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GENERAL INTRODUCTION

"History matters. It matters not just because we can learn from the past, but because the present and the future are connected to the past, by the continuity of a society's institutions." These are the words with which Douglas North's (1990: vii) seminal study on the role of institutions and institutional change in economic theory begins. Since then, claims of "history matters" or, more specifically, path dependence, has become a popular notion for explaining various forms of institutional inertia and resistance to change. The primary objective of the present thesis is not to reaffirm that "history matters", but rather it is to explore the more specific question of "how" history might matter within the comparative political economy of welfare reform. The puzzle motivating this inquiry is to identify those factors that are responsible for linking decision making through time and to test their impact on current welfare policies. Although the title "Regimes, Institutions and Temporality in the Political Economy of Welfare Reform" may give the impression of a heterogeneous collection of topics, the central idea that brings together the three papers is the attempt to integrate temporality in the politico-economic analysis of mature welfare states.

The first part, entitled "Traveling without Moving? Pension regime change in mature welfare states", explores how inherited social and economic arrangements in old age security provision; referred to as pension regimes, restrict reform options available to policymakers today. Results of the explorative multiple correspondence and hierarchical cluster analyses on a sample of up to 18 OECD countries (1988-

2003) indicate that these legacies are responsible for the convergence of pension reform trajectories within regimes, suggesting that pension reforms in mature welfare states are following a logic of “bounded change”, where change takes place but pension regime differences persist. In a nutshell, the classification of old age security systems obtained from the correspondence analysis will be utilized in the second part of the analysis.

The second paper “Are Mature Welfare States on the Path to Gerontocracy? Evidence from 18 OECD countries, 1980-2003”, focuses on concrete pension financing rules rather than regime arrangements. Within the median voter framework it investigates how institutional differences mediate the effect of population ageing on the size and generosity of public pensions. Although mature pension systems face relatively similar challenges, tentative evidence from panel regression analysis indicates that majority voting in aging societies has two opposite effects: it increases overall pension spending as a percentage of GDP but decreases the generosity of pension benefits, depending on whether public pension entitlements are closely linked to contributions. Moreover, while issues of horizontal redistribution appear to matter less in voting on old age security, estimation results indicate that projected growth in population ageing creates momentum for reductions in pension spending and benefit generosity.

The third paper, which has been titled “What Makes Stabilization Reforms Happen? Temporality in the political economy of welfare spending”, takes a different view on temporal aspects in welfare reform policies. Instead of looking at long-term historical legacies, it investigates how temporal contexts influence the effectiveness of political determinants in welfare spending. To do so, the empirical analysis employs interactive dynamic panel regression and event history analysis on a dataset covering 21 OECD countries (1980-2003). While prior research indicates that popular theories of partisanship, electioneering and institutional rigidity are too general to explain recent developments in the dynamics of social expenditure, findings from the empirical analysis suggest that the influence of these political determinants re-emerges if temporal contexts are taken into consideration. Thus, it can be shown that politics still matters in welfare spending, but it matters in a more subtle way than in previous decades.

The underlying thesis makes an attempt to contribute to the comparative political economy literature through a thoughtful integration of time-based factors in the analysis of welfare reform in mature welfare states. Among the various results that can be established within the three papers, the more general insights are the following: First, there is no doubt that history matters in welfare policies, however, "how" exactly history matters is far from being settled. The empirical endeavor to focus on temporal issues shows that historical legacies may restrict policy options, as shown in the case of pension regimes, or help to explain why similar stimuli can result in different welfare policy outcomes, as in the case of voting on old age security. Second, until recently political economy theory underemphasized the role of contextual variation for policy making. The temporal context, such as fiscal stress, however, appears to be among those factors that condition policy-makers' incentives and abilities to manipulate public policy. Third, within the framework of the second and third paper there is evidence that expectations and policy framing play an important role for the conduct of welfare reform. These findings challenge standard rational voting assumptions and make calls for a deeper investigation of expectation mechanism in social policy reform.

In total, a careful integration of "history matters" arguments in the politico-economic analysis can improve the understanding of persistent differences among mature welfare states and offers much promise for resolving empirical anomalies in comparative welfare state research. I hope that the subsequent studies give an impression of quantitative research methods capable of dealing with issues of temporality, provide new insights to the politics of welfare reform and offers suitable information for policymakers interested in the preconditions of politically feasible welfare reform policies.

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Traveling without Moving?

Pension regime change in mature welfare states

Abstract

This paper investigates if pension reform trajectories in mature welfare states are following a logic of "bounded change", whereby change takes place but pension regime differences persist. The explorative empirical analysis employs multiple correspondence, hierarchical cluster and cross-sectional regression analyses on a sample of up to 18 OECD countries (1988-2003). Findings support the notion of regime-specific vulnerabilities towards economic and demographic challenges and suggest that pension reform trajectories in insurance-based Bismarckian pension systems represent a case for "bounded change". Linking reform trajectories to change in old age poverty rates indicates that Scandinavian countries are relatively successful in balancing fiscal need and social equality, while tentative evidence for the liberal countries suggests that pension privatization has already exposed retirees to higher risks of old age poverty. Although the cluster analysis on change in pension reform parameters is not thoroughly convincing, in overall terms, the results of this study indicate convergence of reform trajectories within regimes rather than convergence of pension systems.

Keywords: welfare regimes, public pensions, old age poverty

JEL Classification Numbers: H55, J18

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1. INTRODUCTION

Over the past two decades, reforming old age security has been a dominant issue on the social policy agenda in mature welfare states. With increasing budgetary pressure due to demographic change and new labor market trends, many OECD countries made considerable efforts to redesign public pensions (OECD 2005, 2007). Although comparative welfare state research tends to emphasize the role of policy legacies and institutional rigidity (Esping-Andersen 1990, Pierson 1994), there is increasing evidence on high levels of reform policies (Allan & Scruggs 2004, Hinrichs 2000, Immergut, Anderson & Schulze 2007). The puzzling phenomenon at the heart of this contribution is the simultaneity of welfare regime stability and reform within regimes. The paper investigates if pension reforms are following the logic of "bounded change", arguing that institutional differences between pension regimes do not only create important differences in the degree of vulnerability towards demographic change and economic pressures (Scharpf 2000a,b), but they also create different types of pension reform trajectories. Empirically, "bounded change" simply means that countries within the same regime will adopt similar reform strategies. As a result, the composition of pension regimes remain unchanged while policy shifts may take place within regimes. Finally, linking pension reform strategies to redistributive outcomes allows to test if reforms can be made responsible for the growing risk of old age poverty in aging societies.

In order to trace "bounded change" as a feature of pension policies, the empirical analysis separates into three research questions: First, do pension systems in the OECD cluster into distinct and stable regimes? Second, can we identify common patterns of reform within these regimes? And third, if there are distinct reform patterns, what are their redistributive consequences? For at least three reasons these questions can be considered as relevant: First, the European Commission has repeatedly expressed its concern about the fiscal sustainability of public pensions and its macroeconomic consequences (European Commission 2003). In 2001 EU member states agreed to coordinate old age policies through the so-called Open Method of Coordination, which is designed to induce best-practice learning via benchmark competition. It assumes that an annual adjustment routine among member states will lead to an evolutionary selection of the most efficient pension systems. However, if there are basic incompatibilities between pension regimes, which hamper the possibility of policy transfers, convergence of pension systems is unlikely. Second, recent reforms in developing countries have shown a strong tendency for building up multi-pillar systems that combine funded and unfunded pension provision. Although international organizations such as the World Bank (1994) and the IMF (2005) provide detailed reform templates for such attempts it has become evident that the process of pension privatization in mature welfare states lags behind expectations. It seems that the shift toward a mixed-financed multi-pillar system heavily depends on the preconditions of the existing scheme. In particular, implicit liabilities involved in mature pay-as-you-go (PAYG) pension programs can make a shift toward funding extremely costly. Finally, old age security systems with proportional contribution rates and regressive benefits are known to redistribute income between and within generations. Pension reforms aiming at a reallocation of transfers in either one or both of these dimensions tend to provoke political controversy, often involving highly emotional arguments about generational justice (Kohli 2006). Since any pension reform requires majority electoral support, understanding how different reform options affect the quality of old age security provision is essential for designing politically feasible reforms. This provides an additional reason to take into account the redistributive consequences of pension policies.

The remainder of the study is organized as follows: The second section briefly summarizes contemporary theories of the welfare state to conceptualize "bounded change" and reviews prior empirical work. The third section presents the data, methods and results of the empirical analysis concerning pension regime stability, regime-specific reform trajectories and the redistributive effect of different reform strategies. The last section discusses the limitations and implications of the findings for further research.

2. THEORETICAL FRAMEWORK

Since the seminal publications by Esping-Andersen (1990) and Pierson (1994), much of the comparative welfare state literature has resolved around welfare regimes and the politics of welfare retrenchment. The regime approach has repeatedly been criticized for providing a too static picture of welfare systems (Kasza 2002, Crouch & Farrel 2004), which provoked the issue of how to account for long-term processes of welfare change. This section reviews the regime and retrenchment approach with respect to old age security and suggests “bounded change” as a framework to address issues of change and continuity in mature welfare states.

2.1 Welfare regimes and the politics of retrenchment

Since it is almost impossible to find a paper comparing welfare states that does not refer to Esping-Andersen’s (1990) book “The Three Worlds of Welfare Capitalism”, his typology of welfare regimes serves as a reference point for the analysis of pension regime stability (Table 1). The concept of different welfare regimes however already dates back to Titmuss (1974), who assumes that although no welfare system is truly identical with one another, that there are systematic similarities between groups of welfare states. Taylor-Gooby (1996: 200) defines regimes “as a particular constellation of social, political and economic arrangements which tend to nurture a particular welfare system, which in turn supports a particular pattern of stratification, and thus feeds back into its own stability.” Following this line of reasoning, Esping-

Andersen's (1990) argues that Western welfare states are not a number of unique cases but cluster around three ideal regime types. The liberal regime is marked by means tested benefits and relatively high levels of private insurance. Public insurance programs in the social democratic regime combine universal coverage with generous benefits. And the conservative regime is ideally characterized by contribution financed public insurance schemes that are differentiated by occupational groups. Esping-Andersen (1990, 1999) assumes that each regime reflects a set of principles and values that encompass a distinctive rationale for social security provision. Thus, welfare regimes can be identified either by the sum of programs or by specific welfare programs. In the wake of his work, scholars extensively debated the existence of distinct regimes, the principles on which they are based, and the correct number of welfare regimes (Castles & Mitchell 1993, Huber & Stephens 2001, Hicks & Kenworthy 2003).[1] Gelissen (2002: 140) points out that Esping-Andersen uses the regime types not only to explain cross-national variations in welfare provision but also to explain trajectories in the development of national welfare programs. In this respect, the existing institutional welfare arrangement provides the incentives that encourage individuals and groups to act in ways that reinforce a particular policy. Although Esping-Andersen (1990, 1999) does not specify the mechanisms that are responsible for institutional reproduction, his

[1] The number of regimes that have been identified fluctuates between two (Bonoli 2003) and eight (Gough 2001).

approach suggests that three stable and distinct clusters of old age security provision can be found within the OECD world.

Table 1. Pension regimes by decommodification scores

Liberal	Conservative-corporatist	Social democratic
Australia	Finland	Austria
Canada	France	Belgium
Ireland	Germany	Denmark
New Zealand	Italy	Netherlands
United Kingdom	Japan	Norway
United States	Switzerland	Sweden

Source: Esping-Andersen (1990: 52) Table 5.2.
Note: Decommodification is defined in terms of social rights; it captures the degree to which a person can maintain a livelihood without reliance on the market.

The "new politics" or "retrenchment" literature focuses on the politics of welfare reform. The debate was launched by Pierson's (1994) book "Dismantling the Welfare State", in which he analyzed why Reagan in the United States and Thatcher in the United Kingdom were relatively unsuccessful in rolling back the welfare state compared to the aims of their political agenda. Pierson concludes that the politics of welfare retrenchment are fundamentally different from the politics of welfare state expansion. Welfare expansion was a game of claiming electoral credit, while retrenchment is a game of avoiding electoral punishment for cutting welfare program entitlements. Political conflict over reform therefore plays out less along lines of class, skill or ideology but more along those who benefit and those who pay for existing programs.[2] The financing logic of unfunded pension systems may present an additional obstacle for retrenchment. Even though a transition toward a mixed financed pension system might get majority electoral support, such attempts are confronted with the double payment problem: current contributors would have to finance both the benefits for pensioner entitled under the current public pension scheme as well as the contributions necessary to build up their own retirement fund (Scharpf 2000a). Myles & Pierson (2001) find that, depending on whether pension reforms have taken place before the mid 70s, marking the end of the "golden era" of welfare expansion, countries fall into one of two groups - latecomers or mature pension systems. Latecomers are countries with a relatively young and small flat-benefit pension system. The second group, however, consists of countries with mature and comprehensive public insurance schemes. The latecomer's pension reform is characterized by more or less radical pension privatization, while countries with mature insurance schemes aims to adjust contribution rates, benefit generosity or strengthen the link between contributions and benefit entitlements. Although the "new politics" approach does not explain why these reforms took place even at the cost of electoral punishment, the overall empirical findings suggests two things:

[2] The retrenchment debate is still far from being settled. Empirical studies investigating reductions in the level of social expenditure per GDP suggest that reforms have been incremental rather than radical retrenched (Stepehens, Huber & Ray 1999, Huber & Stephens 2001, Green-Pedersen 2002). Researchers using indicators on the generosity of welfare entitlements, however, claim that there have been considerable changes in replacement rates for sickness and unemployment programs in the last decade (Allan & Scruggs 2004). Qualitative orientated studies support the latter (Bonoli 2000, Bonoli 2003, Immergut et al. 2007).

First, systems of old age security have been subject to more substantial changes during the last decades than the welfare regime approach would predict and second, it seems that reform options available to policymakers are very much restricted by the existing pension system.

2.2 Welfare Reform as "Bounded Change"

The regime and the retrenchment approach provide two alternative perspectives on welfare policies; while the former emphasizes stability of regimes, the latter focuses on the political conditions and ability to reform welfare programs. The task is to find a reasonable framework that accounts for both institutional change and stability. In this regard, Pierson (2001) points out that scholars in the field of comparative welfare research disagree less about how much change has happened than they do about adequate concepts to capture different kinds of change. "Bounded change" stems from the path dependence literature, which has become increasingly popular in social sciences. The number of journal articles referring to path dependence which are listed in the Social Sciences Citation Index (SSCI) increased exponentially from 3 in 1992 to 62 in 2005.[3] A minimal definition of path dependence requires that latter events are not completely independent from those that occurred in the past, or, put simply "history matters". While there is common consent that history is likely to

[3] Obtained from the Social Sciences Citation Index quick search on "path dependence" (15. September 2007). The number of citations in each year increased during the same period from 25 to more than 650.

play a role in the development of welfare states, a vibrant theoretical debate emerged about "how exactly" history matters (Crouch & Farrell 2004, Mahoney 2000, Pierson 2000, Thelen 2000, Schwartz 2001). North (1990: 98) understands path dependence simply as "a way to narrow conceptually the choice set and link decision making through time." Important efforts to apply path dependence explanations to political science are made by Pierson (2000, 2004) who conceptualizes path dependence as a three-phase process driven by increasing returns. In his view increasing returns are the source of path dependence as they induce self-reinforcing processes. Following North (1990), he argues that institutions per se generate increasing returns through learning effects, coordination effects and adaptive expectations.[4] Within the three-phase analogy, a path dependent process starts with a critical juncture in which contingent events trigger a move toward one of at least two alternatives (Mahoney 2000), followed by a period of institutional reproduction in which increasing returns keep pushing things along the same path. A new path, however, can only be initiated by a new critical juncture. In order to avoid determinism, Pierson (2000: 265) argues that his conceptualization of path dependence does not need to imply frozen social landscapes. In contrast, "Change continues, but it is *bounded change* – until something erodes or swamps the mechanisms of reproduction that generate continuity". Thus, he predicts that relatively long periods of stability are followed by short periods of wide-ranging changes.

[4] See Pierson (2000) for a detailed classification of self-reinforcing mechanisms.

More recently, scholars in comparative welfare research tried to build on path dependence to explain gradual processes of institutional change (Deeg 2001, Ebbinghaus 2005, Hering 2003).[5] Thelen (1999, 2003) suggests a two-stage approach comprising on the one hand the analysis of institutional stability by identifying mechanisms of increasing returns, and, on the other hand, the analysis of institutional change by identifying mechanisms of institutional layering or conversion. She claims that both stability as well as change can be at work at the same time. Unfortunately, her approach remains vague in specifying the relationship between these conflicting mechanisms and where to locate them analytically. Hence, empirical investigations drawing on Thelen (1999, 2003) had difficulties in distinguishing between the "old path" of institutional stability and the degree of change or innovation that allows speaking of a "new path". Deeg's (2001) investigation of Germany's financial sector, for example, provides as much evidence for institutional stability as for institutional change. Thus, while Pierson's (2000) conceptualization of change as a radical switch in the reproduction mechanisms is quite restrictive to the possibility of slow moving incremental change, further relaxation of path dependence towards processes of change, bears the risk of making it a meaningless concept.

Referring to North (1990: vii) the question is how the continuity of a societies institutional arrangement that connects decision making through time influences

[5] Suggested concepts reach from a "trodden trail" or "branching pathways" (Ebbinghaus 2005) over "path departure" (Hering 2003) to "path switch" (Deeg 2001).

pension reform policies today. "Bounded change" brings together change and stability by considering welfare regimes to inhibit different vulnerabilities toward demographic change and economic pressure (Scharpf 2000a, 2000b). Although old age security systems in mature welfare states face relatively similar reform challenges, their actual policy response depends on their affiliation to a certain welfare regime. In this respect, "bounded change" argues for a middle path between historical scholars who stress the uniqueness of welfare reform and welfare economists searching for the one best way solution (Kato 1996). "Bounded change" appears to be the observable consequence of pension reform within the limitations of a distinct welfare regime. This does not imply that national systems of old age security remain unchanged once they are established. In contrast, it suggests that changes take place while regime differences persist. It implies no "race to the bottom" nor convergence to the mean; rather, "bounded change" implies convergence of reform trajectories within welfare regimes. This argument is not completely new; it is implicit in the welfare regime concept. Thus, it is surprising why the dynamics of welfare regime change in mature welfare states have yet received little attention by empirical researchers.

2.3. Review and Contribution to the Literature

The existence of welfare regimes has already been tested by means of cluster analysis (Obinger & Wagschal 2001, Gough 2001, Powell & Barrientos 2004) from a static perspective. The overall findings suggest that Esping-Anderson's three world typology "neither passes the empirical tests with flying colours, nor dismally fails them" (Arts & Gelissen 2002: 153). However, the majority of studies analyze overall welfare spending. Due to methodological problems of comparing entire welfare states, Kasza (2002) suggests focusing on particular welfare programs. Public pension schemes are by far the largest welfare program in budgetary terms. To the author's knowledge, only two papers investigate the three world typology with respect to public pensions. Ragin's (1994) analysis identifies four pension clusters, a liberal cluster consisting of Australia, Canada, Switzerland and the United States; a corporativistic cluster including Australia, Belgium, Finland, France and Italy and a social democratic cluster which includes Denmark, Norway and Sweden. Moreover, he identifies a spare cluster that entails Germany, Ireland, Japan, Netherlands, New Zealand and the United Kingdom. Shalev (1996) follows a different approach; he conducts a factor analysis on 14 social policy indicators. Although the majority of indicators concern old age security, his analysis also includes total social expenditure, poor relief, active labor market expenditure and private health expenditure in his analysis. There is no obvious reason why these variables should help to identify pension regimes. However, in contrast to Ragin (1994), Shalev's (1996) findings support Esping-Andersen's typology.

Yet there are two important shortcomings: Both studies are cross-sectional and rely on Esping-Andersen's data presented in "The Three Worlds of Welfare

Capitalism". The majority of these indicators are ad hoc measures for the 1980's that have not been updated or reproduced.[6] This study employs recent cross-sectional time-series data, which is independent from Esping-Andersen's measures. The sample compares the development of pension systems in up to 18 OECD countries covering a maximum time span from 1988 to 2003.[7] Hinrichs & Kangas (2003) point out that the observation period can have substantial consequences for the analysis of pension policies, since it is a common practice to suspend the immediate effect of pension retrenchment. Hence, it can take decades until the consequences of a reform fully materialize. Unfortunately, the availability of data prohibits a temporal or cross-sectional extension of the dataset. Nevertheless, with respect to the onset of population aging and increasing reform activity among OECD countries particularly during the last decade, the observation period should be appropriate for an exploration of "bounded change" in old age security provision.

This study differs in an even more important aspect from prior research; it does not end with the identification of pension regimes. Regimes or typologies as such are attempts to systematize descriptions of a phenomenon. They do not provide a causal explanation for the configuration of old age security systems. Nevertheless, typologies are important steps in the process of theory building (Arts & Gelissen

[6] Scruggs & Allan (2006) present a detailed methodological critique on Esping-Andersen's data.

[7] The OECD 18 sample includes Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Sweden, Switzerland, the United Kingdom and the United States.

2002: 140). This study uses the identification of pension regimes as a starting point for the analysis of regime-specific reform trajectories and their redistributive consequences. A deeper investigation of the redistributive effect of pension reforms should provide additional information on the limits and trade-offs involved in such policies.

3. EMPIRICAL ANALYSIS

In order to answer the three research questions brought up in this paper, the empirical analysis draws on graphical-explorative and analytical-statistical methods. First, multiple correspondence analysis is used to identify different qualitative configurations of old age security provision. Second, hierarchical cluster analysis is used to classify countries possessing similar pension systems. Finally, cross-sectional regression analysis is used to investigate the redistributive consequences of different reform strategies.[8]

3.1 Dimensions of Old Age Security

Building on Esping-Andersen (1990, 1996, 1999), the description of old age security systems starts with three indicators: the size of public pensions, the generosity of public pension benefits and the importance of private pension provision. First, public pension expenditure per GDP accounts for the share of public resources devoted to

[8] The MCA, CA and regression analysis is performed with Stata 10.

old age security and is defined as cash benefits for retired persons. Following Castles (2002) it is assumed that, although aggregate spending data is insufficient to describe the full nature of welfare systems, it still provides a reliable source for the analysis of reform trajectories. Data on public pension expenditure is taken from the OECD (2007a) Social Expenditure Database.

Aggregate pension spending might underestimate the impact of pension reform (Hinrichs & Kangas 2003). Not only is it possible that cuts in one program balance the impact of reforms in another program, but, it is also likely that overall pension spending increases with population aging, while at the same time pension benefits per elderly decrease. Societies with a larger share of elderly have to devote more resources to the elderly, but this does not imply more generous benefits. Therefore, the second dimension of old age security aims to capture the generosity of public pension benefits. Income replacement rates are among the most useful indicators to assess the generosity of a welfare program since they provide a measure of the income that is made up by a welfare program (Korpi & Palme 2003, Allan & Scruggs 2005). With respect to public pensions, the replacement rate is defined as the ratio of net pension paid to a person who earned the average productive worker wage in each year of their working career. It provides a proxy for the level of income substitution through public pensions. A second group of generosity measures focuses on the coverage of welfare programs. For example, Flora & Alber's (1982) study on the rise of Western welfare states did not use aggregate spending or replacement rates, but a measure on the proportion of people covered by public insurance schemes. Coverage is an indicator for the universalization of benefits. The pension generosity scores calculated by Scruggs (2005a) are based on both components; pension replacement rates and pension coverage.[9]

The third indicator concerns the division between private and public responsibility for old age security. Although a relatively large body of literature emerged on the determinants of public pension expenditure (Lindert 1996, Breyer & Craig 1997, Mulligan & Sala-i-Martin 1999, Disney 2007), little attention has been paid to the development of private pensions. In order to understand the logic of old age security systems, it is necessary to take into account the interplay between funded and unfunded pension provision (Esping-Andersen 1990: 103). Unfortunately, data on the size and coverage of private pensions is extremely difficult to obtain. Queisser, Whitehouse, & Whiteford (2007) have recently provided detailed cross-national measures on private pensions for various OECD countries. However, testing the "bounded change" argument requires time-series data. Since no cross-sectional time-series data on private pension entitlements is available at the moment, this study uses the size of pension fund assets relative to the GDP as an indicator for the "privateness" of old age security provision. Pension

[9] A limitation that should be considered in the interpretation of pension generosity scores is that they do not take into account intra-generational redistribution. Thus, it is possible that in a more redistributive scheme low average worker replacement rates mask the fact that replacement rates at lower income levels are higher.

funds are defined as a pool of assets for the exclusive purpose of financing pension plan benefits. Davis (1995) distinguishes three factors that determine the size of pension funds. First, since privately funded pension schemes are the main alternative to pay-as-you-go programs, there should be a negative relationship between the size of public and private pension expenditure. Second, the maturity of the fund and the rates of return influence the size of pension funds (Bailliu & Reissen 1997). And third, by making private pension saving mandatory or by setting tax incentives to build up private pension savings, political regulation also plays a role in determining the size of pension fund assets relative to GDP. Thus, although the maturity of pension funds restricts the interpretation of this indicator it still appears to be reasonable to assume that the size of pension funds contains information about the division between private and public pension provision. Results of the correlation analysis presented in Table 2 provide further justification for the use of pension fund assets as an indicator for private pension provision.

Table 2. Pairwise correlation between private pension spending and pension fund assets

	Private pension spending per GDP		
	Coef.	P-value	Obs.
Pension fund assets per GDP	0.70*	0.00	17
	Private pension plan coverage		
	Coef.	P-value	Obs.
Pension fund assets per GDP	0.35	0.17	17

Source: Queisser et al. (2007: 555) Private pension spending as percentage of GDP in 2003: Australia 3.0, Austria 0.6, Belgium 2.3, Canada 4.2, Denmark 2.2, Finland 2.9, France 0.2, Germany 0.7, Ireland 0, Italy 1.3, Japan 3.1, Netherlands 3.2, Norway 0.7, Sweden 2, Switzerland 4.5, United Kingdom 4.7, United States 3.8, New Zealand = missing; Queisser et al. (2007: 549) Private pension plans include mandatory personal, mandatory occupational and voluntary occupational private pension plans (coverage in percentage): Australia 90, Austria 35, Belgium 40, Canada 39, Denmark 90, Finland 7, France 10, Germany 57, Ireland 50, Italy 8, Japan 45, Netherlands 90, New Zealand 20, Norway 90, Sweden 90, Switzerland 90, United Kingdom 43, United States 47, missing = Netherlands.

The pairwise correlation between pension fund assets per GDP and private pension spending per GDP, as presented in Queisser et al. (2007: 555) is 0.70 and statistically significant. The correlation between pension fund assets per GDP and the coverage of private pensions (Queisser et al. 2007: 549) is also positive although the magnitude of the correlation coefficient is smaller and its statistical significance depends on the cross-sectional sample size. The overall findings support the use of pension fund assets per GDP as a proxy for the "privateness" of pension provision. The OECD (2007a) Social Expenditure Database and the OECD (2007b) Pension at a Glance series provide additional information on private pension spending. However, these indicators are cross-sectional or cover only a very short period of time; usually starting after 2004. Combining data from the OECD (2005)

Institutional Investors Database and the OECD (2007c) Global Pension Statistics allows investigating the development of pension fund assets per GDP for a period of up to 15 years.

In addition to the standard Esping-Andersen indicators recent studies on the changing structure of social risks suggest to consider labor market participation rates for marginal groups as an important feature of the configuration of old age security systems (Bonoli 2007). Issues of long-term unemployment, single parenthood, or the inability to reconcile work and family life are connected to pension policies. For instance, early retirement schemes have become a common practice for employers to shed labor and for older workers to exit the labor market. Particularly after the onset of mass unemployment in the 1970s, early retirement schemes have been extensively used in European countries (Ebbinghaus 2006). In fact, these policies reduced the financing base for public pensions while similarly promote new demand for benefits. Increasing labor force participation rates for females and elderly workers can provide a short-term alleviation against fiscal pressure on pension budgets. However, these policies come up against natural limitations and will raise pension entitlements in the long-run (Disney 2003). With growing female and elderly worker participation an increasing number of women and elderly workers will reach pension age with entitlements on their own. Thus, from a fiscal perspective higher labor force participation rates might have a positive short-term budget effect with unfunded pensions.

The last indicator concerns the design and redistributiveness of public pensions. As mentioned before, public pensions involve inter- and intra-generational redistribution, since pension contribution rates tend to be proportional to income, whereas benefits are often regressive. However, the extent to which public pensions are devoted to horizontal income redistribution varies across countries. These institutional differences may be captured with the distinction between Beveridgean and Bismarckian social policy. The Beveridgean model is characterized by universal, tax-financed, flat rate public pension provision, while the Bismarckian model is based on social insurance contributions and earnings-related benefits for employees (Bonoli 1997: 357). Thus, in tax-based Beveridgean public pension systems intra-generational redistribution tends to be more prominent than in Bismarckian insurance systems, where contributions are closely linked to benefit entitlements. Scharpf (2000a,b) suggests that Beveridgean welfare systems are less vulnerable toward demographic and economic pressures as tax-based social payments are likely to be easier targets of welfare retrenchment than contributory funded insurance systems. This is particularly evident in contribution financed pension system, where current worker's contributions enjoy the legal status of property rights which are usually protected against retrospective cuts. Following Kittel & Obinger (2003) the design of public pension schemes will be captured in the variable PAYG financing, which is defined as social security contributions as percentage of GDP divided by total tax revenues as percentage of GDP. To explore the appropriateness of this measure Table 3 shows the pairwise correlation between the PAYG financing measure and the index of public pension progressivity put forward in “Pensions at a Glance” (OECD 2007b: 45). On this index a pure tax financed basic scheme scores 100 and a pure insurance scheme scores zero. The correlation analysis confirms the expected negative relationship between the strength of PAYG financing and horizontal income redistribution in public pensions. Since the index of progressivity is available only as a cross-sectional measure the analysis will proceed using the PAYG financing indicator. To sum up, the six pension system indicators – pension expenditure per GDP, pension generosity scores, pension fund assets per GDP, labor force participation rates for females and elderly workers and the PAYG financing measure – are presented in Table 4.

Table 3. Pairwise correlation between PAYG financing and redistributiveness

	Index of Progressivity		
	Coef.	P-value	Obs.
PAYG financing	-0.7346*	0.00	18

Source: OECD (2007b: 45) Index of progressivity: Australia 73.1, Austria 30.4, Belgium 58.8, Canada 86.6, Denmark 59.3, Finland 7.6, France 24.6, Germany 26.7, Ireland 100, Italy 3.1, Japan 46.9, Netherlands 0, New Zealand 100, Norway 37.4, Sweden 12.9, Switzerland 53.3, United Kingdom 81.1, United States 40.9.

Table 4. Old age security in 18 OECD countries (1988-2003)

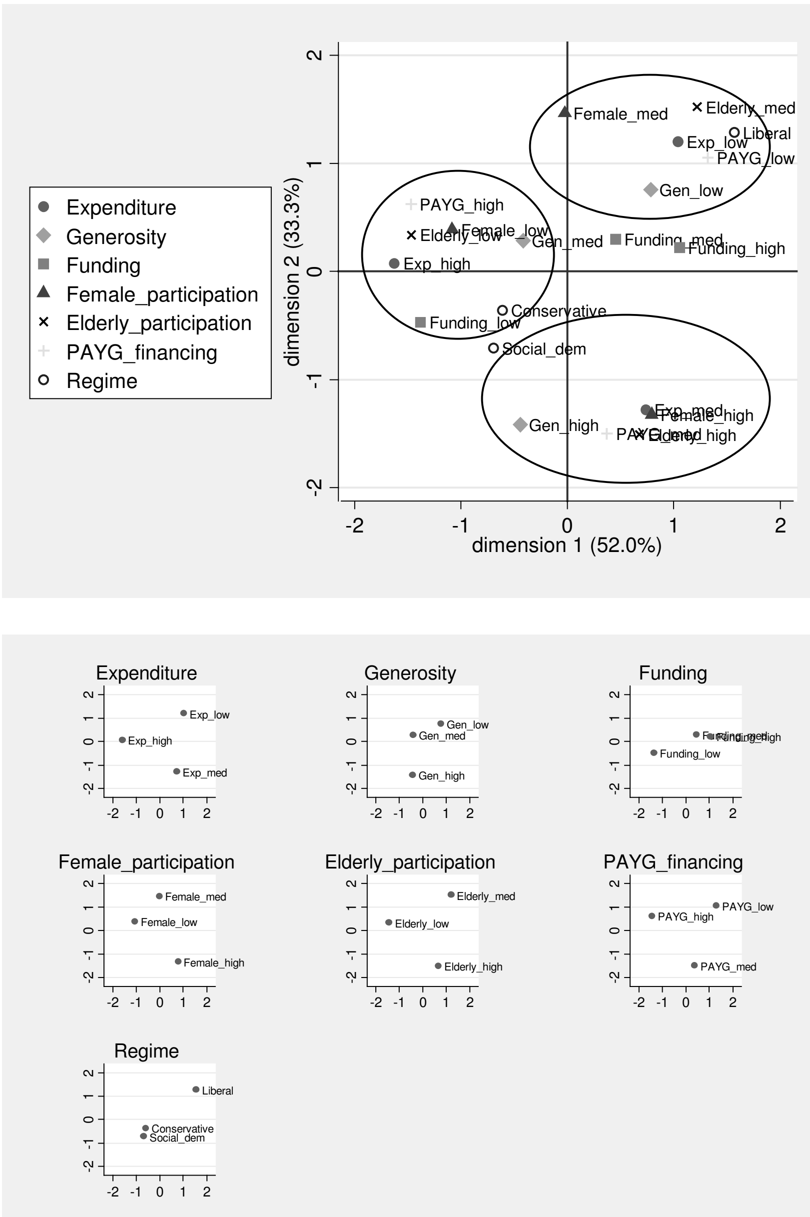
Country	Pension expenditure per GDP	Pension generosity scores	Pension fund assets per GDP	Female labor force participation rate	Elderly labor force participation rate	PAYG financing
Australia	3.12	8.86	42.48	62.92	45.58	0.00
Canada	4.10	13.61	39.38	68.93	49.49	13.43
Ireland	2.72	10.21	20.60	48.85	44.36	14.28
New Zealand*	5.81	15.57	13.11*	65.83	53.59	0.00
United Kingdom	4.95	8.41	68.03	67.67	52.67	17.56
United States	5.24	11.52	58.36	69.34	57.96	24.60
Liberal (av.)	4.32	11.36	40.32	63.92	50.61	11.65
Finland	5.52	13.90	3.23	71.27	44.70	25.65
France	9.91	13.92	3.77	59.77	37.93	41.43
Germany	10.05	7.72	2.94	60.81	41.90	38.05
Italy	10.03	14.49	3.61	44.46	30.82	32.64
Japan	5.12	9.46	14.87	58.56	65.83	31.71
Switzerland	6.09	6.58	79.58	70.58	65.46	24.79
Cons. (av.)	7.79	11.01	18.00	60.90	47.77	32.38
Austria	11.84	13.10	1.78	61.78	30.50	33.08
Belgium	6.88	12.39	3.93	51.56	24.20	31.98
Denmark	5.51	14.43	19.85	75.74	56.16	2.70
Netherlands	4.99	13.57	90.92	59.55	33.80	40.06
Norway	5.05	14.44	5.72	73.44	65.80	23.32
Sweden	7.28	14.90	2.69	78.29	69.20	27.39
Social dem. (av.)	6.93	13.81	20.81	66.72	46.61	26.42

Source: OECD Social Expenditure Database (2007a), Scruggs (2005a) Welfare State Entitlements Dataset, OECD (2005) Institutional Investors Database, OECD (2007c) Global Pension Statistics, OECD (2007d) Labor Force Statistics, OECD (2007e) Revenue Statistics,* only from 1996-2003

The selection of pension system indicators has been motivated by the contemporary literature in comparative welfare state research. If the six indicators adequately capture differences in the configuration of old age security systems, will be explored with the means of Multiple Correspondence Analysis (MCA). The basic idea of this method is to transform a complex data matrix into a new coordinate system in such a way that the greatest variance by any projection of the data lies on the first coordinate, the second greatest variance on the second coordinate, and so on. The results can be represented visually as points within a two or more dimensional space. Categories with similar distributions will be represented as points that are close in space while categories that are very dissimilar in their distribution will be positioned far apart (Clausen 1998: 10). Correspondence analysis starts with the transformation of frequencies in a cross-classification table into a set of row and column profiles. Chi-square distances, which are calculated separately for the row and column profiles, indicate if two profiles are similar or dissimilar. If the distance from the centroid, which is the weighted mean of all row and column profiles, is very large than the profile point is very different from the average profile. The extent to which the points spread around the centroid is measured in terms of intertia, which is analogous to the variance concept. Each dimension is evaluated on the basis of its contribution to the total inertia. The information of a single point's importance for the analysis is measured as its mass.

A very useful feature of correspondence analysis is the possibility of using supplementary, so-called passive variables. The categories of the passive variable are points without mass, hence they do not contribute to the inertia of the dimensions. However, the squared correlation of these points can be calculated and therefore the passive points can be located in the n-dimensional projection room opened up by the MCA (Clausen 1998: 21). For the purpose of this analysis Esping-Andersen's classification of welfare regimes (Table 1) will be employed as the passive variable. In order to use metric variables in MCA, the six pension indicators have been reduced to categorical variables using three percentiles representing the categories low, medium and high.

Figure 1. Multiple correspondence analyses (1988-2003)



Note: Supplementary (passive) variable Regime, coordinates in standard normalization, N=17 (Australia, Canada, Ireland, United Kingdom, United States, Finland, France, Germany, Italy, Japan, Switzerland, Austria, Belgium, Denmark, Netherlands, Norway, Sweden)

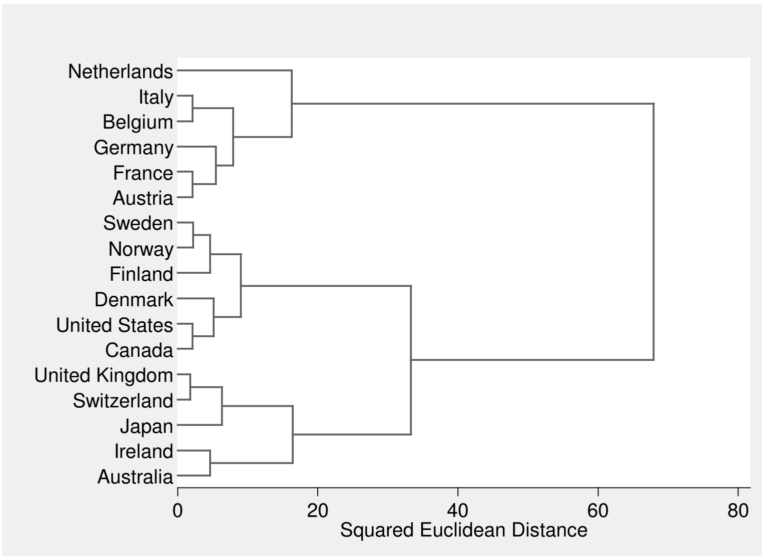
How does the location of points for the six indicators relate to Esping-Andersen's three world typology? If the regime approach properly describes distinct pension system configurations, the categories of the passive regime variables should be close to those points, which are associates with the properties of a certain regime. Figure 1 allows identifying three agglomerations of points. Low pension expenditure, low generosity, low PAYG financing and high pension fund assets are closely located to Esping-Andersen's liberal welfare regime. These systems are also characterized by a medium share of female and elderly labor force participation. The distance between the conservative and social democratic regime is much smaller than the distance of those two clusters to the liberal pension regime. The social democratic regime is relatively close to the agglomerations of points indicating medium spending, medium PAYG financing, high generosity, combined with a high share of female and elderly labor force participation. The conservative regime is relatively close to the clout of points indicating high public pension expenditure, high share of PAYG financing, low private pension spending and medium generosity scores. The points for low female and elderly labor force participation are also relatively close to the conservative regime. This pattern provides some support for Esping-Andersen's (1990) description of welfare regimes and it confirms findings by Scharpf (2000a, 2000b) on the configuration of employment systems in advanced welfare states. The MCA suggests two conclusions: First, the two-dimensional representation of the six pension indicators seems to capture a great deal of variance in old age security systems. Second, the projection reveals three distinct configurations of old age security provision.

3.2 Pension Regime Stability

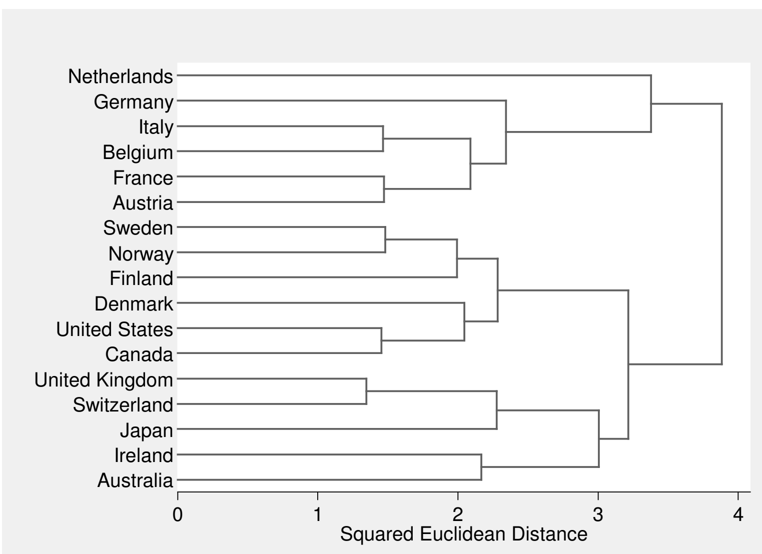
This section investigates which countries possess similar pension systems by means of hierarchical cluster analysis. Cluster Analysis (CA) has become a popular method in the field of comparative welfare research (Kangas 1994, Obinger & Wagschal 2001, Gough 2001, Powell & Barrientos 2004, Bambra 2007). The basic idea is to evaluate similarities between cases by calculating distance measures on a combination of variables that describe the cases. In this context, national pension systems are described on the basis of the six indicators that were presented in section 3.1. The resulting clusters are interpreted as pension regimes. Since variables with different metrics would contribute differently to the distance scores computed in the clustering process, the variables are transformed into standardized z-scores.

At the beginning of the hierarchical cluster process each country represents its own cluster. Step by step, the most similar objects are integrated into a new cluster until all objects are integrated into one cluster. Hence, at the end of the process, all countries are forced into the same cluster. The steps of the integration process are represented in the tree structure of a dendrogram. By a rule of thumb the vertical cut-off line is drawn when the distance measure increases drastically from one cluster step to another. Such a cluster procedure is called hierarchical because once two cases are joined into a cluster they remain joined (Gough 2001). In order to test the robustness of the partitioning, the analysis relies on two alternative algorithms for hierarchical clustering: Ward's linkage and average linkage. Average linkage clustering uses the average similarity of observations between two groups while Ward's linkage uses an analysis of variance approach to evaluate the distances between clusters. The latter attempts to minimize the sum of squares of any two clusters that can be formed at each step. In general, Ward's linkage is regarded as more efficient. However, it is important to note that cluster analysis is a heuristic technique to explore patterns of similarity; it is not capable of testing causal hypotheses. However, with respect to the research question, it provides a transparent and simple method to identify pension regimes by countries (Gough 2001).

Figure 2. Hierarchical cluster analysis (1988-1995)



Note: Ward method. Variables: Public pension expenditure per GDP, Pension generosity scores, Pension fund assets per GDP, Female participation rate, Elderly worker participation rate, PAYG financing

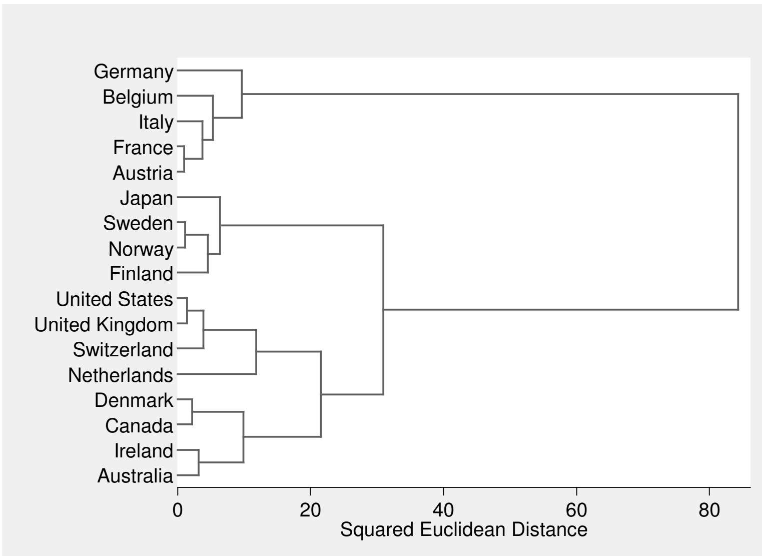


Note: Average linkage. Variables: Public pension expenditure per GDP, Pension generosity scores, Pension fund assets per GDP, Female participation rate, Elderly worker participation rate, PAYG financing

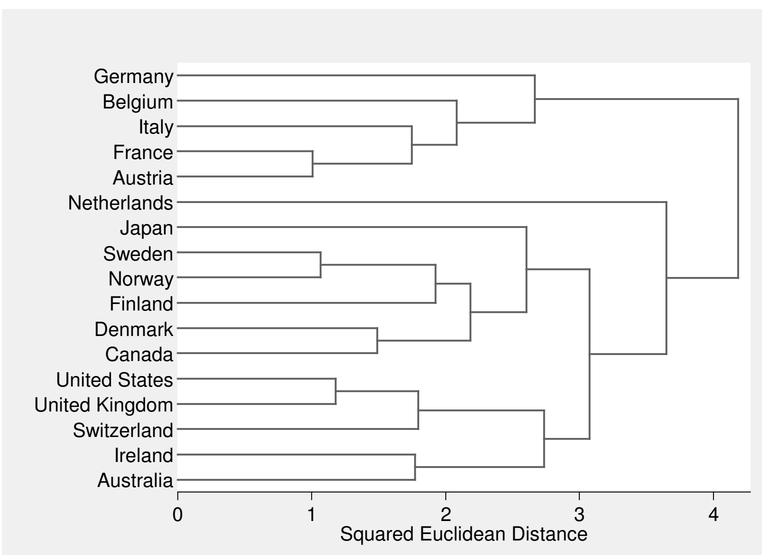
Since this study is interested in structural changes rather than annual fluctuation and reforms are likely to require several years to materialize in pension indicators, the empirical analysis of reform trajectories employs averages for two periods 1988-1995 and 1996-2003.[10] Figure 2 shows the results of the hierarchical cluster process for the early 90's (1988-1995) using Ward's linkage respectively the average linkage method on the six pension system indicators. In both specifications, pension systems in Germany, France, Italy, Belgium and Austria are within the same cluster. The second cluster divides into two subgroups, which include Sweden, Finland and Norway on one side and Denmark, the United States and Canada on the other. The first cluster partly supports Esping-Andersen's (1996) conservative welfare regime, the second cluster, however, locates social democratic and liberal countries within the same cluster. The third cluster is dominated by liberal countries (United Kingdom, Ireland, Australia) but also includes Japan and Switzerland – two conservative countries. The cluster results are robust with respect to the average linkage method.

[10] The selection of periods has also been restricted by the availability of data.

Figure 3. Hierarchical cluster analysis (1996-2003)



Note: Ward method. Variables: Public pension expenditure per GDP, Pension generosity scores, Pension fund assets per GDP, Female participation rate, Elderly worker participation rate, PAYG financing



Note: Average linkage. Variables: Public pension expenditure per GDP, Pension generosity scores, Pension fund assets per GDP, Female participation rate, Elderly worker participation rate, PAYG financing

Figure 3 presents results of the hierarchical cluster analysis for the second period (1996-2003). The first cluster still consists of conservative welfare states (Germany, France, Belgium, Italy and Austria). The second cluster contains three Scandinavian countries (Sweden, Norway, Finland) and Japan. The third cluster is dominated by liberal countries (United Kingdom, United States, Canada, Ireland, Australia) but also includes Denmark, Switzerland and the Netherlands.

Results of the CA in levels suggest that within the last 15 years the conservative pension regime remained unchanged, while the distinction between liberal and social democratic pension systems have faded away. Although the latter appears to be at odds with Esping-Andersen's welfare regime approach, a recent analysis by Lynch (2006: 57) also stresses the historical similarities between universal and means-tested old age security provision in liberal and social democratic countries. Moreover, it seem that the Danish pension system is an outlier within the group of Scandinavian countries. The liberal pension regime is growing in numbers. In the latest period the liberal pension regime incorporates countries such as Switzerland and the Netherlands, which have always been arguable cases in Esping-Andersen's typology. Findings suggest that if the analysis focuses on specific welfare programs instead of the welfare states as such, then Esping-Andersen's typology will have severe difficulties to explain the observed similarities between countries. A more coherent line of demarcation seems to run between insurance-based Bismarckian and tax-based Beveridgean pension systems. This might explain the stability of Esping-Andersen's conservative regime, which consists exclusively of welfare states inclined to the Bismarckian model of public pension provision. To explore this distinction in further detail the analysis shifts attention toward pension reform trajectories. Up to now the cluster analysis presents two snap shots of old age security configurations. To conclude that the stability of pension regimes – particularly the stability of the conservative pension regime – implies stasis might be misleading. Instead, following the logic of "bounded change", pension reforms in fact do take place within regimes.

3.3 Pension Reform Patterns

The six indicators that have been used to describe the properties of different configuration of old age security provision can also be interpreted as the core parameters in pension reform strategies. According to Chand & Jaeger (1996) and Disney (2003) pension reforms are either parametric or structural. Changing the relative size of pension expenditure and the generosity of pension benefits are the main "adjusting screws" in a parametric-reform strategy, which does not affect the dominant contribution-based financing method. Increasing the share of funded pension provision is a substantial part of a structural reform agenda. According to Scharpf (2000a), institutional differences among types of pension systems create differences in the vulnerability toward reform challenges such as new labor market trends, population aging and internationalized market competition. The essence is that despite the similarity of these challenges, policy response is assumed to converge among regimes.

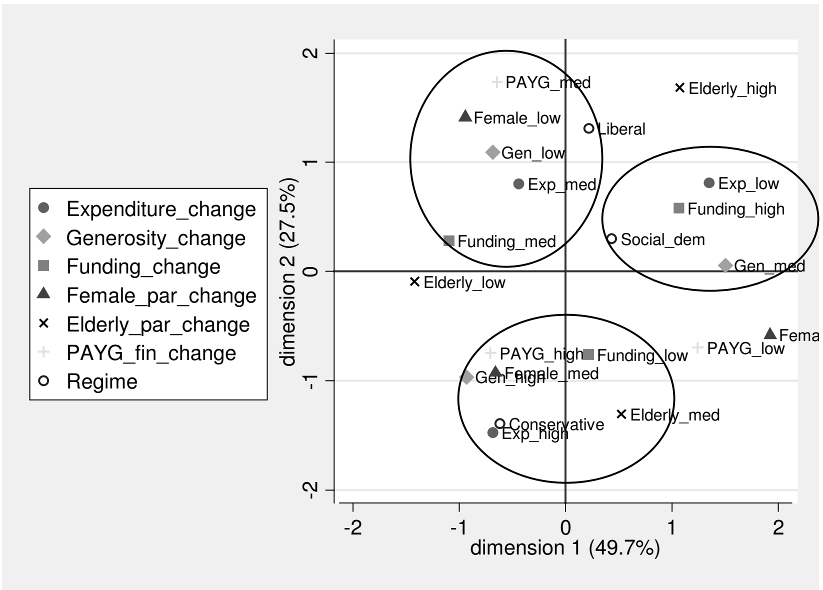
In order to investigate the “bounded change” argument, the MCA and CA is simply re-run using the period differences for each of the six pension system indicators. These differences are defined as the indicators’ values in the second period (1996-2003) minus their values in the first period (1988-1995). If the “bounded change” argument has something to say about pension reform trajectories in mature welfare states, countries within the same pension regime should also cluster into the same group with respect to change in the pension system indicators. Table 5 presents the development of the six pension system indicators within the last 15 years.

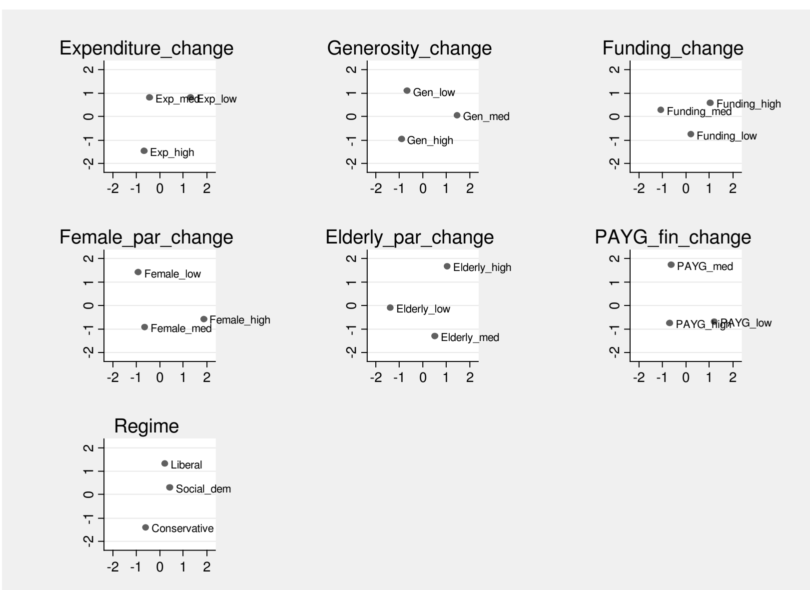
Table 5. Change in old age security in 18 OECD countries

Country	Δ Pension expenditure per GDP	Δ Pension generosity scores	Δ Pension fund assets per GDP	Δ Female labor force participation rate	Δ Elderly labor force participation rate	Δ PAYG financing
Australia	0.38	-0.70	30.93	3.51	4.82	0.00
Canada	0.02	-0.16	13.76	2.34	2.25	0.78
Ireland	-0.84	-0.61	28.89	10.28	3.65	-1.10
New Zealand	-1.65	-1.71	*	4.18	14.34	0.00
United Kingdom	0.46	0.60	16.09	1.72	0.95	-0.46
United States	-0.06	-0.52	22.35	2.11	3.92	-1.05
Liberal (av.)	-0.28	-0.52	22.40	4.02	4.99	-0.30
Finland	-1.57	-1.97	3.61	-0.10	4.58	-0.83
France	0.92	-1.48	5.74	3.43	1.18	-5.93
Germany	0.94	-0.58	0.41	4.57	3.77	2.16
Italy	2.11	0.53	-0.02	2.69	-2.35	-2.04
Japan	1.61	1.19	3.50	2.17	1.07	6.45
Switzerland	0.64	0.64	38.02	3.21	1.81	0.80
Cons. (av.)	0.78	-0.28	8.54	2.66	1.68	0.10
Austria	0.54	0.75	2.27	0.67	-0.04	0.21
Belgium	0.10	1.01	2.81	6.15	2.35	-1.94
Denmark	0.17	-1.38	7.09	-1.35	1.05	0.44
Netherlands	-0.54	0.01	26.46	9.86	7.53	-1.79
Norway	-0.40	-0.22	1.54	4.85	4.78	-3.15
Sweden	-0.47	-3.12	1.72	-3.50	0.99	0.58
Social dem. (av.)	-0.10	-0.49	6.98	2.78	2.78	-0.94

Source: OECD Social Expenditure Database (2007a), Scruggs (2005a) Welfare State Entitlements Dataset, OECD (2005) Institutional Investors Database, OECD (2007c) Global Pension Statistics, OECD (2007d) Labor Force Statistics, OECD (2007e) Revenue Statistics,* no data for the first period.
Note: Δ defined as period_t-period_{t-1}; period_{t-1} refers to averages for 1988-1995, period_t refers to averages for 1996 to 2003

Figure 4. Multiple correspondence analyses on change in pension indicators

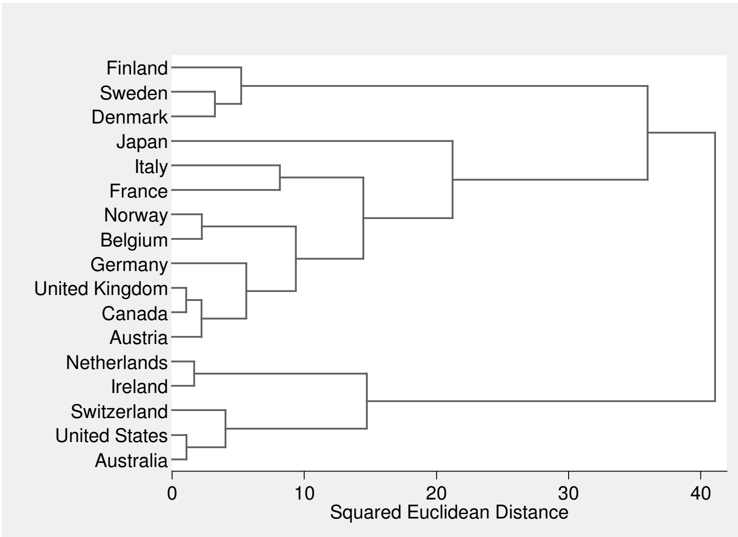




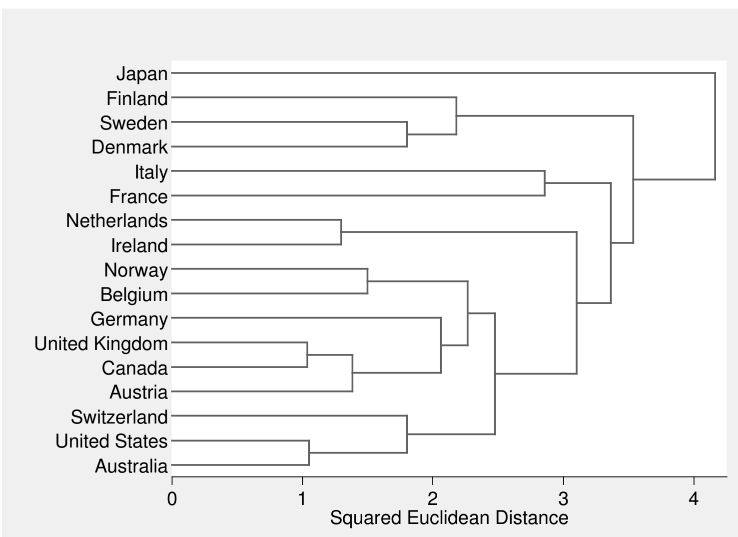
Note: Supplementary (passive) variable Regime, coordinates in standard normalization, N=17 (Australia, Canada, Ireland, United Kingdom, United States, Finland, France, Germany, Italy, Japan, Switzerland, Austria, Belgium, Denmark, Netherlands, Norway, Sweden)

Figure 4 presents results of the MCA using the differences of the six indicators. At first glance, the MCA plot shows that the two-dimensional representation still captures 77.2 percent of variance in the development of old age security systems. Esping-Andersens's (1990) classification of welfare states is once more entered as a passive variable into the MCA. The conservative regime is closely located to those points indicating increased pension expenditure, increased PAYG financing, increased benefit generosity, lower funding and unchanged labor force participation rates. In contrast to the MCA in levels, now the points for the conservative and social democratic regime are located relatively far apart, while it seem that the points for the liberal and social democratic regime have moved closer together. Moreover, the agglomeration of points in these regimes is less dense than in the conservative regime. This might indicate that conservative pension regimes follow a common pension reform trajectory, while pension reform in countries in the liberal or social democratic regimes is less similar. It appears that liberal pension systems maintain labor force participation rates and relatively high levels of funding while they decrease the generosity of public pension benefits. The points indicating lower pension expenditure, unchanged benefit generosity and an increase in the share of pension fund assets are closely located to Esping-Andersen's social democratic regime.

Figure 5. Hierarchical cluster analysis on change in pension indicators



Note: Ward method. Variables: ∆ Public pension expenditure per GDP, ∆ Pension generosity scores, ∆ Pension fund assets per GDP, ∆ Female participation rate, ∆ Elderly worker participation rate, ∆ PAYG financing

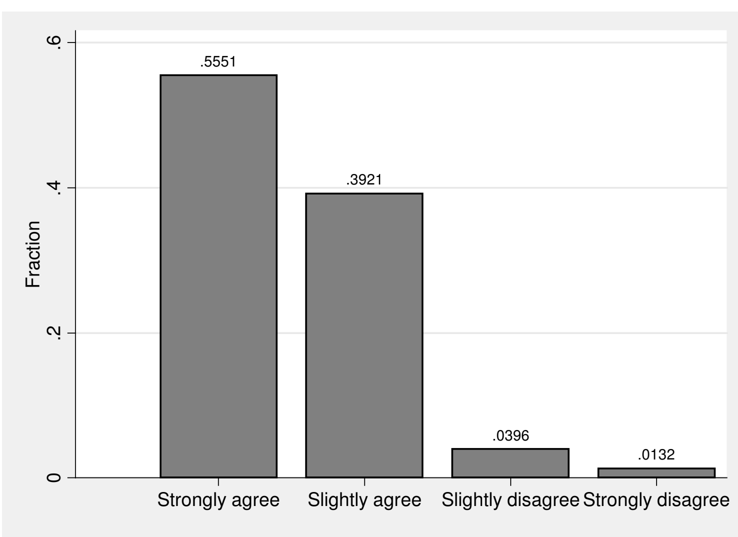


Note: Average linkage. Variables: ∆ Public pension expenditure per GDP, ∆ Pension generosity scores, ∆ Pension fund assets per GDP, ∆ Female participation rate, ∆ Elderly worker participation rate, ∆ PAYG financing

Results of the hierarchical CA for common reform trajectories are presented in Figure 5. The analysis reveals a less coherent clustering of pension systems than in the last section. The interpretation of the dendrogram is limited by the fact that using differences removes all systematic variation in the level of the six pension indicators. One can argue that it makes a difference whether a country with a high level of funded pensions increases funding by 5 percent or whether a country with virtually no pension funds increases the share of funding by the same amount. Thus, it might be justifiable to apply less rigorous criteria in the interpretation of Figure 5. Focusing on the first step of the hierarchical cluster procedure shows that the majority of country-pairs are found within the same regime (e.g. United Kingdom and Canada, France and Italy, Sweden and Denmark). The structure of these pairs is consistent also with respect to the average linkage method. Nevertheless, in contrast to the MCA, there is no doubt that the CA provides weak support for the "bounded change" argument. A careful interpretation of reform trajectories should consider both the level and change in the six pension indicators (see Table 4 and 5).

Concerning the first two research questions on pension regime stability and reform trajectories, the overall findings can be summarized as follows: The results of the MCA and CA in levels suggest that there are at least two clusters of old age security systems within the OECD world – insurance-based Bismarckian pension systems and tax-based Beveridgean systems. The latter separates into Anglo-American countries with low levels of generosity and high levels of funding and countries with a rather mixed system of old age security provisions (primarily Scandinavian countries), which combine tax and insurance-based financing. With respect to reform trajectories, the analysis suggest that maintaining or increasing an already high level of funded pension at the costs of reducing pension generosity characterizes the reform trajectory in liberal welfare states. In an analogy with the "Nixon in China Phenomenon", the Scandinavian reform trajectory represents a case where unlikely countries have implemented substantial policy changes – particularly with respect to pension privatization and fiscal conservatism. However, in contrast to the liberal trajectory, it appears that these systems have succeeded in extending funded pension provision without drastic cuts in benefit generosity. It can be assumed that these reforms benefit from the fact that Scandinavian countries tend to have mixed pension systems that used to organize intra-generational redistribution and pension insurance in separate schemes (Lynch 2006). Thus, instead of introducing a new pension pillar they are able to reallocate the weights of funded and unfunded pensions within the existing old age security system. Meyer, Bridgen & Riedmuller (2007) reached similar conclusions on the basis of qualitative country studies. The comparatively high dependence of Bismarckian pension systems on contribution-based insurance schemes creates a specific vulnerability toward demographic and economic challenges. These schemes seem to reinforce the "male-breadwinner" who gains pension entitlements in the form of protected property rights. The tight binding of pension contributions to entitlements makes Bismarckian systems relatively resistant toward cutbacks in pension spending or measures of means-testing. With the onset of population aging these systems find themselves in a situation of increasing fiscal pressure that can no longer be met by further increases in the contribution rate as this would increase non-wage labor costs and pose the threat of loosing international competitiveness (Scharpf 2000a). On the other hand, steps toward pension privatization might be even harder to achieve than in Anglo-American or Scandinavian systems as insurance-based pension schemes inhibit considerable implicit pension liabilities (Noord & Herd 1994). These liabilities become explicit in any case of privatization, causing a major double payment problem.

Figure 6. Preference for public pensions protecting against old age poverty



Source: Eurobarometer 56.1 (2001) Question 61 “The primary goal of a pension scheme should be to protect elderly people against the risk of poverty.”

3.4 Redistributive Effects

The identification of pension regimes and their specific reform trajectories entails limited information about the effect of pension reforms on the quality of old age security. However, this issue appears to dominate the political conflict over pension reform in mature welfare states. Therefore, the last research question concerns the redistributive consequences of pension reform trajectories identified in the last section. Recent studies focused attention on the redistributive consequences of welfare programs in advanced democracies (Scruggs 2005b, Moene and Wallerstein 2001, 2003). As Hacker (2004, 2006) has shown for the health sector, welfare reforms can expose citizens to new costs and substantial risks. Understanding the trade-offs involved in choosing alternative pension reform strategies might provide insights for developing feasible reform packages. Although public pensions are primarily a saving mechanism for old age income, these schemes also pursue various socio-political goals - protecting against old age poverty is traditionally one of them. The high preference for old age poverty protection through public pension is also confirmed by survey evidence from the Eurobarometer (2001). An astonishing 95 percent of the respondents agree that the primary goal of a pension scheme should be to protect elderly people against the risk of poverty. The redistributive consequence of old age security reform will therefore be approximated in terms of old age poverty rates. Changes in family types and income structure might also affect income inequality and relative poverty among the elderly. However, given the weight of public pensions in the disposable income of elderly people, public pensions are supposed to play a major role in shaping income adequacy and poverty risks among the elderly. To ensure that pensioners achieve some absolute, minimum standard of living compared with the population as a whole is normally the duty of the first tier of public pension schemes (Queisser et al. 2007). The last section already indicates that some OECD countries have substantially changed their system of old age security during the last 15 years. This section aims to evaluate the impact of these changes on old age poverty rates.

Table 6. Old age poverty in 18 OECD countries

Country	Level	Change*
Australia	20.0	8.0
Canada	3.5	1.0
Ireland	26.0	18.0
New Zealand	0.5	-1.0
United Kingdom	13.0	2.0
United States	23.0	4.0
Liberal (av.)	14.3	5.3
Finland	8.5	3.0
France	9.0	2.0
Germany	10.5	-1.0
Italy	15.0	0.0
Japan	22.0	-2.0
Switzerland	9.5	3.0
Conservative (av.)	12.4	0.8
Austria	12.0	-6.0
Belgium	14.5	1.0
Denmark	5.5	1.0
Netherlands	2.0	0.0
Norway	15.5	-7.0
Sweden	6.0	4.0
Social dem. (av.)	9.3	-1.2

Source: * averages in period_t-period_{t-1}; period_{t-1} refers to 1988-1995, period_t refers to 1996-2003, Foerster & d'Ercolei (2005), LIS for Belgium and Switzerland (various waves)

Table 6 presents old age poverty rates for the 18 mature welfare states under observation. The data is taken from Foerster & d'Ercolei (2005) and various waves of the Luxembourg Income Study (LIS). The poverty thresholds are set at 50% of the median income for the entire population and elderly refers to the population aged 66 and above. Foerster & d'Ercolei (2005) already point out that the recent trend of poverty among the elderly needs to be described against the backdrop of the long-term trend which implies a significant improvement of the economic situation for the elderly. However, since the mid-1990s the data suggests a departure from the long-term trend (Foerster & d'Ercolei 2005: 37). Average old age poverty rates per pension regime indicate that liberal countries have the highest level of old age poverty and the largest increase in old age poverty rates within the last decade. Countries in the conservative regime have a moderate level of old age poverty and experienced a moderate increase in old age poverty. On average, the social democratic regime remains to have the lowest level of old age poverty. However, there is considerable inconsistency towards Esping-Andersen's (1990) three world typology if the analysis compares countries and not regimes. For example, with Esping-Andersen's typology it is difficult to explain why old age poverty rates in Canada and New Zealand are up to 5 times lower than in Austria or Norway. Methodically, these inconsistencies, which might be partly due to the quality of the data, suggest focusing on change rather than the level of old age poverty.

The question is, how does change in the pension systems indicators affect income inequality and relative poverty among pensioners. Public pension expenditure per GDP and a higher pension replacement rate are expected to have a negative effect on old age poverty, while a larger share of private pensions is expected to have a positive effect on old age poverty, as it shifts the risk of old age poverty from collective insurance to private responsibility. The relationship between the labor force participation rates and old age poverty should be negative, as work represents the main aid against income poverty. The effect of PAYG financings on old age poverty is undecided. On the one hand, a less progressive pension system might allow more old age poverty, while on the other, a less progressive but on average more generous scheme could decrease the risk of old age poverty. Thus, it remains to be an empirical question how changes in PAYG financing affect change in old age poverty.

Table 7. Determinants of change in old age poverty rates

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
	Δ Old age poverty rate						
Δ Pension ex. per GDP	-0.96						-2.22
	[0.7]						[1.6]
Δ Pension generosity		-1.65					-3.02***
		[0.9]					[0.8]
Δ Pension fund assets			0.21**				0.34***
			[0.09]				[0.07]
Δ Female par. rate				-0.15			0.42
				[0.3]			[0.5]
Δ Elderly par. rate					-0.054		-1.60**
					[0.4]		[0.6]
Δ PAYG financing						0.15	0.21
						[0.4]	[0.3]
Δ Real GDP per capita	0.67	0.93	0.69	0.58	0.43	0.36	1.67***
	[0.8]	[0.9]	[0.8]	[0.9]	[0.8]	[0.7]	[0.5]
Δ ODR	38.2*	43.6**	34.6**	44.6*	40.2*	40.0*	34.6**
	[20]	[20]	[16]	[23]	[20]	[19]	[13]
Observations	17	17	17	17	17	17	17
Adj. R-squared	0.22	0.32	0.46	0.21	0.20	0.21	0.81

Note: OLS estimates, robust standard errors in brackets, constant not reported, N=17 (Australia, Canada, Ireland, United Kingdom, United States, Finland, France, Germany, Italy, Japan, Switzerland, Austria, Belgium, Denmark, Netherlands, Norway, Sweden), Δ defined as period_t-period_{t-1}; period_{t-1} refers to averages for 1988-1995, period_t refers to averages for 1996 to 2003, * $p \leq .10$; ** $p \leq .05$; *** $p \leq .01$

In order to explore the effect of the six reform parameters on old age poverty this study estimates a cross-sectional model in period differences. The literature on welfare retrenchment (Myles & Pierson 2001, Pierson 1994, 2001) and on policy change over time (Pierson 2000, Pierson 2004) similarly indicates that once a social security program is in place, policy makers can, at best, make marginal changes. Table 5 and 6 confirm that the starting level already captures most of the variation in pension expenditure and old age poverty. In order to identify the effect of the reform parameters, there is little to learn from an analysis in levels (Kittel & Obinger 2003), as old age poverty rates can be assumed to change gradually over time. Instead, if changes in the reform parameters play any role, then this should be evident through systematic changes in the dependent variable (Kittel & Winner 2005, Wooldridge 2002).[11] The cross-sectional regression analysis ignores the level of old age poverty and estimates how changes in the six reform parameters affect change in old age poverty rates using OLS estimates with robust standard errors.

Results of the regression analysis are presented in Table 7. In order to account for the macroeconomic development and population aging, each model includes change in the GDP per capita and change in the old age dependency ratio (ODR). Estimation coefficients for pension expenditure per GDP and pension replacement rates are negative and almost statistically significant in the latter case (Model 1 and 2). Change in pension generosity scores, GDP per capita and ODR

[11] A detailed definition of variables is given in Appendix Table A1.

already explains 32 percent of variance in change of old age poverty. The estimation coefficient for pension fund assets per GDP is negative and statistically significant (Model 3). Pedersen (2004) investigates the changing balance between public and private components in the income packages of old age pensioners in nine OECD countries using data from the Luxemburg Income Study (LIS). Although his analysis confirms that in all cases pensioners rely on a mixture of public and private pension provision, he finds weak and inconsistent evidence for the substitution of public pensions through private pensions. This study finds tentative evidence that an increase in pension fund assets can be associated with an increase in old age poverty. Change in pension fund assets, GDP per capita and ODR already explain 46 percent of variance in change of old age poverty. The effect of changing labor participation rates is non-significant and not robust to changes in the specification (Model 4 and 5). Changes in PAYG financing do not affect old age poverty (Model 6). Including all six reform parameters into the regression equation (Model 7) does not change the pattern of the estimation coefficients for the three main reforms parameters: public pension expenditure, public pension generosity and private pensions.

Although, due to the small number of observations (N=17), the estimation results can only have indicative relevance, they offer a possibility to link reform strategies to redistributive outcomes. Different reform strategies are likely to result in different risks of old age poverty. Linking results of the regression analysis with findings form the MCA reveals the following picture: Pension privatization contains a substantial risk of increasing old age poverty. Today, the risk of being poor when old is higher in the liberal than in the conservative or social democratic regime. This finding stands in contrast to Korpi & Palme (1998), who argue that the more a welfare state is devoted to equal public transfers, the less poverty will be reduced. Even after adjusting for macroeconomic and demographic determinants there is a strong positive correlation between funding and old age poverty on the one hand and a less strong negative correlation between benefit generosity and old age poverty on the other, indicating a potential trade-off in balancing fiscal needs against social equality. Moreover, although the relationship between pension spending per GDP and old age poverty is negative, the statistical association is rather weak. Against the backdrop of population ageing a larger share of overall pension spending has to be distributed among a larger share of elderly. Thus, increasing pension expenditure per GDP does no longer need to imply a reduction in the risk of old age poverty.

4. CONCLUDING REMARKS

This paper has examined whether pension reform in mature welfare states follows the logic of "bounded change". To do so, it has proceeded in three steps: First, it has explored the existence and stability of distinct pension regimes, second, it has traced change in pension provision to regime-specific reform trajectories and, finally, has linked these changes to changes in old age poverty. The study suggests that Esping-Andersen's (1990) three worlds typology has limited capacity to explain cross-national differences in old age security provision. This is particularly apparent when it comes to pension reform trajectories. Empirical evidence for the "bounded change" argument is therefore rather mixed. It finds its strongest support in the development of the conservative pension regime respectively countries with an insurance-based Bismarckian pension system. Countries within this cluster indeed follow a similar reform trajectory, which appear to be driven by a distinct vulnerability toward population aging and the double payment problem. These findings are consistent with prior qualitative oriented research on cross-country patterns in pension reform policies (Myles & Pierson 2001, Bonoli 2003). Moreover, it confirms Scharpf's (2000a,b) argument on regime-specific vulnerabilities. With respect to the quality of old age security, empirical evidence indicates that the selection of reform strategies is likely to affect old age poverty. It suggests that an increased risk of old age poverty is (at least partly) due to an increase of funded pensions. The social democratic pension regime has been relatively successful in balancing fiscal constraints and quality of old age security provision. Other authors have reached similar conclusions for social democratic welfare states with respect to their responsiveness to new social risks (Bonoli 2007). However, the empirical analysis also poses questions for further research. For example, are insurance-based Bismarckian welfare states already locked in a vicious cycle of fiscal imbalance and growing benefit generosity, heralding a gerontocracy, as this study might suggest?

And what instigated the reforms in the liberal and social democratic regimes since pension reform is known to be the "third rail" in social policy?

The main findings of this study can be summarized in three points: First, Esping-Andersen's (1990) typology has difficulties to explain the observed similarities between countries. Institutional differences between old age security systems seems to be better captured in terms of insurance-based Bismarckian pensions systems, which primarily aim at old age income maintenance for elderly workers, and tax-based Beveridgean pension systems, which primarily objective is to prevent old age poverty. Second, a naïve convergence thesis in pension reform is simply wrong. Although mature pension systems face similar challenges they respond differently toward these challenges. Thus, the coordination of pension policies may not lead to a single European pension model in the near future. Third, with respect to pension privatization, findings suggest that shifting toward funded pensions may require more careful considerations of its redistributive consequences. Besides fiscal sustainability, the quality of pension provision is an issue of increasing relevance in aging societies that has as yet received little attention in comparative welfare state research.

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Appendix Table 1. Definition and source of variables

Variable	Definition	Source
Pensions exp.	Public pension expenditure per GDP (cash)	OECD (2007a) Social Expenditure Database
Pension gen.	Pension generosity scores	Scruggs (2005a) Welfare State Entitlement Database
Private pension	Pension fund assets per GDP	OECD (2005) Institutional Investors Database OECD (2007c) Global Pension Statistics OECD
Female par. rate	Female labor force participation rate (percentage)	OECD (2007d) Labor Force Statistic
Elderly par. rate	Elderly (55-64) labor force participation rate (percentage)	OECD (2007d) Labor Force Statistic
PAYG financing	Social security contributions as percentage of GDP divided by the total tax revenues as percentage of GDP.	OECD (2007c) Revenue Statistics
Old age pov.	The poverty thresholds are set at 50% of the median income for the entire population. Old age refers to the population aged 66 and above.	Foerster & d'Ercolei (2005), Luxembourg Income Study
Real GDP per capita	Gross domestic product per capita (log)	World-Bank (2006) World Development Indicators
ODR	Old age dependency ratio measured as the share of the elderly (65+) as a percentage of the working age population (15-64)	OECD (2007d) Labor Force Statistics

Appendix Table 2. Preference for public pensions protecting against old age poverty

Country	Strongly agree	Slightly agree	Slightly disagree	Strongly disagree
Ireland	66.03	31.74	2.02	0.21
United Kingdom	64.79	32.15	2.65	0.4
Austria	45.92	43.1	8.58	2.41
Belgium	41.64	51.9	5.13	1.33
France	55.58	41.04	3.07	0.31
Germany (West)	46.56	46.76	4.32	2.36
Germany (East)	53.59	42.09	3.7	0.62
Italy	48.7	48.7	2.07	0.52
Netherlands	53.41	39.35	5.56	1.68
Denmark	65.63	29.64	3.23	1.51
Finland	55.51	40.2	3.67	0.61
Sweden	66.05	25.87	3.99	4.09
Total	55.51	39.21	3.96	1.32

Source: Eurobarometer 56.1 (2001) Question 61 "The primary goal of a pension scheme should be to protect elderly people against the risk of poverty."

Are Mature Welfare States on the Path to Gerontocracy?

Evidence from 18 OECD countries, 1980-2003

Abstract

This paper investigates the effect of majority voting, horizontal redistribution and expected population aging on the size and the generosity of public pensions in mature welfare states. Using recent panel data covering 18 OECD countries the empirical analysis provides tentative evidence that majority voting in aging societies has two opposite effects: it increases overall pension spending as a percentage of GDP but decreases the generosity of pension benefits, depending on the institutional design of the pension system. Moreover, while horizontal redistribution appears to matter less, estimation results indicate that expected population aging decreases both the size and generosity of public pensions. Contrary to popular fear that more pension spending reflects increasing "gray" voting power, this study indicates that aging OECD countries are not yet on the path to gerontocracy.

Keywords: median voter, demographic change, pension institutions

JEL Classification Numbers: H55, D72

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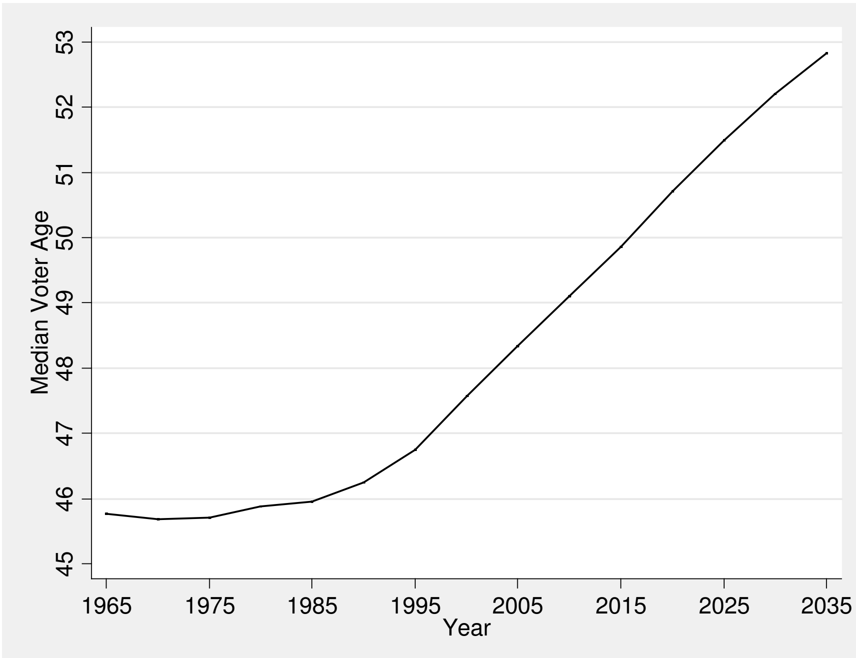
1. INTRODUCTION

In the fifteen years since 1990, the average OECD median voter age has increased three times faster than in the preceding thirty years. Fertility rates have hit post-war lows and elderly cohorts live longer but do not retire later. As a result, unfunded public pension systems in maturing welfare states have come increasingly under budgetary constraints. The economic and social consequences of population aging have risen to the top of the agenda of public policy. International institutions such as the IMF (2004), the OECD (1996) and the World Bank (1994) have devoted major policy documents to address an emerging "old age crisis". In electoral-numerical terms, the political balance between different age cohorts has definitely shifted in favor of the elderly in most rich democracies. This has led many observers to consider that democracies are increasingly likely to reflect a seniority bias, or elderly power (Anderson & Lynch 2007; Esping-Andersen & Sarasa 2002; Goerres 2007; Hinrichs 2002; Lynch 2006; Pampel 1994; Pampel & Williamson 1989; Van Parijs 1998).

Pension voting models in the spirit of Browning (1975) tend to present a gloomy picture of the redistributive consequences of demographic change (Mulligan & Sala-i-Martin 1999, 2003, Sinn & Uebelmesser 2002). As population ages, the median voter is getting older and the pro-public-pension coalition gains political clout. Once the elderly obtain an electoral majority, pension politics are locked-in and systemic reforms or even benefit reductions become much harder, heralding a possible beginning of gerontocracy.[1] Prior empirical research has found that the proportion of elderly citizens in the population increases overall pension expenditure. However, these findings do not discriminate well against alternative explanations. Since aging increases the objective need for pension spending, a hypothetical benevolent dictator would also spend more overall on pensions with a larger number of elderly people in the society. What makes theories of gerontocracy noteworthy is their prediction that population aging significantly affects the generosity of pensions. The more specific question therefore is not addressing program size but rather benefits generosity. Are public pensions more generous, the higher the weight of elderly voters? This paper investigates the effect of population aging on both the size and the generosity of public pensions in mature welfare states using two alternative ways to capture the latter variable. To do so, it uses data from a sample of 18 OECD countries covering a maximum time span from 1980 to 2003.

[1] See for Kotlikoff & Burns (2004) and Sinn & Uebelmesser (2002) for further discussion.

Figure 1. Average median voter age in 18 OECD countries (1965-2035)



Note: Based on United Nations (2006) World Population Prospects for 18 OECD countries, the median voter age *M* is computed by the following formula: $M = \sum_{k=1}^{16} \left(\frac{POP_k}{ELEC} * AGE_k \right)$, where *k* represents 5-year age cohorts (20-24, 25-29,…, 95-99), POP_k is the number of people in cohort *k*, *ELEC* is the total number of people who are eligible to vote and , AGE_k is the average age in cohort *k*.

In order to explore voting on old age security in aging societies, it is clearly desirable to investigate the relevant time period. Figure 1 shows the trend in the median voter age for all 18 countries included in the sample from 1965 to 2035. From 1965 to 1990 the average OECD median voter age increased by 0.7 years; but in half the subsequent time span – from 1990 to 2005, the median voter age actually increased by three times as much (2.1 years). The graph similarly indicates that the median voter age has indeed entered a phase of exponential growth, but only after 1990. Yet none of the political economy articles discussed above uses any data later than the 90's. Although the sample covers only 24 years, which appears to be a relatively short period compared to the emergence of public pension schemes, it might be a particularly important period. Thus, if population aging is indeed shifting mature welfare states on the path of gerontocracy, a sample covering the period 1980 to 2003 should be able to reveal this.

Within the last two decades the majority of OECD countries have started to implement pension reforms, which in some cases have already lead to a substantial reduction of pension benefits (Immergut et al. 2007; Myles & Pierson 2001; Taylor-Gooby 1999). Breyer & Stolte (2001) argue that existing models of rational voter behavior fail to explain why these reforms take place, since they shift the burden of demographic change completely or predominately on pensioners. Among the many public choice models on pension policies (see Breyer 1994, Galasso & Profeta 2002 for review), empirical research has to show which of these models provides a good approximation of the reality of pension politics. Since it would go beyond the scope of a single paper to discuss all theoretical models that link democratic voting and public pensions in aging societies, this paper concentrates on majority voting, horizontal redistribution and the role of expected future aging. Each of them has a concrete and testable empirical implication, which was the criterion for their selection. The intended contribution of this paper is primarily empirical as it complements, refines and partly enlarges prior studies by Breyer & Craig (1997) and Disney (2007).

The paper is organized as follows. The second section presents three alternative pension voting models. Focusing on its empirical implications, each will be compared to the benchmark model of a benevolent dictator. Section three presents the data and the estimation strategy. The fourth section presents the baseline estimation results with respect to the size and the generosity of public pensions, followed by a deeper investigation of the temporal stability of the relationship between population aging and public pensions. The fifth section finishes with some concluding remarks and a discussion of this study's implication for further research.

2. VOTING ON OLD AGE SECURITY

Voting models in the tradition of Browning (1975) stem from the median voter theorem and illustrate how re-election seeking politicians and self-interested voters determine the size of transfers in a system of unfunded social insurance. They assume that each person votes for the pension system that promises the largest lifetime utility. Thus, the size and generosity of public pensions are regarded as intended outcomes of a political decision process (Breyer 1994). The financing logic of public pensions is captured by the standard government budget constrain for unfunded welfare programs.[2] Pension schemes are financed through a payroll tax on

[2] See e.g. Breyer & Craig (1997: 708) for an algebraic representation.

current workers income, which provides a pension benefit to current retired. The retired are former workers who have reached retirement age and exited the labor market with an entitlement to old age pension benefits (Galasso and Profeta 2002: 2). Since the majority of pension systems in the OECD world are to a considerable extent unfunded (OECD 2007c), this appears to be a reasonable framework to start with.

Following Breyer & Craig (1997) the voting models are compared to the benchmark case in which a single benevolent decision maker decides the course of pension policies. Such a dictator with an infinite time horizon chooses an unfunded system if it provides a return on contributions which exceed the returns on savings. This would be the case if the population growth rate plus the growth rate of wages is higher than the interest rate (Aaron 1966). If the returns on contributions do not exceed the returns on savings, the decision of a benevolent dictator is less clear, because the benefits of the first generation of pensioners have to be weighted against the loss incurred by all future participants (Breyer & Craig 1997: 709). Since there is uncertainty about future growth rates and institutional rigidity that hinders radical changes to the pension system, the benevolent dictator is expected to adjust the size of public pensions over time according to the development of economic growth and the interest rate (Breyer & Craig 1997: 709). Uncertainty of future prices might also affect public pensions, as private pension saving has the disadvantage of being vulnerable to unanticipated inflation. Thus, in a world where prices are less flexible downwards than upwards, a benevolent dictator would rely more on unfunded pensions in a high inflation country than in a low inflation country (Breyer & Craig 1997: 709). Finally, the benchmark model also considers the long-term relationship between economic development and pubic pensions. According to Lindert (1996) the notion that social spending will rise with income growth is the most durable black box in comparative welfare state research. Since Wagner (1883), economists have observed that economic development is accompanied by an increased share of public expenditure. Wagner's (1911: 734) empirical "law of increasing state activity" claims "there is both an absolute and relative expansion of the public sector within the national economy". With respect to social security programs there are two perspectives on how economic development affects welfare spending. An optimistic perspective suggest that economic development provides the resources for more comprehensive social insurance, while a more pessimistic view suggests that economic development creates the problems that make social security programs even more urgent. Although there is no consensus why higher income should increase pension spending, both perspectives assume that old age security programs rise with the advancement of economic development (Lindert 1996: 6). To sum up, the benchmark model of a benevolent dictator provides alternative explanations for variance in the size of public pensions against which the predictions of the voting models will be tested.

2.1 Majority voting

In aging societies majority voting on public pensions can have two opposite effects: First, from the perspective of a stylized median voter theorem a larger share of elderly will move electoral support toward larger pensions, which is referred to as the “elderly power” hypothesis. Second, simultaneously low fertility rates and longer life expectancies decrease the profitability of an unfunded system for current contributors, which induce them to vote for smaller pensions. This effect, which is referred to as “fiscal leakage” (Razin, Sadka & Swagel 2002, Breyer & Stolte 2001), has become one of the most disputed issues in the political economy of old age security (see European Journal of Political Economy (2007) – special issue on this topic).

The “elderly power” hypothesis assumes the elderly vote for a revenue maximizing pension contribution rate as they internalize only the benefits and not the costs of higher contribution rates. Browning (1975) suggests that an unfunded pension system will therefore be supported even when its implicit rate of return is lower than the implicit rate of return in a corresponding funded system. This is the case if a high proportion of voters in or near retirement age benefits from the maintenance of the system even though they might be opposed it when they were young (Breyer & Stolte 2001: 410). Population aging is therefore predicted to generate “gray” voting power, as the elderly use their increasing clout to push for larger pension benefits. The costs of more generous pensions will be shifted toward the young and unborn generations that are not entitled to vote. In this respect, population aging would have a twofold effect; it decreases the implicit rate of return from an unfunded pension system and increases the share of voters favoring larger public pensions. This would lead to a paradox described by Marquardt & Peters (1997) as "collective madness", a situation in which the negative implicit rate of return makes the system politically more stable instead of less.

Recently, however, it has been argued that the second effect – the "fiscal leakage" from the median voter to the net beneficiaries of pubic pensions – might outperform the elderly voting majority. Razin et al. (2002: 901) argue that "the greater number of retirees increases the demand for benefits but at the same time reduces the willingness of the working-age population to accede to higher taxes and transfers, since current workers are net losers from the welfare state". The crucial factor determining the size of the pension system is whether the median voter is a net contributor or a net beneficiary of the pubic pension scheme. A higher share of elderly creates a higher tax or contribution burden on the people around the median voter (Razin et al. 2002: 916). Those people for whom the costs of higher pension contributions outweigh pension benefits vote for smaller pensions. Thus a larger share of elderly might decrease the size of pensions until the median voter is retired. At this point there would be a discontinuous jump up in the share of transfers toward the elderly (Razin et al. 2002: 916). The "fiscal leakage" hypothesis predicts that as long as the voting bloc of the retired does not represent the electoral majority, a larger share of elderly will actually reduce pension benefits (Razin & Sadka 2002).

2.2 Horizontal redistribution

Tabellini (2000), Casamatta, Cremer & Pastieu (1999) and Persson & Tabellini (2000) assume that some individuals may vote for larger public pensions because they suppose to benefit from its intra-generational redistribution component. By definition public pension systems redistribute income between young and old generations. However, in many systems of old age security, contribution rates are proportional to income, whereas benefits tend to be regressive. Therefore, redistribution within generations is also known to be a common feature of public pensions. The basic redistribution hypothesis in a unidimensional median voter conception of democratic voting was put forward by Meltzer and Richard (1981). It predicts that the income of the median voter relative to the mean influences the level of intra-generational redistribution: greater inequality ex ante should lead to more redistribution ex post (Scruggs 2005). This argument can be applied to public pensions, assuming that voters base their preferences on both age and income. A flat benefit pension scheme with proportional contribution rates does not only involve a transfer from workers to pensioners but also from high income to low income earners. In such a scheme the benefit of larger pensions is the same for all contributors, but the costs are higher for the richer. Poorer individuals will therefore prefer larger public pensions, as they benefit more from intra-generational redistribution. In this simple form the horizontal redistribution model predicts that greater income inequality makes the decisive voter more willing to exploit the pension system for intra-generational redistribution and thus increase the size of public pensions.[3]

The extent however to which public pensions are committed to intra-generational redistribution differs among countries (Queisser, Whitehouse & Whiteford 2007). This difference is generally referred to as Bismarckian or Beveridgean social security systems.[4] Contribution-based Bismarckian pension systems, which are present for example in countries such as Italy, France and Germany, are characterized by a relatively tight link between contributions and benefits. The pension system in countries such as Great Britain and Australia belong to the Beveridgean welfare state tradition. These schemes tend to provide flat benefits whereas taxation for public pensions is proportional to earning, thus Beveridgean pension schemes involve a rather larger share of within cohort redistribution (Conde-Ruiz & Profeta 2002). These institutional differences are likely

[3] For the empirical treatment the Persson & Tabellini (2000) model has been presented only with respect to intra-generational redistribution. In fact, the model is more complex than that and implies a voter coalition among the elderly and the poor.

[4] The term Bismarckian refers to Otto von Bismarck (1815-1898), first chancellor of the German Empire, under whose administration the Health Insurance Bill (1883), Accident Insurance Bill (1884) and Old Age and Disability Insurance Bill (1889) were implemented. The term Beveridgean refers to William Henry Beveridge (1879-1969), whose report on "Social Insurance and Allied Services" (1942) served as the basis for Britain's welfare legislation after the Second World War.

to affect voting for horizontal redistribution. It suggests that the positive effect of income inequality on the size of pensions should be stronger in systems with a loose binding of contributions and entitlements, where voters from low income groups have a higher chance to exploit intra-generational redistribution (Galasso & Profeta 2002: 9).

2.3 Expected population aging

While majority voting and horizontal redistribution have attracted the interest of scholars in the field for several years now, little attention has been paid to the role of expected population aging in voting on public pensions. In public pension reform debates, however, population projections seem to play an important role. Since the population in affluent OECD countries is forecast to age dramatically, projections raise doubts over the fiscal sustainability of public pension systems. Razin & Sadka (2005) assume that such doubts over the viability of public pension systems are creating momentum for reform. In this respect, expected future aging might provide a second mechanism by which population aging reduces the size of public pensions (Razin & Sadka 2005). According to the implicit generational contract underlying unfunded pension schemes, the young are willing to pay contributions in order to finance old-age benefits only because they expect the next generation to do the same for them. Just as population aging starts to decrease the profitability of an unfunded pension system, contributors realize that there are fewer individuals who will share in their contributions when they are retired. A growing awareness of continuing demographic change might make current contributors to expect smaller pension benefits for themselves than they are paying to the currently old (Razin & Sadka 2007: 565). Since they are not able to commit future voters and taxpayers to maintain the pubic pension system they will vote for smaller pensions (Breyer & Stolte 2001). This would imply that "the expected future aging of the population triggers the current young to vote for lower benefits (and taxes) for the current old" (Razin & Sadka 2007: 565). Interestingly, if such an effect exists, it is likely to persist for the next generations, since its self-validating character fuels the expectation that the following generation will also deviate form the implicit generational contract that underlies unfunded pension schemes. To sum up, the hypothesized effects of the three voting models on public pensions in maturing welfare states are compared in Table 1.

Table 1. Hypothesized effects on public pension spending and generosity

Model	Dependent variable	GDP capita	Growth rate	Interest rate	Infl. rate	ODR	Gini coef.	ODR proj.
Benevolent Dict.	Spending	+	+	-	+			
	Generosity	+	+	-	+			
Majority Voting	Spending					+		
	Generosity					+		
Horizontal Redis.	Spending						+	
	Generosity						+	
Expected Aging	Spending							-
	Generosity							-

2.4 Review and contribution to the literature

The majority of empirical studies suggest that the size of social security increases with the share of elderly people in the society (Tabellini 2000, Perotti 1996, Breyer & Craig 1997, Disney 2007). However, this result is strong only when the size of public pensions is measured as pension expenditure per GDP (Tabellini 2000, Perotti 1996, Breyer & Craig 1997). When the dependent variable is pension expenditure per elderly person, as in Breyer & Craig (1997) and Mulligan and Sala-i-Martin (1999), the proportion of elderly turns out to have no significant effect. Razin et al.'s (2002) empirical analysis finds that the old age dependency ratio has a statistically significant negative effect on the labor tax rate and on total social transfers. However, they apply standard OLS procedures on a sample of 13 countries over 28 periods, reporting 330 observations. This kind of cross-sectional time-series data is likely to violate the classical assumptions of the error term (groupwise heteroskedasticity, cross-sectional and serial correlation). Without testing the assumptions and applying necessary correction, OLS estimates are likely to produce invalid results. A detailed discussion of an estimation strategy, which is assumed to be more appropriate, will be presented in the following section.

Findings on horizontal redistribution suggest that social expenditure per GDP is (slightly) larger with greater inequality in pre-tax income (Tabellini 2000, Perotti 1996). However, weaker results are obtained by Lindert (1996) and Breyer & Craig (1997). Breyer & Craig (1997) find that the Gini coefficient has no statistically significant effect on pension spending. Scruggs (2005) assumes that empirical findings for horizontal redistribution tend to be disappointing because of persistent specification issues. Public pension spending is not necessarily redistributive. In fact, the redistributive share within public pension schemes varies widely among countries. To account for institutional differences Breyer & Craig (1997) employ a "Gini coefficient for a flat benefit rate", which is the product of the Gini coefficient and a dummy variable that takes the value 1 when the pension program provides a flat benefit and 0 otherwise. The "Gini coefficient for a flat benefit rate" is slightly significant and positive. However, Breyer & Craig (1997) assume that only four countries in their sample, namely Australia, Ireland, New Zealand and the Netherlands, would have flat rate pension schemes. Finally, concerning expected population aging there is no doubt that population projections play an important role in the public debate on pension reforms, however, its effect on voting for public pension has yet not been explored.

Disney (2007) argues that the design of social security programs should be taken into account when considering the political economy of public pensions. Referring to Razin et al. (2002) he points out that the "fiscal leakage" hypothesis is based on the implicit assumption that "the average worker does not see any link between his present tax (or the present old age pension) and his future benefit" (Simonovitis 2007: 536). If pension contribution where perceived as taxes rather than savings for benefits entitlements, working age voters are more likely to favor smaller pensions in the wake of population aging. The closer the pension program is to an "actuarial" scheme, the less pertinent is the "fiscal leakage" hypothesis (Disney 2007: 549). In a system where benefits are exactly linked to contributions, current contributors would not necessarily vote for smaller pension in an aging society as he or she stands to benefit proportionally from the program (Disney 2007: 549). The effect of population aging should therefore be weighted against the link between contributions and benefits. To capture the strength of this link Disney (2007) employs two proxy variables: the social security replacement rate calculated by Blöndal & Scarpetta (1998) and the average internal rate of return for an individual at age 55. Both variables are multiplied by either the old age dependency ratio or the inequality measure. Findings suggest that the old age dependency ratio has a positive effect on labor taxes per GDP. However, the coefficient for the product term (ODR x replacement rate) is negative, indicating that the "standard" positive relationship between population aging and the size of the welfare program is stronger in systems where benefits are closely linked to contributions.

Concerning the benevolent dictator model prior empirical work suggest that the size of public pensions is larger with a higher growth rate of the economy (Perotti 1996, Breyer & Craig 1997) and a higher inflation rate (Breyer & Craig 1997). Breyer & Craig (1997) provide evidence that the long-term interest rate is negatively correlated with the size of public pensions. Confirming "Wagner's law" they find that the size of public pensions as a fraction of GDP is larger, the higher GDP per capita is (Breyer & Craig 1997, Mulligan & Salia-i-Martin 1999). Finally, Breyer & Craig (1997) estimation results suggest a positve and significant time trend in pension spending.

On the whole, the empirical literature on the effect of aging on the size of pension programs tilts in the direction of a positive effect (see Table 2). The more specific question of whether and how aging affects the generosity of individual pension benefits remains, by large, unsettled. The finding that the demographic composition is an important determinant for the size of public pensions as measured by expenditure per GDP does not discriminate well against alternative explanations, since the variables do not precisely capture the predictions of the voting models. For instance, Persson & Tabellini (2000: 130) specifically predict, "that pensions per retiree will be higher, the higher the weight of old voters, as this shifts the median-voter equilibrium toward a more generous pension system". From a comparative welfare state research perspective, the case for studying benefit generosity, rather than program spending, is equally strong. Pointing out that "it is difficult to imagine that anyone struggled for spending per se," Esping-Andersen (1990: 21) criticized macro-social spending studies for their inability to indicate much about the impact of welfare programs on the well being of individual citizens or households. One core reason, imminently applicable to pensions and population aging, is that changes in macro-social needs can mask real cuts in individual benefits. As many observers have noted, whenever the percentage growth of welfare program dependents exceeds the percentage per capita reduction in welfare benefits, aggregate spending data will misleadingly indicate higher welfare effort even despite benefits cuts.[5] This study therefore differs in a fundamental point from prior empirical research - in the dependent variable. It employs three alternative measures: pension expenditure per GDP, pension replacement rates and pension spending per elderly, capturing the size and the generosity of public pensions. Table 2 summarizes the existing evidence on the political economy of public pensions.

[5] See furthermore Allan & Scruggs (2004: 498); Hinrichs & Kangas (2003); Kitschelt (2001); Korpi and Palme (2003); Scruggs and Allan (2006).

Table 2. Existing evidence on the political economy of social/old age security

Author(s)	N	Period	Dependent Variable	Population aging	Growth rate	Income	Interest rate	Inflation	Inequality
Lindert (1996)	19	Panel (1960-1981) 5*4 yr averages	Pension expenditure per GDP	Pop65 (+)		In GDP per adult (-)			Lower income gap (-)
Perotti (1996)	49	Cross sectional (1970-1985)	Average marginal tax rate	Pop65 (+)		GDP per capita (-)			Share of income in the 4th and 3rd quartile (-)
Breyer, Craig (1997)	20	Panel (1960-1990) 4*10 yr averages	Pension expenditure per GNP	Median voter age (+)	Real growth rate (+)	In GNP per capita (+)	Real interest rate (-)	Inflation rate (+)	Gini coefficient (-) Gini flat benefit (+)
Tabellini (2000)	40	Cross sectional (1978-1982)	Social security contributions per GNP	Pop65 (+)		In GNP per capita (+)			Top5 income share (+)
Razin, Sadka, Swagel (2002)	13	Panel (1965-1992) annual data	Labor taxes per GDP	ODR (-)	Per capita GDP growth (-)				Rich/middle income share (-)
Disney (2007)	21	Panel (1975-1995) 3*10 yr averages	Labor taxes per GDP	ODR (+) ODR * Replacement rate (-)	Growth of GDP (-)				Income share top 20%/share 20-79% (-) Income share top 20%/share 20-79% * Replacement rate (+)

Note: (+) positive association with the dependent variable (-) negative association with the dependent variable.

3. EMPIRICAL ANALYSIS

3.1 Description of the data

The empirical analysis relies on econometric methods to discriminate among the three pension voting models. To do so, it employs data from a sample of 18 OECD countries covering a maximum time span from 1980 to 2003.[6] Unlike the previous empirical literature, the data also includes the very recent period after the year 1990, when population aging – measured by the median voter age – starts growing exponentially (see Figure 1). Covering both the decade before and the decade after this "take-off" should offer a clear picture on the effect of population aging. Cash pension expenditure as a percentage of GDP (OECD 2007a) provides a measure for the size of public resources devoted to old age security. To test the effect of population aging on the generosity of public pensions and simultaneously to check the robustness of findings, generosity is measured with two different dependent variables. Following Breyer & Craig (1997) and Mulligan and Sala-i-Martin (1999), real pension expenditure per elderly person is used as a first proxy for benefit generosity. This variable is computed by dividing data from OECD (2007a) on real cash public pension expenditure by the number of individuals aged 65 and older (OECD 2007b). Pension generosity scores calculated by Scruggs (2005) provide an alternative measure for pension generosity. Pension generosity scores take into

[6] The sample includes Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Sweden, Switzerland, the United Kingdom and the United States.

account public pension coverage and public pension replacement rates. Replacement rates are among the most useful indicators for the generosity of a welfare program, since they provide a measure of the income that is made up by a welfare program (Korpi & Palme 2003, Allan & Scruggs 2005, Scruggs & Allan 2006). The pension replacement rate is defined as the ratio of net public pension paid to a person earning the average productive worker wage in each year of their working career. A second group of generosity measures focuses on the coverage of welfare programs. For example, Flora & Alber's (1982) study on the rise of Western welfare states does not use aggregate spending or replacement rates, but a measure on the proportion of people covered by public insurance schemes. The pension generosity scores calculated by Scruggs (2005) are based on both components: pension replacement rates and pension coverage.

The three pension voting models require that a number of explanatory variables should enter the statistical model. The set of control variables is based on the benevolent dictator model (Breyer & Craig 1997) and is mostly consistent with the control variables that have been in use in prior empirical studies (see Table 2). These are the log of GDP per capita, the real growth rate, the long-term interest rate and the inflation rate. With respect to the majority voting model the variable of main interest is the old age dependency ratio (ODR). The ODR is defined as the number of people aged 65+ divided by the number of people aged 15-64. Since this measure considers the relative strength of different generations, it is used to investigate the effect of population aging on pension spending. The number of individuals aged 65 and older divided by the total population is less useful regarding theories of the politics of gerontocracy as the denominator includes even more citizens below official voting age (Disney 2007). In order to test for horizontal redistribution the statistical model includes the Gini coefficient. Since cross-sectional measures of income inequality tend to be highly imperfect, the variable is taken from the World Income Inequality Database (WIDER 2005), which provides a quality measure for the Gini coefficient based on a consistent income concept. Finally, expected future aging is captured via population projection data. The variable indicating expected demographic change is defined as the projected ODR in 20 years minus the ODR in the observation year. Thus, a positive value for this measure indicates that a society's population is aging. Population projection data is taken from the United Nations (2006) World Population Prospect Population Database.

The design of public pension programs is captured via a progressivity indicator for the share of income redistribution involved in public pensions. Formally, the index of pension progressivity is calculated as 100 minus the ratio of the Gini coefficient of pension entitlements divided by the Gini coefficient of earning (OECD 2007c: 44). "Pure basic" pension systems pay the same flat rate amount to all pensioners regardless of their work and contribution history. The other extreme is a "pure insurance" pension system, where future pension benefits are closely linked to contributions. These two extreme designs underpin the index of progressivity so that "pure basic" schemes would score 100 and "pure insurance" schemes would score zero. Generally, "Basic" or flat rate pension schemes are associated with Beveridgean countries such as Great Britain, Ireland, Australia or New Zealand, while insurance-based schemes are more likely to be found in countries with a Bismarckian welfare state tradition such as Germany and France. In contrast to Breyer & Craig (1997), who use a dummy to account for Bismarckian and Beveridgean social security systems, the index of pension progressivity provides an enhanced measure for within cohort redistribution. The design of the pension scheme is taken into account by calculating the product of the index of progressivity and the ODR respectively the Gini coefficient. To do so the index has been transformed into an appropriate [-1,1] scale. By using this simple product term population aging and income inequality in countries with a more progressive pension scheme obtains higher values compared to countries with less redistributive schemes.[7] A detailed description and the source of each variable is presented in Appendix Table 1.

[7] Unfortunately the index of progressivity is available only as a cross-country variable. Thus, it is not possible to investigate its conditional effect within the fixed effects model. Nevertheless, using the index of progressivity as a weight on the ODR and inequality measure provides some indicative evidence on how the pension design might affect the effect of aging and inequality on pension spending and pension generosity. Moreover, this proceeding is consistent with prior empirical studies by Breyer & Craig (1997) and Disney (2007) who also draw their conclusions from a simple product terms.

3.2 Statistical model

Variables on pension spending and population aging are characterized by low volatilities, they change gradually over several years. In this case, annual observations and standard OLS estimates are likely to be problematic unless averaging procedures are applied (Disney 2007: 552). Following prior empirical studies by Lindert (1996), Breyer & Craig (1997) and Disney (2007) the annual observations have been grouped into eight-year averages. Using averages is because this study is interested in structural changes rather than annual fluctuation. Lindert (1996: 8) argues that this aggregation procedure also simplifies the adjustment for serial correlation. The variables for each country are calculated as averages of the period 1980 to 1986, 1987 to 1995 and 1996 to 2003, which makes a sample of up to 54 observations (T=3, N=18). One could argue that choosing eight-year periods for the averaging procedure is somehow arbitrary. Nevertheless, this approach is consistent with Breyer & Craig (1997), who use four ten-year periods, Disney (2007), who uses three ten-year periods, and Linder (1997), who uses five four-year periods. Acknowledging that electoral changes need time to materialize in pension spending or generosity measures, this study employs three eight-year periods. Each period represents two average legislation periods, which should be long enough for voters to influence the course of pension politics. The statistical model takes the following form:

$$y_{i,t}= \beta_0 + \beta_1 (x_{i,t}) + \beta_2 (z_{i,t}) +\gamma_i + \varepsilon_{i,t}$$

where $y_{i,t}$ denotes pension expenditure per GDP, pension generosity scores or real pension benefit per elderly in country i at time t. The variable x denotes the set of control variables (log GDP per capita, real GDP growth, long-term interest rate and the inflation rate) derived from the benevolent dictator model. The variable z denotes the independent variable of main interest, namely, the ODR, the Gini coefficient or the projected change in the ODR. The datasets panel structure suggests using the random or fixed effects estimator. The key question is whether the unit effects should be treated as random or fixed. The random effect estimator is heavily influenced by cross-sectional variance and depends on the assumption that unobserved heterogeneity is mean independent of the causal variable (Halaby 2004: 511). This assumption would be defensible under randomized assignment but not in a sample consisting of 18 OECD countries, where each unit is having a distinct set of social security institutions. The fixed effects estimator, which exploits within unit variation as a mean of purging unit heterogeneity, offers to dispense the random effects assumption and still obtains unbiased and consistent estimates when unit effects are arbitrarily correlated with explanatory variables (Halaby 2004: 516). Nielson & Andersen (1995: 686) argue that the fixed effect estimator can be interpreted as "throwing away" all between unit variations in the data. This is true, however it protects against biased and inconsistent parameter estimates, since the possible efficiency advantage of the random effects estimator depends on the random effects assumption. Without plausible theoretical grounds or empirical evidence for the random effects assumption, bias and consistency considerations alone would lead to a fixed effects model (Halaby 2004: 521). Allison (1994: 181) asserts that the fixed effects estimator is nearly always preferable to the random effect estimator with non-experimental data. Nickell (1981: 1418) takes a similar position on this issue saying that "if one takes the view that, in any particular model, the individual effects are likely to be correlated with all the observed exogenous variables, then one is lead inexorably to the fixed effects model." Moreover, the pension voting models also strongly suggested employing the fixed effects (or within country) estimator, since their predictions concern how electoral outcomes change pension spending and generosity within and not between countries. Even Breyer & Craig (1997: 717), who rely on the random effect estimator, admit that "our prior belief was that within-country estimates are more reliable tests of (…) public choice models (…)".

Disney (2007) employs feasible generalized least squares (FGLS) to estimate the effect of population aging on pension spending. However, as shown by Beck & Katz (1995), this method has serious problems with cross-sectional time-series data when T is relatively small compared to N. In this case, the FGLS method yields standard errors that are too small (up to 600 percent) and therefore produces overconfident results. Beck & Katz (1995) recommend using FGLS only when T is very large relative to N. However, this is obviously not the case with Disney's

(2007) dataset with N=21 and T=3. To be on the safe side this study employs a rather conservative estimation strategy and relies on a fixed effect estimator to obtained unbiased and efficient estimates.

Table 3. Pension spending and generosity in 18 OECD countries

	Pension expenditure per GDP		Pension generosity scores		Pension benefits per elderly		Index of progressivity
	Mean	Delta	Mean	Delta	Mean	Delta	
Australia	3.08	0.29	9.03	-0.85	9.57	0.01	73.10
Austria	11.33	1.81	12.70	1.56	9.83	-0.19	30.40
Belgium	6.70	0.60	12.17	1.18	9.25	-0.34	58.80
Canada	3.83	0.84	13.14	1.32	9.68	-0.04	86.60
Denmark	5.21	0.97	14.54	-1.02	13.33	-0.15	59.30
Finland	5.58	-0.97	14.53	-2.87	9.14	-0.55	7.60
France	9.30	2.29	14.28	-1.81	9.58	-0.15	24.60
Germany	10.00	0.62	7.81	-0.55	9.60	-0.17	26.70
Ireland	3.22	-1.91	10.26	-0.46	8.66	-0.44	100.00
Italy	9.52	2.59	13.65	2.79	9.45	-0.37	3.10
Japan	4.61	2.34	9.41	0.76	19.03	-0.39	46.90
Netherlands	5.11	-0.63	13.65	-0.25	9.19	-0.43	0.00
New Zealand	6.22	-2.04	16.07	-2.35	10.42	-0.84	100.00
Norway	4.76	0.68	14.40	0.00	13.59	0.09	37.40
Sweden	7.19	0.05	16.10	-5.14	13.64	-0.18	12.90
Switzerland	5.94	0.77	6.71	-0.06	10.60	-0.42	53.30
United Kingdom	4.67	1.08	8.33	0.56	8.11	-0.08	81.10
United States	5.29	-0.15	11.59	-0.47	9.62	-0.18	40.90

Source: OECD (2007) Social expenditure database, Scruggs (2005) Welfare States Entitlement Dataset, OECD (2007) Labor Force Statistic, OECD (2007: 45) Pensions at a Glance

4. RESULTS

4.1 Descriptive analysis

Table 3 compares the development of the measures on the size and generosity of public pensions between 1980 and 2003 and shows considerable cross-sectional variance in the levels of pension effort on all three measures; for instance, the widely varying levels of pension spending per GDP and pension generosity scores for countries such as Austria and Australia. Italy, France and Austria, all of which are Bismarckian welfare states, stand out with respect to all three effort measures, most strikingly so with respect to pension spending per GDP and pension generosity scores. Only five countries in the sample reduced pension expenditure per GDP during the observation period. While reductions in pension generosity scores took place somewhat more frequently, real pension benefits per elderly (log) decreased in each single country. Finally, taking into account the index of progressivity confirms that public pension schemes with a Bismarckian tradition in countries such as Germany, France and Austria involve a relatively low amount of intra-generations redistribution.

Table 4. Pairwise correlation analyses

	ODR	Gini coefficient	ODR projection	Progres. Index
Pension expenditure per GDP	0.49	-0.25	0.07	-0.53
	(0.00)	(0.07)	(0.60)	(0.00)
Pension generosity scores	0.08	-0.30	-0.20	-0.26
	(0.57)	(0.03)	(0.15)	(0.06)
Real pension benefits per elderly	0.13	-0.25	0.33	-0.10
	(0.33)	(0.06)	(0.02)	(0.49)

Note: Partial correlation coefficient, significance levels in brackets.

Table 4 shows the partial correlation between the three dependent variables and the independent variables of main interest. First, although the correlation analysis suggests a positive relationship between ODR and the three dependent variables, the correlation is significant only with respect to pension expenditure per GDP. Second, there is a negative and statistically significant correlation between the Gini coefficient and the size and generosity of old age spending. Third, with respect to change in the ODR there is a positive correlation with real pension benefits per elderly. Although a pairwise correlation analysis can only provide indicative evidence, it suggest at least that the relationship between the independent variables of main interest and public pensions depends on how public pension are measured – in terms of spending or generosity. This pattern gives reason to proceed using three alternative dependent variables in the panel regression analysis.

Table 5. Determinants of pension expenditure per GDP

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
	Pension expenditure per GDP					
	Random	Fixed	Fixed	Fixed	Fixed	Fixed
GDP per capita (log)	0.79	0.97	0.54	0.93	0.98	-0.55
	[0.6]	[0.6]	[0.3]	[0.6]	[0.5]	[-0.4]
GDP growth	-0.43***	-0.45***	-0.51***	-0.47***	-0.45***	-0.39***
	[-3.5]	[-3.4]	[-4.1]	[-3.4]	[-3.3]	[-3.8]
Interest rate	-0.16**	-0.16*	-0.18**	-0.16*	-0.16*	-0.16**
	[-2.2]	[-1.9]	[-2.1]	[-2.0]	[-1.9]	[-2.3]
Inflation rate	0.12	0.12	0.13	0.13	0.12	0.19**
	[1.6]	[1.2]	[1.5]	[1.3]	[1.2]	[2.2]
ODR	0.13*	0.093		0.094	0.093	
	[1.7]	[1.4]		[1.4]	[1.3]	
ODR*Prog.			-0.15			
			[-1.1]			
Gini coefficient				-0.04		
				[-0.8]		
Gini*Prog					0.00	
					[0.02]	
ODR 20 year dif.						-0.16***
						[-3.5]
1988-1995	-0.03	-0.07	0.23	0.02	-0.08	1.46*
	[-0.04]	[-0.09]	[0.3]	[0.03]	[-0.09]	[1.8]
1996-2003	-0.38	-0.44	0.057	-0.39	-0.45	2.47*
	[-0.4]	[-0.3]	[0.04]	[-0.3]	[-0.3]	[1.8]
Observations	54	54	54	54	54	54
Number of id	18	18	18	18	18	18
R-squared	0.33	0.54	0.56	0.55	0.54	0.66
Breusch-Pagan	39.23***					
Hausmann	7.99					
P-value	0.15					

Note: N=18 countries, T= averages for 3 periods (1980-1987, 1988-1995, 1996-2003) t-statistics in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance. Breusch-Pagan = Breusch Pagan lagrange multiplier test (random effects vs. pooled) Hausman = Hausman test (random effects vs. fixed effects) P-value = P-value for the Hausman test.

4.2 Determinants of pension spending

Since Breyer & Craig (1997) draw their main conclusion from a random effect specification, Table 5 starts with the random effect error scheme (Model 1). The Breusch-Pagan lagrange multiplier test, presented in the lower block of the table, indicates that a random effect model is indeed preferred over the pooled OLS specification. The Hausman (1978) test is used as an additional test to evaluate the significance of the random effect estimator over the fixed effects estimator. Its result is presented in the lower block of the regression table and indicates that the random effect specification is slightly more efficient than the fixed effects estimator (p=0.15). However, the test requires that both estimators are consistent. As this assumption is clearly violated for the random effect estimator, the analysis follows the intuitive argument and uses fixed effects. Nevertheless, to cross check estimation results with Breyer & Craig (1997) both estimators are presented in regression Table 5. The overall pattern of the control variables is consistent with the benevolent dictator model and confirms prior findings by Breyer & Craig (1997). The interest rate has a negative effect on pension expenditure per GDP, while the inflation rate and GDP per capita have a positive effect. The real growth rate has a negative effect on pension spending, which confirms findings by Disney (2007) but is difficult to interpret. There are two explanations for this finding: First, since GDP is used as a denominator for pension expenditure, GDP growth decreases pension expenditure per GDP if pension expenditure remains unchanged. Second, following the formal argument by Breyer & Craig (1997), pension expenditure per GDP represents a proxy for the pension contribution rate. In this respect, GDP growth may reduce the relative burden of social insurance contributions on labor income (Disney 2007: 547).

Model 2 in Table 5 supports the notion that population aging increases overall pension spending, although the coefficient (0.10) is slightly non-significant. Disney reports an estimation coefficient of 0.17*** for the effect of the ODR on labor taxes as a share of GDP applying FGLS estimation procedures. Using the fixed effect estimator and after controlling for non-demographic influences the effect of the ODR is less strong than expected. Nevertheless, the estimated coefficient is robust with respect to different model specifications. A larger share of elderly is therefore likely to increase pension spending per GDP, as societies with more elderly should have to devote more public resources to pensions. The estimated coefficient for the product of ODR and the index of progressivity, however, turns out to be negative, which indicates that in more redistributive pension systems population aging might decrease pension spending per GDP. Disney finds a similar pattern for his product terms, although using FGLS estimates on a T=3 and N=21 sample yields statistically significant coefficients. In total, findings provide indicative support for the idea that the "fiscal leakage" effect is more pertinent if intra-generation redistribution matters more in the design of the pension system.

Model 3 in Table 5 includes the Gini-coefficient in order to test the horizontal redistribution model. The estimated coefficient for the Gini-coefficient is negative. Although in both instances the effect is statistically non-significant, it confirms results from the pairwise correlation analysis, indicating that countries with less income inequality spend more on old age security. Disney (2007) also finds that income inequality has a negative effect on labor taxes as a % of GDP. This is just the opposite of what the Meltzer and Richard (1981) model implies. The relationship between income inequality and pension spending has also been explored using an alternative measure for income inequality. Based on WIDER (2005), income inequality has been defined as the income share top 20 percent divided by income share 20-79 percent. The alternative measure equally suggests a negative relationship between income inequality and pension spending.[8] Prior studies on horizontal redistribution by Bénabou (1996), Perotti (1996), Lindert (1996), Alesina & Glaeser (2004) and Moene & Wallerstein (2001) also find that equality in market income is more likely to be associated with higher redistribution. Alesina & Glaeser (2004: 59) suggest two possible explanations for this so-called "Robin Hood paradox". First, countries use very different means to redistribute income. The before-tax Gini coefficient might be a poor indicator if redistribution has taken place before earning occur at all (e.g. through education). Second, in countries with greater income inequality, the poor may have not enough political influence. With respect to this

[8] The estimation coefficient for income inequality measures as the income share of top 20 percent divided by the income share 20-79 percent is -0.03 [-1.7].

study's sample the latter explanation may not be convincing. Other authors (Iversen & Soskice 2006, Persson & Tabellini 2005) have therefore stressed the importance of the electoral system, whether proportional or majoritarian representation, as a determinant for the ability of voters from low income groups to increase horizontal redistribution. If institutional differences between Bismarckian and Beveridgean systems are taken into account through the product term of the Gini-coefficient and the index of progressivity, the estimated coefficient is positive, providing support for the horizontal redistribution argument. This finding is also consistent with Disney (2007: 550) and Breyer & Craig (1997). However, applying the more appropriate fixed effects error scheme, estimation coefficients lose their statistical significance.

Finally, Model 6 includes the difference between the projected $ODR_{i,t+20}$ and the current $ODR_{i,t}$. Larger positive differences between these two measures indicate a steeper increase in expected population aging. The estimated coefficient for change in the ODR within the next 20 years is negative and significant at the highest margin. It suggests not only that voters and politicians seem to be responsive toward long-term demographic change; it also indicates that population projections heralding further population aging tend to decrease pension spending. This finding can be interpreted as support for Razin & Sadka (2005) who assume that expected future aging fuels doubts over the viability of public pensions and thereby creates a momentum for reductions in the size of pensions. On a more fundamental level, this result speaks for a refinement of Hicks & Zorn's (2005) "paradox of self-limiting immoderation” hypothesis which predicts that factors promoting welfare spending simultaneously building up pressures for fiscal adjustment. Thus, population aging might switch over to pressures for reductions in pension spending. Hicks & Zorn’s (2005) study on welfare retrenchment indeed suggests that an increasing ODR increases the likelihood of welfare retrenchment; however, they had to release an erratum (Hicks & Zorn 2007) which nullifies their findings. Nevertheless, with respect to the findings in this study, the “paradox of self-limiting immoderation” could be reinterpreted. Thus, the current share of elderly voters matters less for pension retrenchment polices; it is the expected demographic imbalance that frames the pension debate and creates a push for pension retrenchment.

Table 6. Determinants of pension generosity scores

	Model 1	Model 2	Model 3	Model 4	Model 5
			Pension generosity scores		
	Fixed	Fixed	Fixed	Fixed	Fixed
GDP per capita (log)	5.93	6.08*	5.84	7.36*	3.52
	[1.6]	[1.7]	[1.7]	[1.8]	[1.2]
GDP growth	-0.28	-0.45**	-0.31	-0.3	-0.21
	[-1.4]	[-2.2]	[-1.4]	[-1.5]	[-1.2]
Interest rate	-0.004	-0.04	-0.012	-0.002	-0.01
	[-0.03]	[-0.3]	[-0.1]	[-0.01]	[-0.09]
Inflation rate	0.047	0.059	0.072	0.008	0.17*
	[0.5]	[0.6]	[0.7]	[0.08]	[1.8]
ODR	0.19		0.19	0.13	
	[1.2]		[1.2]	[1.1]	
ODR*Prog.		0.01			
		[0.03]			
Gini coefficient			-0.09		
			[-1.1]		
Gini*Prog.				0.22	
				[1.3]	
ODR 20 year dif.					-0.26***
					[-3.0]
1988-1995	-2.54	-2.37	-2.32	-3.19	-0.029
	[-1.5]	[-1.5]	[-1.5]	[-1.6]	[-0.02]
1996-2003	-4.74	-4.4	-4.62*	-5.79*	0.024
	[-1.7]	[-1.7]	[-1.7]	[-1.8]	[0.01]
Observations	54	54	54	54	54
Number of id	18	18	18	18	18
R-squared	0.24	0.20	0.28	0.29	0.4

Note: N=18 countries, T= averages for 3 periods (1980-1987, 1988-1995, 1996-2003) t-statistics in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance.

Table 7. Determinants of real pension benefits per elderly

	Model 1	Model 2	Model 3	Model 4	Model 5
	Real pension benefits per elderly				
	Fixed	Fixed	Fixed	Fixed	Fixed
GDP per capita (log)	0.16	0.21	0.16	0.18	0.12
	[0.5]	[0.5]	[0.4]	[0.4]	[0.3]
GDP growth	-0.073***	-0.042	-0.073**	-0.073**	-0.037
	[-2.8]	[-1.3]	[-2.7]	[-2.7]	[-1.1]
Interest rate	-0.038	-0.031	-0.038	-0.038	-0.030*
	[-1.7]	[-1.7]	[-1.7]	[-1.7]	[-1.7]
Inflation rate	0.028	0.024	0.028	0.027	0.026
	[1.2]	[1.0]	[1.2]	[1.2]	[1.1]
ODR	-0.039**		-0.039**	-0.040**	
	[-2.7]		[-2.6]	[-2.6]	
ODR*Prog.		0.023			
		[0.9]			
Gini coefficient			-0.001		
			[-0.1]		
Gini*Prog.				0.002	
				[0.1]	
ODR 20 year dif.					-0.002
					[-0.1]
1988-1995	-0.35**	-0.42**	-0.35**	-0.35*	-0.37*
	[-2.2]	[-2.4]	[-2.1]	[-2.0]	[-1.7]
1996-2003	-0.31	-0.43	-0.31	-0.32	-0.35
	[-1.2]	[-1.5]	[-1.2]	[-1.2]	[-1.0]
Observations	54	54	54	54	54
Number of id	18	18	18	18	18
R-squared	0.78	0.75	0.78	0.78	0.74

Note: N=18 countries, T= averages for 3 periods (1980-1987, 1988-1995, 1996-2003) t-statistics in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance.

4.3 Determinants of pension generosity

Tables 6 and 7 present estimation results with respect to the generosity of public pensions. In Table 6 pension generosity scores are used as the dependent variable; in Table 7 real pension benefits per elderly are used as the dependent variable.

Concerning the benchmark model GDP per capita, real GDP growth, the inflation and interest rate remain to have almost the same effect as if the dependent variable measures pension spending. Although the effect of the long-term interest rate is no longer statistically significant, the overall pattern for the control variables implies a certain degree of robustness of results obtained for the benchmark model. However, if pension generosity is measured in terms of generosity scores, the benchmark model explains only 20 percent of within variance.

Model 1 in Table 6 indicates that ODR has a positive and non-significant effect on pension generosity scores. With respect to real pension benefits per elderly the estimation coefficient is statistically significant and negative (Model 1 Table 7) which is consistent with the notion that with population aging the same amount of absolute pension spending is now distributed among a larger share of elderly. The estimated coefficient for the product of the ODR and the index of progressivity is positive and non-significant with respect to both generosity measures. The same applies to the Gini coefficient and the corresponding product term. However, in both instances the estimation coefficient for the product of the Gini-coefficient and the index of progressivity change their direction. This could be interpreted as weak support for the horizontal redistribution argument that in more progressive pension systems income inequality is indeed likely to increase pension generosity. Finally, Model 6 in Tables 6 and 7 test the effect of expected population aging on the generosity of public pensions. In both instances a “bad” projection decreases pension generosity. This effect is particularly strong if pension generosity is measured in terms of generosity scores and is overall consistent with the estimation results obtained for pension spending. It provides further support for the idea that projected population aging is likely to cause a reduction in both the generosity and size of pensions. With the refinement of measuring expected aging rather than the current share of elderly in the society, findings tilt in the direction of Razin & Sadka (2005) and the yet more general "self-limiting immoderation" thesis by Hicks & Zorn (2005).

In sum, tentative evidence on the majority voting model suggest that population aging increases pension spending, while its effect on real pension spending per elderly is negative, indicating a relative decline of pension expenditure per elderly in times of population aging. The effect of aging on pension generosity scores remains mostly unsettled; the statistically non-significant estimation coefficient is positive but the model has little explanatory power. Findings for horizontal redistribution are very weak. Estimates provide, if anything at all, further evidence on the "Robin Hood paradox". Relatively robust results are obtained with respect to the effect of expected population aging. Estimates are statistically significant with respect to pension spending and generosity scores.

Table 8. Non-linearity

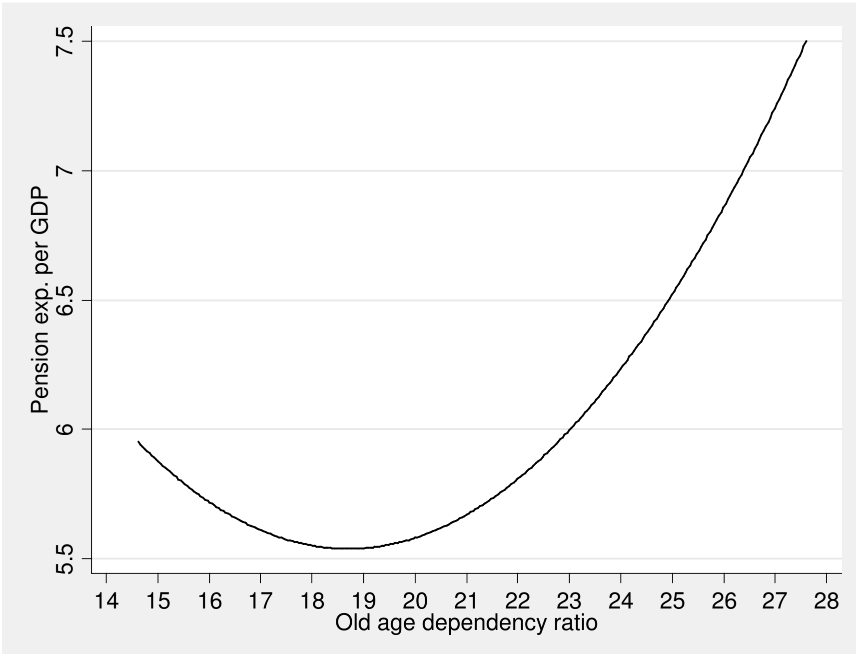
	Model 1 Pension exp. GDP Fixed	Model 2 Pension gen. Scores Fixed	Model 3 Real benefit per elderly Fixed
GDP per capita (log)	0.7	5.85	0.15
	[0.4]	[1.6]	[0.4]
GDP growth	-0.43***	-0.28	-0.072**
	[-3.3]	[-1.3]	[-2.7]
Interest rate	-0.17**	-0.0062	-0.038*
	[-2.5]	[-0.05]	[-1.7]
Inflation rate	0.14*	0.054	0.028
	[1.7]	[0.5]	[1.2]
ODR	-0.93*	-0.14	-0.078
	[-2.0]	[-0.1]	[-0.6]
ODR squared	0.025**	0.0079	0.00093
	[2.2]	[0.4]	[0.3]
1988-1995	0.17	-2.46	-0.34**
	[0.2]	[-1.4]	[-2.2]
1996-2003	-0.2	-4.66	-0.3
	[-0.2]	[-1.6]	[-1.2]
Observations	54	54	54
Number of id	18	18	18
R-squared	0.62	0.25	0.78

Note: N=18 countries, T= averages for 3 periods (1980-1987, 1988-1995, 1996-2003) t-statistics in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance.

4.4 Non-linearity, time trend and robustness analysis

According to Lindert (1996), the relationship between population aging and pension spending could be non-linear. Lindert (1996) identifies a hump-shaped relationship between pension expenditure and the proportion of elderly. Using three age ratios (persons 5-19 to persons 20-64, the share of adults 20-64 who are 39 or younger, and the ratio of population 65+ to population 20-64), he shows that pension spending increases with the share of elderly until the share of elderly per working age population reaches 30.5 percent. Statistical models presented in Table 8 reproduce the majority voting model (Model 1) presented in Tables 5-7 which now includes a quadratic term for the ODR, to allow for either acceleration or reversal of the aging effect. With respect to pension expenditure per GDP (Model 1) the estimated coefficients for ODR and ODR squared are statistically significant and increase the share of explained variation by a considerable 12 percent compared to Model 2 in Table 5. With respect to the generosity measures (Models 2-3 in Table 8), the quadratic term is non-significant in both instances; and neither does the non-linear specification increase the share of explained variance. Thus, the linear models seem to be better able to describe the relationship between population aging and pension generosity. However, with respect to pension spending the non-linear model suggest a convex relationship between pension expenditure per GDP and population aging. The quadratic function is at its lowest point if the ODR equals 18.5. This implies that raising the ODR from its minimum to 18.5 will cause a decrease in; public pension expenditure. After this point population aging will increase pension spending. This finding stands in contrast to Linder (1997) who finds a hump-shaped relationship between aging and spending. However, it is important to bear in mind that Lindert's data refers to a period of welfare expansion and population growth. Using very recent data, where population aging has gained momentum and focusing on within country variation, the analysis suggests a different picture. Figure 2 graphs the non-linear relationship between predicted pension expenditure per GDP and the ODR, when all other independent variables are set to their mean. The graph indicates, that with a low level of public pension spending and a relatively low ODR, population aging will decrease pension spending, while having a relatively high share of pension expenditure per GDP will experience increasing pension expenditure if the population ages. Bearing in mind that those countries with a relatively low share of pension expenditure per GDP are those countries with high scores on the index of generosity, the non-linear relationship between pension spending and population aging might be interpreted to provide further indicative evidence for the idea that the effect of aging on pension spending depends on the institutional features of the pension system. In contrast to Beveridgean pension systems, Bismarckian systems tend to have not only a higher share of spending per GDP they are also characterized by a closer link between contributions and entitlements. In times of population aging voters in a Bismarckian pension system are less likely to votes for smaller pensions as they benefit proportionally from the pension system (Conde-Ruiz & Profeta 2003). The opposite would apply to voters in Beveridgean systems. Moreover, even if a majority of voters in a Bismarckian system would prefer smaller pensions, a medium-term reduction of spending might be very difficult to achieve, since pension entitlements in these systems tend to be protected property rights that are relatively resistant toward retrospective discretionary retrenchment (Scharpf 2000).

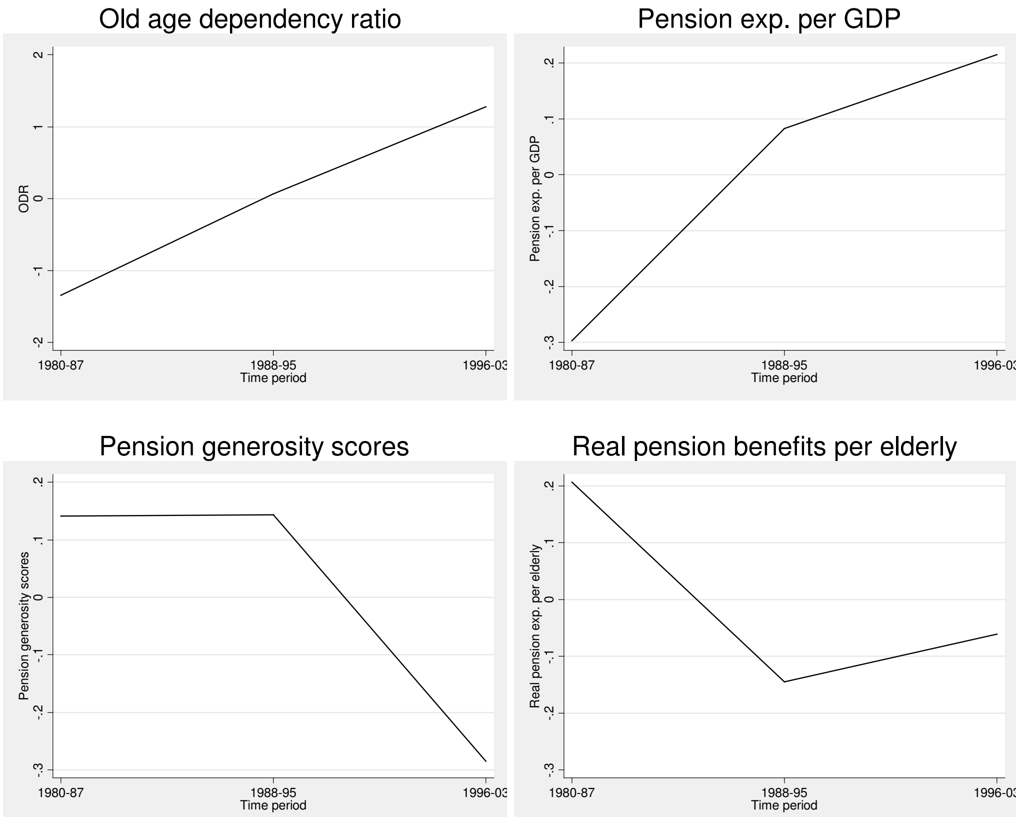
Figure 2. Predicted non-linear effect of population aging on pension spending



Note: Based on Model 1 in Table 8. Holding all other explanatory variables constant at their mean.

Concerning the general time trend in pension spending and generosity, the period dummies presented in Tables 5, 6 and 7 reveals a particularly interesting pattern; in each instance, the estimated coefficients for the second period (1988-1995) is larger than the estimated coefficients for the latest period (1996-2003). This finding is particularly strong with respect to pension expenditure per GDP and with pension generosity scores. It stands in contrast to Breyer & Craig (1996) who use a single continuous time trend variable and find a strong and positive effect of this measure on pension spending for the period ranging from 1960 to 1990. They conclude that ceteris paribus in each decade pension expenditure per GNP increase by 2 percent making it the "kudzu" of government spending (Breyer & Craig 1997: 721). The pattern of the period dummies suggests that the positive effect of population aging on pension spending has at least slowed down in the last period.

Figure 3. Temporal stability



Note: Expressed as deviation form the sample mean.

Table 9. Cross-sectional regressions

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9
	Pension exp. per GDP			Pension generosity scores			Pension benefits per elderly		
Period	1980-87	1988-95	1996-03	1980-87	1988-95	1996-03	1980-87	1988-95	1996-03
GDP per capita (log)	1.91	0.01	-1.74	3.86	3.18	-3.42	-0.67	2.27	3.43
	[0.7]	[0.00]	[-0.4]	[0.9]	[0.5]	[-0.8]	[-0.1]	[0.5]	[0.6]
GDP growth	-1.90***	-0.34	-0.59	0.31	-0.74	-0.4	1.63	-0.16	0.03
	[-3.9]	[-0.6]	[-1.6]	[0.2]	[-1.0]	[-1.3]	[1.3]	[-0.3]	[0.09]
Interest rate	-0.41*	0.00	0.32	0.58	1.86***	1.12*	0.05	-0.13	-0.57
	[-2.0]	[0.00]	[0.6]	[1.7]	[3.4]	[1.9]	[0.1]	[-0.2]	[-0.6]
Inflation rate	0.55***	-0.22	0.13	0.00	-2.29**	0.11	-0.36	-0.82	-1.52
	[3.2]	[-0.2]	[0.1]	[0.00]	[-2.6]	[0.09]	[-0.8]	[-0.9]	[-1.5]
ODR	0.07	0.34*	0.42**	0.14	0.15	0.00	0.15	0.20	0.04
	[0.7]	[1.8]	[2.4]	[0.4]	[0.6]	[0.01]	[0.7]	[0.9]	[0.2]
Observations	18	18	18	18	18	18	18	18	18
R-squared	0.57	0.22	0.48	0.29	0.61	0.25	0.21	0.19	0.43

Note: t-statistics in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance.

This pattern gives cause for a deeper investigation of the temporal stability of the relationship between public pensions and population aging. In order to do so, Table 9 presents cross-sectional regressions for each period with respect to pension spending and generosity. Results for the benchmark model turn out to be less robust than in the panel models although one has to take into account that the cross-sectional regression suffers from data restrictions. Focusing on the majority voting, Models 1 to 3 show that the ODR has an increasing positive effect on pensions spending. Once more, this is consistent with the assumption that an aging society has to devote more resources to the elderly. Concerning the generosity measures, the effect of the ODR is still positive but non-significant. More importantly, the magnitude of the coefficient decreases in both instances in the third period. Since this pattern is consistent among the measures for pension generosity, it suggests that population aging continues to increase the absolute size of pension expenditure but not the generosity of pension benefits.

In order illustrate this point in further detail Figure 3 presents the development of pension expenditure per GDP, pension generosity scores, real pension expenditure per elderly and ODR measured as deviations for the sample's period mean. The first graph shows that the ODR increased drastically over the whole observation period. The second graph indicates that pension expenditure per GDP follows the same path; although its growth slowed down in the last period. The two generosity measures develop in the opposite direction. Pension spending per elderly increases in the first period and turns negative in the second period. Pension generosity scores appear to be stickier than pension expenditure per elderly and decrease only in the last period. Taking both into account, Figure 3 and the cross-sectional regression analysis presented in Table 9, population aging seems to increase pension expenditure in overall terms while the relative generosity of pension benefits remains constant or starts decreasing.

5. CONCLUDING REMARKS

This study tests three voting models on old age security in aging societies using a recent dataset drawing on both the size of public pensions and benefit generosity. Since the mid 1990s, most OECD democracies have experienced accelerating growth in population aging. The empirical analysis suggests that a larger share of elderly voters increases the share of resources devoted to old age security, which is consistent with common economic wisdom that an elder society is likely to devote more resources to the elderly. However, there are two refinements, which can be made on the basis of this study. First, there is tentative evidence that the effect of population aging on pension spending depends on the design of the pension system. In less redistributive schemes, where contributions and entitlement are closely linked, population aging will translate into higher pension spending, which supports the “elderly power” hypothesis. In Beveridgean pension systems with flat benefit schemes, population aging is likely to lower pension spending, which supports the “fiscal leakage” hypothesis. This pattern is consistent with prior findings by Disney (2007: 551) although applying a more conservative estimation strategy causes the estimation coefficients to lose their statistical significance.

Second, the argument that “elderly power” will lead to more generous benefits cannot, thus far, be confirmed empirically. Yet, at the same time however, the findings of a negative relationship between real pension spending per elderly and the old age dependency ratio provides support for the less frequently proposed negative benefit effect or "fiscal leakage" hypothesis. Evidence on majority voting suggests that growing expenditure needs have been followed by growing overall spending commitments. At the same time, budgetary pressures have forced governments to cut smaller slices out of larger cakes, by reducing individual benefits generosity. This result also fits together with Oksanen’s (2003) projection for the EU-12 countries that between 2005-2050 pension expenditure will increase proportionally far less than the ODR because the ratio of average pensions to average wages is predicted to fall.

Empirical evidence for the horizontal redistribution model remains to be very weak. The negative sign of the inequality measure is difficult to interpret as it seems that income inequality reduces pension spending and benefits generosity. Taking into account the design of pension schemes – particularity the degree of redistributiveness – findings tilt in the direction of support for the horizontal redistribution thesis as equally indicated by Breyer & Craig (1997) and Disney (2007). However, none of these effects reaches statistical significance, what might suggests that issues of horizontal redistribution still play a role, but have been less influential than inter-generational issues.

Relatively strong and robust results are obtained for the effect of expected population aging on pension policies. In each instance "bad" population projections – which indicate upcoming fiscal pressure for unfunded pension schemes – decrease the size and generosity of public pensions. Population aging is not only a determinant for growing pension expenditure; but it also provides a rationale to vote for smaller pensions as suggested by Razin & Sadka (2005). Moreover, this result refines the "paradox of self-limiting immoderation" by Hicks & Zorn (2005), in saying than not the current share of elderly voters, but rather the expected threat to the fiscal sustainability of the pension system creates momentum for pension retrenchment policies. Thus, the perceptual frame within population aging takes place might be an underestimated factor in voting on public pensions.

On a more fundamental level, econometric and descriptive evidence alike suggest that the positive time trend in public pensions slowed down with respect to spending and even reversed with respect to pension generosity. Although it remains puzzling to explain the causes for these changes, it might be a first subtle indication that public pensions are starting to get a grip on issues of long-term fiscal sustainability. On the other hand, this study indicates a development in old age security provision that might be equally critical. If population aging continues to be associated with less generous pension benefits, avoiding old-age poverty may soon become a more salient issue on the social policy agenda. While there is no doubt that demographic pressures will exert yet more severe pressures on public pensions in the decades ahead, this study finds tentative evidence that concomitant budgetary pressures have thus far restrained governments from overcompensating pensioners. In this respect, it can be concluded that aging populations notwithstanding, mature OECD welfare states are not yet on the path to the politics of gerontocracy.

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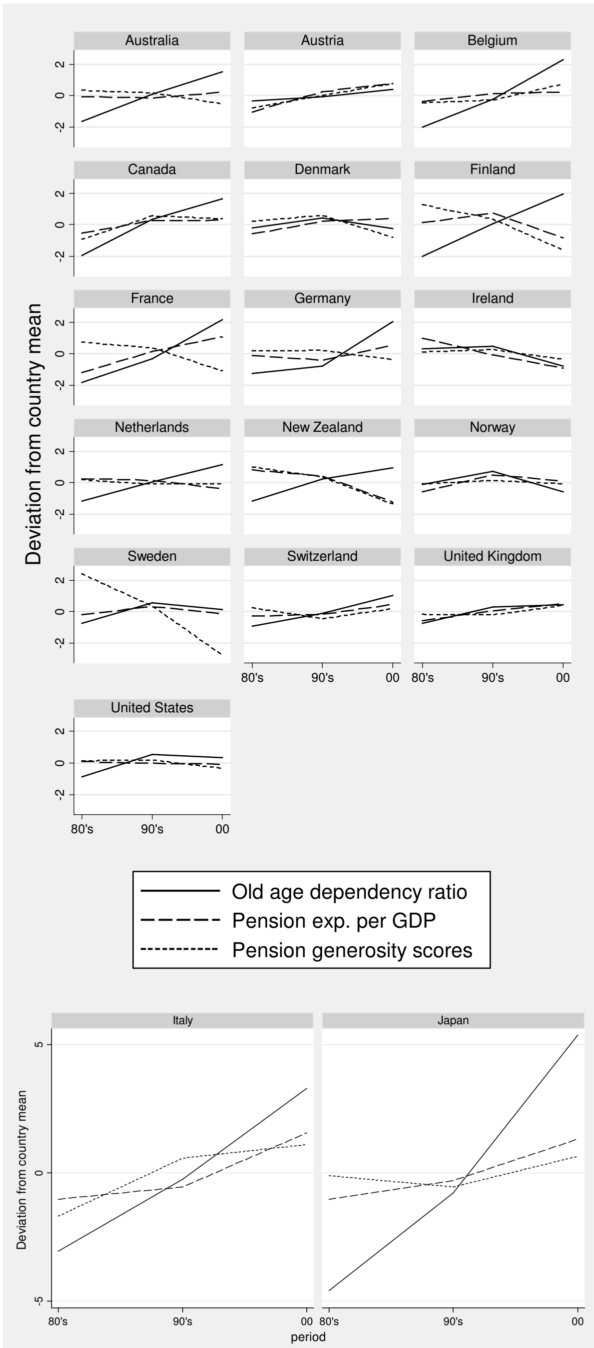
Appendix Table 1. Definition and source of variables

Variable	Definition	Source
Pension expenditure per GDP	Cash pension benefits per GDP	OECD Social Expenditure Database (2007)
Pension generosity scores	Standard pension replacement rate measured as the ratio of net public pension paid to a person earning the average productive worker wage in each year of their working career upon retirement in the year in question weighted by coverage.	Scruggs (2005) Welfare State Entitlements Dataset
Real pension benefits per elderly	Real cash pension expenditure in $ US at constant prices (2000) per working age population (15-64)	OECD Social Expenditure Database (2007), WDI
Log GDP per capita	Log of real Gross Domestic Product per capita	Penn World Tables (2007)
Real growth rate	Real Gross Domestic Product growth	World Bank (2007) World Development Indicators
Interest rate	Long-term interest rate	Armingeon et al. (2007) Comparative Political Dataset
Inflation rate	Consumer price index growth	World Bank (2007) World Development Indicators
ODR	Old age dependency ratio measured as the share of the elderly (65+) as a percentage of the working age population (15-64)	OECD (2007) Labor Force Statistics
Gini coefficient	Gini coefficient in percentage points as calculated by WIDER. Only quality 1 rated data employed.	WIDER (2005) World Income Inequality Database
ODR 20 year dif.	Projection for ODR in year t+20 minus ODR in year t	United Nations (2006) World Population Prospect

Appendix Table 2. Summary statistics

Variable	Obs.	Mean	Std. Dev.	Min	Max
Pension expenditure per GDP	54	6.20	2.43	2.30	12.11
Pension replacement rate	54	57.52	15.22	29.99	88.44
Pension benefits per elderly	54	10.68	2.60	8.01	19.32
GDP per capita (log)	54	9.77	0.33	8.89	10.39
GDP growth	54	2.56	1.20	0.65	8.74
Interest rate	54	8.46	3.25	1.87	15.57
Inflation rate	54	4.27	3.02	-0.02	13.31
Old age dependency ratio	54	21.01	3.31	14.62	27.60
Gini coefficient	50	29.79	4.58	20.83	39.96
ODR 20 years dif.	54	7.14	5.09	-2.16	21.69

Appendix Figure 3. Population aging in 18 OECD countries



Note: Expressed in deviation form the period mean.

What Makes Stabilization Reforms Happen?

Temporality in the political economy of welfare spending

Markus Tepe

Abstract

Theories of partisanship, electioneering and institutional rigidity have been shown to be too general and all encompassing to explain recent developments in welfare spending. This study uses panel data on 21 OECD countries (1980-2003) to explore how temporal contexts affect policy-makers' abilities and incentives to conduct stabilization reforms in social expenditure. Drawing on Alesina & Drazen's (1991) "war of attrition" framework, it examines two closely related questions: First, is the effect of political determinants in social expenditure growth conditional on the fiscal situation? Second, do political determinants influence the passage of time until stabilizations occur? The empirical analysis of the first question is performed by using an interactive model specification in a cross-sectional time-series setting. Although the conditional effect is not thoroughly convincing, it gives tentative evidence that patterns of partisanship and electioneering are more pronounced in periods of fiscal stress. Employing event history analysis to explore the second question indicates that leftist governments and institutional rigidity are likely to delay substantive stabilization reforms. In overall terms, empirical evidence tilts in the direction that the capacity of political determinants to explain welfare spending reemerges if issues of temporality are taken into consideration. Contrary to claims of depoliticization of welfare policies, this study indicates that politics still matters; however, it matters in more subtle ways than in previous decades.

Keywords: political business cycles, institutional rigidity, social expenditure

JEL Classification Numbers: H53, D72

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1. INTRODUCTION

Demographic change and international economic integration are considered as putting increasing pressure on the fiscal sustainability of welfare spending, while at the same time survey research finds that public support for the welfare state remains at constantly high levels (Boeri, Börsch-Supan & Tabellini 2001). Reforms aiming to contain social expenditure have therefore become a persistent source of political conflict in affluent democracies. Drawing on the "war of attrition" framework (Alesina & Drazen 1991), this study examines how temporal contexts affect policy-makers' abilities and incentives to conduct stabilization reforms in social expenditure. This set-up allows the organization of various empirical hypotheses presented in the political economy literature into a coherent framework (Alesina, Ardagna & Trebbi 2006: 2).

Partisanship, electioneering and institutional rigidity are among the most extensively discussed theories in comparative welfare state research. However, empirical evidence on their capacity to explain welfare spending is rather mixed and inconclusive.[1] Franzese (2002) and Franzese & Jusko (2006) try to make sense of existing inconsistencies by arguing that partisan and electoral cycles should always emerge, but crucially, the degree, character, and effectiveness of these cycles is structured by contextual variation. Franzese (2002) emphasizes the importance of the institutional and structural context. This study argues that the fiscal situation in

[1] For a review of the empirical literature see Table 2.

which welfare politics takes place may also constitute a contextual factor that is likely to influence the effectiveness of political manipulations in social expenditure.

Focusing on the temporal rather than the institutional context is also consistent with Pierson (2004), who criticizes contemporary welfare state research for completely ignoring the temporal dimension. He argues that disputes between competing political theories center around the question of which "variable" generates political outcomes, while the significance of such "variables" is likely to be distorted when they are removed from their temporal context. Pierson (2004) concludes that a proper understanding of public policies requires placing "politics in times". Empirical research should therefore try to identify the circumstances under which political processes evolve. Politico-economic theories can justifiably be criticized for being too general and insensitive towards the temporal context. Pierson's (2004) approach, however, runs the risk of providing detailed process descriptions without facilitating any generalizable arguments. Thus, the works of Franzese (2002), Franzese & Jusko (2006) and Pierson (2004) provide equally strong reasons to consider the temporal context when exploring the role of political determinants in welfare spending. However, with reference to Pierson (2004), this study is not about placing "politics in time" but rather it is an attempt to place temporality in politico-economic research.[2]

[2] The political fortune of the former German Bundeskanzler Gerhard Schröder may serve as an example for the importance of timing in public policies. In 2001, at the beginning of an economic recession, he announced a policy of the steady hand ("Politik der ruhigen Hand"). Schröder received

In economics and political science alike the term “reform” is used frequently and can mean very different things.[3] It is therefore necessary to begin with a definition of what is meant by stabilization reform. Alesina et al. (2006) distinguish two types of fiscal policy reforms: structural reforms and stabilization reforms. With respect to welfare politics, structural reforms predominantly concern the financing method of welfare programs which may include deregulation, privatization or other market oriented policy measures (Williamson 1994). The replacement of an unfunded pension system through private schemes would be an example of a structural reform. Stabilization reforms refer to politically intended measures to reduce a budget deficit or to stop imbalanced public spending. The quantitative assessment of the latter is, particularly challenging as it can take several years before a structural reform fully materializes (Hinrich & Kangas 2003). The reform definition utilized in this study is therefore based on public budget decisions and assumes that stabilization reforms can be measured as adjustments in social

harsh critique for this statement from leftist and rightist interest groups. Trade unions demanded a program for the stimulation of the economy, while the opposition simply declared that the government lacked of all competency in dealing with economic issues. Two years later, with the non-appearance of an economic rebound, the public pressure for political action was at its height. Eventually, in 2003, Schröder announced a comprehensive reform program, the AGENDA 2010, which implementation led to the most drastic restructuring of the German welfare state since its foundation. Later, these reforms have been made out to be responsible for losing the re-elections in 2005.

[3] The Oxford English Dictionary defines reform as the amendment, or altering for the better, of some faulty state of things.

expenditure. Although structural and stabilization reforms may take place simultaneously, the empirical analysis focuses exclusively on the latter.

Since stabilization reforms have been defined as adjustment in social expenditure rather than institutional changes, one can think of a second layer of stabilization reforms: short-term incremental adjustment vs. long-term substantial adjustment. Short-term incremental adjustment refers to year-to-year changes in welfare spending, assuming that political determinants affect the growth rate rather than the level of social expenditure, which is consistent with the literature on path dependent welfare systems (Esping-Anderson 1996, Pierson 1996). Alternatively, long-term substantial adjustments capture reductions in the level of social expenditure within a given period of time. The retrenchment literature, which has dominated comparative welfare state research for several years (see Starke 2006 for review), sometimes refers to reductions as being part of programmatic attempts to cut back the welfare state. However, in accordance with Green-Pedersen (2002: 8), "there is no such thing as retrenchment per se". Even though it might be possible that with a certain time lag programmatic retrenchment will lower levels of social expenditure – this is not the main concern here.

In order to explore the role of political determinants for the timing of stabilization reforms in social expenditure, it is clearly desirable to investigate the relevant time period. The end of the post-war economic boom in the early 80's has been considered as a "watershed for social policy" (Kittel & Obinger 2002: 5). Since then structural unemployment, population aging and international economic integration are among those forces that have been identified as marking the end of an expansionary phase of the welfare state (Pierson 1994, 1996, 2001). These forces are predicted to generate considerable fiscal pressure on social budgets and they are unlikely to diminish over the next decades. Initially, many European countries tried to respond to these challanges with early retirement policies which soon after turned out to amplify fiscal pressure (Korpi & Palme 2003, Ebbinghaus 2006). Pierson (2001) characterizes welfare polices in the post expansionary era as taking place under the condition of "permanent austerity". Rising dependency ratios, a limited capacity of the welfare state to increase revenues, and continuously high public preferences for maintaining current levels of welfare provision have created enduring political conflict over welfare spending. In this context international oragnizations, such as the IMF (2004), the OECD (1996) and the World Bank (1994) devoted major policy documents to address the issue of fiscal sustainability of maturing welfare states – with little success. The puzzling question is, why is it that stabilization reforms do not take place immediately, since it is apparent that stabilization measures will have to be adopted eventually.

A potential explanation can be found in Alesina & Drazen's (1991) model on delayed fiscal stabilization. If a stabilization reform has significant distributional implications – as is likely to be the case with welfare spending, the political groups that are necessary to agree on a reform possess equally strong veto powers, and ex-ante it is clear that one of them will have to pay a larger amount of the adjustment burden, so each group has an incentive to delay stabilization. The process leading to fiscal stabilization becomes a "war of attrition" which will not end until one group can impose its desired policy on its opponent who has exhausted the ability to resist an undesired reform. This framework is used to examine two closely related questions: First, is the effect of political determinants within the dynamics of social expenditure conditional on periods of fiscal stress? And second, do political determinants influence the passage of time until stabilization takes place?

The paper is organized as follows: The second section reviews the "war of attrition" model and links it with testable implications from the comparative welfare state literature – namely, partisanship, electioneering and institutional rigidity. The third section presents the empirical analysis which divides into two subsections: The first presents the interactive model specification used to analyze short-term incremental stabilization reform. The second employs event history analysis to investigate the influence of political determinants on the passage of time until long-term substantial adjustments happen. The last section summarizes the results, discusses its limitations and presents the concluding remarks.

2. TEMPORALITY IN POLITICO-ECONOMIC THEORY

Franzese (2002) points out that contemporary theories underemphasize variation in the contexts in which incumbents make policy. How variation in the temporal context interferes with the incentives and abilities to exploit mechanisms of partisanship, electioneering or institutional rigidity, is a pending theoretical question. The “war of attrition” model (Alesina & Drazen 1991) brings together temporality and political determinants by focusing on fiscal stabilizations. Alesina & Drazen (1991) consider an economy in which the government is running a positive deficit, implying growing government debt. Stabilization would require an increase in tax revenues or a cut in spending while further delay would accumulate the fiscal burden. Although there is agreement on the need for fiscal stabilization among competing political groups, conflict over how the burden of higher taxes or expenditure cuts should be allocated prevents immediate stabilization (Alesina & Drazen 1991: 1171). Ex ante the competitors know that one of them has to pay more than half of the costs of a stabilization reform. The dilemma is that none of them will accept to pay the higher fraction when both groups have a veto power. Each group knows how costly it is for them to delay a stabilization reform, but is uncertain about the waiting costs of the opponent. Competitors will choose to postpone stabilization as long as the marginal cost of waiting is lower than the marginal benefit of waiting (Alesina et al. 2006: 4). Thus, resolution is in general not immediate; the passage of time is needed to reveal which of the two groups is the weakest, i.e. has the highest costs of waiting. Eventually, the group with higher waiting costs will concede. The essence is that, although all might suffer while reforms are delayed, each group has an incentive to resist immediate stabilization in the hope that the opponent will capitulate first and agrees to bear a disproportionate burden of the reform (Martinelli & Escorza 2007: 1224). This logic of intentionally delayed stabilization can be applied to explore political conflict over welfare spending in times of permanent austerity.

Alesina & Drazen's (1991) original model does not consider changes in external circumstances, which may lead to changes in the capacity to play the "war of attrition". This idea is put forward by Alesina et al. (2006: 2), arguing that "the nature of political institutions influences the distribution of political power between competing social groups, and this is the connection between the model [war of attrition] and testable implication (…)." Exploring the interaction between political determinants and a country-year dummy for deficit respectively inflation crisis, they find that fiscal adjustments are more likely at the beginning of a government's term in office and in countries where political power is concentrated at the executive level. Partisanship turned out to have no significant effect on fiscal stabilization. Likewise Alesina et al.'s (2006), short-term incremental adjustments in social expenditure are studied within an interactive model specification; referring to the question whether the effect of political determinants changes in times of fiscal stress.

The "war of attrition" framework, however, provides a second argument on how temporality might matters in the political economy of reform that has not been addressed by Alesina et al.'s (2006) estimation approach. This argument is not about context, it refers to the passage of time until stabilization occurs. The passage of time is needed to reveal which of the opponents have higher costs of fighting the “war of attrition”. Thus, although nothing observable might have happened, stabilization occurs after one of the opponents realizes that the marginal costs of waiting are higher than the marginal benefits of waiting. Only then, the group with higher waiting costs concedes. The empirical question is whether political determinants lengthen or shorten the period until a substantial adjustment in welfare spending takes place. This question will be explored with the means of event history analysis.

At this point, it should be noticed that the empirical analysis does not provide a straightforward test of the Alesina & Drazen (1991) model; neither do Alesina et al. (2006). The “war of attrition” provides the conceptual framework to explore the role of political determinants for the timing of stabilization reforms in social spending. With precise knowledge of the waiting costs of the opponents it would be possible to fully describe the process of conflict resolution. However, this information is typically unknown, nor are real political actors willing to reveal their waiting costs. Thus, an explicit test of Alesina & Drazen’s (1991) model based on waiting costs would require an experimental setup; a task that will be left to future research. The advantage of the “war of attrition” framework is that it allows the organization of conflicting empirical hypotheses into a coherent framework (Alesina et al. 2006: 2). Since it would go beyond the scope of a single paper to address all theories that link politics and welfare spending, this study concentrates on three prominent theoretical strands: partisanship, electioneering and institutional rigidity.

2.1 Partisanship

Delayed stabilization reforms in social expenditure might be the result of ideological differences. Partisan budget cycles in the spirit of Hibbs (1977) and Tufte (1978) assume that politicians are not opportunistic but decide the course of public policy according to their political preferences. Once elected, the governing party translates the preferences of their constituencies into policies that benefit their supporters. Or as Tufte (1978: 104) puts it, "Party platforms and political ideology set priorities (…). The electorate, by choosing the governing political party, influences the choice of macroeconomic priorities (…)." The partisan theory predicts leftwing governments to pursue a Keynesian macroeconomic strategy to hold unemployment down and to increase the income of the worst-off, while rightwing governments primarily aim to run a balanced budget (Boix 1997, Franzese 2002). Concerning welfare spending, leftist governments are predicted to be more inclined to redistributive policies and therefore follow a more expansionary welfare policy than rightist governments.[4] Empirical support for the partisanship hypothesis however is

[4] Advocates of the rational-choice approach take a different view. They assume re-election seeking incumbents to be rational, utility-maximizing actors that primarily care about staying in office rather than representing the preferences of their constituencies. Downs (1957) seminal median voter model predicts that party competition in a two party system with one-dimensional preferences will end up in policy convergence since each party aims to meet the preferences of the median voter. This perspective of partisan competition is more consistent with current accounts on an end of political ideologies. However, as Alesina & Rosenthal (1995) show, even within a strict rational-choice

rather mixed. Cameron (1978), Stephens (1979), Castles (1982) and Schmidt (1982) all find that variation in welfare spending depends on the strength of leftist and rightist parties, while Wilensky (1981) and van Karsbergen (1985) point out that leftist and rightist parties are not different with respect to the overall level of welfare spending but with respect to the allocation of welfare spending. Subsequent studies by Swank (2001), Castles (2001) and Kittel & Obinger (2002) suggest that the statistical association between partisanship and social expenditure has broken down in the last decade.

The "New Politics of the Welfare State" debate initialized by Pierson (1996, 2001) spread further doubt about the role of partisanship in welfare spending. Pierson's (1996) study on welfare reform policies in the USA and Great Britain in the early 80's reveals a remarkable resilience of the welfare state toward retrenchment. The central argument is that as a result of policy feedbacks, welfare programs have generated new interest groups defined in terms of benefit recipients which prevent a dismantling of the welfare state. While welfare state expansion has been a game of claiming electoral credit, retrenchment becomes a game of avoiding electoral punishment. Therefore, conflict over retrenchment plays out less along lines of class, skill or partisan ideology – as it was the case in the era of welfare state expansions – but along those who benefit from social expenditure and those who finance welfare programs. In general, retrenchment policies are assumed to be highly

perspective it is possible that partisanship continues to affect macroeconomic outcomes in the short-run if prices and wages are sticky.

unpopular and therefore politically risky endeavors as they create concrete losses on relatively large groups and rather diffuse and uncertain benefits. Weaver (1986) hypothesizes that politicians adapt to this new situation with blame avoidance strategies. In order to avoid electoral backslashes politicians try to make cuts less transparent or to share responsibility for retrenchment measures among key actors. Assuming that the "New Politics of the Welfare State" (Pierson 2001) is based on the logic of blame avoidance, it can be argued that the relevance of partisanship in social expenditure is likely to fade away (Kittel & Obinger 2002). This question has become a hotly debated issue in comparative welfare state research and scholars agree that its resolution is ultimately empirical (Garrett 1998, Siegel 2002, Swank 2001, Allan & Scruggs 2004).

Referring to this debate, the present study tests if patterns of partisanship matter for the timing of stabilization reforms in social expenditure. In order to survive electoral competition in representative democracies, political parties must develop, adapt and maintain ideological representation (Franzese 2002: 373). In times of fiscal stress and austerity increasing issue salience may forces them to sharpen their ideological positions on social spending. This assumption is consistent with the directional voting literature (Rabinowitz & MacDonald 1989, MacDonald, Rabinowitz & Listhaug 1995, 1998, 2001), which predicts voters to have a dichotomous view in policy assessment. Instead of utilizing a continuum of policy positions to evaluate social policy, voters focus on the agreement or disagreement with a certain policy. Thus, the incumbents' incentive to apply ideological budget strategies in order to appeal to their constituencies may rise with issue salience. In consequence, there would be a polarization of policy positions in times of fiscal stress, leading to the reemergence of partisanship in welfare spending. Thus, leftist governments are predicted to increase welfare spending in times of growing budget imbalance, whereas rightist governments see the welfare state itself as an important reason for over-spending and cut expenditure in fiscally "bad" times. In this respect, leftist governments delay stabilization in social expenditure while rightist governments are predicted to stabilize immediately.

2.2 Electioneering

Electioneering provides an alternative explanation for delayed stabilization reforms in social expenditure. In contrast to partisanship, it assumes that a government's political ideology is unimportant for public spending. It predicts that budget decisions follow the electoral cycle, as the incumbent tries to buy votes trough public transfers. There are at least two perspectives on how elections effect welfare spending: First, the classical electoral cycle hypothesis put forward by Nordhaus (1975) and McRae (1977) predicts that governments implement the most popular budget measures immediately before an election and the most unpopular budget measures immediately after an election. Incumbents increase social expenditure at the end of their term in office and cut back social expenditure at the beginning of their legislative period (Blais & Nadeau 1992). Rogoff & Sibert (1988) and Rogoff (1990) come to a contrasting conclusion. They assume that governments exploit temporary information asymmetries to signal fiscal competence immediately before an election takes place. Thus, the incumbent is predicted to reduce public expenditure at the end of the legislative period in order to demonstrate to the electorate that it can provide a reasonable level of social security with scant use of public resources.

Both variants of the electioneering hypothesis face theoretical and empirical critique. Nannenstad & Paldam (1994) point out that systematic manipulations of budget decisions before elections require backward looking voters. If voters base their electoral decision strictly on future benefits there would be no room for electioneering. Moreover, particularly with respect to budget decisions, there is no guarantee that the effect of the manipulation will materialize at the desired point in time. Even if politicians conceive a strategy based upon electoral cycles, they may not succeed in implementing it due to the rigidity of the political process (Blais & Nadeau 1992: 390). Bearing in mind these restrictions it is not surprising that the empirical evidence on electioneering is inconclusive (Drazen 2000).

Taking into account temporal context, it is assumed that growing fiscal stress is not only associated with increasing popular demand for political action, it is also likely to increase the incumbents' expected risk of electoral punishment. Franzese (2002: 372) argues that the incumbents' incentive to exploit mechanism of electioneering will rise with the closeness of elections. If the incumbent assumes that stabilization reforms in social expenditure are politically unpopular, the incentive to delay stabilization in times of fiscal stress increases. Accordingly, electioneering in the spirit of Nordhaus (1975) and McRae (1977) should be more pronounced in times of fiscal stress. On the other hand, it is equally plausible that incumbents benefit from a social expenditure crisis. In times of fiscal stress and permanent austerity the willingness to accept unpopular cuts in social expenditure in order to overcome a crisis might increases. This would speak for electoral budget manipulations in the spirit of Rogoff & Sibert (1988). The incumbent stabilizes immediately before the election in order to claim electoral benefit for making a determined effort of welfare reform.

2.3 Institutional rigidity

Alesina et al. (2006) find that constitutional arrangements play an important role for immediate fiscal stabilization. Written constitutions are the manifestation of rules that determine the process and the actors involved in political conflict resolution. They thereby codify how governmental power can be exercised once a political party is in office. Particularly, federalism and bicameralism are considered to be among those features that determine the separation of political power. Riker (1975) defines federalism as a form of government in which the activities of government are divided between the regional and the central level in such a way that that each kind of government has some activities on which it makes the final decision. Bicameralism refers to a legislature that is composed of two-chambers, usually termed the lower and upper house. Federalism and bicameralism have been supposed to slow down the legislative process and to render abrupt policy changes (Riker 1992). Or as Lijphart (1999: 272) puts it “Federalism, second chambers, rigid constitutions, strong judicial review, and independent central banks can all be assumed to inhibit the decisiveness, speed and coherence of the central government’s policy making compared with unitary systems, unicameralism, flexible constitutions, weak judicial review, and weak central banks”. Referring to the “war of attrition”, federalism and bicameralism are considered to produce dispersed political power and multiple points of influence which will make it easier for the opposition to delay stabilization. In countries with relatively concentrated executive powers the group in office should be able to stabilize earlier (Alesina et al 2006: 5) as the political opponent faces relatively high waiting costs. Tsebelis (2000) makes a similar argument in which he associates a larger number of veto players – whose agreement would be necessary for a reform – with a high degree of power separation. Empirically, Huber, Ragin & Stepehens (1993) find that their constitutional index, building on various features of the state structure, has a strong negative influence on social expenditure. Subsequent studies by Bonoli (2001), Huber and Stepehen (2001a), Swank (2001) and Kittel & Obinger (2002) provide additional support for this hypothesis.

Other scholars point out that concentrated political power tends to be coupled with concentrated electoral accountability. The more pronounced the separation of powers, the easier it is for voters to associate political outcomes with political actors. With lower institutional constraint, voters can easier identify who is to blame for cutbacks in social expenditure. Or to put it the other way around, with a higher concentration of political power there is less room for "blame avoidance" through the politics of blame sharing (Pierson 1994, Pal & Weaver 2003). Higher institutional rigidity therefore might provide political actors with the possibility to "share" or to "pass" blame for social expenditure stabilizations between different levels of government (Starke 2006: 109) and thereby makes stabilization happen more immediately. Thus, the effect of concentrated governmental power has to be weighted against the risk of electoral punishment (Starke 2006: 109). For the empirical treatment the implications of institutional rigidity for the magnitude and timing of stabilization reforms are straightforward. Referring to Alesina et al. (2006) this study tests if institutional rigidity hinders immediate and rigorous stabilizations in social expenditure.

Table 1. Hypothesized effect on stabilization reform

Theory	Variable	Social expenditure growth in times of fiscal imbalance	Occurrence of cuts in the level of social expenditure
Partisanship	Leftist gov.	+	-
	Rightist gov.	-	+
Electioneering	Years left in term	-	+
Institutional rigidity	Institutional rigidity	+	-
Temporal perspective		Short-Term incremental	Long-Term substantial

3. EMPIRICAL ANALYSIS

As shown in the last section, the comparative welfare state literature provides conflicting hypotheses about the impact of partisanship, electioneering and institutional rigidity on welfare spending. Table 1 summarizes the hypothesized effects of the political variables for short-term incremental adjustments and for the timing of long-term substantial stabilizations in social expenditure. Before turning to the empirical analysis it is worth taking a look at pervious research. Table 2 presents an updated version of Kittel & Obinger's (2002: 12) review of empirical studies on the determinants of welfare spending published in the last 15 years. The overall findings suggest, that early studies drawing on the „golden era" of welfare expansion (around 1960-1980) find a significant influence of political determinants on welfare spending, while studies drawing on the "post golden era" (around 1980-2000) tend to provide weaker evidence for the relevance of partisanship and institutional rigidity. On the whole, the recent empirical literature tilts in the direction that politics matter less in explaining variation in social expenditure. Castles (2001: 210) concludes that political factors “are still reflected in levels of expenditure, but do not appear to have been nearly so influential in determining the trajectories of spending in recent years”.

This paper is less pessimistic about the impact of partisanship, electioneering and institutional rigidity for the conduct of welfare politics as it assumes that these mechanisms are not equally apparent at each moment in time. Indeed, the majority of empirical studies ignore issues of temporality and test for an unconditional linear relationship between political determinants and welfare spending (see Table 1). Scharpf (2000a) doubts that comparative quantitative welfare state research offers any adequate tools to cope with the complexity of the decision-making processes in welfare politics. Referring to this critique and to the theoretical arguments put forward by Franzese (2002) and Pierson (2004), the statistical analysis focuses on issues of temporality. The following sub-sections examine what econometric research might contribute to explore the role of partisanship, electioneering and institutional for the timing of stabilization reforms in social expenditure.

Table 2. Studies on social expenditure

Author(s)	N	Period (design)	Dependent variable	Impact of politics
Hicks & Swank (1992)	18	1960–1982 (pool)	Social expenditure/GDP	Left parties (+), centrist parties (+), electoral turnout (+), bureaucratic traditionalism (+), state centralization (+) and leftist corporatism
Hicks & Misra (1993)	18	1960–1982 (pool)	Social expenditure/GDP	Left parties (+), centrist parties (+), bureaucratic paternalism (+), strikes (+), voter turnout (+), left corporatism (+), state centralization (+), electoral competition (-)
Huber et al. (1993)		1960–1989 (pool)	Social expenditure/GDP	Left parties (+), Christian democratic parties (+), institutional veto points (-)
Schmidt (1997)	18	1960–1992 (pool)	Social expenditure/GDP	Left parties (+), Christian democratic parties (+), liberal parties (+), age of democracy (+), conservative parties (-),institutional veto points (-), single party government(-)
Garrett (1998)	14	1966–1990 (pool)	Income transfers	Interaction of left parties (+), powerful labor market institutions (+), high exposure to trade (+)
Hicks & Kenworthy (1998)	18	1960–1989 (pool)	Government transfers, decommodification, active labor market policy (ALMP)	Transfers, decommodification: neocorporatism (+) Active labor market policy: left (+) Social transfer payments: Christian democracy (+)
Castles (1999)	16-17	1960–1993 (cross-section)	Change in total social expenditure/GDP	Decentralization (-)
Kittel et al. (2000)	18	1960–1995 (pool, cross-section)	Social expenditure/GDP	Institutional veto points (-) [1960–1989], no consistent partisan effects
Wagschal (2000)	21	1980–1995 (cross-section)	Change in social expenditure/GDP	No partisan effects, fractionalization of party system (+)
Siegel (2001)	22	1980–1995 (pool)	Social expenditure/GDP	Left parties (+), single party government (-), veto points (Schmidt index) (-) and single party government (-)
Swank (2001)	15	1965–1995, 1979–1995 (pool)	Social expenditure/GDP	Social corporatism (+), inclusive electoral institutions (+), federalism (-), no partisan effects

Table 1. (cont.)

Author(s)	N	Period (design)	Dependent variable	Impact of politics
Garrett & Mitchell (2001)	18	1961–1991 (pool)	Social security transfers/GDP	No effect of different partisan portfolios when lagged dependent variable and a battery of country and time dummies are included
Castles (2001)	19	1984–1997 (cross-section)	Change in social expenditure/GDP	No partisan effects
Iversen (2001)	15	1961–1993 (pool)	Total government spending, government transfers and government consumption as % of GDP, unemployment replacement rate	Replacement rate, government consumption: left parties (+) Government consumption: voter turnout (+) Government consumption: concentration of union power (+)
Armingeon et al. (2001)	22	1960–1998 (repeated cross-section)	Social expenditure/GDP	Golden age (1960–1984): corporatism (+), consociational democracy (+), leftist and center parties (+/-), veto points (-) 1985–1998: leftist parties (-), consociational democracy (-), no effect of institutional constraints, center parties (+/-)
Huber & Stephens (2001a)	18	1958/61–1989/95 (pool, cross-section)	Social security benefits/GDP	Golden age (until 1985): leftist parties (+) and Christian democratic parties (+), institutional veto points (-) 1980s, early 1990s: no partisan effects
Kittel & Obinger (2003)	21	1982/97 (pool, cross-section)	Social expenditure/GDP	No consistent partisan effects, institutional rigidity (-), partisanship is conditional on institutional rigidity
Korpi & Palme (2003)	18	1976/95 (event history analysis)	Failure events defined via replacement rates	Initial benefit level (-), Unemployment (+), Veto Points (-), Trade openness (+), Leftist gov. (-), Rightist gov. (+)
Hicks & Zorn (2005)	18	1978/94 (event history analysis)	Failure events defined via real social expenditure per capita	Lagged social expenditure/GDP (+), Unemployment (+), Elderly ratio (+), Trade openness (-), GDP per capita (-), GDP growth (-), Leftist gov. (+), Rightist gov. (+/-)

Note: (+) = positive association (-) = negative association
Source: Kittel & Obinger (2002: 12), updated by the author

3.1 Variables and data

The sample includes 21 OECD countries covering a maximum time period ranging from 1980 to 2003. Although the sample covers only 24 years, which appears to be a relatively short period compared to the emergence of the welfare state as such, it is likely be a particularly relevant period for the study of stabilization reforms in social expenditure (Pierson 2001). Following prior empirical studies in comparative welfare state research, welfare spending is defined as total social expenditure per GDP (OECD 2007a). This variable provides an indicator for the overall size of public resources devoted to welfare programs. The independent variables of main interest capture the partisan complexion of the government, the electoral cycle and institutional rigidity. Data on the ideological composition of the cabinet are taken from Armingeon et al. (2007). Leftist government is defined as the share of cabinet seats belonging to social-democratic and other left wing parties measured in percentage of total cabinet posts, weighted by days in office. Rightist government is measured analogous. The number of years left in the incumbent's current term is used to account for electoral cycles. Prior studies separated phases of the electoral cycle using dummies for post, pre and election years (Golden & Poterba 1980, Alesina 1988, Blais & Nadeau 1992). However, with respect to the short-term conditional effect model it is advantageous to employ a continuous variable to measure the electoral cycle. The rationale behind this decision will become more apparent with the discussion of the estimation strategy for short-term stabilization reforms. Alesina et al. (2006) also use the number of years left in the incumbent's term to account for electoral cycles. The variable is taken from the World Bank's Database of Political Institutions, compiled by Beck, Clark, Groff, Keefer & Walsh (2001) and updated in 2004.[5] The operationalization of institutional rigidity follows the approach suggested by Kittel & Obinger (2002). Institutional rigidity is measured via an artificial index, which is based on Lijphart's (1999) index of bicameralism and federalism. In order to get an easy interpretable index the two indices are transformed to range from [-1,1].

Two sets of control variables account for economic and socio-economic factors which are expected to be important factors of welfare spending. The first set of control variables reflects internal causes of social expenditure growth; these are the real GDP growth rate, the unemployment rate and the old age dependency ratio (ODR), defined as the share of 65 plussers divided by the share of people aged 15 to 64. The second set of control variables concerns the dominant financing method of welfare programs and the impact on international economic integration. Scharpf (2000b) argues that tax-based social payments are likely to be easier targets of government discretion than contributory funded insurance systems. This is particularly evident in unfunded contribution financed pension system, where current worker's contributions enjoy the legal status of property rights. Usually, these protected welfare entitlements are relatively resistant against retrospective

[5] A limitation of this measure concerns the possibility of early elections, an exception that is not considered by this measure.

retrenchment policies. The variable measuring contribution-based financing of welfare programs is defined as the ratio of social security contributions relative to total tax revenue (OECD 2007b). Kittel & Obinger (2002) find that the level of this variable has a significant negative effect on the growth rate of social expenditure. This effect is rather difficult to explain since Scharpf's (2000b) argument on the resilience of insurance-based welfare states as well as the mere financing logic of these programs would suggest that a larger share of contribution financing should positively correlate with social spending.

International economic integration – frequently referred to as globalization – is measured in terms of trade openness. The dominant view on globalization assumes that it forces governments to reduce welfare spending. The majority of empirical studies tend to support this so-called "competitive" perspective (Krugman 1995, Greider 1997, Friedman 2000). Other scholars point out that international economic integration also increases the need to "compensate" citizens for the economic risk entailed by globalization (Garrett & Mitchell 2001, Korpi & Palme 2003, Swank 2002, Rodrik 1997, 1998). Thus, globalization would lead to more social expenditure instead of less. If the effect of globalization is measured in terms of trade openness, which is defined as the ratio of imports and exports per GDP, empirical findings are rather mixed. Garrett (1998), Hicks (1999), Swank (2002), and Ploemper, Troeger & Manow (2005) provide evidence in support for the "compensation" hypothesis, that is trade openness has a positive effect on welfare spending, whereas Garrett &

Mitchell (2001), Burgoon (2001), and Korpi & Palme (2003) findings support the "competitive" hypothesis of a negative effect on welfare spending. In sum, although the selection of control variables always contains a certain degree of arbitrariness, the identification of control variables is primarily driven by theoretical considerations and is consistent with prior empirical studies on social expenditure dynamics (see Table 2). Appendix Table 1 presents the definition and source of the variables.

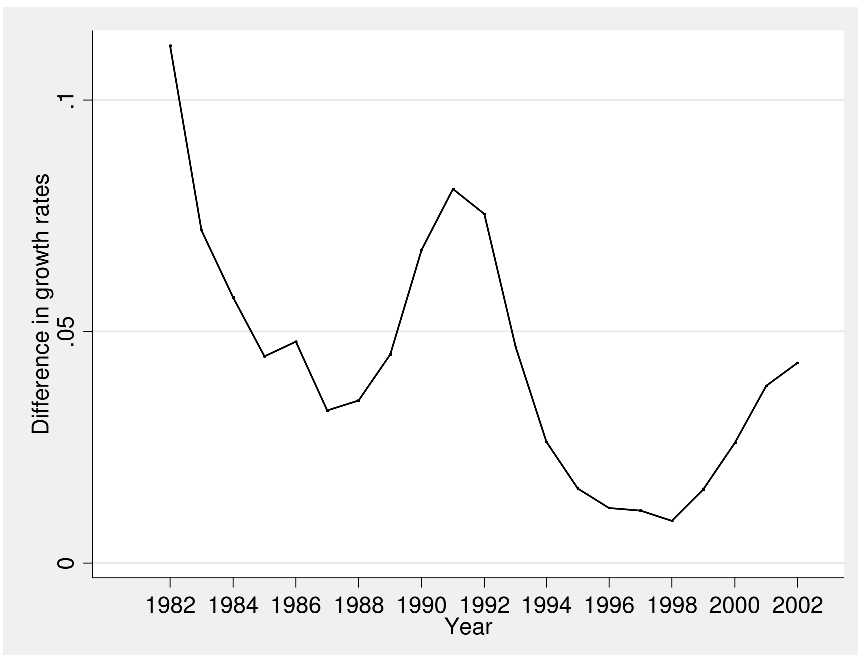
3.2 Short-term incremental stabilization

This section draws on Alesina et al. (2006) to explores if the effect of political determinants on social expenditure growth is conditional on the fiscal situation; thus, if the impact of political variables in fiscally "bad" times is different from their impact in normal times. Alesina et al. (2006) investigation of delayed deficit and inflation stabilization measures fiscal stress via a dummy variable that indicates country-years in which the deficit ratio or inflation rate exceeds a certain critical value; defined as the 75th percentile of the deficit per GDP ratio or the 75th percentile of the inflation rate.[6] The critical-value approach, however, is not applicable to social expenditure since there are persistent differences in the level of social spending across countries. For example, if the critical values for social

[6] The inflation crisis dummy accounts for country-years with an inflation rate equal or above 14.07 percent, and the debt crisis dummy for country-years with a deficit per GDP ratio equal or above 4.75 percent.

expenditure per GDP would be set at a level of 10 percent, countries such as Austria, Germany and France would be defined as being in a state of fiscal stress in almost each year. In order to obtain a country specific measure indicating fiscally "bad" times this study draws on the government budget constrain for unfunded welfare programs.

Figure 1. Measuring fiscal imbalance in pay-as-you-go social security systems



Note: Differences in growth rates are defined as the growth rate of social expenditure per head minus the growth of real GDP per worker. Positive values indicate imbalanced growth in social spending.

3.2.1 Measuring fiscally "bad" times

Virtually all affluent democracies belonging to the OECD finance a large amount of welfare spending via the pay-as-you-go method, by which current contributors pay the expenses for current recipients. Fiscal conservatism under the pay-as-you-go rule implies that any increase in expenditure has to be matched by an increase in contributions. The government's budget restriction is balanced if the growth rate of social expenditure per head equals the growth rate of GDP per worker. If the growth rate of social expenditure per head exceeds the growth rate of GDP per worker, welfare spending is, by definition, fiscally imbalanced. Figure 1 shows the difference between the two growth rates on the whole sample. There are two insights; first, imbalanced social expenditure growth is a cyclical phenomenon. The gap between social expenditure per head growth and real GDP per worker growth is larger in times of an economic recession. And secondly, during 1981 and 2003 unbalanced social expenditure growth has been the normal case in the 21 OECD countries under observation. This finding supports the "permanent austerity" condition and gives further reason to focus on this period for the analysis of long-term substantial stabilization reforms. Concerning the analysis of short-term incremental stabilization reforms, it is necessary to identifying which country-years in a period of imbalanced spending are likely to create issue salience and thereby increases popular demand for reform policies. In order to identify these country-years, the dummy variable only takes into account country-years of growing budget imbalance. It is calculated by applying the following formula:

$$\Delta\left(\Delta\ \frac{SOC_{it}}{POP_{it}} - \Delta\ \frac{GDP_{it}}{WOK_{it}}\right) > 0$$

where $\Delta(SOC_{it}/POP_{it})$ is the growth rate of social expenditure per head and $\Delta(GDP_{it}/WOK_{it})$ is the real GDP per worker growth rate. Periods of growing fiscal imbalance shorter than two years have not been taken into account, since it can be assumed that the period is too short to create relevant political pressure. The resulting dummy variable for 21 countries is shown in Table 3. In order to crosscheck this operationalization of fiscally "bad" times Appendix Table 1 presents the social expenditure "growth gap" for each country. As a preliminary test of the validity of this measure, Table 4 presents the pairwise correlation between imbalanced social expenditure and political pressure. The notion of issue salience assumes a positive relationship between fiscal stress and political pressure. The latter is measured via two alternative variables: government crisis and anti-government demonstration, both of which are taken from Banks (2005) cross-national time-series archive. Government crises is defined as any rapidly developing situation that threatens to bring the downfall of the present regime – excluding situations of revolt aimed at such overthrow. Anti-government demonstrations is defined as any peaceful public gathering of at least 100 people for the primary purpose of displaying or voicing their opposition to government policies or authority, excluding demonstrations of a distinctly anti-foreign nature (Banks 2005: Codebook). In both instances there is a positive and statistically significant correlation. This relationship provides indicative support for the assumption that the imbalance dummy can be associated with times in which governments are under political pressure. Thus, popular demand for political action appears to be more urgent in fiscally “bad” times than in normal times (Franzese 2002).

Table 3. Periods of growing fiscal imbalance (Dummy coding)

Country	Periods of growing fiscal imbalance
Australia	1981-82, 1988-90, 1997-00
Austria	1991-93, 1997-99
Belgium	1989-93, 1999-01
Canada	1990-92, 1994-97, 2000-02
Denmark	1986-88, 1990-93, 1997-99
Finland	1988-91, 1998-01
France	1983-85, 1989-93, 2000-03
Germany	1983-85, 1989-91, 1997-02
Greece	1988-90
Ireland	1990-92, 2000-02
Italy	1986-88, 1990-92, 1999-02
Japan	1985-87, 1991-95, 1997-02
Netherlands	1985-90, 1999-02
New Zealand	1986-89, 1994-97
Norway	1997-99
Portugal	1984-86, 1989-91, 1997-02
Spain	1986-90, 1997-02
Sweden	1986-92, 2000-03
Switzerland	1990-92, 1994-97, 2000-02
UK	1988-92, 1999-01
USA	1987-91, 1999-02

Note: Periods of growing imbalance in social expenditure growth shorter than 2 years have not been taken into account.

Table 4. Growing fiscal imbalance and political conflict

	Social expenditure per head growth minus real GDP per worker growth
Government crisis	0.14
	(0.00)
Anti government demonstrations	0.08
	(0.09)

Note: Pairwise correlation, p-values in parenthesis. Government crisis and anti government demonstrations are defined as in Banks (2005) Cross-National Time-Series Archive.

3.2.2 Statistical model

Estimating determinants of social expenditure in a cross-sectional time-series data set-up with N=21 and T= max. 24 is plagued with various methodological problems. Non-stationarity and autocorrelation are particularly important issues with this kind of political economy data. Stationarity implies that there is no systematic change in either the mean or the variance of a time series. If the dependent variable is non-stationary the OLS estimator produces invalid results, which Granger & Newbold (1974) called "spurious regressions". Results of the augmented Dickey-Fuller test presented in Appendix Table 2 confirms that social expenditure per GDP is a non-stationary variable. There are at least two approaches to deal with non-stationary data; using a dynamic specification in first-differences, or to explore the possibility of co-integrating relationships. Since time series with T=23 are too short for the estimation of reliable parameters in a cointegration framework (Maddala & Wu 1999) this study follows Kittel & Winner (2005) and switches to a model in first-difference. First-differences focus on systematic variation between annual changes in the variables – the short-term effects – while it removes level variation – the long run effect – from the data (Kittel & Winner 2005). This perspective is consistent with the idea of short-term incremental stabilization in social expenditure, where political determinants are expected to affect growth rather than the level of social expenditure.

The second specification issue concerns autocorrelation. Autocorrelation in the residuals is known to be a source for inefficient estimates. It can be regarded as nuisance in the residuals that has to be corrected; or autocorrelation can be seen as an indicator of persistency in the dependent variable. The first view suggests to model an AR(1) process in the residuals, which can be done with the Prais-Winston transformation. Concerning the political budgeting process this study assumes that the budget decision on social expenditure depends on the budget decision in the previous year. In line with Kittel & Winner (2005) the autoregressive process is therefore modeled by including a lagged dependent variable into the fixed effect specification in first-differences. If the time dimension of the panel is small, a fixed effect estimator or Least-Square Dummy Variable (LSDV) estimator including a lagged dependent variable generates biased estimates, the so-called Nickell-bias (Nickell 1981). In this case it is often suggested to use more consistent estimation procedures such as IV or GMM (Baltagi 2001: 131, Wawro 2000). However, Judsen & Owen's (1999: 13) simulation study has shown that for an unbalanced panel, as employed in this analysis and T=30, the fixed effects estimator performs as well or even better than these alternatives. Thus, this study proceeds using the dynamic LSDV estimator.

The core element of the short-term model specification is the interaction between political determinants and county-years of imbalanced social expenditure growth. The majority of studies in comparative welfare state research employ linear-additive models. These models assume that a dependent variable has a constant, unconditional relationship with a set of independent variables in the way that each unit increase in the independent variable causes a response in the dependent variable in the same way under any condition (Kam & Franzese 2007). The inclusion of interaction terms adds a second layer of complexity to the analysis, asking not simply whether some relationship exists between an independent and dependent variable, but under what conditions and what manner such relationship exists. Specifically, the interactive model tests if the effect of partisanship, electioneering and institutional rigidity is greater or lesser in times of imbalanced social expenditure growth.[7] In order to obtain easily interpretable coefficients and to minimize multicollinearity all independent variables are expressed as deviations from the mean (Friedrich 1982). Variants of the following model are estimated:

[7] The use of multiplicative interaction terms has been criticized for not being the appropriate way of addressing the presence of interaction among independent variables (Althauser 1971, Zedeck 1971). Their critique focused on two issues: interpreting estimation coefficients for the interacted variables and collinearly among independent variables caused by the multiplication of terms.

$$\Delta y_{it} = \beta_0 + \beta_1 \Delta y_{t-1} + \beta_2 \Delta X_{it} + \beta_3 POL_{it-1} + \beta_4 IB_{it-1} + \beta_5 (POL * IB)_{it-1} + \Delta \varepsilon_{it}$$

where Δy is the first-difference of social expenditure per GDP in country i at time t. ΔX is a vector of covariates, including GDP per capita, the unemployment rate, the old age dependency ratio (ODR), trade openness and the share of social insurance contribution as a percentage of tax revenues. Since GDP in period t is the denominator of the dependent variable, GDP per capita as an explanatory variable is entered with one-year lags. *POL* is the variable of main interest – accounting for partisanship, electioneering or institutional rigidity. *IB* is the temporal dummy indicating country-years of growing imbalanced social expenditure. The interaction term *POL*IB*, *IB* and time varying *POL* variables are lagged by one year. Using lags is based on a presumed lag between a change in the political determinants and its substantive effect. All political determinants are entered in levels since taking the first difference of these variables would remove the required information. First differencing of the ideological composition of the government for example would remove any information on the strength of the partisan power. Entering the political variables in levels is also consistent with prior research by Kittel & Obinger (2002) and Alesina et al. (2006).[8]

Table 5 presents the baseline model without interaction effects. Results for the Lagrange-multiplier test for first-order residual serial correlation (Baltagi 2002:

[8] The cross-sectional time-series analyses were performed with Stata 9 using the xtpcse procedure.

95) indicate that the specification including a lagged dependent variable is free of autocorrelation. Model 1 includes country and time fixed effects since it is an empirical question whether first differencing removes country and time fixed effects. The F statistic of the joint Wald test presented in the lower block of the table confirms that country effects are not statistically significant after taking first-differences. Accordingly, Model 2 is estimated without country effects while year effects remain significant. Furthermore, results of the modified Wald test for groupwise heteroskedasticity suggest to employ panel corrected standard errors (Beck & Katz 1995). In a correctly specified model, the standard errors of the OLS estimates should not deviate considerably from panel corrected standard errors. Comparing Model 2 and 3 in Table 5 shows that differences between standard errors are marginal. Thus, it can be assumed that a dynamic LSDV specification estimated with panel-corrected standard errors (Beck & Katz 1995) is robust enough to be used for the interaction analysis.

Table 5. Baseline short-term model (1982-2003)

	Model 1	Model 2	Model 3
		Δ Social exp.	
	FE(CT)	FE(T)	FE(T) PCSE
Δ Social exp. $_{t-1}$	0.10	0.13**	0.13*
	[0.07]	[0.06]	[0.07]
Δ GDP per capita $_{t-1}$	0.89	1.01	1.01
	[1.8]	[1.7]	[1.9]
Δ Unemployment	0.28***	0.27***	0.27***
	[0.07]	[0.07]	[0.05]
Δ ODR	-0.044	-0.002	-0.002
	[0.1]	[0.1]	[0.1]
Δ Trade openness	-0.039***	-0.041***	-0.041***
	[0.01]	[0.01]	[0.01]
Δ Social ins. as % of taxes	0.10***	0.10***	0.10***
	[0.03]	[0.03]	[0.03]
Observations	444	444	444
Adj. R-squared	0.31	0.32	0.32
Number of id	21	21	21
Wald (country effects)	1.09		
Wald (time effects)	3.12***		
Mod. Wald (GH), chi2(16)	474.77***		
LM(AR1), chi2(1)	0.012		

Note: FE(CT) = fixed country & time effects, FE(T) = time fixed effects, PCSE = Panel corrected standard errors, robust standard errors in brackets, Wald(time effects) = F statistic of joint Wald test for the inclusion of year dummies, Wald(country effects) = F statistic of joint Wald test for the inclusion of country dummies, Mod. Wald (GH) = Modified Wald test for groupwise heteroskedasticity (Greene 2000: 598), LM(AR1) = Lagrange-multiplier test for first-order residual serial correlation in panel data (Baltagi 2002: 95) *** p<0.01, ** p<0.05, * p<0.1, constant, country and time effects included but not reported.

Before examining the interaction terms in greater detail, findings for the control variables should be considered. The coefficient for the lagged dependent variable is stable and around 0.12, suggesting a persistent growth in social expenditure equaling about 12 percent per year. Assuming that the decision to increase welfare spending depends on the previous year's budget, the magnitude of the coefficient is considerable. Lagged change in GDP per capita also has a positive but statistically non-significant effect. Change in the unemployment rate has a strong positive and significant effect on social expenditure growth confirming prior research by Kittel & Obinger (2002), while change in the ODR has a statistically non-significant and slightly negative effect on social expenditure growth. Although this appears to be a rather unexpected finding at first, it can be linked to recent works that investigate if population aging can be associated with less welfare spending (Breyer & Stolte 2001, Razin et al. 2002, Hicks & Zorn 2005, Tepe 2008a). Concerning the level of welfare spending it might still be true that societies with a larger share of elderly devote more public resources toward the elderly. With respect to growth rates, however, the statistical association is no longer significant, which indicates that a growing share of elderly does not translate into an equally growing share of social expenditure.

Change in the share of social insurance as a percentage of taxes has a statistically significant and positive effect on social expenditure per GDP growth. The estimated coefficient indicates that an increase in social security contributions relative to tax revenues is associated with a 10 percent increase in social expenditure growth. Thus, countries with contribution-financed insurance systems do not only have a higher level of social expenditure, they also continue to increase welfare spending. In this context, Scharpf (2000b: 222) argues that insurance-based continental welfare states are vulnerable to a vicious cycle in which rising unemployment will create a need to raise social security contributions which causes an increase in non-wage labor costs and will thereby further reduce employment opportunities. Although this study does not provide a proper test of his argument, evidence leans in the direction that insurance-based countries, which are characterized by a large share of social insurance contributions per tax revenues, have difficulties maintaining the current level of social expenditure.

Trade openness has a statistically significant and negative effect on social expenditure growth of about 4 percent. This finding provides further support for the "competitive" hypothesis on the relationship between globalization and welfare spending, suggesting that market integration puts a downward pressure on the welfare state. Rodrik's (1998) idea of a growing governmental sector that protects against the risks of economic integration might be true with respect to other sectors of public spending, however, with respect to social expenditure it can thus far not be confirmed. In sum, although the estimation coefficient for the ODR shows an interesting deviation from the expected pattern, findings for the baseline model are generally consistent with prior research (see Table 2).

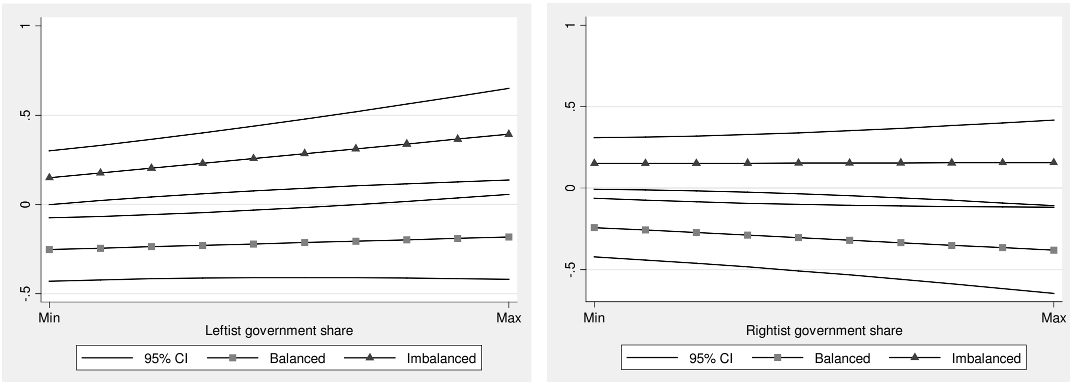
Table 6. Partisanship and short-term incremental stabilization

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7	Model 8	Model 9	Model 10
					Δ Social expenditure per GDP					
Δ Social exp. $_{t-1}$	0.13*	0.11*	0.11*	0.062	0.22**	0.13*	0.11*	0.11*	0.062	0.22**
	[0.07]	[0.07]	[0.07]	[0.1]	[0.09]	[0.07]	[0.07]	[0.07]	[0.1]	[0.10]
Δ GDP per capita $_{t-1}$	1.02	1.86	1.87	5.69	-1.35	1.05	1.88	1.89	5.62	-1.34
	[1.9]	[1.9]	[1.9]	[4.7]	[2.8]	[1.9]	[2.0]	[2.0]	[4.8]	[2.8]
Δ Unemployment	0.27***	0.26***	0.26***	0.37***	0.052	0.27***	0.26***	0.26***	0.36***	0.058
	[0.05]	[0.05]	[0.05]	[0.09]	[0.07]	[0.05]	[0.05]	[0.05]	[0.09]	[0.07]
Δ ODR	0.0058	0.012	0.0091	-0.28	0.16	0.0041	0.0088	0.012	-0.26	0.15
	[0.1]	[0.1]	[0.1]	[0.2]	[0.1]	[0.1]	[0.1]	[0.1]	[0.3]	[0.1]
Δ Trade openness	-0.040***	-0.034***	-0.034***	-0.034	-0.041***	-0.040***	-0.034***	-0.034***	-0.037	-0.038***
	[0.01]	[0.01]	[0.01]	[0.03]	[0.01]	[0.01]	[0.01]	[0.01]	[0.03]	[0.01]
Δ Soc. ins. as % of tax	0.10***	0.097***	0.095***	0.097*	0.12***	0.10***	0.097***	0.097***	0.100*	0.12***
	[0.03]	[0.03]	[0.03]	[0.06]	[0.04]	[0.03]	[0.03]	[0.03]	[0.06]	[0.04]
Leftist gov. $_{t-1}$	0.1	0.14*	0.071	0.12	-0.1					
	[0.08]	[0.08]	[0.10]	[0.2]	[0.1]					
Rightist gov. $_{t-1}$						-0.058	-0.073	-0.14	-0.19	0.066
						[0.09]	[0.09]	[0.1]	[0.2]	[0.2]
Imbalace $_{t-1}$		0.40***	0.40***	0.51***	0.22**		0.39***	0.39***	0.50***	0.21**
		[0.07]	[0.07]	[0.2]	[0.09]		[0.07]	[0.07]	[0.2]	[0.09]
Left Imbal. $_{t-1}$			0.17	0.12	0.40*					
			[0.1]	[0.2]	[0.2]					
Right Imbal. $_{t-1}$								0.14	0.18	-0.077
								[0.1]	[0.3]	[0.2]
Observations	444	444	444	234	210	444	444	444	234	210
Number of id	21	21	21	21	21	21	21	21	21	21
Period	1982-2003	1982-2003	1982-2003	1982-1993	1994-2003	1982-2003	1982-2003	1982-2003	1982-1993	1994-2003
Wald (Pol. & Pol.*Imbal.)			4.90*	1.49	3.34			1.88	1.09	0.18
Adj. R-squared	0.32	0.35	0.35	0.35	0.33	0.32	0.35	0.35	0.35	0.32

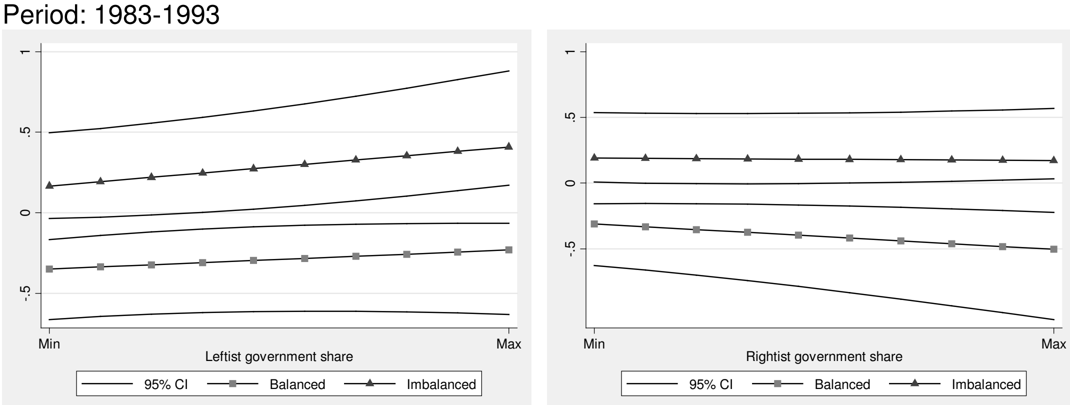
Note: FE(T) = Fixed time effects, panel corrected standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance, constant and time effects included but not reported, Wald (Pol. & Pol.*Imbal.) = F statistic of joint Wald test for the political variable and the interaction term.

Figure 2. Partisanship conditional on fiscal imbalance

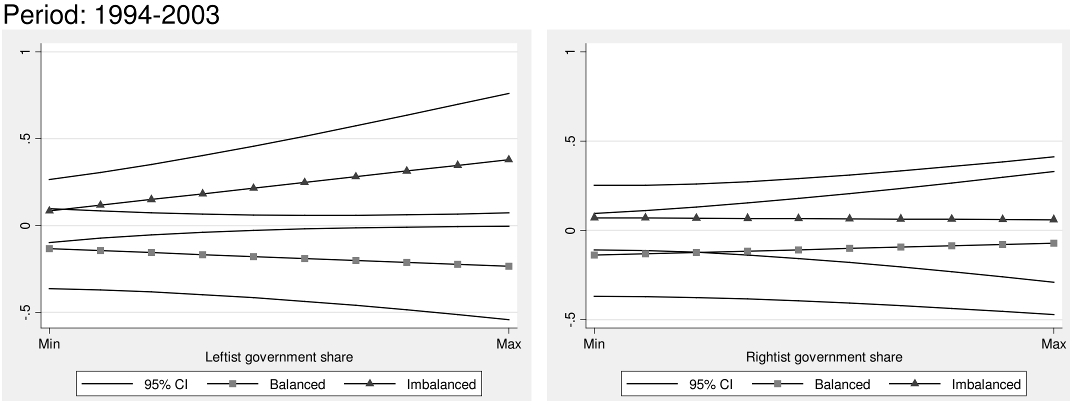
Period: 1983-2003



Period: 1983-1993



Period: 1994-2003



Note: The x-axis represents the political variable ranging form its minimum to its maximum, the y-axis represents predicted values for social expenditure growth, 95% CI represent the lower and the upper bound of the 95% confidence interval.

3.2.3 Empirical evidence and interpretation

Table 6 presents results of the interaction analysis exploring if partisanship matters for the timing of short-term incremental stabilization reforms in social expenditure growth. Model 1 and 6 represent the baseline specifications including the variable for the ideological composition of the government. If leftist government is included (Model 1), rightist und central governments are the omitted categories; if rightist government is included (Model 6) central and leftist governments represent the omitted categories. Although the direction of the estimation coefficients indicates that social expenditure growth increases with a larger share of leftist ministers in the cabinet and decreases with a rightist cabinet, none of the coefficients reaches statistical significance. Models 2 and 7 include the lagged dummy variable for country-years of growing fiscal imbalance in social expenditure. The estimated coefficient shows in the expected direction and is statistically significant at the highest margin.

Models 3 and 8 include the interaction between the ideological composition of the cabinet and country-years of growing budget imbalance. The estimation coefficient and its standard error presented in the regression table is inadequate to interpret the effect and the statistical significance of the interaction. Interactive terms require a more careful interpretation, since the effect depends upon two coefficients, and so too does the standard error and confidence interval. This study follows the suggestions by Kam & Franzese (2007: 73) and employs predicted effect plots to evaluate the conditional effect.[9] Figure 2 (first row) compares the predicted effect plots for the partisan composition of the cabinet based on Models 3 and 8.[10] The two tagged lines represents the predicted social expenditure growth if the ideological cabinet composition ranges from its minimum to its maximum in either times of balanced or imbalanced spending. The hourglass curves indicate the 95 percent confidence interval around the predicted values in times of imbalanced social expenditure growth. Both graphs indicate that there is no overlap in the 95 percent confidence interval for the predicted values of social expenditure growth in the presence and the absence of unbalanced expenditure when the leftist or, respectively, rightist cabinet share reaches from its minimum to its maximum.[11] However, differences in the slope of the predicted effect are rather small. In times of imbalanced social expenditure growth rightist governments seem to contain social expenditure while leftist governments continue to increase social expenditure. This pattern provides some indicative support for the partisanship approach (Hibbs 1977, Tufte 1978). It suggests that leftist governments follow a deficit spending strategy in times of fiscal stress, while rightist governments are more likely to hold on to fiscal

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[9] Brambor, Clark & Golder (2006) find that although conditional hypotheses have become popular in political science and multiplicative interaction models capture this intuition quite well, a survey of the top three political science journals from 1998 to 2002 suggests that the execution and presentation of these models is often flawed with inferential errors.

[10] The author is very grateful to Bernhard Kittel who provided a Stata ado file to compute conditional effect plots based on Kam & Franzese's (2002: 126) syntax.

[11] In order to enhance readability of the predicted effect plots, the confidence interval for the predicted effect in times of balanced social expenditure growth has not been included.

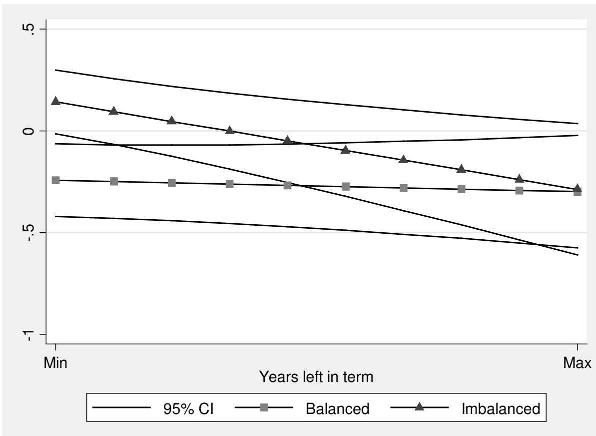
conservatism in times of fiscal stress. In order to test the robustness of the partisanship pattern and to explore if partisanship faded away in recent years, Model 4-5 and 9-10 re-estimate the interaction model for two sub periods, 1982-1993 and 1994-2003. The corresponding predicted effect plots are represented in Figure 2 (middle and last row). Although differences in the slope of the conditional effect still appear to be rather small, the partisan pattern is robust with regard to the two observation periods. In substantial terms, however, including the interaction between the ideological composition of the cabinet and country-years of growing budget imbalance does not increase the share of explained variance (adj. R-squared). Moreover, the Wald test for the inclusion of the political variable and the interaction term is statistically significant only with respect to the leftist government specification. In overall terms, partisanship contributes marginally in explaining the dynamics in social expenditure growth.

Table 7. Electioneering and short-term incremental stabilization

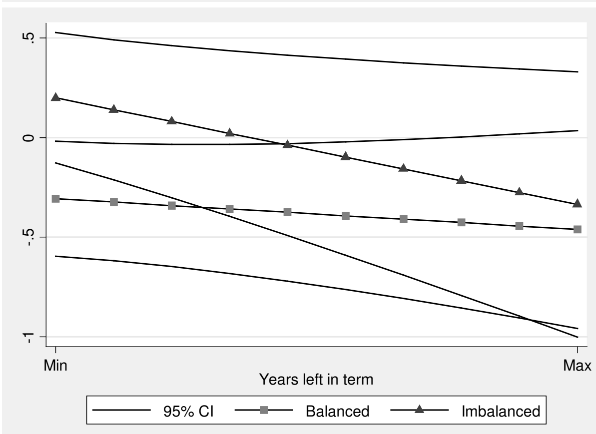
	Model 1	Model 2	Model 3	Model 4	Model 5
			Δ Social expenditure per GDP		
Δ Social exp. $_{t-1}$	0.13*	0.11	0.11	0.052	0.22**
	[0.07]	[0.07]	[0.07]	[0.1]	[0.09]
Δ GDP per capita $_{t-1}$	1.07	1.87	1.96	5.63	-1.37
	[1.9]	[2.0]	[1.9]	[4.6]	[2.8]
Δ Unemployment	0.27***	0.27***	0.26***	0.38***	0.048
	[0.05]	[0.05]	[0.05]	[0.09]	[0.07]
Δ ODR	-0.003	0.000	0.010	-0.23	0.15
	[0.1]	[0.1]	[0.1]	[0.2]	[0.1]
Δ Trade openness	-0.040***	-0.034***	-0.034***	-0.037	-0.036***
	[0.01]	[0.01]	[0.01]	[0.03]	[0.01]
Δ Social ins. as % of taxes	0.10***	0.095***	0.097***	0.096*	0.12***
	[0.03]	[0.03]	[0.03]	[0.06]	[0.04]
Years left in term $_{t-1}$	-0.050**	-0.047**	-0.014	-0.039	0.023
	[0.02]	[0.02]	[0.03]	[0.06]	[0.04]
Imbalance $_{t-1}$		0.39***	0.38***	0.51***	0.21**
		[0.07]	[0.07]	[0.2]	[0.09]
Years left in term*Imb. $_{t-1}$			-0.093*	-0.095	-0.1
			[0.05]	[0.09]	[0.07]
Observations	444	444	444	234	210
Number of id	21	21	21	21	21
Period	1982-2003	1982-2003	1982-2003	1982-1993	1994-2003
Wald (Pol. & Pol.*Imbal.)			7.66**	5.00*	2.28
Adj. R-squared	0.32	0.36	0.36	0.36	0.33

Note: FE(T) = Fixed time effects, panel corrected standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance, constant and time effects included but not reported, Wald (Pol. & Pol.*Imbal.) = F statistic of joint Wald test for the political variable and the interaction term..

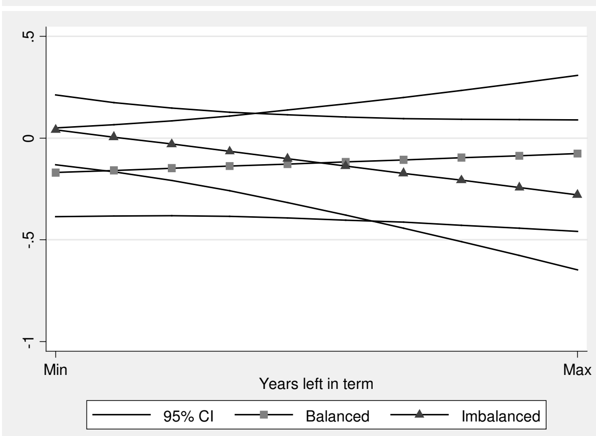
Figure 3. Electioneering conditional on fiscal imbalance



Period: 1983-2003



Period: 1983-1993



Period: 1994-2003

Note: The x-axis represents the political variable ranging form its minimum to its maximum, the y-axis represents predicted values for social expenditure growth, 95% CI represent the lower and the upper bound of the 95% confidence interval.

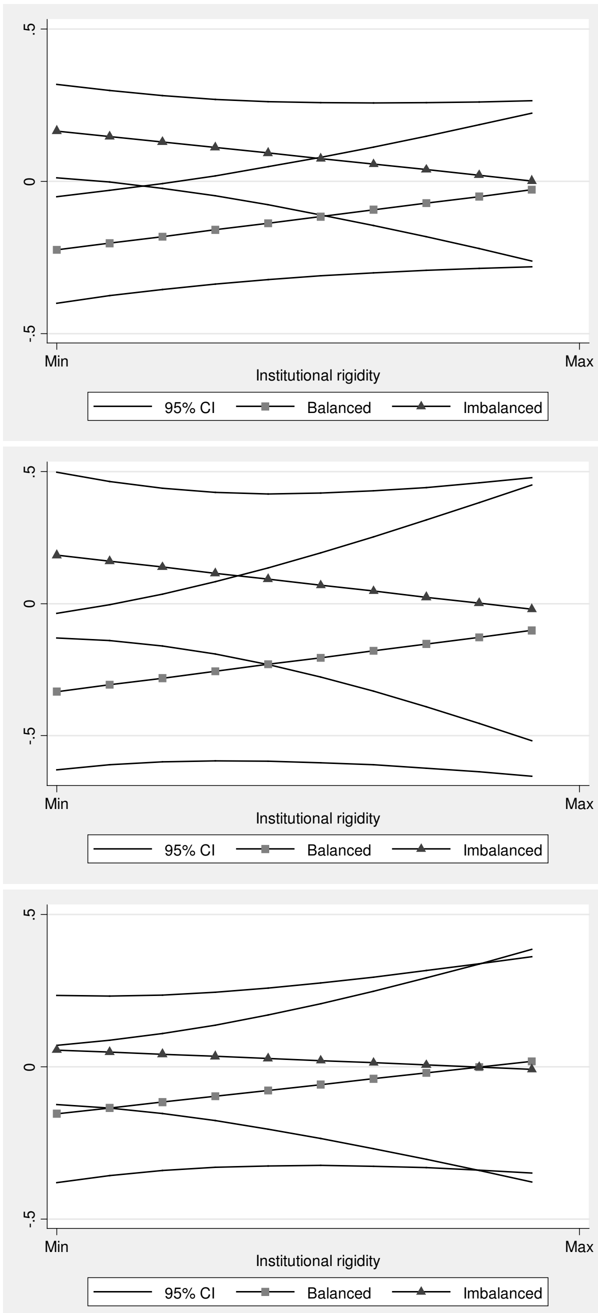
Table 7 presents estimation results with respect to electioneering. The order of entering the independent variables of interest into the short-term interactive model is the same as in Table 6. Model 1 indicates that more years left in term are statistically associated with significantly lower social expenditure growth rates, providing evidence on electioneering in the spirit of Nordhaus (1975) and McRae (1977). Governments indeed seem to fear electoral punishment as a consequence of reductions in welfare spending and adjust spending immediately after they get into office, hoping that voters have a rather steep forgetting curve. The predicted effect plots in Figure 3 suggest that this type of electioneering is even stronger in times of imbalanced budget growth. The graph indicates that there is a substantial overlap in the 95 percent confidence interval for the predicted values of social expenditure per GDP growth in the presence and the absence of unbalanced expenditure in the last years of the legislative period, while less overlapping occurs in the confidence interval at a higher number of years left in term. This suggests that the impact of the temporal context on social expenditure growth becomes more discernible as the number of years left in term increase. Testing the robustness of this patter, Models 4 and 5 show estimation results for the two sub-periods. Predicted effect plots presented in Figure 3 (middle and last row) indicate that electioneering has been less strong in the last decade. In general, findings tilt in the direction of reemerging electoral cycles if in fiscally "bad" times. The Wald test on the joint significance of years left in term and the interaction term supports this interpretation. However, figures for the adj. R-squared indicate that the inclusion of the interactive term does not improve the explanatory power of the statistical model.

Table 8. Institutional rigidity and short-term incremental stabilization

	Model 1	Model 2	Model 3	Model 4	Model 5
			Δ Social expenditure per GDP		
Δ Social exp. $_{t-1}$	0.13*	0.12*	0.11*	0.073	0.21**
	[0.07]	[0.07]	[0.07]	[0.1]	[0.10]
Δ GDP per capita $_{t-1}$	1.13	1.88	1.79	5.31	-1.25
	[1.9]	[2.0]	[1.9]	[4.7]	[2.7]
Δ Unemployment	0.27***	0.27***	0.26***	0.37***	0.054
	[0.05]	[0.05]	[0.05]	[0.09]	[0.07]
Δ ODR	-0.0075	-0.0019	0.024	-0.27	0.16
	[0.1]	[0.1]	[0.10]	[0.2]	[0.1]
Δ Trade openness	-0.041***	-0.035***	-0.034***	-0.039	-0.036***
	[0.01]	[0.01]	[0.01]	[0.03]	[0.01]
Δ Social ins. as % of taxes	0.10***	0.095***	0.094***	0.091	0.12***
	[0.03]	[0.03]	[0.03]	[0.06]	[0.04]
Institutional rigidity	0.05	0.027	0.11	0.13	0.095
	[0.06]	[0.06]	[0.07]	[0.1]	[0.1]
Imbalance $_{t-1}$		0.39***	0.39***	0.52***	0.21**
		[0.07]	[0.07]	[0.2]	[0.09]
Inst. rigidity*Imbal. $_{t-1}$			-0.20*	-0.24	-0.13
			[0.1]	[0.2]	[0.2]
Observations	444	444	444	234	210
Number of id	21	21	21	21	21
Period	1982-2003	1982-2003	1982-2003	1982-1993	1994-2003
Wald (Pol. & Pol.*Imbal.)			3.80	1.31	0.90
Adj. R-squared	0.32	0.35	0.35	0.35	0.33

Note: FE(T) = Fixed time effects, panel corrected standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance, constant and time effects included but not reported, Wald (Pol. & Pol.*Imbal.) = F statistic of joint Wald test for the political variable and the interaction term.

Figure 4. Institutional rigidity conditional on fiscal imbalance



Period: 1983-2003

Period: 1983-1993

Period: 1994-2003

Note: The x-axis represents the political variable ranging form its minimum to its maximum, the y-axis represents predicted values for social expenditure growth, 95% CI represent the lower and the upper bound of the 95% confidence interval.

Table 8 presents estimation results of the interactive model specification with respect to institutional rigidity. The coefficient for institutional rigidity is positive and statistically non-significant (Model 1), which stands in contrast to findings by Kittel & Obinger (2002) who use the same indicator, however, with a different estimation strategy and find a negative effect. The pairwise correlation between the level of social expenditure and institutional rigidity is indeed negative (-0.18***). However, exploring the effect of rigid institutions on social expenditure growth after controlling for macroeconomic determinants nullifies this statistical association. Model 3 indicates that the estimated coefficient of the interaction term changes its direction with respect to the fiscal situation. This is confirmed by Figure 4 which shows that there is already a substantial overlap in the 95 percent confidence interval for the predicted values of social expenditure per GDP growth in the presence and the absence of imbalanced social expenditure growth. However, less overlap occurs in the confidence interval when institutional rigidity is low, suggesting that the impact of periods of unbalanced growth on predicted social expenditure becomes more discernible as institutional rigidity decreases. Concerning the direction of the conditional effect, Figure 4 indicates that in times of growing fiscal stress higher institutional rigidity leads to reductions in social expenditure growth. This pattern is particularly strong in the first period and fades away in the last (Models 4 and 5). The overall impression is that in times of imbalanced spending, institutional rigidity decreases social expenditure growth, whereas in normal times higher institutional rigidity increases social expenditure growth. Substantively, the interaction should be evaluated on the basis of the Wald test and the variables contribution to explain variance in social expenditure growth. Both figures are presented in the lower block of the regression table and indicate that institutional rigidity contributes marginally in explaining social expenditure growth. Yet there is no unambiguous explanation for the changing effect of institutional rigidity for social expenditure growth. It could be possible that with higher institutional rigidity there are more veto players involved in the political process, which opens up the possibility to “share blame” for adjustments in social expenditure growth. However, this apparently contradicts Alesina et al.’s (2006) findings that governments with strong executive power adjust more rigorously in times of deficit or inflation crisis.

Findings from the analysis of short-term incremental stabilization reforms can be summarized as follows. Social expenditure growth is mainly driven by macroeconomic determinants. It is positively affected by the last year’s budget, growing unemployment and a growing share of social contribution as a ratio of tax revenue, while growing market integration decreases social expenditure growth. Political determinants contribute little in explaining social expenditure growth and so does the interaction term. However, concerning partisanship and electioneering the predicted effect plots suggest that politically motivated manipulations in social expenditure are at least more pronounced in times of fiscal stress. Findings on institutional rigidity are rather difficult to explain. If at all, evidence from the predicted effect plot on institutional rigidity support Pal & Weaver (2003), suggesting that governments opt for blame sharing in times of fiscal stress.

3.3 Long-term substantial stabilization

The last section explored the role of political determinants for short-term incremental stabilizations in social expenditure and ignored the possibility of long-term adjustments in the level of aggregate welfare spending. This section explores the second argument derived from the "war of attrition" framework. Alesina et al. (2006: 5) assume that the passage of time is needed to reveal which of the opponents have the higher costs of waiting. Thus, although nothing observable might have happened, eventually, the group with higher waiting costs concedes. This idea will be tested with the means of event history analysis. The question is whether political determinants lengthen or shorten the passage of time until a stabilization occurs. Reductions in social expenditure are therefore analyzed as adjustment events that are traceable to political determinants. Thus, in contrast to the short-term analysis which used a continuous measure for social expenditure growth, event history analysis measures stabilization reforms in terms of discrete policy events. To the knowledge of the author, thus far only a few studies have taken up this approach in comparative welfare state research. In this respect Korpi & Palme (2003) and Hicks & Zorn (2005) are of particular interest.

Korpi & Palme (2003) investigate cuts in sickness, work accident and unemployment insurance in 18 OECD countries from 1976 to 1995. Defining retrenchment events as 10 percent reductions in replacement rates, they find 19 of such events. Korpi & Palme (2003) employ intensity regressions for repeated events and show that unemployment and trade openness have a positive and statistically significant effect, indicating that the risk for cuts in the replacement rate are higher with higher unemployment and in more open economies. With respect to partisanship, their findings suggest that cuts are statistically less likely with leftist governments. The initial benefit level remains to have a negative and insignificant effect. There are two methodological issues concerning their study. First, Korpi & Palme's (2003) results cannot be reproduced, since they are using data from the Social Citizenship Indicator Program (SCIP) which is not available to the scientific community. More importantly, the use of intensity regressions, which do not include covariates, provides a preliminary indication of the effect of political determinants for the timing of cuts in the replacement rate.

Hicks & Zorn's (2005) study goes a step further; they conceptualize retrenchment events as cuts in real social expenditure per capita and analyze these events via a Cox hazard model for repeated events including various covariates. Using a dataset covering 18 OECD countries from 1978 to 1994 they find no partisan effect. In contrast to Korpi & Palme (2003), trade openness has a negative effect on the risk of cuts. Moreover, their estimation results suggest that unemployment and a larger share of elderly has a significant positive effect on the likelihood of retrenchment events. They interpret this pattern as evidence for the "paradox of self-limiting immoderation" (Hicks & Zorn 2005: 632). The self-limiting immoderation hypothesis predicts that factors which promote welfare spending simultaneously build up pressures for fiscal adjustment. Thus, population aging and growing unemployment might switch over to pressures for reductions in welfare spending. Unfortunately, Hicks & Zorn were unable to reproduce their published findings and eventually had to announce errata (Hicks & Zorn 2007) to their 2005 article. Using a new dataset nullified their findings. New estimations do not indicate that pressures from unemployment, population aging and globalization stressed in Hicks & Zorn (2005) drive retrenchment.

3.3.1 Measuring stabilization events

This study aims to measure substantial stabilization events on relatively objective grounds aided by data on social expenditure. Korpi & Palme (2003: 432) argue that social expenditure per GDP is a misleading indicator for the study of retrenchment policies. As pointed out earlier, this study regards social expenditure per GDP as a measure for the share of public resources devoted to welfare programs not as an indicator of welfare effort. In this respect social expenditure per GDP is the appropriate measure for the study of delayed fiscal adjustment in welfare spending.[12] Following Hicks & Zorn (2005: 642), substantial stabilization reforms in the level of social expenditure are defined as discrete policy events coded according to the following formula:

$$\frac{SOCEX_{it+k} - SOCEX_{it}}{SOCEX_{it}} < c$$

where *SOCEX* is social expenditure per GDP for country *i* at time *t*, *k* is the parameter for the length of the stabilization period and *c* is the cutoff at which a stabilization event has occurred. Substantial stabilization reforms are defined as reversals in the level of social expenditure (by the amount of *c*) within a certain adjustment period (by the length of *k*). Since the theoretical literature does not suggest any concrete values for the amount of reversal or for the corresponding time period, values for *k* and *c* are chosen with respect to prior empirical research (Hicks & Zorn 2005, Korpi & Palme 2003). In order to obtain a robust indicator for major stabilization events *k* ranges from 3 to 4 years and *c* ranges from -0.05 to -0.10. For example, if *k*= 3 and *c*= -0.10 the stabilization event dummy captures a 10 percent reduction in social expenditure per GDP within a three-year period.

[12] There is no doubt that replacement rates, as used by Palme & Kopri (2003) or Allan & Scruggs (2004), are more appropriate for the study of programmatic retrenchment.

Table 9 presents four alternative measures for stabilization events. In contrast to Hicks & Zorn (2005), who let *c* vary slightly (-0.04, -0.05, -0.06), this study focuses on 10 percent and 5 percent cuts. The issue with slightly varying values for the parameter *c* is that the number of events remains almost unchanged. Using more extreme parameter values for *c* does affect the number of events; ranging from 28 events with c=-0.05 and k=3 to 12 events with c=-0.10 and k=3. Changes to the adjustment period *k* instead have a marginal effect on the number of events. Thus, following Hicks & Zorn (2005) *k* is set to 3 years. Although even focusing on 10 percent reductions remains somehow arbitrary, the approach is also consistent with Korpi & Palme (2003) who focus on 10 percent reductions in replacement rates. Moreover, comparing the resulting dummy for stabilization events in social expenditure per GDP with Korpi & Palme's (2003: 444) and Hicks & Zorn's (2005: 644) indicator of welfare retrenchment events shows, that although the measures are not identical, there is a considerable amount of overlapping. Finally, to cross-validate the stabilization event dummy, Table 10 presents the pairwise correlation between stabilization event dummies (with c=-0.05 and c=-0.10) and the dummy for country-years of growing budget imbalance, which has been used in the analysis of short-term incremental stabilizations. The correlation analysis confirms the expected negative relationship, however, only if c=-0.10 the negative correlation coefficient reaches statistical significance. Although the pairwise correlation analysis can have indicative character at best, it provides additional support to set c=-0.10.

Nevertheless, in order to explore the robustness of statistical findings, results for both event measures (c= -0.10 and c= -0.05) will be presented in the regression tables.

Table 9. Measuring stabilization events

Country	k=3, c= -0.05	k=4, c= -0.05	k=3, c= -0.10	k=4, c= -0.10
Australia	1985	-	-	-
Austria	-	-	-	-
Belgium	1986	1985, 1993, 1996	1987	1986
Canada	1992, 1996	1991	1993	1992
Denmark	1982, 1994, 1997	1982, 1993	1983	
Finland	1992	1991	1992, 1994	1992
France	-	-	-	-
Germany	1987	-	-	-
Greece	1985, 1987, 1989	1986	-	1987
Ireland	1985, 1993	1985, 1992	1986, 1994	1985, 1993
Italy	1987	-	-	-
Japan	-	-	-	-
Netherlands	1982, 1992	1982, 1991	1983, 1993	1982, 1992
New Zealand	1989, 1998, 2000	1990, 1998	1991,	1990
Norway	1993,1998	1992	-	-
Portugal	1981	1981	-	-
Spain	1993	1993	-	1993
Sweden	1992	1992	1993	1993, 1996
Switzerland	-	-	-	-
UK	1985, 1994	1984, 1993	1986	1985
USA	1982	1982, 1995	-	-

Table 10. Growing fiscal imbalance and stabilization events

	Country-years of growing fiscal imbalance
Stabilization event (k=3, c=-0.05)	-0.07
	(0.12)
Stabilization event (k=3, c=-0.10)	-0.13
	(0.00)

Note: Pairwise correlation, p-values in parenthesis.

3.3.2 Statistical model

With growing interest in path dependence and easy access to longitudinal data, event history analysis has recently become more popular in political science. Event history analysis considers not only if something happens but also when something happens (Box-Steffensmeier & Jones 2004). In this respect, the dependent variable measures the duration of time that a country spends in a state of imbalanced welfare spending before experiencing a substantial stabilization reform. A Cox hazard model for repeated events is used to investigate the influence of political determinants on the passage of time until a stabilization event happens. The Cox model takes the following form:

$$h(t) = h_0(t)exp(\beta X_{it})$$

where X_{it} is a vector of time varying and time invariant covariates and $h_0(t)$ is an unspecified baseline hazard on which the independent variables act in a multiplicative fashion. The advantage of this specification is that it does not assume an underlying form for the baseline hazard. A feature which made the Cox model particularly popular in applied research (Hicks & Zorn 2005: 647). Moreover, the Cox hazard model can account for the fact that several nations experience more than one stabilization event within the observation period. Applying the conditional risk-set approach by Prentice, Williamson & Peterson (1981), these multiple failures can be estimated in elapsed or interevent time. The elapsed time specification uses the time from each unit entry, while the interevent time refers to the duration since the previous event. The elapsed time specification has a "carry over effect" in which the total time of the second risk interval includes the first interval, and the third risk interval contains the first and so on (Kelly & Lim 2000: 28). Following the suggestion by Box-Steffenmeister & Zorn (2002) the interevent duration time is modeled through the use of the stratified variant of Cox's model for the analysis of repeated events.[13] Box-Steffenmeister & Zorn (2002: 1079) claim that this approach "will offer the best combination of characteristics for addressing questions of interest to political scientist." Since Hicks & Zorn (2005) apply a similar estimation approach it should be less difficult to compare estimation results.

In contrast to Kopri & Palme (2003) the analysis includes a full set of covariates; these are virtually the same as in the interactive model specification used in the last section. To be consistent with prior research and to avoid simultaneity

[13] All event history analyses were performed with Stata 9 using the stcox procedure.

bias, the independent variables are lagged according to Hicks & Zorn (2005: 645). GDP per capita, unemployment and the old age dependency ratio and the political variables are assumed to have an immediate effect. One-year lags are used for social expenditure per GDP, GDP growth, trade openness and social insurance contributions as a percentage of tax revenues.

Table 11. Event history analysis of stabilization events (k=3; c= -0.10)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
Social exp. $_{t-1}$	0.48**	0.44*	0.48**	0.52*	0.40*		0.38
	[0.2]	[0.2]	[0.2]	[0.3]	[0.2]		[0.3]
GDP per capita log	-0.17	-1.3	-1.6	-0.89	3.03*		3.62*
	[1.3]	[1.6]	[1.4]	[1.5]	[1.7]		[1.9]
GDP growth $_{t-1}$	0.068	0.18	0.12	0.14	0.081		0.27
	[0.2]	[0.2]	[0.3]	[0.3]	[0.3]		[0.6]
Unemployment rate	0.21**	0.29**	0.26***	0.23**	0.29**	0.45**	0.46***
	[0.08]	[0.1]	[0.09]	[0.1]	[0.1]	[0.2]	[0.2]
ODR	-0.46**	-0.39	-0.43	-0.49*	-0.67*	-0.088	-0.84**
	[0.2]	[0.3]	[0.3]	[0.3]	[0.4]	[0.2]	[0.4]
Trade openness $_{t-1}$	0.016	0.007	0.015	0.019	0.028*	0.025***	0.043***
	[0.01]	[0.01]	[0.01]	[0.01]	[0.02]	[0.008]	[0.01]
Social ins. as % of taxes $_{t-1}$	-0.057**	-0.036*	-0.031	-0.071**	-0.023	-0.011	0.035*
	[0.02]	[0.02]	[0.03]	[0.03]	[0.02]	[0.06]	[0.02]
Left		-0.032***				-0.053**	-0.012
		[0.007]				[0.02]	[0.03]
Right			0.025***			0.012	0.042
			[0.005]			[0.01]	[0.03]
Years left in term				0.51		0.31	0.2
				[0.3]		[0.2]	[0.3]
Institutional constraints					-2.02**	-1.29***	-3.99***
					[1.0]	[0.5]	[1.1]
Observations	447	447	447	447	447	449	447
Number of id	21	21	21	21	21	21	21
Wald chi2	23.9***	32.5***	183***	69.7***	60.4***	31.3***	1444***
Wald (Pol. Vars.)						21.75***	33.71***
Log Likelihood	-14.2	-12.9	-13.0	-13.4	-12.8	-14.7	-9.79

Note: Failures = 12, cell entries are coefficient estimates, robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance, Wald (Pol. Vars.) = Wald test for the inclusion of political variables.

3.3.3 Empirical evidence and interpretation

Results of the Cox repeated event hazard analysis for substantial stabilization events defined by k= 3 and c= -0.10 are presented in Table 11. The regression table reports robust standard error estimates, grouped by country. For an easier interpretation the coefficients represent log hazard ratios, whereas a positive coefficient indicates an increase in the hazard (risk) of a stabilization event. Model 1 is the baseline model; it does not include the political variables of main interest. The lagged level of social expenditure has a statistically significant and positive effect on the occurrence of stabilization events. Substantively, a one percent increase in the lagged level of social expenditure is associated with a 61 percent increase in the hazard (risk) of a major stabilization event. This is consistent with Hicks & Zorn (2005) who find a similar effect of lagged social expenditure per capita for the occurrence of retrenchment events. In contrast to popular claims that large welfare programs create influential constituencies that prevent reductions in welfare spending (Pierson 2001), estimation results presented in Table 11 suggest that a substantial reduction in social expenditure happens more immediately in larger welfare states. Although this might provide support for the notion that fiscal stress promotes stabilization, the estimation coefficient is also likely to reflect the simple fact that in larger welfare states there is more scope for expenditure cuts.

GDP per capita and GDP growth have no statistically significant effect on the hazard (risk) of a stabilization event, nor is the estimation coefficient robust toward changes in the specification. In contrast to the unemployment rate, which has a positive and statistically significant effect on the hazard (risk) of a stabilization event. The estimation coefficient for the unemployment rate is consistent with Hicks & Zorn's (2005) "self-limiting immoderation" hypothesis. The corresponding figure for a one percent increase in the unemployment rate is a 23 percent higher hazard (risk) of a stabilization event. However, contradictory to the "self-limiting immoderation" hypothesis, the ODR has a negative effect on the hazard (risk) of stabilization. Model 1 associates a one percent increase in the ODR with a 37 percent lower hazard (risk) of a stabilization event. This pattern might be explained by the role of expectation in stabilization policies. Empirical studies on the vote-and-popularity function confirm that the unemployment rate is a key figure to predict an incumbent's chance of getting re-elected (Nannestad & Paldam 1994). A higher unemployment rate signalizes that the economy is in bad shape. Thus, if the government expects voters to be risk adverse and sees unemployment as an indicator for increasing issue salience, higher unemployment would create momentum for substantial stabilization reforms. This type of expectation mechanism is less likely to apply to the current level of the ODR, since ageing might appears to be a long-term policy issue that is not in the direct sphere of governmental action. However, considering expected future aging instead of the level of ODR, it can be shown that expected population aging indeed induces a reduction in pension spending (Tepe 2008a). These considerations might help to understand why estimation results support the “self-limiting immoderation” prediction with respect to unemployment but not with respect to share of elderly people in the society.

Consistent with the “competitive” view, trade openness shows signs of promoting substantial stabilization reforms. Although the substantive effect is rather small (2 percent increase in the risk of stabilization) and statistically non-significant, the direction of the estimated coefficient is consistent with findings from the short-term interactive model specification where trade openness was associated with lower social expenditure growth rates. Coefficient estimates for social insurance contributions measured as a share taxes indicates that counties with contribution financed welfare programs require a longer passage of time until substantial stabilization reforms occur. Although the statistical significance of the estimation coefficient is sensitive to the model specification, taken as a whole, social insurance contributions as a percentage of tax revenues tend to decrease the hazard (risk) of stabilization (by 6 percent in Model 1). This result provides further indicative evidence on Scharpf (2000b) who claims that contribution-financed welfare states, in which welfare entitlements tend to enjoy the status of protected property rights, have limited capacity to cut back social expenditure. Explorative analyses of public pension systems in 17 OECD countries equally suggest that the development of contribution-based insurance systems within the last decade is characterized by growing fiscal commitment to welfare spending (Tepe 2008b).

Models 2 and 3 include the variables measuring the ideological composition of the cabinet. Consistent with the partisanship approach (Hibbs 1977, Tufte 1978) leftist governments show significant signs of delaying substantial adjustments events in social expenditure, while rightist governments show significant signs of promoting such events. Thus, the passage of time until stabilization occurs is likely to increase with a leftist government. In substantive terms, a one percent increase in the leftist cabinet composition decreases the hazard (risk) of a stabilization event by 3 percent while a rightist cabinet increases the hazard (risk) of a stabilization event by 2 percent. Although the substantive effect is rather small and the estimation coefficients do not exert statistical significance in each model, the partisanship pattern remains in different specifications (see Model 6 and 7).

Estimation results of the Cox model provide weak evidence on electioneering. Although the estimated coefficients shows in the direction of electioneering (Nordhaus 1975, McRae 1977, Tufte 1978), indicating that substantial stabilization events are more likely to occur at the beginning of a legislation period, the coefficients are non-significant. Models 1 to 4 in Table 12 re-estimate the electioneering model using period dummies for the single phases of the electoral cycle instead of using the number of years left in term. The pattern of estimation coefficients also leans in the direction of electoral cycles, indicating that stabilization events are getting less likely with elections approaching. Nevertheless, estimation results still provide no convincing proof of electioneering.

Table 12. Event history analysis and electioneering

	Model 1	Model 2	Model 3	Model 4
		(10 % in 3 Years)		
Social exp. $_{t-1}$	0.46**	0.54**	0.48**	0.50*
	[0.2]	[0.2]	[0.2]	[0.3]
GDP per capita log	-0.11	-0.87	-0.22	-1.52
	[1.3]	[1.6]	[1.4]	[2.2]
GDP growth $_{t-1}$	0.059	0.13	0.073	0.15
	[0.2]	[0.2]	[0.2]	[0.3]
Unemployment rate	0.22**	0.22***	0.21***	0.20**
	[0.09]	[0.08]	[0.08]	[0.09]
ODR	-0.47**	-0.47**	-0.46**	-0.43
	[0.2]	[0.2]	[0.2]	[0.3]
Trade openness $_{t-1}$	0.018	0.014	0.017	0.019
	[0.01]	[0.02]	[0.02]	[0.01]
Social ins. as % of taxes $_{t-1}$	-0.056**	-0.068***	-0.057**	-0.072**
	[0.02]	[0.02]	[0.02]	[0.03]
Years left in term = 0	-0.66			
	[0.8]			
Years left in term = 1		-0.83		
		[0.9]		
Years left in term = 3			-0.32	
			[1.5]	
Years left in term = 4				1.37
				[1.0]
Observations	447	447	447	447
Number of id	21	21	21	21
Wald chi	54.9***	98.8***	39.5***	125***
Log Likelihood	-14.1	-13.9	-14.2	-13.1

Note: Failures = 12, cell entries are coefficient estimates, robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance

Finally, with Model 5 in Table 11 institutional rigidity enters the Cox hazard model. Estimation coefficients indicate that larger institutional rigidity appears to delay adjustment events in social expenditure. A one percent increase in institutional rigidity is associated with a remarkable 87 percent decrease in the hazard (risk) of major stabilization events. This implies that with a larger share of institutional veto players substantial stabilization events are less likely to occur immediately. This result supports Alesina et al. (2006) who find that governments with concentrated executive power stabilize more rigorous. Thus, it appears that concentrated political power hinders the incumbents' opponents from delaying substantial fiscal stabilization in social expenditure.

As a part of the robustness and sensitivity analysis, Model 6 in Table 11 withholds the variables on lagged social expenditure, GDP per capita and GDP growth from the baseline model, since social expenditure has the same denominator as the criterion for defining stabilization events. Therefore these measures might have artificial relations with stabilization events. Excluding these variables form the statistical model effects the significance levels but not the direction of the estimation coefficients. The Wald test of the joint effect of the political determinants (see Model 6 and 7) indicates that the four political determinants exert a statistically significant impact on the occurrence of substantial stabilization events. Finally, results of the Likelihood-ratio test comparing Model 7 and Model 1 from Table 11 also suggest that the political determinants are significantly different from zero (LR chi2(4) = 22.91***).

Table 13. Event history analysis of stabilization events (k=3; c= -0.05)

	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6	Model 7
Social exp. $_{t-1}$	0.096	0.1	0.1	0.096	0.088		0.079
	[0.07]	[0.07]	[0.07]	[0.07]	[0.08]		[0.09]
GDP per capita log	-1.31	-1.47	-1.35	-1.31	-1.16		-1.19
	[1.0]	[1.0]	[1.0]	[1.0]	[1.4]		[1.4]
GDP growth $_{t-1}$	-0.11*	-0.096	-0.11	-0.11*	-0.11*		-0.089
	[0.06]	[0.07]	[0.07]	[0.06]	[0.07]		[0.07]
Unemployment rate	0.052	0.058	0.053	0.052	0.058	0.13**	0.073
	[0.05]	[0.05]	[0.05]	[0.05]	[0.05]	[0.05]	[0.06]
ODR	-0.022	-0.0069	-0.023	-0.022	-0.028	0.023	-0.004
	[0.1]	[0.1]	[0.1]	[0.1]	[0.1]	[0.07]	[0.1]
Trade openness $_{t-1}$	0.002	0.001	0.002	0.002	0.002	0.000	0.000
	[0.006]	[0.007]	[0.006]	[0.006]	[0.006]	[0.008]	[0.006]
Social ins. as % of taxes $_{t-1}$	-0.036**	-0.039**	-0.034**	-0.036**	-0.034*	-0.043*	-0.041*
	[0.02]	[0.02]	[0.02]	[0.02]	[0.02]	[0.03]	[0.02]
Left		-0.0047				-0.012*	-0.008
		[0.005]				[0.006]	[0.006]
Right			0.002			-0.008	-0.005
			[0.006]			[0.006]	[0.007]
Years left in term				0.001		-0.079	-0.034
				[0.2]		[0.2]	[0.2]
Institutional constraints					-0.11	-0.74**	-0.22
					[0.5]	[0.3]	[0.5]
Observations	461	461	461	461	461	463	461
Number of id	21	21	21	21	21	21	21
Wald chi2	20***	19.8***	20.9***	20.1***	29***	25.5***	32.8***
Wald (Pol. Vars.)						12.69**	2.01
Log Likelihood	-55.2	-54.9	-55.2	-55.2	-55.2	-56	-54.7

Note: Failures = 27, cell entries are coefficient estimates, robust standard errors in brackets, *** p<0.01, ** p<0.05, * p<0.1 levels of significance, Wald (Pol. Vars.) = Wald test for the inclusion of political variables.

In order to further check the sensitivity of estimation results, Table 13 presents the result of the Cox repeated event hazard analysis of major stabilization reforms with *k*= 3 and *c*= -0.05, which makes 27 stabilization events. Compared to Table 11, the control variables point in the same direction, although the substantive effect tends to be smaller and many estimation coefficients now do not reach statistical significance. However, a one percent increase in social insurance contributions as a share of tax revenues still reduces the hazard (risk) of a stabilization event by 4 percent. Estimate coefficients of the political variables also show in the same direction although none of them reaches statistical significance in the single specification (Models 2 to 5). The Wald test of the joint effect of the political determinants (see Models 6 and 7) indicates that the four variables still exert a statistically significant impact on the occurrence of substantial stabilization events. In all, although the descriptive analysis (see Table 10) and Korpi & Palme (2003) suggest to rely on c=-0.10, estimation results presented in Table 13 provide indicative support for the robustness of findings obtained form the event history analysis (Table 11). Even if stabilization events are defined as relatively frequent and "small" adjustments, the overall pattern of estimation coefficients for the political variables remains.

To sum up, focusing on 10 percent reductions over a three year period indicates that stabilization is likely to occur more immediately in countries with a relatively large welfare state, high unemployment and an open economy. Contribution-financed welfare provision, in contrast, has been associated with delayed stabilization. Compared to these macroeconomic determinants, politics plays a minor role for the timing of substantial stabilization events. However, results from the event history analysis tilt in the direction that leftist governments and institutional rigidity foster a delay in substantial stabilization events, providing support for the partisanship approach (Hibbs 1977, Tufte 1978) and the hypothesis that multiple institutional veto points make it easier for the opposition to block immediate stabilization (Alesina et al. 2006, Tsebelis 2000). Nevertheless, these results should be evaluated carefully since the robustness analysis already indicates that findings are sensitive to the definition of stabilization events.

4. CONCLUDING REMARKS

Theories of partisanship, electioneering and institutional rigidity have been shown to be too general and all encompassing to explain recent developments in welfare spending. Kittel & Obinger (2002: 50) argue the breakdown of the statistical association between political determinants and social expenditure “does not mean that politics does not matter any more. It means that politics matters in more subtle ways.” Taking into account recent theoretical claims in comparative welfare state research and political economy (Franzese 2002, Pierson 2004), this study has explored how temporal contexts structure policy-makers’ incentives and abilities to manipulate welfare spending. For this endeavor, the “war of attrition” model by Alesina & Drazen (1991) has been used to conceptualize stabilization reforms in social expenditure and to link the timing of these reforms to politico-economic theories of partisanship, electioneering and institutional rigidity.

First and foremost both the dynamic cross-sectional time-series regression and event history analysis suggest that stabilization reforms in social expenditure are primarily driven by macroeconomic developments. With regards to the inhibited size of the welfare state, the short-term analysis indicates constantly growing social budgets, while the event history analysis suggests that substantial reductions in welfare spending are more likely to occur in countries with high levels of spending. This pattern speaks for discontinuous adjustments in the level of social expenditure followed by periods of continuing social expenditure growth. Equally interesting are findings demonstrating the effect of rising dependency ratios. While growing unemployment leads to significantly higher social expenditure growth, population aging is no longer positively associated with a significantly growing welfare state. Although societies with a larger share of elderly will have to devote more public resources towards the elderly, a growing ODR does not translate into equally strong growing social expenditure, suggesting that mature welfare states policy response to fiscal pressures emerging from population aging is expenditure containment. The event history analysis suggests that higher unemployment rates create momentum for substantial stabilization events, supporting the “self-limiting immoderation” hypothesis put forward by Hicks & Zorn (2005). These findings may challenge standard rational choice assumptions. In any case they signal that expectations and policy framing have yet been underappreciated in the analysis of welfare reform.

Concerning international economic integration, this study provides further evidence for the "competitive" hypothesis (Krugman 1995, Greider 1997, Friedman 2000). Both types of models – short-term incremental and long-term substantial stabilizations – indicate that economic integration forces governments to reduce spending. Equally consistent estimation results are obtained with respect to contribution-based welfare financing. Providing further support to Scharpf (2000b); welfare states with a large share of contribution-financed services seem to have difficulties in containing social expenditure growth and fail to stabilize immediately.

The role of political determinants for the conduct of stabilization reforms in social expenditure can be summarized as follows: In general, political variables contribute little in explaining the dynamics of social expenditure growth and the occurrence of stabilization events. With respect to the first research question on the conditionality of political determinants, the empirical analysis attests to the fact that patterns of partisanship and electioneering are more pronounced in periods of fiscal stress, albeit the conditional effect is not thoroughly convincing. With respect to the second research question concerning the effect of political determinants on the passage of time, findings from the event history analysis suggest that leftist governments and institutional rigidity are likely to delay substantive stabilization events. A vanishing partisan effect, which Pierson's (2001) approach might suggest, could not be confirmed by the statistical analysis. Moreover, contrary to recent claims of a reversal of partisanship in social policy reform (Kitschelt 2001), empirical evidence on the timing and rigorousity of stabilization reforms in social expenditure suggest a persistent partisan orientation.

In total, findings confirm prior research that "politics matters" less compared to macroeconomic factors (Castles 2001). However, if time-based contextual factors, which are predicted to affect policymaker's abilities and incentives to manipulate welfare spending, are taken into account, there is tentative evidence that the old patterns of partisanship and electioneering reemerge. These findings support Franzese's (2002) and Franzese & Jusko's (2006) argument that the degree, character, and effectiveness of political-budget-cycles is structured by contextual factors. The ambivalent role of institutional rigidity for the timing of stabilization reforms is difficult to explain. Yet, there is no unambiguous explanation for the finding that institutional rigidity tends to delay substantial stabilization events in the level of social expenditure, but promotes short-term incremental stabilizations in social expenditure growth. Methodological limitations of this study should also be taken into consideration. First, the definition of fiscal stress and stabilization events is essential. This study has tried to develop a thoughtful indicator for country-years of growing fiscal imbalance as well as an indicator for stabilization events in social expenditure that are driven by quantitative data. Subsequent studies could apply alternative measure to explore the role of temporal context for the conduct of welfare reforms. Second, this study does not provide a straightforward empirical test of the Alesina & Drazen (1991) model. Nevertheless, the "war of attrition" framework has proven to provide a useful conceptual basis to integrate issues of temporality into the politico-economic analysis. Testing the "war of attrition" model in an experimental setup could offer valuable insights to the micro processes of political conflict resolution.

Although this study has examined a relatively new approach that clearly needs further theoretical and empirical advancement, it has attempted to show that a careful integration of the temporal dimension in politico-economic analysis can improve not only the understanding of welfare politics in mature welfare states, but also offers much promise for resolving the empirical anomalies present in the comparative welfare state literature. Contrary to popular concerns of depoliticization of welfare policies, this study indicates that politics still matters; however, it matters in more subtle ways than in previous decades.

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Appendix Table 1. Definition and source of variables

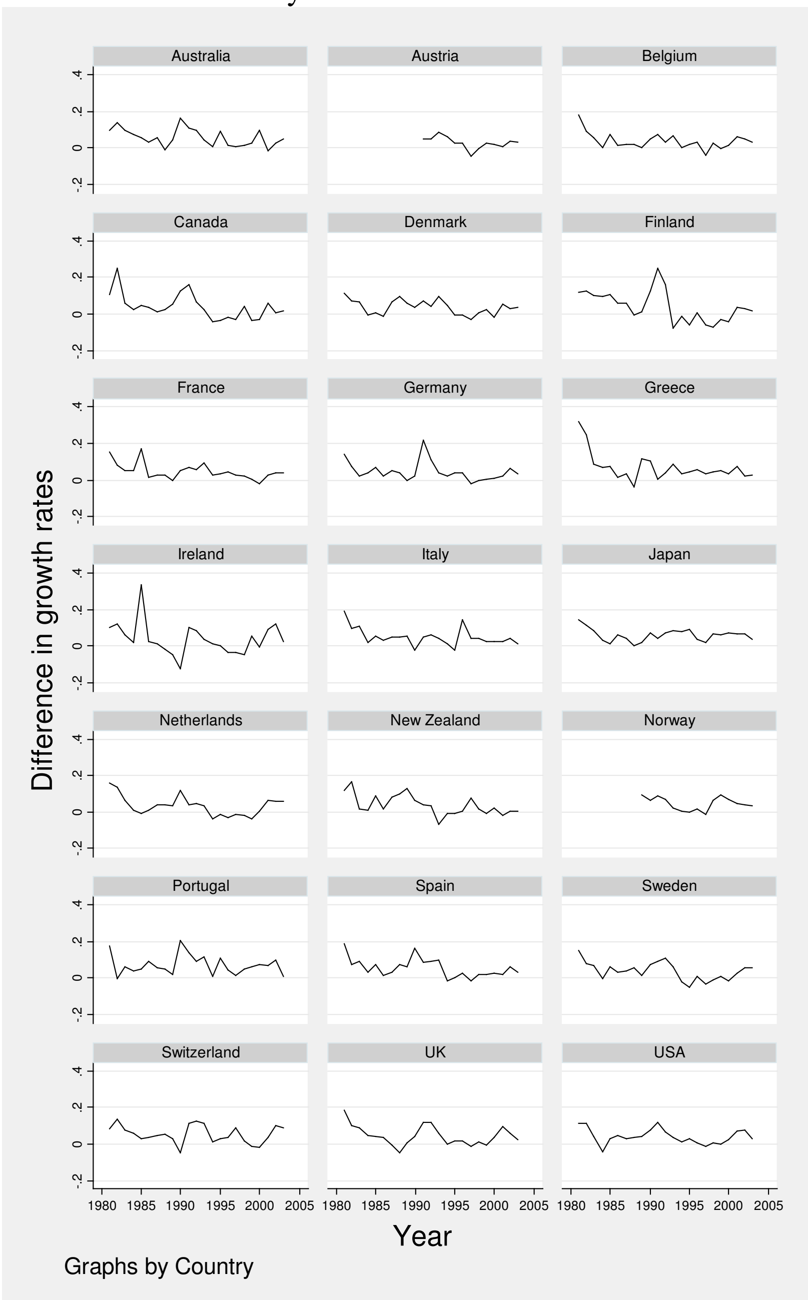
Variable	Definition	Source
Social expenditure per GDP	Total social expenditure as percentage of GDP	OECD Social Expenditure Database (2007)
Social expenditure per head	Total social expenditure in current prices and current PPP's per head	OECD Social Expenditure Database (2007)
Real GDP per Worker	Real GDP per Worker	Penn World Tables (2007)
GDP per capita growth	Gross Domestic Product per capita growth	Penn World Tables (2007)
Unemployment	Unemployment rate as a percentage of civilian labor force.	OECD Labor force statistics via Armingeon et al. (2007) Comparative Political Dataset
ODR	Elderly population (65+) as a percentage of the working age (15-64) population.	OECD Labor force statistics via Armingeon et al. (2007) Comparative Political Dataset
Trade openness	Openness of the economy in current prices, measured as total trade (sum of import and export) as percentage of GDP.	OECD Labor force statistics via Armingeon et al. (2007) Comparative Political Dataset
Social insurance as % of taxes	Social security contributions as percentage of GDP divided by the total tax revenues as percentage of GDP.	OECD Revenue Statistics (2007)
Leftist gov.	Left-wing parties in percentage of total cabinet posts.	Armingeon et al. (2007) Comparative Political Dataset
Rightist gov.	Right-wing parties in percentage of total cabinet posts.	Armingeon et al. (2007) Comparative Political Dataset
Years left in term	Years left in current term	World Bank Database of Political Institutions (2005)
Institutional rigidity	Sum of index of bicameralism (1971-1996) and index of federalism (1971-1996)	Lijphart (1999: 313-314)

Appendix Table 2. Panel unit root test for social expenditure per GDP

		Level	Difference
No Trend			
	chi2(42)	41.11	230.85
	Prob > chi2	0.51	0.00
Trend			
	chi2(42)	30.91	167.69
	Prob > chi2	0.90	0.00

Note: Fisher Test for panel unit root using an augmented Dickey-Fuller test

Appendix Figure 1. Measuring fiscal imbalance in pay-as-you-go social security systems



Note: Differences in growth rates is defined as the growth rate of social expenditure per head minus the growth of real GDP per worker. Positive values indicate imbalanced growth in social spending.

GERMAN SUMMARY

Die vorliegende Dissertation besteht aus drei Aufsätzen, die im Rahmen des DFG Graduierenskollegs „Pfade organisatorischer Prozesse“ verfasst wurden.

Am Beispiel der staatlichen Altersvorsorge untersucht der erste Aufsatz „Traveling without Moving? Pension regime change in mature welfare states”, ob und in welchem Umfang führzeitige institutionelle Entscheidungen die gegenwärtige Verfügbarkeit von Reformoptionen einschränken. Der Vergleich der Rentensysteme in 18 OECD Staaten von 1980-2003 legt den Schluss nahe, dass die Wahl der Reformoptionen einem Muster folgen, das als “Bounded Change” beschrieben wird. Demnach konvergieren Reformenstrategien innerhalb eines wohlfahrtsstaatlichen Regimes während Differenzen zwischen Regimes fortbestehen.

Der zweite Beitrag trägt den Titel „Are Mature Welfare States on the Path to Gerontocracy? Evidence from 18 OECD countries”. Vor dem Hintergrund des Median Wähler Paradigmas wird die Wirkung der gesellschaftlichen Alterung auf den Umfang und Generosität der staatlichen Rentenausgaben untersucht. Die Ergebnisse der Regressionsanalyse für 18 OECD Staaten von 1980-2003 zeigen, dass bei steigendem Rentnerquotienten die absoluten Rentenausgaben steigen, während die individuellen Rentenausgaben unverändert bleiben oder sinken. Die Intensität dieses Effekts hängt jedoch von institutionellen Faktoren ab – je stärker der Zusammenhang zwischen Beiträgen und zukünftiger Rente desto stärker kann der gesellschaftliche Alterungsprozess mit steigenden Rentenausgaben in Verbindung gebracht werden.

Der dritte Beitrag „What Makes Stabilization Reform Happen? Temporality in the political economy of welfare spending", untersucht den Einfluss von Wahlzyklen und parteipolitischen Differenzen auf das Wachstum der Sozialausgaben im Kontext fiskalischer Krisen. Auf Basis von 21 OECD Staaten (1980-2003) und mit Hilfe einer interaktiven Modellspezifikation können empirische Belege für diesen Konditionaleffekt identifiziert werden. Betrachtet man den Einfluss von Wahlzyklen und parteipolitischen Differenzen auf die Dauer einer politischen Anpassungsreaktion im Rahmen einer Event History Analyse wird die zunehmende Relevanz politischer Determinanten in Krisenzeiten ebenfalls bestätigt. Insgesamt stützen die Befunde die These das "politics still matters", sofern der temporale Kontext in der Analyse berücksichtigt wird.