

# A race to liberalization? Diffusion of economic policy reform among OECD-economies

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**Abstract** Most OECD economies witnessed a liberalization of economic policies over the past thirty years. The present paper examines to what extent this development is caused by domestic political and economic factors on the one hand, and international policy diffusion via competitive interaction of governments on the other. Employing a comprehensive index of economic reform it can be shown that policy diffusion is a driving factor for economic liberalization. Especially in the fields of regulatory, monetary and trade policies we find significant interdependence of policy choices, as suggested by theories of policy diffusion.

**Keywords** Market-oriented reform · Policy diffusion · Yardstick competition

**JEL:** D78, P11, P21

## 1 Introduction

Over the past three decades, many countries in all regions of the world joined a global trend towards market-friendly economic policies. One of the main puzzles in Political Economy is to give reasons for this trend towards liberalization. While Public Choice theory has been very successful in explaining the appearance and the persistence of politically imposed restrictions on economic freedom, only recently a newly emerging Political Economy of Reform provided further theoretical insights on the economic and political preconditions to fundamental policy changes.

Research has focused primarily on two major determinants of market-oriented reforms. On the one hand, theories of crisis and reform (Rodrik, 1996) emphasize that fundamental policy changes are enacted as a response to severe economic downturns. Drazen and Easterly (2001), and Pitlik and Wirth (2003) report empirical evidence in support of the crisis

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hypothesis. On the other hand, institutional approaches consider that governments which are less constrained by checks and balances may be more decisive when it comes to enacting policy changes. The idea is that reform-minded politicians do not depend on the consent of numerous reform-opposing groups and their representatives (e.g. Cox & McCubbins, 2001). De Haan and Sturm (2003), Pitlik and Wirth (2003), Herz and Vogel (2005), and Pitlik (2005) yet find no empirical support for this hypothesis. Contrary to conventional wisdom, institutional constraints for executive action often appear to be positively related to liberalization.

A further impulse for a policy change may be that incentives to undertake reforms can be shaped by the adoption of market-friendlier policies by other governments. Weyland (2002, p. 2) claims that “one of the most striking phenomena in the area of public policy are the waves of diffusion that sometimes sweep across important regions of the world . . . A bold reform adopted in one nation soon attracts attention from other countries, which come to adopt the novel approach.” Put differently, economic policy reforms may be influenced by a diffusion of policies, i.e. “the process by which an innovation is communicated through certain channels over time among members of a social system” (Berry & Berry, 1999, p. 171). Economic policies of countries are linked, e.g. by a competitive strive to attract internationally mobile firms and capital, or by a spread of new economic ideas, which may be promoted by the international community of economists as policy advisors.

Surprisingly, international policy diffusion as a potential source of policy liberalization has not found much attention in the empirical literature on broad reforms, with a few exceptions. Investigating causes of a liberalization of the current account and the capital account, as well as a unification of the exchange rate, Simmons and Elkins (2004, p. 171) find that in a world-wide country sample economic liberalization clusters in time and space. Brune and Garrett (2000) study determinants of privatization for developing countries over the period 1988–1997. They show that privatization in other countries of a nation’s peer group has substantial positive impact on SOE sales in the respective country. Supporting a central notion in theories of tax competition, some evidence can also be found for a strategic interdependence among OECD-members in corporate income taxation (Devereux et al., 2002) as well as for capital taxation in general among EU 15-countries (Pitlik, 2006). Governments respond to tax reforms in competing countries by adjusting tax policies in similar directions.<sup>1</sup> For a sample of industrial countries International Monetary Fund (2004) reports that domestic reform efforts are often linked with economic policies of main trading partners. Policy interdependence is reported for labor and product market reforms, financial and trade liberalization, but not for tax reforms. Most recently, Meseguer (2006) reports preliminary evidence for policy learning from successful countries with respect to trade liberalization, privatization, and the decision to enter IMF agreements.

These studies, however, only focus on a limited set of economic policies. This paper aims at filling the gap in the empirical literature on comprehensive policy reforms by exploring economic policy diffusion among 23 OECD-countries over the time period 1970–2000. In contrast to the above mentioned studies, diffusion of economic policies is analyzed by employing a distinctive index of market-oriented policy, developed by the Fraser Institute (Gwartney & Lawson, 2004). In particular, we examine if there is a linkage between overall economic policy reforms among the countries in our sample.

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<sup>1</sup> Yet, strategic interaction in capital taxation may be conditional on the inclusion of political variables which affect the ability of governments to respond to tax policy changes in competing countries. See Basinger and Hallerberg (2004) and Pitlik (2006).

In Section 2, stylized facts about economic policy diffusion among OECD-countries are presented. Section 3 reviews very briefly some theoretical concepts of international policy diffusion and discusses explanations why reform decisions may cluster. Section 4 provides empirical tests of the main hypotheses. Section 5 offers conclusions.

## 2 Market-oriented policy reforms in OECD-countries

According to Solomon (1999, p. 1), “the world has undergone dramatic changes since the end of the 1970s . . . The principal one was a shift toward greater reliance on market forces . . . Governments became less involved in economic processes: markets were deregulated, import barriers were lowered, income taxes were reduced, state-owned enterprises were privatised and central planning, where it existed, gave way to direction by the price system.” While most researchers probably agree to Solomon’s statement in general, a quantification of his hypothesis is far from being a trivial exercise. Measuring comprehensive market-oriented reforms requires a clear understanding of (i) which policies matter most and should be included in an indicator for broad reforms, (ii) how to assign weights to these policy areas in constructing such a measure, and (iii) what kind of policies do constitute a market-friendly environment. Although we know about the detrimental effects of many interventions, *some* regulations and public expenditures are crucial for the functioning of a market economy. It is not yet clear *how much* government involvement exactly should be qualified as market-supporting. Despite these caveats, the measurement of market-oriented reforms has been the subject of an increasing number of papers. See Loayza and Soto (2003) for an overview.

One of the most ambitious projects in this field of research is a broad index developed by the Fraser Institute (Gwartney & Lawson, 2004). The Economic Freedom of the World-index (*efw*) estimates the degree of market-friendliness of policies on a 0 to 10 scale by a multidimensional set of indicators, with higher values indicating more economic freedom. The overall *efw*-index is based on five policy areas, all weighted equally, including indicators for (i) government size, (ii) legal structure and security of property rights, (iii) sound monetary policies, (iv) openness to international trade and capital movements, and (v) regulation of business and labor markets. For the time period 1970 to 2000 the index provides data in five year intervals.

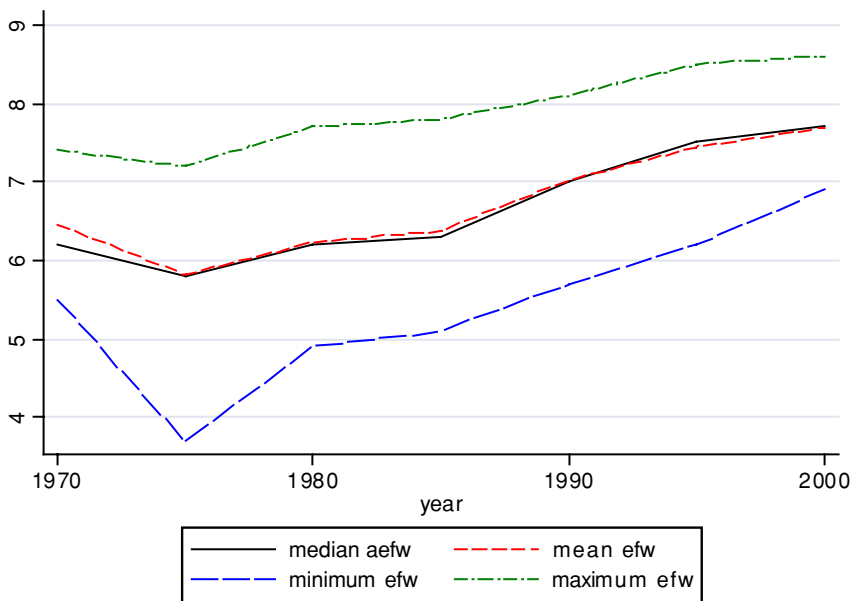
Figure 1 depicts the development of the comprehensive *efw*-index in a sample of 23 OECD-economies.<sup>2</sup> Both the mean and the median, as well as the minimum and the maximum values follow a similar time path. During the 1970–1975 period, a deterioration of economic freedom is observed. Since then, all countries witness a liberalization of economic policies. The coefficient of variation drops from 0.146 in 1975 to 0.061 in 2000, indicating a significant convergence of policies as judged by the *efw*-index.

Table 1 shows that the nations included in the sample followed parallel patterns of economic policy changes, measured by changes in the *efw*-index ( $\Delta efw$ ) over the respective periods. During 1970–1975, 22 out of 23 countries restricted liberties. Since then, in every time interval at least 16 nations liberalized economic policies. In sum, the data indicate a convergence of policies among OECD members towards economic liberalization. Can corresponding changes in economic policies be explained? Are similar policy changes the result of isolated reform decisions, or do they originate from explicit harmonization and coordination,

<sup>2</sup> The sample contains all EU 15-countries plus Australia, Canada, Iceland, Japan, New Zealand, Norway, Switzerland, and the U.S.

**Table 1** Change in economic freedom ( $\Delta efw$ ) in 23 OECD-countries (1970–2000)

Period	Mean	Std. Dev.	Minimum	Maximum	Observations with	
					$\Delta efw > 0$	$\Delta efw \leq 0$
1970–1975	-0.64	0.536	-2.3	+0.1	1	22
1975–1980	+0.41	0.369	-0.1	+1.8	21	2
1980–1985	+0.15	0.285	-0.5	+0.9	16	7
1985–1990	+0.65	0.352	+0.1	+1.5	23	0
1990–1995	+0.42	0.396	-0.2	+1.2	19	4
1995–2000	+0.25	0.225	-0.3	+0.7	20	3

**Fig. 1** Economic freedom in 23 OECD-countries, 1970–2000

or is some kind of policy diffusion at work? The remainder of the paper aims at providing some answers for these questions.

### 3 Policy convergence, policy diffusion, and economic reform

Although there are strong signs of convergence of economic policies among OECD members, that does not necessarily mean that governments respond to each other when they decide on liberalization. Elkins and Simmons (2005) consider three sources of a clustering of similar policies, namely (i) similar responses to similar problems and conditions, (ii) coordinated policy making, and (iii) uncoordinated interdependence. The quintessence of policy diffusion is that the probability of adoption of a certain policy increases the more other governments have adopted that policy. Policy diffusion is therefore characterized by interdependence of

policy decisions of different countries, working through different channels of communication (cf. Levi-Faur, 2005).

One source of a policy clustering is that countries react in the same way, but independently to similar changes in domestic conditions. Above all, a crisis explanation of economic reforms fits this category. If economic crises facilitate policy reforms, economic shocks which hit a majority of countries simultaneously cause similar policy reactions. One may think of oil crises in the 1970s and 1980s that troubled almost every developed nation.

Comparable policies in a cluster of nations may also be facilitated by coordinated policy decisions, e.g. by international treaties enforced by supra-national organizations, or by a more decentralized process of coordinated policy making through communication and sharing of expertise.<sup>3</sup> Coordinated policy making in international organizations may help explaining trends towards trade liberalization. E.g., trade barriers in the European Union have been abolished simultaneously in member states in order to expand markets. These conjectures do not necessarily carry over to other policy areas. International treaties regularly remained silent about policy harmonization and international standards on economic regulation, monetary policy, and a coordination of fiscal policies, until recently.<sup>4</sup> Evidence on the impact of membership in international organizations on trade reforms is yet not too conclusive. In a recent study, Rose (2004) shows that WTO members do not have a more liberal trade policy than non-members.

Finally, an adoption of similar policies may as well be a result of uncoordinated, yet interdependent policy making. According to Brueckner (2003) a resource flow model and a spillover model can be distinguished as important sources of policy diffusion. Both models capture the idea that policies are linked by (international) externalities. Strategic policy interdependence then might lead to an uncoordinated adoption of similar policies.<sup>5</sup> In the resource flow framework, externalities stem from policy decisions which have an impact on residential choice of mobile resources. According to this perspective, competition over scarce resources links policies. For example, theories of tax competition conjecture that governments try to attract mobile capital by reducing the capital tax burden. A tax cut by one jurisdiction generates pressures on other actors to respond similarly with tax reductions (Wilson & Wildasin, 2004). A similar logic of a competitive race to more liberal policies is often applied to explain deregulation, a liberalization of international trade and factor movements, or a cutback of the welfare state (Sinn, 2003). Lower costs of international transactions in goods and factor markets contribute to an increasing strategic interdependence of national economic policies.

<sup>3</sup> There certainly exist various ways to institutionalize coordination among independent governments. E.g., the “Open Method of Coordination” is a new form of governance within the European Union, which aims at providing information on “Best Practices” in policy areas with no competences for the Union. See Eckardt and Kerber (2005).

<sup>4</sup> In developing countries reform programs have been heavily demanded, but less rigorously monitored or enforced, by Washington Financial Institutions. Developing nations often had little choice but to accept proposals and to adopt certain policies to gain access to international aid. Yet, with respect to economic reforms in most OECD countries it is hardly plausible that this is a major source of policy convergence.

<sup>5</sup> Simmons and Elkins (2004, 2005), provide a slightly different typology. In their view, a first channel is the impact of economic policies of one country on costs and benefits of other governments’ policies. A second channel is that policy choices of a country provide new information about the suitability of certain policies which may then impact other governments’ decisions to change their own course of action. In that case, one may think of policy diffusion via information externalities. See Eckardt and Kerber (2005, pp. 141–142). Both channels thus also reflect the spillover and the resource flow models. Despite minor differences, Brueckner (2003), Simmons and Elkin (2004), and Eckardt and Kerber (2005) agree broadly on the driving forces of policy diffusion.

In the spillover model, each jurisdiction chooses policies autonomously, but parameter choices in other jurisdictions have a direct impact on optimal choices of the respective government. For example, the process of yardstick competition (Besley & Case, 1995) connects policy decisions through information provided by certain tax policies even in the absence of capital mobility. If voters judge competence of home authorities by comparing tax rates, tax parameter choices of other jurisdictions directly affect domestic tax policies. To the extent that tax rates in neighboring nations serve as a benchmark, yardstick competition generates incentives to cut taxes at home as a reaction to tax reductions abroad (Salmon, 2003).

However, since competition for mobile factors and yardstick contests are not restricted to a particular policy field, it is not obvious that externalities unequivocally cause an adoption of similar policies. In a seminal paper, Tiebout (1956) develops the idea of fiscal competition among local jurisdictions. According to his view, governments compete for heterogeneous citizen-taxpayers by offering different bundles of public services and taxes, establishing a fiscal competition in both spending and tax policies. Likewise, rational voters do not compare only taxes, but judge overall quality of public services as well, to gather information whether the home government is competent. High tax jurisdictions, providing a high quality of public services, and low tax-low quality jurisdictions might co-exist in equilibrium. Competition does not necessarily imply that policies strictly converge as long as voters and owners of mobile resources judge competence of governments and attractiveness of jurisdictions by the entire set of fiscal and monetary policies as well as the legal and regulatory environment.

From a theoretical perspective it is not clear-cut that a rivalry for scarce resources and yardstick competition cause a comprehensive liberalization of policies. Competing for capital surely requires an abolition of restrictions on capital flows, stable monetary conditions, and security of property rights, all of which are vital elements of market-oriented economic reforms. Governments yet often try to attract investment by means of distortionary subsidies (Wildasin, 1988) or further policies which are not conducive to economic freedom. One might also conjecture that it is hardly plausible that voters are informed about policy packages as a whole supplied by nations. Even if information costs are reduced if voters can compare policies of jurisdictions, they generally remain rationally ignorant about policy contents in detail (Downs, 1957). That does not rule out the impact of benchmark competition, however. Voters may decide by comparing easily available performance indicators. Economic successful policies, as revealed by income growth, unemployment rates etc., may then serve as a yardstick for politicians. Many studies, surveyed by Berggren (2003) and de Haan, Lundström, and Sturm (2006), strongly support the view that liberalization contributes positively to GDP growth. Thus, authorities may update their beliefs about benefits and costs of reforms and imitate policies to improve re-election chances (Menseguer, 2006). Knowledge about successful strategies – presumably communicated by economic advisors – then supports policy learning (Eckardt & Kerber, 2005).

The resource flow and the spillover model leave open the question, which countries belong to the main reference groups, i.e. who are direct competitors and peers whose policy choices affect domestic governments' decisions. Are economic policies more connected among neighboring jurisdictions, or among countries with a similar economic or legal environment, or among nations with a similar historical background? One might suppose that the cost of capital movements and the information cost of voters and policymakers about policies in other nations fall substantially with geographical closeness, and with increasing social, cultural and economic similarity of the countries. It can also be assumed that reform decisions of economically powerful countries have a greater impact on other agents than policy choices of smaller countries. Reforms in 'leader' or 'hegemonic' states may affect willingness to reform in follower countries. While one can speculate about the potential strength of policy

reactions depending on several indicators of connectivity or distance, these are essentially empirical questions.

## 4 Empirical investigation

### 4.1 Model specification and estimation procedure

Theories of policy diffusion claim that policy choices of one government depend on policy choices of other governments. As regards market-oriented reforms, an initiation of reforms in a number of countries might enhance the probability of similar policy changes in other countries. Hence, the estimating equation can be written

$$efw_{i,t} = \alpha \cdot efw_{i,t-1} + \beta \cdot Wefw_{i,t} + \gamma \cdot X'_{i,t} + \eta_i + \mu_t + \varepsilon_{i,t}, \quad (1)$$

with  $i$  and  $t$  representing countries and time periods. To capture persistence of policies a lagged dependent variable  $efw_{i,t-1}$  is included, the coefficient  $\alpha$  expected to be positive.  $X'_{i,t}$  is a vector of control variables, to be discussed below. Hausman tests clearly support use of a fixed effects-estimator. Unit effects  $\eta_i$  are therefore included to control for unobserved country heterogeneity. A deterministic time trend  $\mu_t$  is also included.<sup>6</sup>

The central focus of the study is on the coefficient  $\beta$  of the explanatory variable  $Wefw$ . As seen from a country  $i$ ,  $Wefw_{i,t}$  represents a weighted average of the economic freedom indicator of all other 22 OECD-countries  $j$  ( $j \neq i$ ) in the sample, such that

$$Wefw_{i,t} \equiv \sum_{j \neq i} \varpi_{ij} efw_{j,t} \quad \text{and} \quad \sum \varpi_{ij} = 1. \quad (2)$$

The specification of a weighting scheme is based on an *a priori* choice which reflects the *presumed* interdependence of reform decisions in country  $i$  with other countries  $j$ . The lower the weights  $\varpi_{ij}$ , the less we expect policy decisions in  $i$  to depend on policy choices in  $j$ . Choosing different weighting schemes and comparing estimates for  $\beta$  enables us to find out more about the international relationship of reform policies in the sample. A positive relation of  $Wefw$  and  $efw$ , i.e.  $\beta > 0$ , indicates that governments revise policies in the same direction, as suggested by theories of policy diffusion.

In the following estimations, a number of weighting schemes are employed. The simplest procedure is to assign equal weights to each country  $j$ . In this case it is assumed that policies of all other nations  $j$  in the sample matter the same for country  $i$ . One may however conjecture that some kind of geographical or socio-economic distance between OECD-members makes a government react more or less intense to policy choices in other countries. The further away a country is (by different definitions of distance), the weaker will be its impact on home country decisions. In that case, each element of the weighting matrix is defined as

$$\varpi_{ij} = \frac{1}{d_{ij}} \bigg/ \sum_j \frac{1}{d_{ij}}, \quad (3)$$

where  $d_{ij}$  is a distance measure between countries  $i$  and  $j$ .

<sup>6</sup> In principle, we would prefer to use time dummies to control for period specific shocks to reform activities. Devereux et al. (2002) however show that period dummies can almost completely absorb the effect of policy diffusion variables, especially in the case of persistent covariates. Therefore, we include a time trend variable in all models instead. Except for the regulation area, results hold when period dummies are included.



A first measure is geographical distance ( $\Delta GEO$ ) between capital cities of the two respective nations.<sup>7</sup> This reflects the simple assumption that, e.g., a German government cares more about political decisions in France than in Australia. If one assumes that the transaction cost of capital relocation depend positively on spatial distance, geographical proximity should be particularly important in an intergovernmental competition for mobile firms. Also, voters are better informed about policies in neighboring countries. Yardstick competition may be more intense, the closer the nations lie to each other.

Beck et al. (2004) note that in Political Economy space is more than geography. We also expect that according to socio-economic indicators of similarity, policy choices of more similar countries are of a higher importance for own decisions. To test this hypothesis, we construct two further (time-variant) economic distance parameters by calculating the absolute value of country differences in real GDP size ( $\Delta GDP$ ) and in real GDP per capita ( $\Delta GPC$ ), and generate the weighting matrix according to (3).

La Porta et al. (1999) have shown that the legal tradition of a country has a considerable influence on economic policy choices. In particular, nations with a British legal origin observe a higher quality of several policy measures than countries with a French tradition. Hence, we test whether policy links among nations with a similar legal origin are stronger than across countries of different traditions. To do this, we construct a similarity measure ( $\Delta LEG$ ), based on La Porta et al. (1999), in which countries with the same origin of the legal system (British, French, German, and Scandinavian) enter with equal weights. A weight of zero is assigned to countries with different legal traditions.

As noted in Section 3, the impact of one government's policy choice on others' decisions is presumably higher, the more important that nation is for the world economy. In order to capture a leadership effect, a further measure is constructed by giving higher weights to countries with higher real GDP ( $GDP$ ). Time varying weights are calculated as

$$\varpi_{i,j,t} = \frac{GDP_{j,t}}{\sum_j GDP_{j,t}}. \quad (4)$$

Moreover, in a decision to liberalize policies, governments might be influenced by success of other countries. Average GDP growth rates during the respective time period indicate the performance of economic policies. Weights for relative growth performance ( $GW$ ) are calculated by the same procedure as in (4). Alternatively, success of economic policies can also be measured by levels of real GDP per capita ( $GPC$ ), which are used to compute further weights.

Theory suggests that economic policy reforms are also driven by a number of domestic factors. To control for the impact of economic crises on liberalization, e.g. Herz and Vogel (2005) calculate a country's misery index, i.e. the averaged sum of inflation rates and standardized unemployment rates over a respective time period. Using inflation rates however possibly produces biased results as inflation performance is also included in the economic freedom index. Moreover, it is not clear which levels of the misery index constitute a severe crisis. To avoid these problems, we simply use (lagged) GDP per capita growth ( $growth$ ) over a five year period as an indicator for economic performance. If crises promote reforms, the sign of  $growth$  is negative. Furthermore, trade openness ( $open$ ), measured by the period averaged sum of import and export quotas, and log of real GDP per capita ( $GPC$ ) to control for the impact of economic development, are added to the set of explanatory variables. All economic covariates enter regressions lagged one period.

<sup>7</sup> Data are from Byers (2005).



Domestic political factors are also supposed to have an impact on reform. Institutional restrictions on government action, as established by a federalist system, an independent judiciary, or a bicameral legislature with opposing majorities, significantly increase the need for consensus to change policies. On the one hand, this may impede reforms as executive decisiveness seems to be reduced. In a static perspective, a government unconstrained by veto actors is certainly better able to execute policy changes when facing opposition of interest groups harmed by a reform. On the other hand, the need to bargain may also facilitate reforms. In a dynamic perspective the political cost of reform increase if a policy is based on unilateral action rather than on compromise, as conflicts over the distribution of the cost and benefits of reform are never truly finished. Institutional restrictions can make future policy reversals less likely and enhance credibility of reforms, thereby reducing the political cost of a policy change (Stiglitz, 1998; Spiller et al., 2003; Pitlik, 2005).

To assess the impact of institutional factors on reform policies, we follow previous studies and use an index of restrictions on executive behavior (*polcon*) from Henisz (2000). The index varies with the number of veto points in a polity and is normalized on a 0 to 1-scale, higher values indicating more restrictions. In a sample of OECD states *polcon* however shows only very little variation both over time and across countries. In addition we employ a measure of legislative fragmentation which may have an effect of its own on the ability of governments to implement reforms. The *raison d'être* is that in multiparty-systems the likelihood of single-party dominance decreases and the need for bargaining to execute policy reforms increases. A simple indicator for legislative fragmentation is the effective number of political parties in parliament (*enopp*) from Laakso and Taagepera (1979), which varies considerably in the cross country and time series dimension. The index is defined as

$$enopp = \frac{1}{\sum_k p_k^2}, \quad (5)$$

where  $p_k$  is the share of total seats in the legislature held by party  $k$ . Figures are from Armingeon et al. (2004) and are averaged over the respective five year periods. We also include *govleft* as an indicator for political orientation of a government, which measures the five-year averaged fraction of left party cabinet positions of total cabinet seats (Armingeon et al., 2004). If left-wing politicians are less favorable towards market-oriented reforms, we should expect a negative sign of *govleft*. Reform may additionally be facilitated by international organizations and multilateral arrangements. Therefore we control for membership in the European Communities/European Union (*EC*), *EFTA*, *APEC*, and *NAFTA*. The set of political variables enters with contemporaneous values. Summary statistics of all variables are shown in the appendix.

To calculate the relation between the reform variable and various covariates, we have to take into account that estimates of (1) are plagued by serious endogeneity problems, which stem from the autoregressive parameter in a fixed effects regression, as well as from *Wefw* variables. If it is true that one agent's reform decision depends on policy choices of other governments, *Wefw* is clearly endogenous. To cope with these problems in a large cross-section/small time-dimension setting, an instrumental variable GMM estimator of Arellano and Bond (1991), employing a first differences-procedure with lagged levels as instruments, is often used. Blundell and Bond (1998) yet report that the first difference GMM-estimator performs poorly in the presence of persistent variables because lagged levels of the series provide weak instruments for subsequent changes.<sup>8</sup> In that case, which is

<sup>8</sup> As a result, the first difference-GMM-estimator produces a serious downward bias of estimated coefficients for the autoregressive parameter, similar to a within-estimator, which confirms that a weak instrument bias is

clearly relevant in our context, Blundell and Bond (1998), and Bond (2002) recommend a system GMM-estimator, combining equations in first differences with equations in levels, using lagged first differences as instruments in the levels equations.<sup>9</sup> Estimates in the next sections are based on the one step-system GMM estimator, with standard errors corrected for heteroskedasticity. In order to keep the number of instruments as small as possible, the use of instrumental variables for endogenous variables is restricted to just one observation with a lag of two periods. The validity of (additionally) included instruments is tested by means of a Hansen-Sargan test of overidentifying restrictions and a difference-Sargan-test. As consistency of our estimates requires that error terms are not second-order serially correlated, we report results of AR(2)-tests.

#### 4.2 Results for overall economic reform

Table 2 presents the results of dynamic panel regressions of the *efw*-index on a number of control variables and the weighted *efw*-score of all other countries (*Wefw*), calculated with several weighting schemes. *P*-values of tests for overidentification and presence of second order serial correlation in the bottom lines indicate validity of all specifications.

Column (1) shows estimates for *Wefw* when all other countries' economic freedom scores are weighted equally. A coefficient  $\hat{\beta} \approx 0.8$  indicates a strong and highly significant positive interdependence of policies between countries in our sample. Using geographical proximity to calculate the weighting matrix (column (2)) leads to even better results. This suggests that making their policy choices governments in particular look at economic policies in nearby countries. Columns (3) and (4) report results when the weighting matrix is calculated on measures of economic similarity, i.e. differences in real GDP and in real GDP per capita. Although coefficients show the expected positive sign, results are much weaker. Hence, we cannot conclude that policymakers adjust policies primarily looking at policy choices of economically similar countries. Similarity of legal traditions performs better in explaining policy reactions (column (5)). Equation (6), using a GDP weighted index, shows that size matters for policy reactions, though estimated  $\hat{\beta} \approx 0.55$  is not exceptionally high. Hence, there is some weaker evidence in favor of a hegemony hypothesis. Columns (7) and (8) report results for performance based weighting schemes. Estimates suggest that policy decisions of successful countries, as measured by average annual GDP growth, matter a good deal for domestic reform decisions. The effect is weaker, but still highly significant, if weights are computed on a GDP per capita basis. Learning from successful countries appears to be an important source of policy diffusion.

Turning to the control variables,  $efw_{i,t-1}$  is highly significant in every specification. It shows the expected positive sign with coefficients varying between 0.39 and 0.47, indicating a considerable policy persistence. A higher lagged per capita GDP level also appears to lead to more liberalization, although the coefficient of *GPC* is not always significant. Pre-period GDP per capita *growth* has a negative sign, confirming the idea that crisis is conducive

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relevant in our case. In pooled OLS-estimates without country fixed effects coefficients of the lagged dependent variable are biased upwards, as expected. Dropping country effects in combination with estimation in levels is however inappropriate, as predictions on policy diffusion are more about reaction to *changes* of other countries policies. To check robustness of our findings we re-estimated all equations without unit effects using a simple 2SLS-procedure, where the *Wefw* (*Wgov*, *Wreg* etc.) variables entered in changes, instrumented by the lagged levels. Results clearly confirmed our estimates.

<sup>9</sup> The system GMM estimator requires stationarity of the series. Inspection of Figure 1 indicates that nonstationarity may be a problem. A test suggested by Im, Pesaran, and Shin (2003) however shows that the time series are trend stationary. Hence, we include a deterministic time trend in the regressions.

**Table 2** Determinants of overall economic reform

Weights	Dependent variable: overall economic freedom (efw)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
efw ( $t-1$ )	0.433 (0.000)	0.464 (0.000)	0.438 (0.000)	0.438 (0.000)	0.386 (0.000)	0.469 (0.000)	0.474 (0.000)	0.450 (0.000)
Wefw	0.781 (0.000)	1.007 (0.001)	0.192 (0.171)	0.310 (0.024)	0.581 (0.000)	0.555 (0.010)	0.886 (0.000)	0.789 (0.000)
GPC ( $t-1$ )	0.374 (0.030)	0.109 (0.564)	0.358 (0.087)	0.159 (0.438)	0.245 (0.195)	0.293 (0.083)	0.293 (0.062)	0.345 (0.041)
growth ( $t-1$ )	-0.067 (0.067)	-0.061 (0.055)	-0.084 (0.029)	-0.084 (0.009)	-0.084 (0.005)	-0.062 (0.061)	-0.079 (0.033)	-0.069 (0.049)
open ( $t-1$ )	0.246 (0.075)	0.283 (0.064)	0.312 (0.029)	0.287 (0.050)	0.440 (0.003)	0.237 (0.059)	0.242 (0.031)	0.232 (0.048)
polcon	1.730 (0.000)	1.554 (0.000)	1.562 (0.000)	1.545 (0.000)	1.326 (0.001)	1.780 (0.000)	1.666 (0.000)	1.727 (0.000)
govleft	-0.282 (0.000)	-0.247 (0.000)	-0.325 (0.000)	-0.315 (0.000)	-0.260 (0.000)	-0.300 (0.000)	-0.289 (0.000)	-0.296 (0.000)
enopp	-0.072 (0.007)	-0.071 (0.009)	-0.076 (0.009)	-0.061 (0.015)	-0.007 (0.837)	-0.080 (0.002)	-0.070 (0.005)	-0.069 (0.008)
EU	-0.063 (0.645)	-0.181 (0.182)	-0.026 (0.866)	-0.089 (0.554)	-0.012 (0.935)	-0.079 (0.545)	-0.053 (0.692)	-0.061 (0.646)
EFTA	-0.095 (0.488)	-0.095 (0.478)	-0.056 (0.704)	-0.127 (0.329)	-0.016 (0.915)	-0.077 (0.541)	-0.073 (0.604)	-0.083 (0.532)
NAFTA	-0.259 (0.015)	-0.338 (0.001)	-0.236 (0.043)	-0.189 (0.135)	-0.370 (0.030)	-0.174 (0.144)	-0.242 (0.039)	-0.269 (0.006)
APEC	0.181 (0.537)	-0.025 (0.920)	0.238 (0.384)	0.197 (0.477)	0.080 (0.745)	0.135 (0.636)	0.193 (0.505)	0.182 (0.527)
Obs./Countries	134/23	134/23	134/23	134/23	134/23	134/23	134/23	134/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(1.000)	(0.999)	(1.000)	(1.000)	(1.000)	(0.998)	(1.000)	(1.000)
AR2-test	(0.674)	(0.869)	(0.539)	(0.471)	(0.943)	(0.910)	(0.409)	(0.701)

*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend and a constant (not reported). Robust *P*-values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

to reform. The impact of slow growth on economic liberalization is however weak, as a 1 percentage point decrease in the growth rate leads to an increase of the *efw*-score of only 0.07 points in the following period. We also find that (lagged) openness (*open*) is positively associated with reform.

As regards political variables the estimates show a robust negative relation between left wing-cabinets (*govleft*) and liberalization, confirming our expectations. Procedural restraints on executive action are positively related to reform in all equations, indicated by highly significant positive signs of *polcon*. The effect is not too strong, however. A one standard deviation increase of *polcon* in our sample (0.08) leads to a higher *efw*-score of about 0.12. Dropping four outlier observations for Greece and Spain does not change the results. In contrast, legislative fragmentation (*enopp*) is almost always negatively associated with liberalization at a 1%-level. This result confirms expectations of bargaining theories of reform

**Table 3** Determinants of reform in the 'size of government' area

Weights	Dependent variable: size of government (gov)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
gov ( $t-1$ )	0.663 (0.000)	0.736 (0.000)	0.705 (0.000)	0.699 (0.000)	0.647 (0.000)	0.713 (0.000)	0.639 (0.000)	0.744 (0.000)
Wgov	0.073 (0.742)	-0.135 (0.596)	0.014 (0.900)	0.128 (0.466)	0.233 (0.068)	-0.158 (0.476)	0.126 (0.577)	-0.056 (0.830)
GPC ( $t-1$ )	-0.386 (0.333)	-0.683 (0.036)	-0.546 (0.141)	-0.478 (0.168)	-0.522 (0.159)	-0.229 (0.460)	-0.161 (0.682)	-0.342 (0.308)
growth ( $t-1$ )	-0.144 (0.000)	-0.145 (0.000)	-0.144 (0.000)	-0.116 (0.003)	-0.126 (0.000)	-0.140 (0.000)	-0.129 (0.007)	-0.153 (0.000)
open ( $t-1$ )	0.110 (0.584)	0.263 (0.045)	0.135 (0.369)	0.169 (0.284)	0.269 (0.138)	-0.016 (0.928)	0.035 (0.862)	0.237 (0.013)
polcon	1.524 (0.008)	2.089 (0.003)	1.796 (0.007)	1.653 (0.002)	1.258 (0.018)	1.650 (0.004)	1.219 (0.022)	1.500 (0.004)
govleft	-0.400 (0.003)	-0.363 (0.009)	-0.398 (0.006)	-0.362 (0.008)	-0.355 (0.016)	-0.345 (0.019)	-0.423 (0.002)	-0.370 (0.013)
enopp	-0.099 (0.050)	-0.091 (0.089)	-0.093 (0.060)	-0.092 (0.067)	-0.059 (0.349)	-0.100 (0.022)	-0.104 (0.038)	-0.118 (0.008)
Obs./Countries	134/23	134/23	134/23	134/23	134/23	134/23	134/23	134/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(1.000)	(0.992)	(0.998)	(1.000)	(1.000)	(0.990)	(0.998)	(0.874)
AR2-test	(0.944)	(0.946)	(0.945)	(0.887)	(0.807)	(0.954)	(0.975)	(0.937)

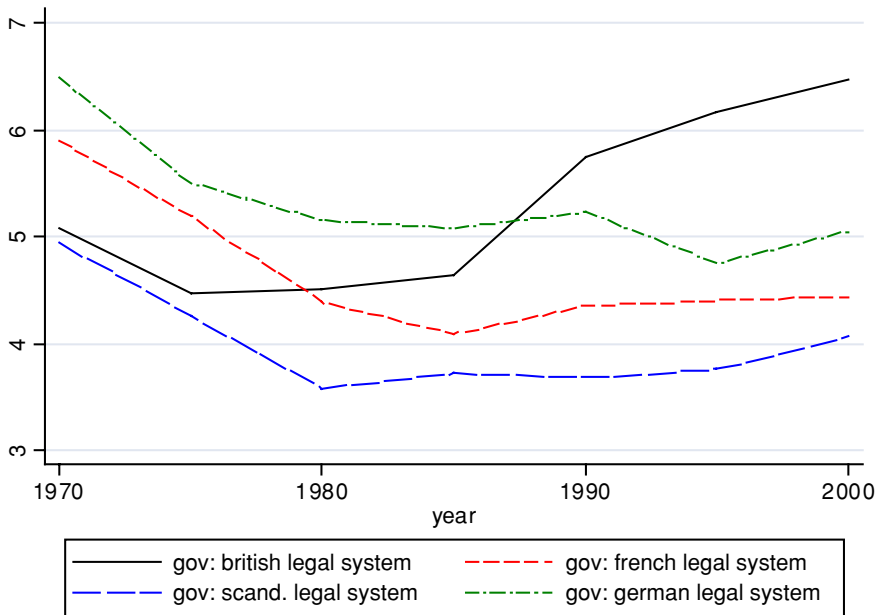
*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend, trade bloc-dummies and a constant (not reported). Robust  $P$ -values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

(Alesina & Drazen, 1991). The only exception is model (5), where weights are generated according to legal tradition. Seeing that legislative fragmentation is highly correlated with the voting system and legal history, this may be explained by collinearity. Note that these are not necessarily contradictory findings, as *polcon* is primarily a measure of policy credibility determined by several political institutions, whereas *enopp* depicts the impact of transaction cost only in legislative bargaining. Hence, if some policies are not in the major responsibility of the legislature, we should not expect *enopp* to be related to economic liberalization in that certain policy field (see below).

Finally, throughout all specifications no impact of membership in EC, EFTA, NAFTA, or APEC on overall liberalization can be found. We checked whether results change if EC-and EFTA-membership are not separated, but found no effects. Even if the weighted policy scores are eliminated, no impact can be found. Hence, it can be concluded that a coordination of policies in these organizations has not been the driving force of parallel economic reforms.

#### 4.3 Results for separate policy areas

In the next step we investigate whether these results hold for all five policy areas of the *efw*-index. As noted above in Section 3, from the theoretical models of policy diffusion it is not quite clear that competition among states unambiguously causes a reduction of government intervention and a deregulation of markets. In contrast, the resource flow model implies



**Fig. 2** Government size-score in different legal traditions, 1970–2000

that policy interconnectedness is pronounced in liberalization of trade in goods and factors, monetary stability and security of property rights.<sup>10</sup>

Considering government size (*gov*), it should be noted first that we observe no common trend of liberalization over the period 1970–2000. In 1970, the average *gov*-score of the 23 nations is 5.6, and it declined to 5.0 in 2000. The coefficient of variation remained stable (0.29). Table 3 shows results of system GMM-estimations.

Coefficients of  $W_{gov}$  are never significant. Only in model (5), with weights generated by similarity of legal history, a positive impact can be found. Figure 2 shows underlying developments in countries with different legal traditions. On average, states with a British tradition improved on the ‘government size’-score, while countries with other legal traditions did worse in 2000 as compared to 1970.

As expected, the lagged dependent variable is positive and highly significant, indicating a considerable policy persistence. Estimates also show that GDP per capita is negatively related to *gov*, indicating that an increase in per capita income leads to an expansion of government size. Yet, the results are typically insignificant. No stable impact of trade openness *open* can be found, but slow pre-period *growth* appears to facilitate a reduction of government size. A stable and significant impact of left wing cabinets on government size is also in line with theoretical predictions. The impact of *polcon* and *enopp* on government size shows a similar pattern as for the composite *efw*-index. Institutional constraints for the executive and a limited effective number of legislative parties reduce government size. Membership in an

<sup>10</sup> As one referee criticized, an aggregation of policy categories into a composite index is therefore somewhat suspicious. Although we generally agree, we are not aware of a superior measure to assess the overall quality of a bundle of policies.

**Table 4** Determinants of reform in the regulation policy area

	Dependent variable: regulation policy (reg)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
reg ( $t-1$ )	0.568 (0.000)	0.532 (0.000)	0.631 (0.000)	0.510 (0.000)	0.438 (0.000)	0.596 (0.000)	0.606 (0.000)	0.579 (0.000)
wreg	0.839 (0.000)	0.840 (0.000)	0.457 (0.003)	0.600 (0.000)	0.579 (0.000)	0.757 (0.000)	0.848 (0.000)	0.813 (0.000)
GPC ( $t-1$ )	0.552 (0.024)	0.569 (0.044)	0.714 (0.009)	0.325 (0.375)	0.347 (0.239)	0.727 (0.001)	0.483 (0.049)	0.508 (0.031)
growth ( $t-1$ )	-0.043 (0.228)	-0.040 (0.259)	-0.038 (0.290)	-0.042 (0.149)	-0.010 (0.762)	-0.046 (0.110)	-0.053 (0.116)	-0.024 (0.414)
open ( $t-1$ )	0.132 (0.514)	0.049 (0.793)	-0.046 (0.754)	0.209 (0.357)	0.238 (0.348)	0.005 (0.972)	0.088 (0.585)	0.143 (0.284)
polcon	0.511 (0.194)	0.082 (0.841)	0.330 (0.502)	0.452 (0.235)	0.276 (0.572)	0.130 (0.774)	0.425 (0.256)	0.527 (0.163)
govleft	-0.262 (0.042)	-0.257 (0.031)	-0.194 (0.181)	-0.230 (0.068)	-0.190 (0.047)	-0.264 (0.044)	-0.255 (0.036)	-0.250 (0.047)
enopp	-0.092 (0.009)	-0.090 (0.002)	-0.083 (0.053)	-0.096 (0.012)	-0.004 (0.885)	-0.096 (0.020)	-0.083 (0.020)	-0.093 (0.009)
Obs./Countries	133/23	133/23	133/23	133/23	133/23	133/23	133/23	133/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(0.337)	(0.962)	(0.467)	(0.729)	(0.183)	(0.564)	(0.637)	(0.138)
AR2-test	(0.569)	(0.570)	(0.918)	(0.689)	(0.676)	(0.500)	(0.178)	(0.518)

*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend, trade bloc-dummies and a constant (not reported). Robust *P*-values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

international organization is in no case associated with liberalization, so we refrain from reporting results.

Estimates for government regulation (*reg*) are reported in Table 4. We find a high level of policy persistence as indicated by the lagged regulation index. In all specifications the weighted *reg*-scores show the expected positive sign at a high level of confidence. The coefficients are, yet, somewhat lower in models (3), (4), and (5). The per capita GDP level (GPC) shows a positive relation to deregulation, while *growth* is negatively associated to *reg*, though not significant. We find again a positive impact of executive constraints on reforms, but *polcon* is not significant at conventional levels. A left wing orientation of government and a high degree of legislative fragmentation are negatively related to deregulation of markets. Membership in EU/EC, EFTA, NAFTA, or APEC is not associated with *reg* (results not shown).

Table 5 reports results for trade liberalization (*trade*). In these estimates we disregard the openness variable. We find a highly significant positive coefficient of *Wtrade* when employing equal weights (column (1)), geographical distance weights (2), GDP weights (6), growth weights (7) and GDP per capita-weights (8). Weighted by GDP difference, per capita GDP, and legal tradition estimates show no significant relation to trade liberalization. Hence, relative success, relative country size and geographical proximity matter most for governments when deciding to do away with trade barriers.

**Table 5** Determinants of reform in the trade policy area

	Dependent variable: trade liberalization (trade)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
trade ( $t-1$ )	0.674 (0.000)	0.639 (0.000)	0.567 (0.000)	0.597 (0.000)	0.613 (0.000)	0.637 (0.000)	0.624 (0.000)	0.668 (0.000)
wtrade	1.190 (0.000)	1.050 (0.000)	-0.127 (0.524)	0.061 (0.752)	0.339 (0.116)	0.918 (0.006)	1.234 (0.000)	1.118 (0.000)
GPC ( $t-1$ )	-0.090 (0.793)	-0.121 (0.707)	-0.000 (1.000)	-0.140 (0.683)	0.036 (0.915)	0.028 (0.939)	-0.029 (0.928)	-0.079 (0.784)
growth ( $t-1$ )	0.003 (0.931)	0.016 (0.594)	-0.034 (0.320)	-0.017 (0.625)	-0.021 (0.520)	0.020 (0.555)	-0.017 (0.593)	0.005 (0.865)
polcon	1.641 (0.003)	1.647 (0.002)	1.502 (0.009)	1.617 (0.002)	1.559 (0.005)	1.714 (0.004)	1.705 (0.001)	1.575 (0.008)
govleft	0.018 (0.889)	0.075 (0.546)	-0.028 (0.821)	-0.092 (0.446)	0.059 (0.640)	0.052 (0.658)	0.014 (0.917)	0.022 (0.863)
enopp	-0.046 (0.156)	-0.071 (0.019)	-0.034 (0.360)	-0.032 (0.283)	-0.025 (0.345)	-0.058 (0.089)	-0.042 (0.197)	-0.045 (0.166)
Obs./Countries	134/23	134/23	134/23	134/23	134/23	134/23	134/23	134/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(0.920)	(0.932)	(0.998)	(0.521)	(0.720)	(0.700)	(0.969)	(0.871)
AR2-test	(0.666)	(0.775)	(0.170)	(0.296)	(0.545)	(0.973)	(0.720)	(0.726)

*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend, trade bloc-dummies and a constant (not reported). Robust  $P$ -values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

GDP level and GDP growth rates are not related to trade liberalization. Though the sign of *enopp* is always negative, it is considerably smaller than in the previous estimates, and not always significant. We also find no influence of the ideological orientation of the cabinet on *trade*. In contrast, executive constraints measured by *polcon* are always significant and positively related to freedom of international trade. We also observe a small positive impact of EC membership on trade liberalization, which is near a 10%-level of confidence in some specifications. According to our estimates, membership in EFTA, NAFTA, and APEC is not related to more liberal trade policies (not shown). This surprising result holds if weighted *trade-score* are eliminated from the regressions.

Determinants of reform in monetary policies (*money*) are reported in Table 6. In all equations *Wmoney* is positively related to monetary reforms. Our results show clear evidence in favor of policy diffusion as regards monetary stability. An exceptionally high value of *Wmoney* in the GDP weighted model (6) additionally points to a special importance of monetary policies in the U.S., Germany, and Japan for other countries, supporting the notion of political leadership and hegemony.

Per capita GDP level and GDP growth do not seem to matter for monetary policy reforms. International trade openness is positively related to *money* but not at usual levels of significance. As could possibly be expected, legislative fragmentation and government ideology are also irrelevant for monetary policy in our sample. Constraints for executive action are always significant and positively related to *money*, confirming the importance of institutionalized credible commitment for monetary policy, recently emphasized by Keefer and Stasavage (2003).



**Table 6** Determinants of reform in the monetary policy area

	Dependent variable: monetary policy (money)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
money ( $t-1$ )	0.601 (0.000)	0.603 (0.000)	0.606 (0.000)	0.597 (0.000)	0.587 (0.000)	0.635 (0.000)	0.608 (0.000)	0.602 (0.000)
Wmoney	0.467 (0.068)	0.817 (0.023)	0.115 (0.336)	0.399 (0.035)	0.297 (0.025)	1.153 (0.023)	0.467 (0.024)	0.536 (0.043)
GPC ( $t-1$ )	0.557 (0.276)	-0.106 (0.850)	0.254 (0.622)	-0.068 (0.884)	0.101 (0.866)	0.428 (0.285)	0.598 (0.262)	0.570 (0.333)
growth ( $t-1$ )	-0.053 (0.563)	-0.065 (0.461)	-0.043 (0.616)	-0.041 (0.641)	-0.080 (0.380)	-0.062 (0.465)	-0.036 (0.688)	-0.053 (0.564)
open ( $t-1$ )	0.328 (0.288)	0.390 (0.180)	0.359 (0.150)	0.363 (0.155)	0.574 (0.046)	0.274 (0.175)	0.206 (0.399)	0.391 (0.192)
polcon	0.328 (0.288)	0.390 (0.180)	0.359 (0.150)	0.363 (0.155)	0.574 (0.046)	0.274 (0.175)	0.206 (0.399)	0.391 (0.192)
govleft	2.062 (0.005)	1.996 (0.006)	2.039 (0.001)	1.538 (0.077)	2.208 (0.009)	2.298 (0.003)	1.992 (0.003)	2.062 (0.008)
enopp	-0.101 (0.654)	-0.060 (0.775)	-0.144 (0.519)	-0.126 (0.506)	-0.101 (0.652)	-0.099 (0.662)	-0.084 (0.711)	-0.126 (0.607)
Obs./Countries	134/23	134/23	134/23	134/23	134/23	134/23	134/23	134/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(0.957)	(0.218)	(0.145)	(0.219)	(0.660)	(0.314)	(0.828)	(0.918)
AR2-test	(0.596)	(0.506)	(0.815)	(0.674)	(0.817)	(0.381)	(0.529)	(0.527)

*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend, trade bloc-dummies and a constant (not reported). Robust  $P$ -values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

What makes reforms likely when it comes to reforms in legal structures (*legal*)? Table 7 shows regression results. We find a positive relation with *Wlegal* in all specifications, which is especially strong with geographical proximity and GDP size based weights. Higher per capita GDP and lower GDP growth rates are positively associated with improvements in this policy area. We also find a significant positive sign of *polcon*, indicating that restrictions on executive action go hand in hand with security of property rights and higher quality of legal structures. A negative sign of *govleft* and *enopp* is surprising, because we do not have a stringent theory relating ideology and legislative transactions cost to the quality of legal systems. As before, membership in international organizations is not related to improvements. Note however that the autoregressive parameter has no explanatory power.

## 5 Summary and conclusion

The present paper examined to what extent domestic political and economic factors, and international policy diffusion via competitive interaction of governments are responsible for a liberalization of economic policies in OECD. As regards overall liberalization, captured by a comprehensive freedom index, results of dynamic panel estimates show that there is a strong and significant positive interdependence of policy reforms between countries in the sample. The findings suggest that geographical proximity and learning from success have a

**Table 7** Determinants of reform in the legal structure area

Weights	Dependent variable: legal structure and security of property rights (legal)							
	(1) Equal	(2) $\Delta$ GEO	(3) $\Delta$ GDP	(4) $\Delta$ GPC	(5) $\Delta$ LEG	(6) GDP	(7) GW	(8) GPC
legal ( $t-1$ )	0.167 (0.308)	0.178 (0.202)	0.026 (0.833)	0.080 (0.475)	0.067 (0.548)	0.128 (0.402)	0.154 (0.326)	0.165 (0.311)
wlegal	0.741 (0.017)	0.936 (0.000)	0.083 (0.667)	0.467 (0.022)	0.580 (0.001)	1.090 (0.057)	0.577 (0.018)	0.813 (0.020)
GPC ( $t-1$ )	1.021 (0.014)	0.625 (0.142)	1.310 (0.003)	0.966 (0.024)	1.072 (0.021)	1.388 (0.007)	1.014 (0.011)	0.941 (0.017)
growth ( $t-1$ )	-0.094 (0.091)	-0.076 (0.086)	-0.103 (0.056)	-0.144 (0.000)	-0.035 (0.504)	-0.102 (0.081)	-0.119 (0.017)	-0.100 (0.059)
polcon	0.541 (0.077)	0.443 (0.101)	0.694 (0.065)	0.755 (0.020)	0.815 (0.022)	0.475 (0.112)	0.546 (0.076)	0.575 (0.051)
govleft	1.920 (0.010)	2.391 (0.003)	1.650 (0.019)	1.484 (0.057)	1.780 (0.022)	2.066 (0.034)	1.891 (0.009)	2.022 (0.010)
enopp	-0.586 (0.002)	-0.499 (0.005)	-0.629 (0.003)	-0.625 (0.001)	-0.479 (0.012)	-0.623 (0.002)	-0.593 (0.001)	-0.606 (0.001)
Obs./Countries	132/23	132/23	132/23	132/23	132/23	132/23	132/23	132/23
Hansen-test	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)	(1.000)
DIF-Sargan	(0.455)	(0.219)	(0.802)	(0.724)	(0.803)	(0.990)	(0.604)	(0.970)
AR2-test	(0.353)	(0.409)	(0.246)	(0.440)	(0.688)	(0.535)	(0.331)	(0.350)

*Note.* Results obtained from one step-system GMM estimator. All regressions include a deterministic time trend, trade bloc-dummies and a constant (not reported). Robust *P*-values in parentheses. With the exception of *enopp*, *govleft* and the trade bloc-variables, all covariates are treated as endogenous

powerful impact on reform decisions. A decomposition of the economic freedom index into its sub-components shows that interconnectedness is important in trade policies, regulatory policies and monetary policies. With respect to government size interdependence appears to be much weaker. This is surprising, since most theories suggest that competition among states leads to a downsizing of the public sector. A further unexpected result is that membership in international and supranational organizations appears to have no impact on liberalization. As Rose (2004, p. 230) notes, these institutions might matter for economic liberalization independent of membership.

Our estimates, admittedly, cannot discriminate between different sources of economic policy diffusion. The impact of geographical proximity, for example, can be traced back to both a resource flow model and an information spillover model. To separate the impact of competition for mobile factors from yardstick competition, we need to know whether internationally mobile factors of production are attracted by (a combination of) certain policies, and to which extent domestic election results and voting-popularity functions are influenced by policy decisions in other countries. This requires a lot of information and sophisticated empirical analyses, which is well beyond the scope of the present paper. It is even more difficult to test empirically the impact of a spread of certain economic ideas and the influence of policy experts on political decision making. Thus, a lot of further work needs to be done in order to assess the presumably different impact of competition for mobile factors and yardstick competition on economic policy reform decisions.

## Appendix: Summary statistics and data sources

Variable	Obs	Mean	Std. Dev.	Min	Max	Source
<i>efw</i>	134	6.81	0.91	4.2	8.6	Gwartney/Lawson
<i>gov</i>	134	4.65	1.30	2.1	7.6	Gwartney/Lawson
<i>legal</i>	133	7.73	1.19	3.9	9.6	Gwartney/Lawson
<i>money</i>	134	8.32	1.68	2.0	9.8	Gwartney/Lawson
<i>trade</i>	134	7.55	0.90	3.5	9.1	Gwartney/Lawson
<i>reg</i>	134	6.14	0.91	4.4	8.8	Gwartney/Lawson
<i>enopp</i>	134	4.17	1.52	2.00	9.84	Armingeon et al.
<i>growth</i>	134	2.24	1.49	−1.51	8.48	Heston et al.
<i>open</i>	134	0.68	0.40	0.14	2.13	World Bank
<i>govleft</i>	134	0.34	0.32	0	1	World Bank
<i>GPC (log)</i>	134	9.82	0.27	8.75	9.05	Heston et al.
<i>polcon</i>	134	0.77	0.08	0.35	0.89	Henisz
<i>EU</i>	134	0.49	0.50	0	1	
<i>EFTA</i>	134	0.28	0.44	0	1	
<i>NAFTA</i>	134	0.03	0.17	0	1	
<i>APEC</i>	134	0.07	0.25	0	1	

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