Economic Freedom and Institutional Convergence

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Abstract

Fukuyama (1992) argues that liberal democracy is the "final form of human government" and will become more and more prevalent in the long term. If this prediction is true, countries should converge in their political and economic systems toward liberal democracy, i.e. institutional convergence. In this paper, we examine whether there is economic institutional convergence using the Economic Freedom of the World-index (EFW) over the period 1970-2005. We find that countries with lower institutional quality experience faster institutional change than countries with higher quality, i.e., we observe institutional convergence. However, countries with lower institutional quality have higher variability of institutional change. Using distributional analysis, we examine institutional transition probabilities, and find that the probability of ending up with high-quality institutions is high in the long-run. These findings support Fukuyama's prediction.

Keywords: Economic Freedom; Institutional Convergence; The Law of Proportionate Effect

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1 Introduction

Fukuyama (1989, 1992) argues that the advent of Western liberal democracy may signal the end point of humanity's sociocultural evolution and the arrival of the final form of human government:

"What we may be witnessing is not just the end of the Cold War, or the passing of a particular period of post-war history, but the end of history as such: that is, the end point of mankind's ideological evolution and the universalization of Western liberal democracy as the final form of human government." (Fukuyama 1992)

A conservative intepretation of Fukuyama's argument is that liberal democracy will become more and more prevalent as form of government in the long term, although the development may suffer temporary setbacks. Yet for this prediction to have any merit, there should be an observable catching-up effect across the globe. Can we observe such institutional convergence, i.e. that countries converge in their political and economic systems towards liberal democracy? The purpose of this paper is to answer this question with regards to a subeset of the institutions underlying liberal democracy, namely the economic institutions.

The term Washington Consensus has come to signify a general orientation in recent decades among governments towards a strongly market-based approach, supposedly promoted by international organizations such as the International Monetary Fund (IMF), and the World Bank, and the enforcement of various multilateral free-trade agreements (Chang, 2010). Yet the dominant view today, that economic institutions are the ultimate determinants of economic performance (cf. Acemoglu et al 2005, North, 2006), is an echo of that of the classical economists, who argued that market-oriented policies were important to foster economic growth (de Haan et al, 2006). In the economics discipline, the recent surge in interest regarding economic institutions is largely thanks to the introduction of several indices that quantify institutional quality, with respect to e.g. corruption and economic efficiency.

The Economic Freedom of the World-index (EFW) is arguably one of the most employed indices measuring the quality of a country's economic institutions. The evidence suggests a positive relationship between economic institutional quality, quantified by EFW, and important variables such as economic growth.

In this paper we employ EFW to answer the following two questions: (i) Do we observe economic institutional convergence (henceforth institutional convergence), i.e. do countries with economic institutions of relatively poor quality at the outset experience faster institutional change than countries with high quality institutions? (ii) Furthermore, if there is institutional convergence, what are its long run implications?

We answered the first of these questions by testing whether the Law of Proportionate Effect (LPE) holds for EFW. LPE is a well-known empirical law in the industrial organization literature. In its most basic form, the law entails that growth in an entity should be independent of its size. In this context, there should be independence between changes in EFW and the level of EFW if the law holds. We focused on three formal propositions both necessary and sufficient for the law to hold: (A) That changes in the EFW index are independent of its level. If this proposition holds, countries ranked with lower institutional quality have equal probability to experience institutional change as countries with higher institutional quality, and we do not observe institutional convergence; (B) That index changes over two consecutive periods are independent of each other. This concerns the stickiness of institutional change. If it holds, then changes in one period will not influence the prospect of further changes in following periods; (C) That the variability of changes are independent of the index level. This proposition concerns the magnitude of institutional change. If it holds, countries have the same variability in institutional change irrespective of their institutional quality.

We tested LPE on EFW for all available countries on a five-year basis over the time-period 1970-2005. For the overall sample, proposition (A) was rejected, suggesting that we do observe institutional convergence. This convergence seems to have commenced in the latter part of the 1980s. Proposition (B) was not rejected, i.e. we do not find evidence of a negative feedback-loop on institutional change. Proposition (C) was rejected, suggesting that countries with lower EFW-levels are likely to see more turbulence in their pace of reform, which may hamper the convergence process.

We turned to the second question, and tried to assess the long-run implications of the observed convergence. To do this, we employed distributional analysis. The steady state distribution, implied by a Markov Chain model, showed that the long-run probability of a country having high institutional quality is quite high, while that of having low institutional quality is low. As for the time of convergence, it is far from infinite: countries have a fairly high probability of substantially improving their institutional quality in a matter of decades. As such, this paper provides support for Fukuyama's prediction for that subset of instititutions underlying liberal democracy that are labeled economic.

The rest of this paper is structured as follows: In section 2, we discuss the literature on EFW. In section 3 we outline a theoretical framework to explain various outcomes from testing LPE in the context of institutional change.

Section 4 presents data and methodology, while in section 5 we answer the two questions empirically. In section 6 we summarize and draw conclusions.

2 The Economic Freedom of the World-Index

The most widely used index measuring institutional quality is The Economic Freedom of the World-index (EFW), jointly published by the Fraser Institute and the Cato Institute.[1] The first major publication of EFW was made by Gwartney et al. (1996). Today, the index has data points for every five years from 1970 to 2000, and annual data 2001-2009. The most recent edition covers 141 countries. However, most countries do not have time series stretching all the way back; only 54 countries have index-values in 1970.

EFW is a comprehensive measure for the institutional quality with respect to a functioning market economy. It is the unweighted average of five components, reflecting a country's institutional quality with respect to: (i) Size of Government: Expenditures, Taxes, and Enterprises (henceforth government); (ii) Legal Structure and Security of Property Rights (henceforth legal); (iii) Access to Sound Money (henceforth money); (iv) Freedom to Trade Internationally (henceforth trade); (v) Regulation of Credit, Labor, and Business (henceforth regulation).

These five components are in turn constructed from several subcomponents, in total 42 in the most recent edition. EFW is normalized on a scale from 0 to 10, where higher values reflect better institutional quality. Any increase in EFW can be interpreted as an institutional change in a free-market direction, while a decrease is an institutional change in the opposite direction (Pitlik, 2011).

The evidence points to a positive effect of economic freedom on variables such as wealth and economic growth (Berggren 2003; Doucouliagos and Ulubasoglu 2006) and that institutional change in a free-market direction stimulates economic growth (de Haan et al. 2006).[2]

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[1] The Index of Economic Freedom (IEF), compiled by The Heritage Foundation and the Wall Street Journal, is similar but less widely used. EFW has received the most attention in economic literature, probably because it goes as far back in time as the 1970s. As Berggren (2003) notes, the two indices are very similar in their overall implications.

[2] EFW has nonetheless been critically scrutinized in a number of studies. The relationship between EFW and growth has been found to depend on the measure used (de Haan and Siermann 1997); the weighting procedure in the construction of EFW is said to be arbitrary (Heckelman and Stroup 2000); their has been criticism of the inclusion of some subcomponents, or of the manner of this inclusion (de Haan and Sturm, 2000; Heckelman and Sturm 2000; Leschke 2000); the underlying methodology of many studies has been critizised, no-

Table 1: Descriptive statistics of efw during the period 1970-2005.

Efw	Obs	Mean	Std.Dev.	Min	Max	Kurt.	Skew.
1970	54	6.113	1.047	3.9	8.4	2.24	-0.136
1975	72	5.506	1.078	3.4	8.3	2.859	0.216
1980	104	5.475	1.081	3.2	8.6	2.865	0.202
1985	110	5.488	1.158	2.3	8.3	3.003	-0.0728
1990	113	5.667	1.176	2.9	8.2	2.632	-0.0912
1995	123	6.046	1.22	3.4	9.1	2.481	0.0472
2000	123	6.468	0.99	4	8.8	2.744	-0.0748
2005	141	6.646	0.967	3.4	9	3.946	-0.641

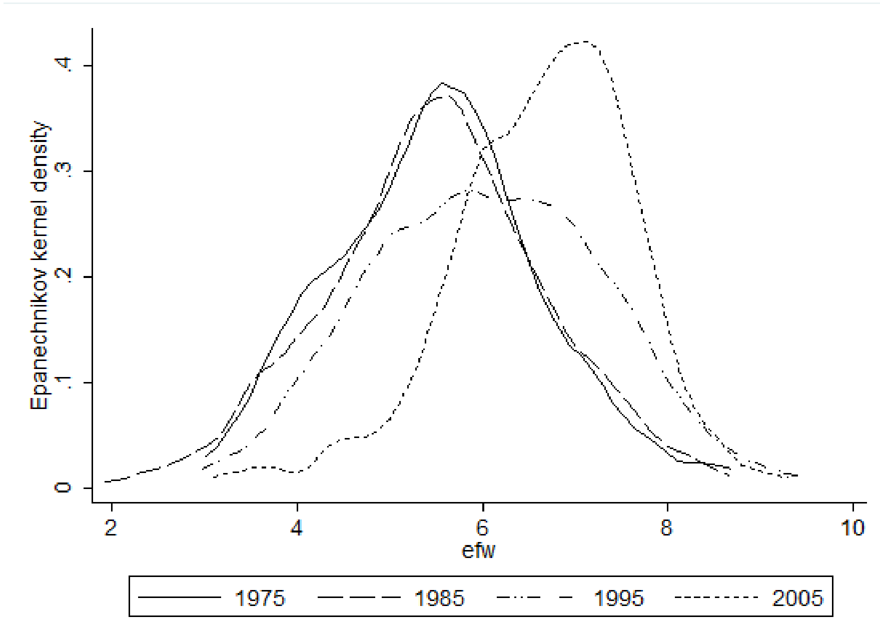
De Haan et al (2006, p.180-181) summarize twelve studies that try to explain (changes in) the EFW in terms of different independent variables. Most find a positive effect from democracy but Lundström (2003) argues that it depends on what components you examine. It appears that low economic growth stimulates economic reform (Gwartney et al, 2004, De Haan and Sturm, 2003, Pitlik and Wirth 2003, Pitlik 2011). Two studies find a positive relationship between foreign aid and EFW (De Haan and Sturm 2003, Pitlik and Wirth 2003), but Heckelmann and Knack (2005) find that upon disaggregation, aid turns out to have a negative effect on some components.

To our knowledge, little attention in the EFW-literature has thus far been given to the question of how previous institutional quality fosters further institutional change. A mere look at the data nonetheless suggests a trend that could be interpreted as institutional convergence. Table 1 shows the descriptive statistics for EFW 1970-2005. While the mean has steadily increased since 1980, the standard deviation increased until 1995 after which it has declined. In 1975 and 1980 the EFW-distribution was right-skewed. Thereafter, the distribution has mainly been left-skewed (1995 is an exception).

The descriptive statistics suggest that at least in the past 10 years some form convergence has occured. Figure 1, showing non-parametric kernel den-

tably the assumption of a linear relationship between EFW and growth (Doucouliagos and Ulubasoglu 2006); the relationship between EFW and growth is not robust to alternative aggregation methods of the index (Heckelman and Stroup 2005, Lundström 2003); the relationship between various subcomponents of the index and economic growth has been found to differ (Heckelmann and Stroup 2000, Carlsson and Lundström, 2002) a clear publication bias has been found in the literature concerning the relationship between EFW and growth (Doucouliagos 2005); it has been found that since the overall index only represents the average for those areas for which data are available, it is measured inconsistently both across nations and over time (Heckelman and Knack 2008).

Figure 1: Kernel density plot of Economic freedom of the world



Note: The density is computed with the Epanechnikov kernel function evaluated at 50 points in 1975, 1985, 1995 and 2005. Here, the x-axis shows the non-normalized values of the the economic freedom of the world for all countries.

sity plots for the EFW at 10 year intervals from 1975 to 2005, supports this interpretation. Until 1985 the distribution appears static, but after that it has gradually become more left-skewed. Between 1985 and 1995, however, a number of countries experience fast institutional change, a development that continues into the 21st century.

This motivates our attempt to examine the convergence hypothesis for the EFW-index. However, de Haan et al (2006) assess that while some subcomponents of EFW are institutional measures in the sense of measuring the rules of the game, others have more of a policy character, and can even be said to measure outcomes of the game. As Glaeser et al (2004) state, an essential aspect to institutions is their durability. Property rights may for example change more slowly than freedom to trade. For this reason, we will also examine the convergence hypothesis separately for each subcomponent of EFW.

3 Theoretical framework

In the literature regarding economic growth, the convergence hypthesis states that the per capita income of poorer economies will grow at a faster rate than that of richer economies, eventually resulting in convergence of income across countries. The theoretical justification is supposedly greater diminishing returns in capital rich countries captured by the concept of β-convergence (Barro 1991; Barro and Sala-i-Martin 1992; Mankiw et al 1992; Gundlach 2007). In an authoritative survey, Sala-i-Martin (1996) concludes that there is a uniform β-convergence across countries of about 2% per year. Quah (1996 p.1046) states that the convergence hypothesis "directly applies [...] to convergence and growth in other economic units" as well. Therefore, the notion of β-convergence should be applicable to examine convergence in institutional quality across-countries.[3] The Solow model predicts growth-convergence conditional on a set of covariates (Romer and Weil 1992). In contrast, the analysis in this paper is based on unconditional convergence, as we are primarily interested in the dynamics of institutional quality in and of it self.

We use the so called Law of Proportionate Effect to test the institutional convergence hypothesis. LPE, or Gibrat's law, is not so much a law as an empirical regularity. It suggests that growth (for firms) is a purely random phenomenon and therefore should be independent of size (Gibrat, 1931). In other words LPE describes a situation where small and large entities share the same propensity to increase or decrease their proportional size. The law has received a great deal of interest in the international organization literature, as attested by several authoritative surveys (Sutton, 1997; Caves, 1998; Geroski, 1995; Lotti et al, 2003).

Applied to EFW, the law describes a situation where countries with high or low levels of institutional quality share the same propensity to increase or decrease their quality. The institutional LPE model can be formulated as

$$Y_{i,t} = AY_{i,t-1}^{\beta}, \tag{1}$$

where $Y_{i,t}$ refers to the index value of country$i$ in period t. In this simple model, institutional quality depends on previous quality in period $t-1$ up to a constant β. Here the parameter A captures unobservable time invariant factors affecting institutional quality. In a stochastic version of (1), A can be decomposed into $A_{i,t} = e^{(1-\beta)\alpha+\epsilon_{i,t}}$, now containing a deterministic component

[3] While borrowing this vocabulary to the field of institutional study, we should also note that institutions in themselves lie at the heart of the debate regarding economic convergence; institutions can be said to be a core prerequisite for economic convergence or divergence to occur (Olson, 1998).

$(1-\beta)\alpha$ and a random component $\epsilon_{i,t}$. Equation (1) can be linearized by taking logarithms

$$y_{i,t} = (1-\beta)\alpha + \beta y_{i,t-1} + \epsilon_{i,t}, \tag{2}$$

resulting in a stochastic difference equation with an AR(1) structure, where $y_{i,t}$ refers to the logarithm of EFW. The disturbance term denotes the logarithmic institutional change, from period $t-1$ to t. Note that (2) can be expressed as,

$$\Delta y_{i,t} = (1-\beta)\alpha + (\beta-1)y_{i,t-1} + \epsilon_{i,t}, \tag{3}$$

which inversely relates institutional change,$\Delta y_{i,t}$, with institutional quality in the previous time period, $y_{i,t-1}$, through $(\beta-1)$.

Following Del Monte et al (2003), we extend (3) by adding to the set of regressors the one period lagged rate of institutional change, $\Delta y_{i,t-1}$. This makes the model equivalent to the convergence model suggested by Dobson et al. (2011) that is used to analyze income convergence in developing countries. The model, which can be seen as a general data generating process for EFW, thus becomes,

$$\Delta y_{i,t} = (1-\beta)\alpha_i + \delta_t + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t.} \tag{4}$$

The parametrization in (4) allows for simultaneous testing of three formal propositions that, according to Tschoegl (1983), are both necessary and sufficient for LPE to hold. Below, we state the three propositions as testable hypotheses.

(A)[4] First, we test whether changes in the EFW-index are independent of its level

$$H_o^{(A)} : (\beta-1) = 0,$$
$$H_1^{(A)} : (\beta-1) < 1.$$

If the nullhypothesis is true, then institutional quality and institutional change can be said to be independent of each other. Institutional change is in

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Since the analysis for $\beta > 1$ is similar to when $\beta = 1$, we concentrate on the one sided test. In the case when $\beta > 1$ institutional change is explosive in the sense that institutional change increases with higher institutional quality. This development is, however, unlikely to continue for long.

other words equiproportional for countries with varying degrees of institutional quality under the nullhypothesis. If the nullhypothesis is rejected, institutional quality is taken to be converging.

Equiproportional growth across the size distribution of firms has been attributed to constant costs, i.e. the absense of scale economies (cf. Tschoegl, 1983 p.189). Thus under the null hypothesis, the political costs to perform institutional change should be constant, while under the alternative hypothesis we should see diseconomies of scale. The outcome could also be driven by some notion of historical dependence (cf. David 1985; Arthur 1988, 1989; Sewell 1996). Path dependence was briefly discussed with regards to EFW by Feng (2005, p.262), who states that if it exists, the lagged EFW variable should take a positive sign and be statistically significant, i.e. $\beta > 0$.[5] The outcome could also be driven by institutional competition (cf. Tiebout, 1956; Karlsson et al. 2007; Bergh and Höijer 2008; Ochel 2003). Assuming free movement of production factors, capital and labor should move where the institutional quality is better, thus prompting countries with low-quality institutions to make reforms.

(B) Secondly, we test whether changes over two consecutive periods are independent

$$H_o^{(B)} : \theta = 0,$$
$$H_1^{(B)} : \theta \neq 0.$$

If the nullhypothesis is true, then changes in one period will not influence the prospect of further changes in the following periods. Institutional changes from one period to another are thus stochastic. This basically means that there is no persistence of reform, and that autocorrelation, $cov(\Delta y_{i,t}, \Delta y_{i,t-1})$, is zero. By contrast, a positive value for θ implies positive autocorrelation and persistence, and vice versa.

Why some institutional changes are more likely to take hold than others can be explained by increasing returns, increasing the probability of further steps

[5] Page (2006) defines four different kinds of historical dependence, of which three can be can be associated with the parameters parameters α and β taking different values. Standard path-dependence refers to the property where the order of historical events matter for the configuration of current outcomes. Phat-dependence, in contrast, means that the *set* rather than the *order* of historical events impact current outcomes. In addition, so called early-dependence denotes a process where early events have a disproportionate impact on current outcomes (Arrow, 2000). Lastly, equilibrium-dependence. This is the case when more than one equilibria exists. If the null hypothesis holds, institutional change can be said to be phat- or early-dependent, while it should be standard path-dependent if the null hypothesis is rejected (cf. Freeman 2010).

along the same path with each move down that path so that some institutional arrangements become "locked-in" (Pierson 2000, Boettke et al 2008). If the nullhypothesis holds, past change is not locked-in, i.e. it has no impact on subsequent change.

If $\theta > 0$, institutional change is positively autocorrelated, i.e. advantages acquired in the previous period increase the probability that changes continue in the same direction in the next, while disadvantages acquired in the previous period create a negative feedback loop. Positive auto correlation should hence be evidence of a gradual adjustment path. If $\theta < 0$, changes in the previous period in one direction increase the probability for later changes in the opposite direction. Negative persistence could occur if institutional changes are associated with high fixed costs, so that the opposition becomes strong enough to have a reverse effect in the preceeding period (Pierson 2000). Negative autocorrelation should thus describe an oscillating adjustment path.

(C) Third, we test whether the variability of changes are independent of the index level

$$H_o^{(C)} : E[\epsilon_{i,t}^2] = \sigma^2,$$
$$H_1^{(C)} : E[\epsilon_{i,t}^2] = \sigma^2(i).$$

$\epsilon_{i,t}$ is assumed to be independent and identically distributed, where $E(\epsilon_{i,t}) = 0$ and $E(\epsilon_{i,t}^2) = \sigma_\epsilon^2$. The third proposition thus tests for homoskedastic disturbances. If the nullhypothesis is true, countries have the same variability in their institutional change regardless of their institutional quality. If the nullhypothesis is rejected, on the other hand, then countries starting from a lower institutional quality should see a more turbulent pace of change if the variance is negatively related to institutional quality, and a less turbulent pace of reform if the variance is positively related to institutional quality. One might expect that countries with lower institutional quality, once having commenced upon political reform, see a more turbulent pace of reform, possibly due to learning factors, compared to countries with higher institutional quality. It has been shown that volatile institutional change depresses growth even if the long-run trend is toward market liberalization (Pitlik 2002). Berggren et al (2010), using the political risk index ICRG, find that institutional instability is positively related to growth in rich countries, but negatively related to growth in poor countries. This underlines the importance of focusing on the third proposition.

4 Methodology

The following section describes our empirical strategy. We answer the first question - whether we observe institutional convergence - using regression analysis to test propositions (A)-(C) in section 3. The method will be outlined in section 4.1.

To answer the second question, i.e. the long run implications of any observed convergence, we use so-called Markov-chain analysis. This is because regression analysis may suffer from a type of "Galton's fallacy" (Quah 1993), i.e. we risk to falsely conclude that there is convergence in the cross-sectional distribution of countries, when in fact there is none. The method will be outlined in section 4.2.

4.1 Regression analysis

We estimate (4) by lest squares to test propositions (A) and (B). We test proposition (C) by applying a standard test for heteroskedasticity to the residual in each regression. The motivation for these choices is given below.

There are a number of problems when estimating dynamic models such as (4). First, different estimators are valid under $H_0^{(A)}$ and under $H_1^{(A)}$. When $\beta < 1$, the process y_t is stationary, whereas $\beta = 1$ is synonymous with y_t having a unit root. OLS has been found to be consistent and unbiased for $\beta = 1$ but not for $\beta < 1$, i.e. under mean stationarity. The opposite is the case for IV-estimators, such as the system GMM-estimator which requires the process to be stationary, and therefore breaks down when $\beta = 1$ (Breitung and Meyer 1994; Ribeiro 2007).

A second potential problem has to do with country fixed effects when $\beta < 1$. If the effects are heterogenous, i.e. $\sigma^2_{\alpha_i} > 0$, the least squares estimate is upward biased as the variation of α_i is absorbed in the error term, making $y_{i,t-1}$ correlated with $\epsilon_{i,t}$.[6] If $\sigma^2_{\alpha_i} > 0$, a full set of controls x_i can be included to capture the variance of α_i where $\alpha_i = \gamma x_i$. Alternatively, dummy variables can be introduced for each country, which is equivalent to using fixed effects. Contrary to the upward bias in OLS when $\beta < 1$, the FE-estimate exhibits a downward bias that only dissappears when $T \to \infty$, and can be as large 60 percent for short panels (Greene 2010, p.341).

Although the FE-estimator handles heterogenous country effects, it does not resolve the endogeneity problem (Greene 2010 p.245). Difference GMM- or system GMM-estimators use endogenous lags from more than two periods

[6]In fact, it can be shown that $\beta^{OLS} \to 1$ as $\sigma^2_{\alpha_i}$ becomes sufficiently large (Nickell, 1981), resulting in a potential type II-error.

as instruments and at the same time correct for heterogenous country effects (Arrelano and Bond 1991; Blundell and Bond 1998). The estimators are suitable for small T, large N panels provided $\beta < 1$. If $\beta = 1$, the correlation between $y_{i,t-1}$ and potential instruments like $y_{i,t-2}$ and $\Delta y_{i,t-2}$ is weak, which may result in severe finite sample bias (Bond 2002).

On the contrary, if $\beta = 1$ least squares is unbiased and has an assymtotically normally distributed t-statistic (Breitung and Meyer 1994). This holds even under the assumption of heterogeneity. This is because the country effects cancel out, i.e. $(1-\beta)\alpha_i = 0$ when $\beta = 1$.

As a matter of fact, even if OLS is likely to be biased when $\beta < 1$, under certain conditions its performance is superior to the alternative estimators. This has to do with the magnitude of $\sigma^2_{\alpha_i}$. As long as the variation of α_i is small, the local power of OLS has been found to be high and the asymptotic bias of OLS to decreases when $\beta \rightarrow 1$ provided the variances of α_i and $(y_{i,t-1} - \alpha)_i$ are of the same order (Madsen 2010). Hall and Mairesse (2005) find that the power of OLS is superior to alternative estimators provided that $\sigma^2_{\alpha_i} < \sigma^2_{\epsilon_{i,t}}$, which they argue is usually the case for short panels. Should the variance of $\sigma^2_{\alpha_i}$ on the other hand be large, OLS suffers from a loss of power, but so does also GMM. Even when $\beta < 1$, difference-GMM performs poorly if the ratio $\sigma^2_{\alpha_i}/\sigma^2_{\epsilon_{i,t}}$ is large (Bond 2002). In a Monte Carlo study by Bond et al (2002) they find that the rejection frequencies of OLS is higher than both difference GMM and system GMM despite large variance in α_i for values of β close to unity.[7]

Hence, we choose to use least squares to estimate (4), well aware of the potential bias from heterogeneity, because we see the problems with the other techniques as more troublesome. First, fixed effects estimators will also exhibit a bias, which is probably larger. Secondly, GMM assumes stationarity, i.e. $\beta < 1$, and can be discarded for this reason.

[7] In the recent convergence literature, a number of panel unit roots tests have been used to test the convergence hypothesis in (i). Tests like the ones suggested in e.g. Im et al. (2003) usually requires relative long time series and are therefore not suitable for our data.

We also attempted the Breitung and Meyer (1994) unit root test (extended in Dobson et al. 2011) that modifies (3) by subtracting the first observation $y_{i,0}$ from $y_{i,t-1}$. One advantage with this test is that it has good small sample properties even when country effects are heterogenous. However the properties of β^{BM} does not surpass those of least squares for panel attested by Goddard (2002) who found in a Monte Carlo simulation that OLS had more power despite large variance in μ_i when where T=5, . Given the upward bias for β^{OLS} and downward bias for β^{FE}, following Bond (2002) the correct estimator should therefore lie somewhere in between β^{OLS} and β^{FE}. When applied to EFW, the Breitung-Meyer estimateβ^{BM} was as expected found to be outside the interval $\beta^{FE} < \beta^{OLS} < \beta^{BM}$, and was therefore discarded.

The Breusch–Pagan test is used to test for heteroscedasticity in (4). It is a chi-square-test that examines whether the estimated variance of the residuals from the regression are dependent on the values of the independent variables.

4.2 Distributional analysis

To examine the long run implications of institutional change, we complement regression analysis by modeling intra-distributional dynamics across countries directly. In the convergence litterature such attempts have been made by modeling the distributional "law of motion". Follwing Quah (1996), in the simplest form, let $\Im_t$ denote the probability distribution of EFW_t for all countries in time t. Alluding to the autoregressive form in equation (1), consider the model

$$\Im_{t+1} = M \cdot \Im_t. \tag{5}$$

where M, is a stochastic matrix containing the probabilities for transitioning from $\Im_t$ to $\Im_{t+1}$. Appropriately discretized, (5) becomes a Markow chain, where the entries in M describes the probability of transitioning from a state j to a state k. If M is time invariant, (5) can be used to predict the steady state distribution of EFW_{t+s}, as $s \to \infty$. Taking the convolution with respect M, the steady state distribution of EFW is given by M^s,

$$\Im_{t+s} \quad = \quad = (M \cdot M \cdot M \cdot \ldots M) \cdot \Im_t = M^s \cdot \Im_t. \tag{6}$$

One merit with this approach is that it can address a variety of different intra-distributional dynamics, and therefore complement the more constrained regression analysis. We discretize M into a uniform grid of quartiles,

$$M_{j,k} = \begin{bmatrix} p_{1,1} & p_{1,2} & p_{1,3} & p_{1,4} \\ p_{2,1} & p_{2,2} & p_{2,3} & p_{2,4} \\ p_{3,1} & p_{3,2} & p_{3,3} & p_{3,4} \\ p_{4,1} & p_{4,2} & p_{4,3} & p_{4,4} \end{bmatrix}, \tag{7}$$

where $p_{j,k}= p(y_{i,t+1} = k|y_{i,t} = j)$ denotes the probability of transitioning from quartile j in period t to the bounds defined by quartile k in the next period.

5 Results

5.1 Regression results

Table 2 shows the result from estimating equation (4) for EFW over the period 1970-2005. To see how institutional convergence may have varied over time, the model is estimated for each five-year time-period. Since (4) contains a lagged term, it can only be estimated from the 1980-1975 time-period onwards.

With regards to proposition (A), the null hypothesis is rejected if the coefficient for $y_{i,t-1}$ is significantly lower than zero. This is the case for the regression on the overall sample in column 1. However, for the five-year estimates, the magnitude of the deviation differs substantially, and is only significant from 1995 onwards. Hence, the evidence suggests that the convergence for the overall sample is attributable to the development in the second half of the time-period.

The null hypothesis for proposition (B), that changes in the previous period have no impact on changes on changes in the current period, is rejected if the coefficient for $\triangle y_{i,t-1}$is negative and statistically significant. It is however not significant for the overall sample, and only the significant in the first half of the time-period. These results suggest that persistence has disappeared in recent times. Hence countries that increase their institutional quality fast are no more likely to face positive or negative feedback loops than countires that increase their institutional quality more slowly.

Regarding proposition (C), we run the Breusch-Pagan/Cook-Weisberg test for heteroskedasticity and conclude that the null hypothesis of constant variance is generally rejected. This implies that variability of reform should be higher for countries with lower institutional quality.

As mentioned, we also tested LPE separately for each of the five subindices of EFW. Results are available in tables 4-8 in the appendix. Regarding proposition (A), the nullhypothesis is rejected for the overall sample for all five subindeces, but with differences in magnitude, where (i) government and (iii) money see a notably faster convergence than the other three subcomponents. As for (B), the nullhypothesis cannot be rejected for (iii) money and (v) regulation, while it is negative for the three other subcomponents. Proposition (C), finally, is generally rejected, although less often for (ii) legal, than for the other subindices.

Table 2: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + \delta_t + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

EWF	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.195***	-0.113*	-0.312***	-0.304***	-0.099	-0.053	-0.001
	(0.03)	(0.07)	(0.04)	(0.07)	(0.06)	(0.07)	(0.12)
$\triangle y_{i,t-1}$	0.013	0.112	0.015	0.324***	-0.196**	-0.011	-0.492***
	(0.04)	(0.12)	(0.06)	(0.12)	(0.09)	(0.11)	(0.18)
Const.	0.333***	0.238*	0.622***	0.594***	0.211*	0.090	-0.015
	(0.05)	(0.13)	(0.07)	(0.13)	(0.11)	(0.13)	(0.23)
Observations	572	123	113	109	103	71	53
Breusch-Pagan[a]	[0.000]	[0.000]	[0.001]	[0.000]	[0.000]	[0.000]	[0.234]
R^2	0.1755	0.093	0.436	0.257	0.151	0.0109	0.1821

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

5.2 Distributional results

Table 3 shows the transition probabilities for the distribution of EFW across countries. The quartiles are defined using the average location of country EFW-values over 1970-2005 and are hence fixed in the preceeding hypothetical period by bounds that are defined in the note to the table. The diagonal elements correspond to the probability for a country of remaining within the same bounds in period t+1 as in t. Countries located in q_1, with EFW scores between 2.3 and 5.2, face a 0.66 probability of remaining within these bounds five years hence. Countries can transition from q_1 to the bounds defined by q_4 in just one time period - which would correspond to an EFW increase of at least 1.8 over five years, but the probability of that happening is a mere 0.005.

For countries in the upper quartile, with EFW scores between 7.0 and 9.1, institutional quality is highly persistent. The probability for remaining in the bounds defined by q_4 in the next five-year period is 0.86, while that of falling to the bounds defined by the lowest quartile q_1 is 0. The second and third quartiles are less persistent, but there is a large difference in the probability of transitioning from q_3 to the bounds defined by q_4 compared to that of transitioning from q_2 to the bounds defined by q_4. In the former case the probability is 0.34 whereas in the latter it is only 0.028. Considering the mirror image, we notice that both probabilities are smaller: the probability of

Table 3: Transition probabilities in EFW over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.656	0.261	0.078	0.005
q_2	0.151	0.486	0.335	0.028
q_3	0.024	0.171	0.470	0.335
q_4	0,000	0.014	0.126	0.859
Erg. Dist.	0.073	0.129	0.228	0.571

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in EFW that corresponds to: $q_1 \in [2.3, 5.2]$, $q_2 \in [5.3, 6.0]$; $q_3 \in [6.1, 6.9]$; $q_4 \in [7.0, 9.1]$.

falling from q_2 to the bounds defined by q_1 is 0.15, while that of falling from q_3 to the bounds defined by q_1 is 0.024.

The transition probabilities to the right-above the diagonal are larger than those to the left-below, which suggests that it is easier for countries to increase their institutional quality than it is to decrease it.

Assuming the transition probabilities are time-invariant, and that institutional convergence will continue at the same rate, it is possible to compute the steady state vector, which describes the long run distribution of countries within the bounds defined by the quartiles at the onset. The last row of Table 2, shows the ergodic steady state distribution of institutional quality. At the onset, by definition, the distribution of countries over q_j is uniform, with 25 % of the observations in each quartile. In steady state, however, only 7 percent of the countries have an institutional quality between 2.3 and 5.2, while the bulk of countries (58%) have transitioned to the high quality institutions with EFW-levels above 7.0.

One important feature of the Markow property, is that the result relies on the arbitrary selection of the grids. As a test for robustness, alternative grids was choosen containing 3, 5 and 10 quantiles, respectively. Since the results are comparable, they are not included in the paper.

As for the time involved in making a substantial transition, consider for example the fact that a country located in q_1 today has a 0.31 probability of moving to the bounds defined by q_4 over a thirty-year period. Hence, while institutional convergence may appear slow in the short run, the time to make a big leap is surprisingly short, while the probability of making it is non-trivial.

In the same manner, we tested the transition probabilities for each subcomponent of the index. Results are available in tables 9-13 in the appendix. Interestingly enough, for the subcomponents (i) government, (ii) legal, and (iii) money, the ergodic distribution of countries is rather different from that for the overall EFW in that the probabilities are rather evenly divided across the four bounds defined by the quartiles in period t. Here, for example, the probability in the steady-state to end up in the bounds defined by q_4 with regards to (iii) money is only 0.29, compared to 0.25 of ending up in the bounds defined by q_1. The subcomponents (iv) trade and, particularly, (v) regulation, nonetheless see a steady-state distribution more similar to that of the overall EFW. [8]

6 Summary and conclusions

If Fukuyama's (1989, 1992) argument that the advent of Western liberal democracy signals the arrival of the final form of human government, we should be able to observe that countries converge in their political and economic systems towards liberal democracy. The purpose of this paper was to examine whether we observe such institutional convergence with regards to economic institutions, and if so, try to assess the implications in the long-term.

First, we tested whether we observe convergence in economic institution across countries. This was done by examining whether LPE holds for EFW over the period 1970-2005. We did this by testing three propositions that must hold for the law to be in effect. These were:

(A) that changes in the EWF index are independent of its level;

(B) that changes over two consecutive periods are independent of each other, and;

[8] It has been pointed out that while the creators of EFW appear to have created a valid framework to determine which variables can be used to measure economic freedom, this is not a sound argument for arranging the variables in the particular manner they do (Rode 2010). Heckleman and Stroup (2000, 2005) Sturm et al (2006) and de Haan et al (2006) critizise the current composition and aggregation EFW. We adress the problem with an arbitrary weighting scheme by using principal component analysis. Since the number of subcomponents of the index changes over time, we chose to only perform PCA on the five main areas of the index. PCA-results are available in table 14 in the appendix. For each time-period, PCA delivers 2 factors with an eigen-value above 1 (the Kaiser-criterion). The first factor always loads heavily on all components except government, and explains 45-55% of the total variance. It hence lends itself to a fairly straightforward interpretation as an index of economic freedom with little weight on the government variables. As the second factor explained only about 25% of the total variation and is less easily intepreted, it was excluded from the analysis. In view of the similarities from one year to another, we construct a new index based on the first factor created in 2005 to test LPE. Results are reported in table 15 in the appendix, and show strong similarities to EFW-results in table 2, which is reassuring. Transition probabilities using the new index are reported in table 16, and suggest a slightly higher likelihood of a country ending up with high-quality institutions in the long-run than before.

(C) that the variability of changes are independent of the index level.

Results suggested that the null hypothesis with regards to proposition (A) could be rejected. Hence, we do observe a convergence process where countries with lower institutional quality do improve their institutions faster than countries with higher institutional quality. For proposition (B) the null hypothesis was not rejected, i.e., we do not find evidence of a negative or positive feedback-loop on institutional change. As for (C), the null hypothesis was rejected, which we conjecture should imply that the pace of institutional change is more variable for countries with lower institutional quality.

Secondly, we wanted to assess the implications if this institutional convergence continues at the same rate in the future. Distributional analysis showed that the overall trend points towards convergence. The steady-state probability of a country finding itself in the bounds defined by the highest quartile is 0.58, while that of being in the lowest is a mere 0.07. As for the time for convergence, it is far from infinite: countries have a fairly high probability of improving their institutional quality in a matter of decades.

As such, the findings in this paper give some support at least to a conservative interpretation of Fukuyama's claim, in that it appears that market-oriented economic institutions will become more and more prevalent among countries in the long term. Granted, one should be careful using the historical record to make predictions about the future, but the trend of the past twenty years is certainly in the direction of Fukuyama's prediction. Furthermore, we have only examined a subset of the institutions underlying liberal democracy. Future research should be devoted to examine whether a similar development is observable with regards to political institutions, for example using a method similar to the one in this paper.

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7 Appendix

Table 4: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

(i) GOVERNMENT	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.336***	-0.421***	-0.170***	-0.421***	-0.365***	-0.371***	-0.135
	(0.05)	(0.06)	(0.05)	(0.15)	(0.13)	(0.11)	(0.10)
$\triangle y_{i,t-1}$	-0.099**	0.037	-0.061	-0.018	-0.238**	-0.221*	-0.072
	(0.04)	(0.08)	(0.05)	(0.07)	(0.12)	(0.13)	(0.15)
Const.	0.471***	0.824***	0.344***	0.779***	0.682***	0.558***	0.142
	(0.09)	(0.10)	(0.10)	(0.25)	(0.22)	(0.19)	(0.18)
Observations	646	121	114	112	108	102	89
Breusch-Pagan[a]	0.000	0.000	0.002	0.000	0.000	0.000	0.018
R^2	0.31	0.40	0.13	0.29	0.37	0.34	0.05

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 5: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

(ii) LEGAL	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.214***	-0.092*	0.020	-0.462***	-0.067	-0.147	-0.478***
	(0.03)	(0.05)	(0.05)	(0.06)	(0.06)	(0.10)	(0.09)
$\triangle y_{i,t-1}$	-0.220***	-0.066	-0.108**	-0.233**	-0.271**	-0.220**	-0.387**
	(0.05)	(0.13)	(0.05)	(0.10)	(0.12)	(0.08)	(0.18)
Const.	0.423***	0.139	-0.042	0.881***	0.145	0.307	0.880***
	(0.07)	(0.10)	(0.09)	(0.11)	(0.10)	(0.19)	(0.18)
Observations	534	123	111	110	90	50	50
Breusch-Pagan[a]	0.000	0.000	0.417	0.033	0.680	0.009	0.959
R^2	0.30	0.04	0.05	0.44	0.16	0.27	0.52

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 6: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

(iii) MONEY	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.440***	-0.499***	-0.680***	-0.694***	-0.164	-0.225***	0.086
	(0.09)	(0.10)	(0.06)	(0.18)	(0.15)	(0.08)	(0.18)
$\triangle y_{i,t-1}$	0.087	0.043	0.041	0.400**	0.508***	-0.061	-0.449***
	(0.06)	(0.03)	(0.04)	(0.17)	(0.13)	(0.10)	(0.16)
Const.	0.775***	1.048***	1.435***	1.266***	0.256	0.431***	-0.260
	(0.17)	(0.20)	(0.12)	(0.34)	(0.28)	(0.16)	(0.36)
Observations	664	119	117	110	106	106	106
Breusch-Pagan[a]	0.000	0.000	0.000	0.000	0.011	0.049	0.138
R^2	0.24	0.48	0.76	0.33	0.22	0.07	0.13

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 7: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

(iv) TRADE	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.200***	-0.082	-0.256***	-0.306***	-0.260***	-0.044	-0.204*
	(0.04)	(0.08)	(0.10)	(0.08)	(0.09)	(0.07)	(0.10)
$\triangle y_{i,t-1}$	-0.104**	-0.209**	-0.137*	-0.032	-0.182	-0.064	-0.175
	(0.05)	(0.09)	(0.08)	(0.07)	(0.15)	(0.11)	(0.15)
Const.	0.362***	0.131	0.538***	0.651***	0.495***	0.052	0.372**
	(0.07)	(0.17)	(0.19)	(0.15)	(0.17)	(0.13)	(0.19)
Observations	589	116	106	102	100	92	73
Breusch-Pagan[a]	0.000	0.000	0.000	0.000	0.000	0.000	0.400
R^2	0.22	0.12	0.23	0.28	0.32	0.02	0.11

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 8: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta-1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

(v) REGULATION	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.198***	-0.299***	-0.326***	-0.197**	-0.165***	0.004	-0.010
	(0.03)	(0.05)	(0.04)	(0.10)	(0.04)	(0.04)	(0.06)
$\triangle y_{i,t-1}$	0.011	-0.038	0.098**	0.180	-0.235**	0.050	-0.399*
	(0.04)	(0.08)	(0.05)	(0.18)	(0.11)	(0.07)	(0.21)
Const.	0.333***	0.654***	0.617***	0.407**	0.273***	-0.004	0.029
	(0.05)	(0.09)	(0.08)	(0.16)	(0.07)	(0.06)	(0.10)
Observations	567	123	118	106	102	73	45
Breusch-Pagan[a]	0.000	0.020	0.000	0.000	0.001	0.241	0.004
R^2	0.27	0.21	0.39	0.12	0.29	0.00	0.27

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 9: Transition probabilities in (i) GOVERNMENT over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.621	0.251	0.097	0.031
q_2	0.219	0.459	0.296	0.092
q_3	0.094	0.146	0.479	0.281
q_4	0.013	0.065	0.286	0.636
Erg. Dist.	0.203	0.209	0.292	0.296

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in GOV. SIZE that corresponds to: $q_1 \in [0.7, 4.6]$, $q_2 \in [4.7, 5.7]$; $q_3 \in [5.8, 6.8]$; $q_4 \in [6.9, 9.7]$.

Table 10: Transition probabilities in (ii) LEGAL over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.653	0.220	0.116	0.012
q_2	0.230	0.497	0.230	0.042
q_3	0.025	0.248	0.509	0.217
q_4	0.006	0.051	0.158	0.785
Erg. Dist.	0.184	0.240	0.258	0.318

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in LEGAL that corresponds to: $q_1 \in [1.1, 4.0]$, $q_2 \in [4.1, 5.4]$; $q_3 \in [5.5, 6.8]$; $q_4 \in [6.9, 9.6]$.

Table 11: Transition probabilities in (iii) MONEY over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.573	0.282	0.119	0.026
q_2	0.245	0.454	0.240	0.061
q_3	0.131	0.186	0.437	0.246
q_4	0.063	0.057	0.149	0.731
Erg. Dist.	0.248	0.236	0.229	0.288

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in Money that corresponds to: $q_1 \in [0.0, 5.7]$, $q_2 \in [5.8, 6.8]$; $q_3 \in [6.9, 8.7]$; $q_4 \in [8.8, 9.9]$.

Table 12: Transition probabilities in (iv) TRADE over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.639	0.279	0.067	0.014
q_2	0.167	0.532	0.253	0.048
q_3	0.033	0.180	0.520	0.267
q_4	0.006	0.006	0.177	0.811
Erg. Dist.	0.114	0.178	0.271	0.437

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in TRADE that corresponds to: $q_1 \in [1.4, 5.2]$, $q_2 \in [5.3, 6.3]$; $q_3 \in [6.4, 7.2]$; $q_4 \in [7.3, 9.8]$.

Table 13: Transition probabilities in (v) REGULATION over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.653	0.245	0.088	0.014
q_2	0.167	0.529	0.240	0.064
q_3	0.014	0.096	0.466	0.425
q_4	0.000	0.032	0.145	0.823
Erg. Dist.	0.067	0.121	0.225	0.587

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in REGULATION that corresponds to: $q_1 \in [2.5, 5.0]$, $q_2 \in [5.1, 5.7]$; $q_3 \in [5.8, 6.5]$; $q_4 \in [6.6, 8.9]$.

Table 15: Results from testing LPE in the period between 1970-2005

Estimated equation: $\Delta y_{i,t} = (1-\beta)\alpha_i + (\beta - 1)y_{i,t-1} + \theta\Delta y_{i,t-1} + \epsilon_{i,t}$.

PCA-index	All years	2005-2000	2000-1995	1995-1990	1990-1985	1985-1980	1980-1975
$y_{i,t-1}$	-0.275***	-0.243***	-0.320***	-0.380***	-0.144	-0.015	-0.100
	(0.033)	(0.062)	(0.050)	(0.060)	(0.092)	(0.125)	(0.112)
$\triangle y_{i,t-1}$	0.050	0.080	0.022	0.247*	-0.226	0.071	-0.560**
	(0.059)	(0.096)	(0.059)	(0.134)	(0.204)	(0.169)	(0.220)
Const.	0.521***	0.495***	0.659***	0.766***	0.302*	0.032	0.235
	(0.068)	(0.127)	(0.099)	(0.116)	(0.171)	(0.261)	(0.230)
Observations	459	114	102	94	76	42	31
Breusch-Pagan[a]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.808]	[0.028]
R^2	0.326	0.301	0.405	0.487	0.228	0.009	0.484

Note: Standard errors in parentheses and *** p<0.01, ** p<0.05, * p<0.1. The regressions are estimated with robust standard errors.

[a] Breusch-Pagan / Cook-Weisberg test for heteroskedasticity, where brackets refers to p-value for rejecting the null hypothesis of constant variance.

Table 14: Principal component analysis of the five areas composing EFW, component loadings displayed for Factor 1. Uniqueness within parentheses.

Year of PCA	2005		2000		1995		1990		1985		1980		1975		1970	
Variable	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2	F1	F2
government	-0.000	0.988	-0.048	0.974	0.006	0.961	-0.027	0.941	-0.070	0.920	0.007	0.958	0.016	0.981	-0.040	0.935
	(0.0244)		(0.0489)		(0.0759)		(0.1135)		(0.1482)		(0.0815)		(0.0378)		(0.1250)	
property	0.859	-0.195	0.772	-0.462	0.786	-0.446	0.786	-0.332	0.813	-0.208	0.771	-0.043	0.834	-0.164	0.844	-0.234
	(0.2238)		(0.1913)		(0.1841)		(0.2726)		(0.2962)		(0.4044)		(0.2780)		(0.2331)	
money	0.817	0.165	0.801	0.010	0.808	0.021	0.781	0.050	0.758	0.142	0.714	0.097	0.768	0.132	0.702	0.153
	(0.3049)		(0.3577)		(0.3462)		(0.3877)		(0.4060)		(0.4809)		(0.3933)		(0.4836)	
trade	0.787	0.047	0.799	-0.158	0.792	-0.094	0.854	0.082	0.890	0.112	0.836	-0.023	0.870	0.036	0.909	0.087
	(0.3782)		(0.3361)		(0.3636)		(0.2645)		(0.1950)		(0.3001)		(0.2423)		(0.1672)	
regulation	0.831	0.032	0.771	0.131	0.818	0.361	0.710	0.535	0.600	0.626	0.659	0.410	0.719	0.249	0.489	0.396
	(0.309)		(0.3885)		(0.2006)		(0.2096)		(0.2477)		(0.3985)		(0.4212)		(0.6040)	
Variance explained	0.543	0.209	0.495	0.241	0.514	0.252	0.492	0.258	0.478	0.263	0.447	0.220	0.512	0.214	0.454	0.232

Note: The component solution has been rotated using the Varimax technique.

Table 16: Transition probabilities in PCA-index over quartiles during 1970-2005.

	q_1	q_2	q_3	q_4
q_1	0.627	0.319	0.054	0.000
q_2	0.170	0.483	0.333	0.014
q_3	0.008	0.148	0.586	0.258
q_4	0.000	0.000	0.077	0.923
Erg. Dist	0.042	0.083	0.198	0.678

Note: The data is uniformly discretized into quartiles, where q_j refers to the $j = 1...4$ quartile in PCA-index.