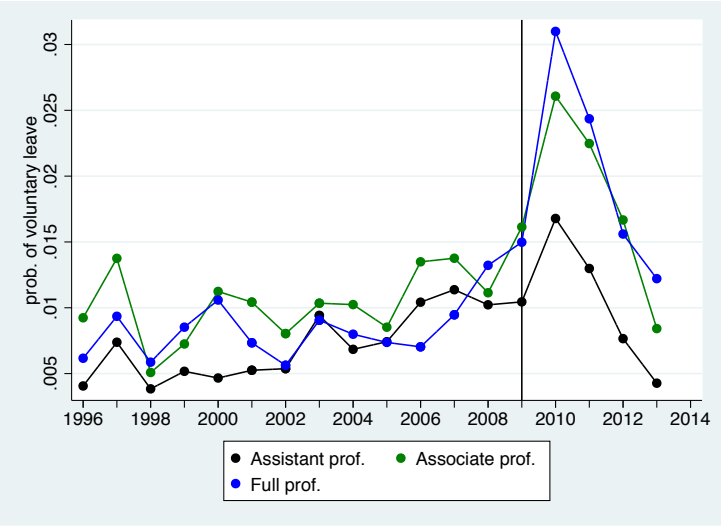
Table 10: Structural break

	(1) baseline	(2) assistant p..	(3) associate p.	(4) full p.
$time\ from\ reform$	0.0070***	0.0030	0.0086*	0.0073*
	(0.0017)	(0.0030)	(0.0034)	(0.0031)
$time\ from\ reform^2$	-0.0013***	-0.0006	-0.0016*	-0.0013*
	(0.0004)	(0.0006)	(0.0007)	(0.0006)
$constant$	-0.0252	0.0282	-0.0288	0.0351
	(0.0348)	(0.0493)	(0.0378)	(0.0412)
% scientific field	Yes	Yes	Yes	Yes
% seniority	Yes	No	No	No
university controls	Yes	Yes	Yes	Yes
year controls	Yes	Yes	Yes	Yes
N	674	657	661	671
R^2	0.518	0.216	0.323	0.388

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Figure 8: Never treated and voluntary leaves by seniority



Note: Weighted averages.

parameters are possibly capturing some other source of variation over time. Finally, were these tests inconclusive, we would still not be able to exclude the presence of confounders responsible for the increase in voluntary leaves around 2009. The difference along the cross section dimension in our diff-in-diff equation 3.1 was indeed in place to control for that. Leaving proper tests on the timing of structural breaks and on its funcional form for future analyses we now move to investigate a method that might be able to control for university-specific trends.

3.4.2 Synthetic control

We started our analysis with a difference in difference approach that exploited many but heterogeneous universities and then possibly overshot by retaining just few of them in the structural break of section 3.4.1. We try now to strike a balance between the two by selecting two relatively homogeneous groups. The first one consists of university with no budget imbalances before 2009 but with personnel expenses above the threshold

in at least one year afterwards (NT/T). The second consists of universities never experiencing the treatment (NT/NT) as in section 3.4.1. The analysis we propose is similar to synthetic control methods developed in Abadie and Gardeazabal (2003). The first group serves as treated unit against which we construct a synthetic control by weighting the characteristics of universities from the second group. The methods is more robust compared to difference in difference in that it allows for different trends in the pre reform period (Abadie et al., 2010). Operatively the construction of the synthetic group passes through the recursive minimization of two quadratic problems:

$$\begin{aligned} &\mathbf{V}^*(\mathbf{w}) = \arg\min_{\mathbf{V}(\mathbf{w})}(\mathbf{x}_1 - \mathbf{X_0}\mathbf{w})'\mathbf{V}(\mathbf{x}_1 - \mathbf{X_0}\mathbf{w}) \\ &\mathbf{w}^* = \arg\min_{\mathbf{w}}(\mathbf{y}_1 - \mathbf{Y_0}\mathbf{w})'(\mathbf{y}_1 - \mathbf{Y_0}\mathbf{w}) \\ &\quad s.t.\ \textstyle\sum_j w_j = 1,\ w_j \geq 0\ \forall j \in J \end{aligned} \tag{3.3}$$

where $\mathbf{x}_1$ is a $k \times 1$ vector of pre intervention determinants of the outcome variable for the treated unit, $\mathbf{X}_0$ is a $k \times J$ matrix obtained by horizontally stacking the same determinants for each unit $j \in J$ in the control group, $\mathbf{y}_1$ is a $T \times 1$ vector of outcomes levels, one for each of the T pre intervention periods and $\mathbf{Y}_0$ is its $T \times J$ analog for the group of controls. The parameters to be estimated are the diagonal matrix $\mathbf{V}$ measuring the relative importance of determinants $\mathbf{x}_1$ in the prediction of pre reform voluntary leaves $\mathbf{y}_1$ and the vector of weights $\mathbf{w}$ measuring the contribution of each control observation $j \in J$ to the synthetic control.

Technically speaking, the first minimization problem along with the constraints on $\mathbf{w}$ serves as a regularization term to avoid overfitting. In this exploratory analysis we simplify the problem and avoid this step by setting $\mathbf{V} = \mathbf{I}$ where $\mathbf{I}$ is the identity matrix.

Finally, synthetic control methods are usually implemented when there exists just one treated unit. In the present setting we have several universities in the NT/T group. Rather than replicating the process for each university in the latter, we use only their average to comply with the usual setting.

Figure 9: Donor pool against treated unit

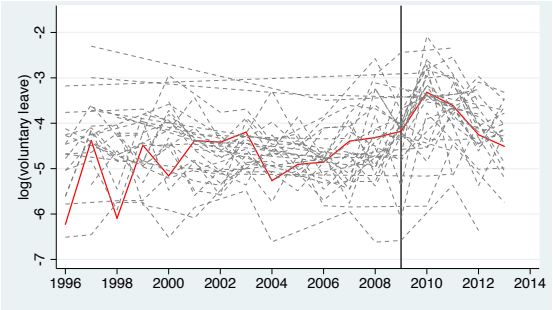


Figure 9 shows the "donor pool" for the synthetic control. The dashed lines represent the trend in log voluntary leaves for NT/NT universities. Their optimal convex combination is used to construct the synthetic analog of the treated unit represented here as the red line. Note the red line falls almost always in the convex hull created by the dashed lines implying that we can expect a good fit in the pre reform period [10].
The optimal weights for the minimization problem 3.3 are reported in table 11. Universities with high weights have voluntary leaves similar to that of the average treated units. These universities are highly heterogeneous in dimensions, geographical location and quality, possibly a consequence of skipping the first minimization problem.

Then voluntary leaves for the synthetic control are computed as:

$$\hat{y}_{1t} = \mathbf{y_{0t}}\mathbf{w}$$

where $\mathbf{y_{0t}}$ is a $1 \times J$ vector of voluntary leaves in the donor poll at time t. Figure 10 plots average voluntary leaves for the treated units against its synthetic analog. Note the two match closely in the pre reform period so that we can leave apart any concern about universities in the

[10] Potentially this is a perfect fit in that we have 40 parameters, one per each university in the donor pool against 14 equations (13 periods to predict plus the constraint on $\sum w$). The small deviations in figure 10 are most likely due to some universities in the donor pool not present in all periods.

Table 11: Synthetic control weights

University	Assistant p.	Associate p.	Full p.
Politecnico di TORINO	0,010	0,146	0,324
della CALABRIA	0,217	0,026	0,006
PALERMO	0,004	0,179	0,003
SASSARI	0,160	0,015	0,004
PARMA	0,009	0,087	0,156
Università IUAV di VENEZIA	0,071	0,126	0,004
MACERATA	0,001	0,001	0,116
TUSCIA	0,102	0,001	0,012
Politecnica delle MARCHE	0,009	0,090	0,017
Scuola Normale Superiore di PISA	0,087	0,000	0,002
VERONA	0,003	0,018	0,087
TRENTO	0,005	0,072	0,085
BERGAMO	0,082	0,005	0,005
Stranieri di PERUGIA	0,069	0,000	0,005
BRESCIA	0,007	0,046	0,004
CAGLIARI	0,009	0,001	0,040
BOLOGNA	0,014	0,005	0,036
ROMA TRE	0,036	0,004	0,004
PADOVA	0,021	0,030	0,006
CAMERINO	0,006	0,025	0,008
CATANIA	0,024	0,002	0,003
Scuola Superiore Sant'Anna	0,006	0,001	0,019
BASILICATA	0,017	0,003	0,003
Politecnico di MILANO	0,006	0,014	0,004
SISSA - TRIESTE	0,001	0,006	0,013
TORINO	0,009	0,004	0,001
MILANO	0,004	0,003	0,009
SALENTO	0,001	0,004	0,008
CHIETI-PESCARA	0,005	0,007	0,007
Stranieri di SIENA	0,001	0,005	0,005
SALERNO	0,004	0,001	0,003
SANNIO di BENEVENTO		0,072	

Note: Green cells represent weights greater than 0.1 while green universities must have at least one green weight.

donor pool not being comparable to those treated. As for the results, we would expect the synthetic control to systematically show lower voluntary leaves than the average treated units if individuals were reacting differently based on the regulatory turn-over cap. Surprisingly, figure 10 show no such pattern with the synthetic control substantially above the average treated one in 2010. In fact, this is in line with results from difference in difference in table 9 where treated units showed on average lower voluntary leaves than universities with higher turn-over cap. Apart from mandatory caution in interpreting results on just few post-

reform periods, these results point to the presence of some confounders that we investigate in the next section.

Figure 10: Synthetic control: assistant p.

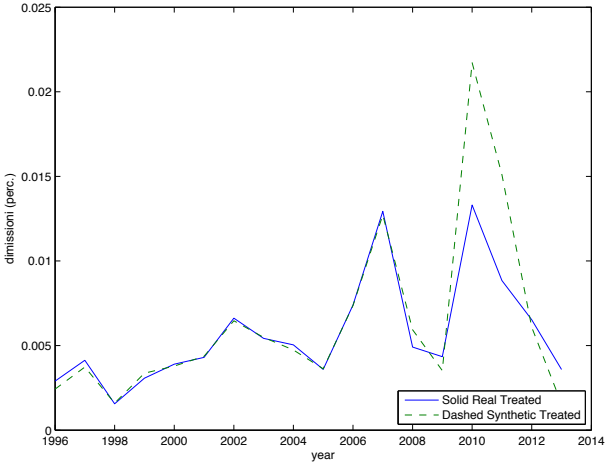


Figure 11: Synthetic control: associate p.

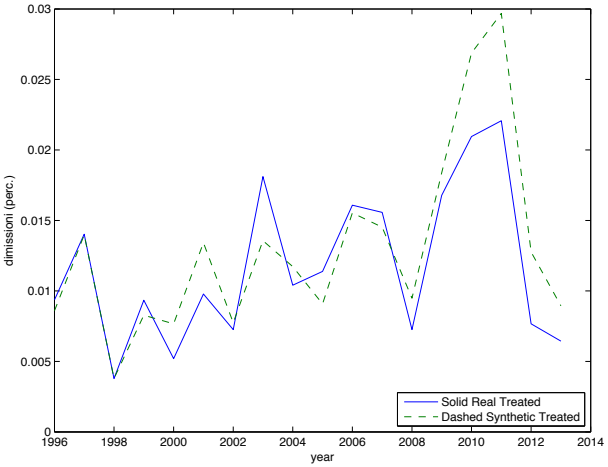
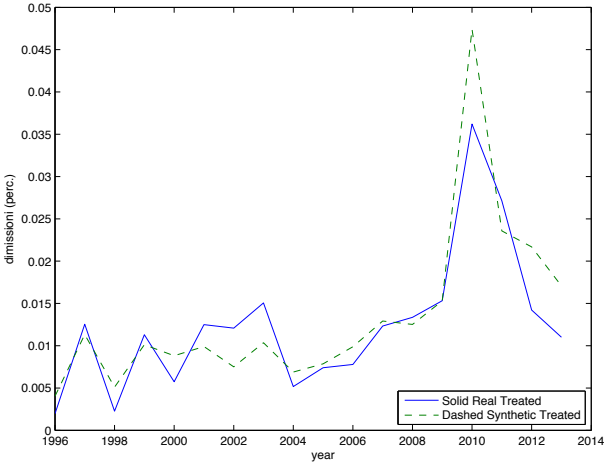


Figure 12: Synthetic control: full p.



3.5 Discussion

Difference in difference and synthetic control methods failed to capture a seizable effect of the reform. Despite the reform appears to increase outflows from treated universities more than untreated ones, such increase still leaves voluntary leaves more frequent on average among "sound" institutions. Since these methods are both based on inter groups (treated vs untreated) and time comparisons (pre vs post reform) they are more robust than structural break models relying solely on time variation. We do observe and measure an increase in voluntary leaves after 2009 but this appear to be a trend common to the whole system with personnel turn-over caps accounting only for little of its variability. On the top of that, such caps appear to be inversely related to researchers dropouts with lower rates occurring in institutions with tighter constraints (table 9). So far we pointed out statistical explanation for these puzzling results, such as variability in treatment sequences, small sample size and few post-reform observations among others.
Nevertheless, were our measures correct how would we explain such

findings? We propose three tentative explanations based on the presence of confounders. First, institutions with persistent personnel imbalances may be of lower quality compared to those with sound budgets. For this reason, their professors face worse outside opportunities and when career expectations deteriorate they are still better off with their current positions. This would explain why both difference in difference and synthetic control reported lower voluntary leaves on average for this group. Remarkably, this is true after controlling for university fixed effects in section 3.4. The second explanation looks at time confounders rather than longitudinal ones. With this respect figure 8 shows an increasing trend in dropouts that started possibly in 2005, a time in which regulation put no constraint on "good" institutions. Such patter is mirrored in figure 6 where expenses for promotions and new hires drop after 2005. Indeed, the legislator introduced a mandatory planning of university activities[11] in 2005. Such planning required universities to communicate, classes, curricula, research goals as well as estimated labour and capital investments to meet them. While this administrative regulation received little attention at the time it might well have affected managerial practice in universities beyond many other policies. Indeed, on-line procedures made any personnel imbalance much more evident to both university and ministerial boards so that the seven years old personnel expenses' cap may practically have come into effect only after that. This would explain why also sound institutions reduced promotions and new hires in the pre reform period when they were totally unconstrained. Finally, voluntary leaves may just be an administrative alternative to retirement. Institutions may push senior members to quit their positions in exchange of mid term temporary contracts covering the last years before actual retirement. This would explain the positive coefficient for age in our micro level regressions (table 8) and a distribution of age increasingly concentrated between 55 and 75 years for those who left in the post reform[12].

[11] art. 1-ter law n. 43/2005

[12] Findings available upon request

3.6 Conclusions

We describe two contemporary trends in Italian universities. The first one is a generalized reduction in new hires and promotions while the second one is a 15% reduction in the number of professors over 2008-2013. We use micro data to investigate the reasons why an increasing amount of professors left academia. While aging and retirement are indeed a driving factor, voluntary leaves are also responsible for a consistent part of the trend. We argue that the increment in voluntary leaves may be a consequence of budget constraints selectively imposed on universities in 2009. Under such claim, the new rules deteriorated the career prospects of Italian professors so that, faced with lower expected earnings, some decided to quit for foreign institutions or for the private sector. We use three econometric models to identify whether constraints on personnel expenses causally determined an increase in voluntary leaves from universities. We exploit the selectivity of the constraint and its time variation to build a difference in difference estimator and a synthetic control evaluation. Difference in difference does indicate an increase in voluntary leaves in universities more severely affected by the new rules. Nevertheless, both methods show that this group still present voluntary leaves below that of less constrained institutions. While both descriptive and parametric analysis (structural break) point to a clear increase in voluntary leaves along with the introduction of the new regulation, the evidence of a selective impact on university with tighter constraints on personnel expenses remains weak. We highlight the possible shortcoming of our methodologies along with some alternative explanations for the results. With this respect, universities targeted by staff limits may be of lower quality and their professors may face lower outside opportunities. The deterioration of career prospects still result in current expected earning higher than any external alternative so that professors from these institutions keep their positions more on average than colleagues elsewhere. Secondly, there may be some anticipation effects due to the introduction of other rules that possibly affected the two groups differently. Finally, voluntary leaves may hide administrative retirements in which

individuals were grated short term contracts until their actual retirement upon resigning.

Chapter 4

Voting Behavior, Coalitions and Government Strength through a Complex Network Analysis

4.1 Introduction

A great deal of recent research has been devoted to explaining political polarization in parliaments (Fiorina et al., 2008; Layman et al., 2006). This literature has been dominated by models where party polarization is either explained by the polarization of the electorate or through the party and ideology of deputies. A new stem of literature has recently adopted tools of Complex Network Science (Caldarelli, 2007; Newman, 2003) to investigate this issue, with a network representation being given to committees and subcommittees who share the same members in the US Congress (Porter et al., 2005), to members of the Congress who co-sponsor bills (Fowler, 2006; Tam Cho and Fowler, 2010) or those who place the same roll-call votes (Waugh et al., 2009a). We follow the latter approach so that deputies are represented as nodes within a network where the number of shared roll-call votes determines the strength of

their links. Similarly to Zhang et al. (2008) and Waugh et al. (2009b) we make use of the network science concept of modularity in order to reconstruct the community structure of the parliament (Newman and Girvan, 2004). We introduce a novel method to characterize the position of each deputy in the community of reference, based on its contribution to the modularity score, proposing a more intuitive interpretation compared to that based on the spectral decomposition developed in Waugh et al. (2009b) and in Porter et al. (2005). The method presented here can be easily generalized on a wider European scale, and replicated across a longer time span or in industry-specific policies. In particular, the analysis can be extended to deal with multiple interdependent networks (Buldyrev et al., 2010) thanks to the interplay between senate and house of representatives or between national and european parliaments and take advantage of recent development in different fields of complex science ranging from critical infrastructures (Panzieri and Setola, 2008; Peerenboom et al., 2001) to epidemics transmission (Zhao et al., 2014,?). Indeed, nowadays political life of european countries is increasingly connected to, and interconnected through, European Parliament decisions. Moreover, European parliamentary acts and documents are semantically classified and organized in a EUROVOC thesaurus [1], that will make it possible to analyze political positions across different and controlled thematic areas. The rest of the paper is organized as follows. In the "Methods" section we present the methodology used to investigate parliamentary polarization, party cohesion, community structure and their time evolution, in the "Results" section the main findings related to the specific case of the Italian Parliament are presented, while in the "Discussion" section we draw our conclusions and sketch the lines of future research.

4.2 Methods

As the first step in our methodology we construct a graph where each node represents one of the n deputies and edges are drawn every time

[1] http://eurovoc.europa.eu/drupal/

two deputies display the same voting behavior[2]. We then normalize edges by the total number of votes in the reference period in order to obtain a weighted graph where weights are $0 \leq w_{ij} \leq 1$. Full weight is given to two deputies i, j if they participated in all sessions and voted exactly the same way in all of sessions. When a deputy quits the parliament, because of incompatibility, resignation etc., and his or her seat is taken by a new person, we consider the two deputies as being just one node[3].
Initially, we look at the topological structure of parties in order to study their cohesion over time. Completely ignoring any *a priori* knowledge of party affiliation, we look at the communities arising directly from voting behavior to see whether they match or not.

4.2.1 Analysis of party cohesion

Consider each party as a subgraph C of the graph G, with n_C being the number of deputies in the party. An intuitive way of measuring party cohesion (i.e. the tendency of the party to vote as a single entity) is to evaluate the intra-cluster density $d_{int}(C)$ defined as the ratio between the total internal strength of the sub-graph C and the number of all possible edges inside that cluster Fortunato (2010)

$$d_{int}(C) = \frac{\sum_{ij \in C} w_{ij}}{n_c \left(n_c - 1\right)/2} \, .$$

Similarly, we can define the inter-cluster density $d_{ext}(C)$ as the ratio between the observed strength of edges running from the nodes of C to the rest of the network and the maximum number of edges connecting internal with external nodes:

$$d_{ext}(C) = \frac{\sum_{i \in C, j \notin C} w_{ij}}{n_c \left(n - n_c\right)} \, .$$

[2]i.e. both vote in favor, against or abstain from vote. No edges are drawn for absent deputies.

[3]We check whether this transition leads to some votes in which none of the two deputies had their chairs without finding any discontinuity.

A party stands out as a specific political group if $d_{int}(C)$ is appreciably larger than the average link density $d(G) = \sum_{i,j} w_{ij} / \frac{n(n-1)}{2}$ of the entire network G and similarly we expect $d_{ext}(C)$ to be appreciably smaller. Searching for the best tradeoff between a large intra-cluster density and a small inter-cluster one is indeed an implicit or explicit goal for most algorithms used in community detection (Fortunato, 2010; Newman and Girvan, 2004).

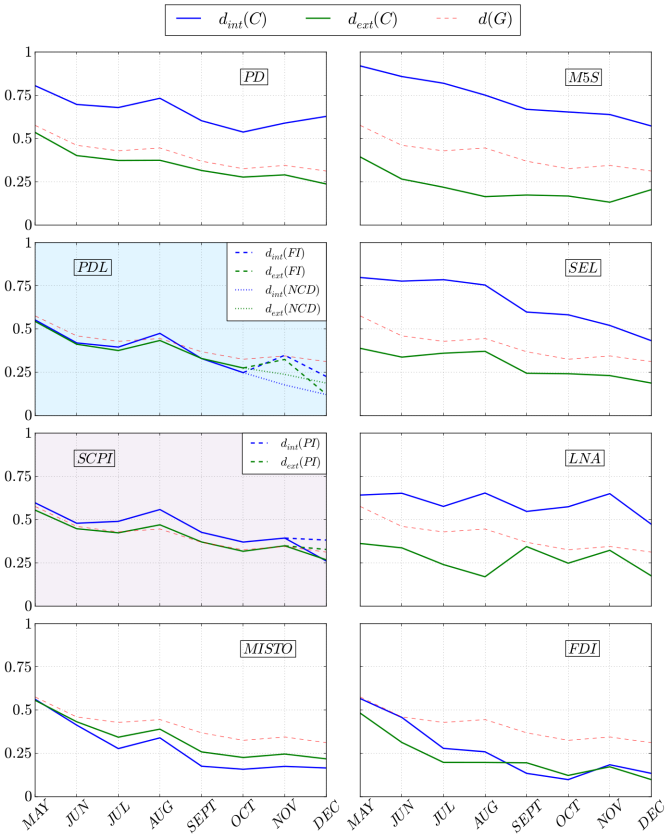
4.2.2 Community and core detection

Modularity optimization is a well-established method for detecting communities (Newman and Girvan, 2004). The idea behind modularity is that a random graph should not have a cluster structure so that communities are revealed maximizing the difference between the density of edges in a sub-graph and that expected if edges were connected at random. Hence the modularity function of a weighted graph (Newman, 2004), where in our case nodes are deputies and edges represent the percentage of votes that two of them have in common, is given by:

$$Q = \frac{1}{2W} \sum_{i,j} \left(A_{ij} - \frac{s_i s_j}{2W} \right) \delta \left(C_i, C_j \right)$$

where A_{ij} gives the fraction of similar votes deputies i and j share in common ($A_{ij} = w_{ij}$), W is the total weight in the network, $\delta \left(C_i, C_j \right)$ is a delta function that yields one if deputies i and j are in the same community ($C_i = C_j$) and 0 otherwise, and s_i, s_j represent the strength of node i and j respectively.
In the general case of modern democracies the typical result of the modularity optimization should be the splitting of the graph into two communities that reproduces the government coalition and the opposition. Moreover each node in its community usually doesn't have the same importance for its stability. Indeed, the removal of a node in the community core should affect the partition much more than the deletion of a boundary node. In other terms, some deputies display such a high degree of

Figure 13: Cluster Densities



Members of a party show cohesion if the links connecting them are stronger than the ties with other deputies. We capture the former by the intra-cluster density $d_{int}(C)$ and the latter by the inter-cluster density $d_{ext}(C)$. The party shows high cohesion when the two lie considerably higher and lower respectively copared to the average link density of the whole parliament $d(C)$.

internal connections so that they can be identified as the bulk of the coalition. As we proceed toward the boundary, deputies display increasing connections to the opposite community.
In order to investigate this structure, we exploit the properties of the modularity function following a new procedure introduced in De Leo et al. (2013). By definition, if the modularity associated with a network has been optimized, every perturbation of the partition leads to a negative variation of the modularity $dQ < 0$.
We compute the effect on the modularity associated with the shift of a deputy from one community to another and we plot the corresponding $dQ's$ distribution in order to check the coreness of each deputy and his party. In case of three or more communities dQ was originally developed in De Leo et al. (2013) to report the minimum variation in modularity, i.e. modularity was compared against a setup in which each node was moved, one per time, to its *closest* community. Here we rather consider movements to the *farthest* community in order to avoid abrupt shifts in the distribution of dQ due to the rise of small temporary (third) communities. Finally the histogram of the $dQ's$ will highlight the different groups that make up the coalition and will show different sub distributions along the support interval of dQ.

4.2.3 Measures of Polarization, Cohesion and Stability

Dealing with roll-call vote's networks as a whole, standard approaches (Moody and Mucha, 2013; Waugh et al., 2009a,b) have adopted the modularity score as a measure for party polarization. However, our methodology gives us the possibility to consider the overall voting behavior on a much finer scale, considering the contribution of every single deputy. In line with this, we have decided to measure the polarization as an average decrease in the modularity score consequent to the substitution of two opposite deputies; the larger the decrease in the modularity score, the larger the current contraposition between the two coalitions becomes. So we define the polarization as the sum of the median of the monthly dQ

distributions of the two communities[4].

When we focus on features of only one community, we still need to account for the community structure of the whole graph. Think for instance of two time frames in which the members of the ruling coalition vote exactly the same while the opposition voted $1/2$ and $1/4$ of the time with the government. Then the government dQ distribution would present more extreme values in the latter case, determining a shift towards more negative values of the mass of the entire distribution, despite the cohesion of the government *per se* not changing at all. Therefore any measure of cohesion should be robust to changes in the location of the distribution. A suitable one is represented by the interquartile difference of the $dQ's$ distribution that we will employ as our standard definition for the party/coalition cohesiveness.

In addition to polarization and cohesion, the stability of the government is directly affected by the number of its loyal deputies; in order for laws to be passed, half plus one of the total number of deputies are needed in the Chamber of Deputies. So as a rough rule of thumb, we can consider a government that keeps up to half plus one deputies on his side to still be *safe*. This measure accounts for the stability of the government comunity in the shape of a safety zone that divides the last critical deputy able to break down the majority from the $dQ = 0$ postion before the oppositon community region.

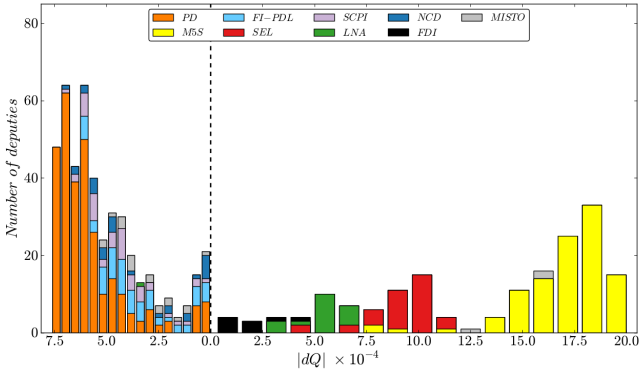
4.3 Results

As a concrete case we analyze the network of deputies in the newly elected Italian Chamber of Deputies (2013). We collect information on the 630 deputies and their voting behavior from the government open data SPARQL endpoint[5]. The reader may refer to table 12 for an outline of the main Italian parties mentioned in this paper. The available data cover parliamentary votes from April 2013, when the new parlia-

[4]The median has been preferred to the mean as a measure of location, because the distributions of both communities are strongly (negatively) skewed.

[5]http://dati.camera.it/sparql

Figure 14: Community Structure of the Italian Parliament



The vertical dashed line separates the two main coalitions/communities (Government/Opposition). Each coalition comprises different parties corresponding to different colors. The quantity 'dQ' is associated to the coreness of each deputy/party. The distributions are obtained computing the coreness of each deputy and then aggregating them in the form of a stacked istogram. The more the distance of the bars from the vertical dashed line, the more the deputies/parties are at the core of their coalition. Notice how the main parties tend to segregate in clusterd distributions with different positions in the 'dQ' axes.

ment was appointed, to the end of December 2013. Despite being quite a short period of time, the dataset covers 2820 parliamentary votes, which implies more than 1,5 million individual votes in our time span. Importantly, the Italian government has made semantic data following W3C standards available, which translates into fast and precise data manipulation through computer based queries. We refer to this source of data for the profiles of deputies and the classification of votes, while data on voting behavior of single deputies was taken directly from institutional web sites[6].

Figure 13 represents the evolution of density measures over time for each party in the Italian Chamber of Deputies. While the structures of the M5S, PD, SEL and LNA parties are recognizable within the graph, the other groups present inter- and intra-cluster densities that are very close to each other, or at times even overlapping. This means that at a certain point their votes proportionally coincide to a greater degree with other groups than with their own members. The plots marked with a colored background report the splitting of two political groups, when the PDL breaks up into the NCD and the FI in November 2013 and the PI exits from the SCPI in December 2013. The inter-cluster density, represented in green, is clearly higher for groups who support the government (PD, PDL and SCPI). Theoretically these groups should vote in compliance with the majority's prescriptions, thereby showing a similar voting behavior. Once we take into account the average monthly levels of edge density $d(G)$ the topological structure of parties becomes very similar to the rest of the graph. As such, parties may not be the most appropriate representation of voting structure, thus leading us to consider the behavioral identification of political groups through the modularity function. Once applied to the graph of deputies, the modularity optimization usually splits the graph into two communities that almost exactly match the government coalition and the opposition as shown in figure 14 where the vertical dashed line separates the two coalitions.

Afterword we compute the effect on the modularity associated with the removal of a deputy from his community computing the corresponding

[6]http://documenti.camera.it/votazioni/votazionitutte/FormVotazioni.Asp?Legislatura=XVII

Table 12: Outline of the main Italian parties

Party	Coalition	%*	Notes
Partito Democratico (PD)	Gov	$46,5\%$	Main center-left party, historically lead by Prodi
Il Popolo della Libertà (PDL)	Gov/Opp	$15,2\%$	Main conservative party, lead by Berlusconi
Forza Italia (FI)	Opp	$10,6\%$	From PDL split, founded and lead by Berlusconi
Nuovo Centro Destra (NCD)	Gov	$4,6\%$	From PDL split, lead by Alfano
Scelta Civica (SC)	Gov	$7,3\%$	Lead by Monti, PM for one year after 2011 crisis
Movimento 5 Stelle (M5S)	Opp	$16,3\%$	Lead by comedian Grillo, form of *direct* democracy
Sinistra-Ecologia-Libertà (SEL)	Opp	$5,6\%$	Left party, former ally of PD
Lega Nord (LN)	Opp	$3,2\%$	Autonomist party of Northern Italy, former ally of PDL

$*$ Shares updated to may 2014, smaller parties omitted.

$dQ's$ and the result is also shown in figure 14.
The histogram shows the $dQ's$ distribution of the government's coalition on the left side of the dashed line and that of the opposition on the right, with alle the dQ associated to different parties in different colors.
Indeed, the core of the coalition appears to be made up by a relatively higher share of deputies from the center-left party PD while relatively more deputies from the center-right party PDL appear to be at the periphery as we move to the right. This provides an interesting insight on the rather different roles played by the two main Italian parties joined by a coalition pact, namely the PD and the PDL, with the latter ultimately quitting the government in mid November 2013. As for the opposition, note that the support of dQ is far more dispersed with each group taking on a limited range of values in the distribution. This is not surprising in that the opposition is not a coalition *per se* but rather a set of groups that might vote with the ruling coalition depending on the subject at hand. In particular, deputies from the M5S make up the core of the opposition with a higher magnitude of dQ, which also holds true when compared to the core of the government coalition. This may be due to a relatively inflexible opposition to the government or in equal measure to the fact that it is the largest group in the opposition community. On the other hand, the SEL and the LNA are progressively closer to the border of the community, which may be reasonable if we consider that these groups used to be allies of two parties in the government coalition, namely the PD and the PDL respectively.

4.3.1 Time evolution of the community structure

The same analysis has been carried out over time, dividing votes per month, building up the corresponding graphs and performing the community and core analysis on each monthly network.

In figure 15 the two main communities present increasingly extreme values of dQ over time, which in turns provides evidence of increasing polarization in the parliament, as it is measured as the sum of the median of the monthly distributions (see Methods section).
This may reflect the change in the political position of M5S which moved from declared openness to the government on a single bill basis Il Fatto Quotidiano (2013) to a very sharp contrast as events unfolded. Two controversial bills that occupied a large fraction of the assembly's sessions over summer 2013 [7] and the worsening of the political climate that led to the repeated demand for resignation of government ministers in the following autumn ANSA (a), might have driven the political debate towards increasingly polarized configurations as it is evident in figure 15 also on a monthly level of aggregation.
In this respect, December noticeably stands out, with a reduction in the extreme values of dQ for the opposition. This is actually driven by the fragmentation of the PDL, which witnessed its deputies loyal to the leader Silvio Berlusconi, withdraw their support of the government Bloomberg (2013) and start to vote with the opposition to the point of being identified as part of it at least in its border. The figure 15 illustrates also the cohesion, or rather its flip side: the heterogeneity of deputies within a single community, along with the government stability represented through a green *safety zone*. This area spans values of dQ smaller (in absolute value) than the monthly critical value $dQ_{critical}$. The latter corresponds to the level of coreness of the deputy which would pose the government in numeric inferiority, were he leaving the coalition. In the specific case of the Italian Parliament, the Chamber of Deputies has 630 representatives and the critical value will correspond to the dQ relative to the $316th$

[7] the shut down of an old iron factory and a bill ANSA (b) containing economic reforms to tackle the crisis culminated with M5S's deputies blocking the assembly and then leaving it once the measures were eventually passed

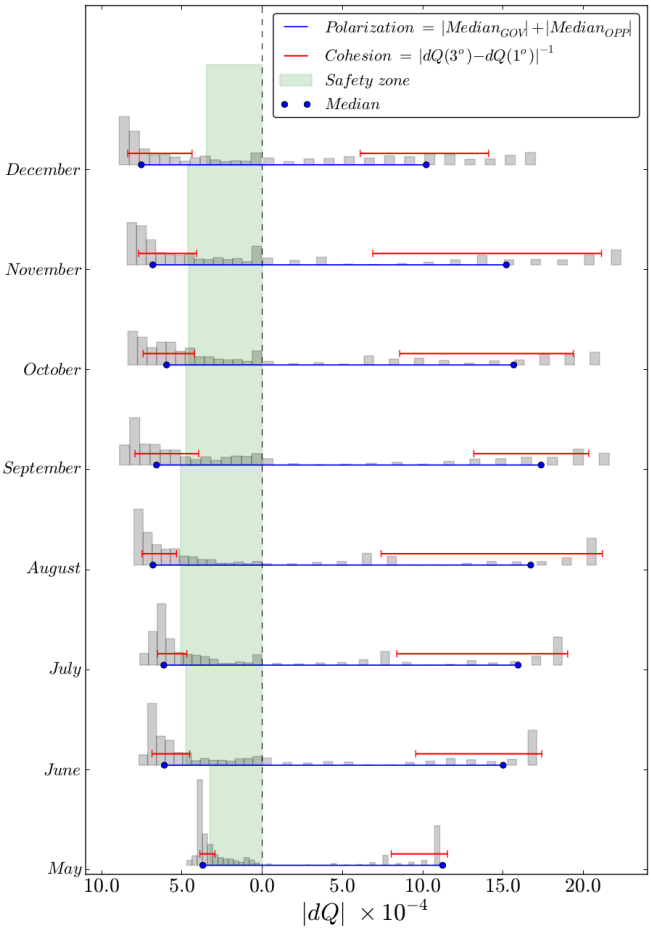
deputy.

With fixed levels of polarization and cohesion, a greater absolute value of $dQ_{critical}$ would widen the *safety zone* in that a relatively more loyal deputy would have to leave the government coalition in order to make it facing the risk of having its laws rejected. Having investigated the peculiar structure of the government coalition, we focus on a political party that may be partly responsible for the variability of the coalitions topology over time. Indeed the PDL, after a long debate regarding whether to support the government or not, eventually split into two different parties. After the split in mid November, deputies from the FI moved into the opposition community. However, surprisingly, those who left moved from the core of the government to relatively core positions in the opposition, as reported in figure 16.
This dynamic may somehow explain the peculiar drop of the polarization observed in December in figure 15, as the FI group switched voting behavior to such a degree as to be recognized as part of the opposition, simultaneously reducing the contraposition between the two communities.

4.4 Discussion

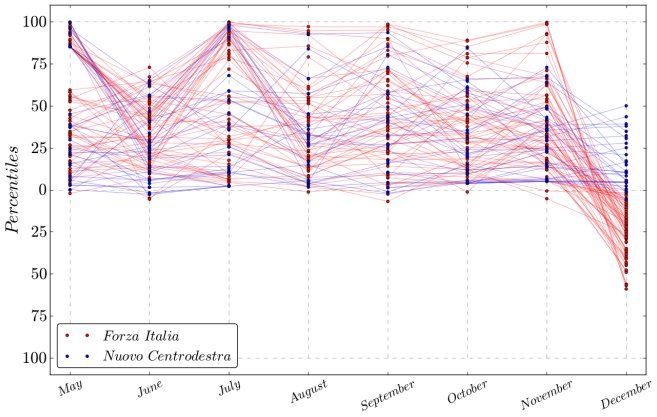
The study of the consensus dynamics in modern parliamentary democracies is of great importance for the validation by citizens of the performance of their representatives. These dynamics are often hidden by complicated voting procedures that prevent the easy identification of these civil representatives. We need new ways to look at the details of the political activities, which go beyond the standard statistical indicators, ways that are able to reveal the dynamics of the general organization of the government, its opposition and even their internal structures, in a format that is intelligible to non-expert users. In this study we introduced a novel procedure to map parliamentary voting trends onto a network structure in which the nodes are the deputies and the edge weights are the strength of their relations. These weights, month by month, quan-

Figure 15: Community structure over time



The evolution of community structure over time provides a way to track the cohesion of the government and the overall polarization in the parliament. The empyrical analog of the cohesion is represented here by the interquantile difference of the dQ distribution, where higher cohesion occurs for lower values of the interquantile. On the other hand higher parliament polarization is captured by the distance between the two medians. Finally the *safety zone* that divides the last critical deputy able to break down the majority from the vertical dashed line ($dQ = 0$) is represented in green.

Figure 16: Community structure during parties split up



The position of single deputies within communities provides insights on what happens when a party splits up. In this particular case the PDL party in mid November 2013 splits into two different parties 'Forza Italia' and 'Nuovo Centrodestra'. Interestigly, nodes at the core of the government coalition become core in the opposition one when the split up occurs. This is evidence of political voting being driven by coalitions' affilitions rather than the policy content of each roll call vote.

titatively measure the degree of closeness between couples of deputies as the number of votes they shared in a specific time frame. Once this network has been built up, using Community Detection techniques borrowed from Complex Network Science, it is possible to reconstruct the main coalitions, the government and the opposition from the bottom up; through a 'Core Detection' analysis it is also possible to uncover the internal structure of these aggregations. Using the leverage of later analyses we were able to quantitatively detect the position of each party, the strength and consistency in its coalition and the level of polarization between government and opposition.

Furthermore, the Open Data movements around the world are pushing public administrations to provide free and open access to massive amounts of data, which can be used by citizens and companies as a starting point for the detailed analysis of public policies. In this study, we relied on a recent service introduced by the Italian parliament that allows the automated extraction of certified information about the votes of the Chamber of Deputies. Through this service we have been able to perform a thorough analysis of the dynamics of the Italian parliamentary factions over nearly a year of legislation, using the aforementioned methodology.

These methods open up new possibilities of bringing citizens closer to their representatives, thereby establishing the foundations for a more transparent democracy.

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