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The Measurements and Determinants of Institutional Change: Evidence from Transition Economies

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Abstract

It is widely appreciated that institution building is at the heart of the transition process. Without functioning institutions markets cannot work effectively and the sustainability of the economic transition process can be undermined. The crisis in Russia provided just one piece of evidence in this regard. While institutions are central to the transition process, institutional reform is not an area that is well understood by researchers and policy makers alike. In this paper we examine the determinants of institutional change using a panel dataset comprising 25 transition economies. One of the defining characteristics of our approach is that we treat institutional change as a multidimensional unobserved variable, accounting for the fact each of our indicators represents a noisy signal.
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The Measurement and Determinants of Institutional Change: Evidence from Transition Economies

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Proleptically, I would say that whether we can measure something depends, not on that thing, but on how we conceptualize it, on our knowledge of it, above all on the skill and ingenuity which we bring to bear on the process of measurement which our enquiry can put to use

Kaplan, Abraham
The Conduct of Enquiry

Let us remember the unfortunate econometrician who, in one of the major functions of his system, had to use a proxy for risk and a dummy for sex

Fritz Machlup (1974)

1 Introduction

Most social scientists agree that institutions play a fundamental role in economic development (see, for example, Keefer and Knack (1995) and Keefer and Knack (1997); and for the transition economies Havrylyshyn and Van Rooden (1999); Stone, Levy, and Paredes (1996)). Despite the recognition of their importance empirical research on the determinants of institutional change remains scarce. To what extent institutions can be actively influenced by policy and what other factors interact in shaping the pattern of institutional adaptation remains unclear. Existing work on the causes of institutional change has concentrated on the build-up of historical case studies, which show that institutions often adapt efficiently to serve new economic opportunities (see, for example, North and Weingast (1996), and Krueger (1996)). At the same time, comparative historical work also stresses the role of path dependence and the actions of the state in shaping the direction of institutional change (North (1990)). In this paper we present what we believe is the first attempt to estimate a structural model of institutional change using both time series and cross-sectional data on transition economies.[4] This allows us to test directly

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[4]This paper is also available as an EBRD working paper. See Raiser, DiTommaso, and Weeks (2000a).

the relevance of different factors in driving institutional change, such as path dependence, changes in the structure of market demand, interaction with the outside world, and the capacity of the state for implementation and enforcement of new rules.

The transition economies seem particularly suited to such an exercise. First, they are characterised by an unprecedented degree of institutional change going largely in the same direction. The move from a system based on state planning and an allocation of resources by government dictate to a system of decentralised market allocation necessitates substantial change in laws and regulations, as well as in norms and expectations. However, the outcome of concerted efforts across the region to redraw the set of laws and regulations governing economic exchange has differed widely between countries, with institutional performance showing significant divergence after the first decade of transition (EBRD (1999)). It is therefore interesting to examine the determinants of such variation in institutional performance and the constraints on institutional change in some of the less advanced transition economies. Second, the transition economies have embarked on this process of institutional change largely at the same time. With EBRD's transition indicators spanning 5 years of institutional change in the transition economies, the opportunity arises for the first time to test the determinants of institutional change in a panel of 25 countries facing a similar worldwide environment (see Table 1). EBRD has recently backdated its transition indicators to 1989.[5]

We adopt an analytical framework that treats institutional change as an imperfectly measured underlying process. We use EBRD's transition indicators as our measures of institutional change. These indicators are designed to rate the outcome of changes in formal rules and resulting changes in economic behaviour. Given this conceptualisation, it is evident that measurement problems are of particular concern. For this reason we adopt a latent variable representation of institutional change which explicitly accounts for measurement error. The Multiple Indicator, Multiple Cause (MIMIC) model adopted in this paper allows us to model institutional change as an imperfectly measured multidimensional process linked to a common set of causal factors.

The MIMIC approach has been applied to a broad set of issues both within and outside of economics. Examples include intelligence, education, management expertise and institutional change. Perhaps the most prevalent example in economics is based upon Friedman's (1957) permanent income hypothesis where there is no directly observable measure of permanent income.[6] In the context of time series applications Watson and

[5] We do not consider this expanded dataset in this paper, focusing instead on developing a methodology to measure and model institutional change. A future more policy oriented paper will review the results of our analysis with the expanded dataset.

[6] For other applications in economics see Harvey (1981), Avery (1979), Lahiri (1976), Ziebart (1987),

Engle (1983) demonstrate the relationship between a broad class of structural equation models including factor and multiple indicator and multiple cause (MIMIC) models.[7] Most directly related to this paper is work by Ohlson (1979), who examines the relationship between political democracy and industrialisation in developing countries, treating both constructs as unobservable latent variables. Measurement error models for political democracy and industrialisation are constructed using a single common factor model and ordinal indicators such as freedom of the press and the effectiveness of the legislature (for democracy), and per capita gross national product for industrialisation. Kaufmann, Kraay, and Zoido-Lobaton (1999) adopt a similar approach in examining cross-country variation in governance.

The choice of causal factors are motivated by the following considerations, which have so far not received analytical treatment in a multivariate framework. First, we test whether market liberalisation and privatisation have had any effect on institutional change. This hypothesis was originally proposed by the World Bank in its 1996 World Development Report and justifies a sequencing of economic reforms that begins with the "easier" tasks of liberalisation and privatisation to generate a constituency for the more demanding institutional reforms (see also Fischer and Gelb (1991)). We condition the impact of these demand factors by controlling for the extent of political rights and civil liberties established in each transition economy, which are expected to influence the constraints of policy makers in responding to market demands. Second, we investigate the impact of structural change on institutional change. For instance a shift in trade patterns towards market economies may lead to the adoption of new rules to satisfy western trading partners. We also look at the effect of changes in the structure of employment. Third, we construct an index of initial conditions that allows us to control for the impact of different starting points on the patterns of institutional change as would be predicted by a model of path dependent institutional adaptation. Fourth, we explicitly control for state capacity as a factor in institutional change using public expenditure as a proxy. The multivariate analysis we develop allows us to test for the significance and quantitative importance of these four set of causal factors simultaneously.

The structure of the paper is as follows. In section 2 we evaluate a number of alternative strategies for modelling institutional change and specifically focus on the way in which we utilise the information within indicators. Section 3 introduces our measures of institutional change. Section 4 turns to the causal factors in our model and motivates the

and Ohlson (1979). The application by Spanos (1984) where the author develops a structural model of the demand for money, treating liquidity as an underlying latent variable, is particularly noteworthy.

[7]Seminal studies using the MIMIC model include Zellner (1970), Hauser and Goldberger (1971), and Goldberger (1972).

above-mentioned hypotheses. Section 5 develops the formal structural model of institutional change and deals with specification and estimation issues. Section 6 presents the main results. All variable definitions are reported in Appendices.

2 Modelling Institutional Change: the MIMIC Approach

Institutional change is a multidimensional process and as such no one single measure is likely to adequately capture this process. At the same time, the various dimensions of change are intrinsically linked. The choice of multiple indicators to measure institutional change therefore raises the question of how to combine them in empirical research on institutional change as an underlying process rather than focusing on just one sub-dimension. The MIMIC approach developed in the paper is one solution to this problem. We begin by evaluating its merits over other strategies.

In evaluating the relative adequacy of the MIMIC approach we first abstract from the existence of causal factors, and focus upon the problem of how to both choose and optimally combine the information in multiple indicators, utilising this to make inference on the unobserved underlying process. For example, one of the most basic strategies would be to pick an indicator we believe is *closest* to the unobserved construct, and ignore any measurement error. Alternately we could use the information in all indicators by either creating a synthetic variable, such as a simple mean indicator, or estimating a system regression over the m indicators. In this context a natural framework is the unobserved component model, or confirmatory factor analysis. Kaufmann, Kraay, and Zoido-Lobaton (1999) utilise a variant of this approach to combine information on a large number of indicators of governance.[8] The authors assume that each observed score for a particular indicator is a linear function of unobserved governance and a disturbance term, which is assumed to be uncorrelated across indicators. A conditional mean function of the form

$$E(y^*|\mathbf{y} = (y_1, ..., y_m)'), \tag{1}$$

where y^* denotes the unobserved variable and $\mathbf{y}$ is a $m \times 1$ vector of indicators, is then estimated using one of the outputs of the unobserved components model, namely the variance of each indicator, to weight the contribution of each indicator. The variance of each indicator may obviously be interpreted as a guide to how informative each indicator is

[8]Note that Kaufmann, Kraay, and Zoido-Lobaton (1999) use indicators to examine three aspects of governance: probity, bureaucratic quality and the rule of law. However, for the sake of exposition we lose nothing by considering undifferentiated governance.

with respect to the unobserved variable governance, and can therefore be used as weights in constructing an optimal combination of observed indicators.

The MIMIC model also yields a set of weights to combine several indicators into a composite measure, but in this case the weights are derived from the relationship between a set of observed endogenous variables and an unobserved latent construct on the one hand, and a set of observed exogenous variables on the other. In our case, this seems more appropriate as the measures of institutional change we shall use come all from the same source, and hence their variance is unlikely to be a good weight. We also note that the MIMIC approach presented here can be viewed as the first stage in a multiple equations model of transition, where institutional change is caused by a set of exogenous factors and in turn influences economic outcomes. If one's interest is in the latter relationship the endogenous generation of weights for the measures of institutional change based on a set of causal relationships may be a particular attraction.

A sufficient condition for parameter identification in both the one factor measurement error model and the MIMIC model is the existence of three indicators. However, with more than three indicators available, the question of whether to include additional indicators is not simply a statistical issue but is very much related to how we conceptualise and represent the unobserved component. In our particular problem, the inclusion of $m > 3$ indicators can be justified insofar as the m indicators represent the main dimensions of institutional change. We elaborate upon this issue in section 4.

Turning to the issue of institutional change in a causal framework with both multiple causes and multiple indicators, we need to decide whether to allow the causal factors to affect each dimension of institutional change in a different way, or whether to focus on their impact on the underlying process. The principal difference between a MIMIC model and more standard regression-based approaches is the fact that MIMIC presupposes the existence of a common factor, institutional change, which is imperfectly measured by m indicators.[9] Thus, although both a MIMIC and a Seemingly UnRelated equation approach (SUR) utilise information in all the m indicators, SUR assumes no measurement error, and related, assigns an equal weight across indicators. In addition it is instructive to note that a standard regression-based model bases inference on the conditional moments $E(y|x)$, where x denotes a vector of observed causes. If y is an imperfect indicator of y^* then the parameters of these moments are *not* the fundamental parameters of the system being investigated. A welcome by-product of the MIMIC approach is that instead of estimating m regression equations for the set of indicators, we estimate the parameters

[9] In this respect a structural equations approach also assumes that the principal focus of investigation is institutional change and not on one or more of the indicators.

of a single structural equation. Ignoring covariance terms, and assuming we have a $s \times 1$ vector of causes, we have a total of $m + s$ estimable parameters. This compares with a total of $m \times s$ parameters if we ignore measurement error and simply estimate a system of equations over the m indicators. A drawback of our approach is that if the causal factors under investigation are expected to have different impact across the various dimensions of institutional change, these differences cannot be captured by the MIMIC model. We suggest that our approach complements rather than substitutes for more detailed investigations in each dimension (see for example Pistor, Raiser, and Gelfer (2000), Dutz and Vagliasindi (2000)).

3 Measuring Institutional Change

In this paper we rely on EBRD's measures of institutional reform as representing the best available data on institutional change in the transition economies. Following Stern (1997), the approach we adopt is to start from an identification of the four basic elements of a market economy[10]. These elements are: enterprises and households which are responsible for decisions concerning production and consumption; markets where agents interact and resources are allocated; financial institutions which determine how transactions occur and budget constraints are enforced; and the legal system which underpins the system of contracts and investment. These four basic elements delineate the main area in which institutional change is likely to be preeminent during the transition. This is not to say that there may not be other dimensions of such change, but to provide a focus that can form the basis of empirical analysis.

The EBRD has constructed five subjective ratings pertaining to institutional change during transition: Governance and Enterprise Restructuring (ER_ST), Competition Policy (CP), Banking Reform and Interest Rate Liberalisation (BR_I), Securities Markets and Non-Bank Financial Institutions (S_NB), and Overall Legal Effectiveness and Extensiveness (OLE).[11] The five dimensions complement EBRD's ratings on market liberalisation and privatisation, which are treated as conceptually distinct in this research. The distinction arises from the fact that while in all of the above dimensions new rules need to be created and credibly enforced by the state, market liberalisation and privatisation of

[10] See, in particular, EBRD *Transition Report* 1994.

[11] Note that the OLE index is a unweighted average of ratings on the extensiveness of commercial law covering the areas of secured transactions, bankruptcy and investor protection and the effectiveness with which these laws are enforced. This indicator suffers a break between 1996 and 1997 when the scope of coverage was changed from foreign investment laws only to a broader set of commercial legislation. The results in this paper are robust to excluding this dimension of institutional change.

state assets require predominantly that the state relinquish control.[12]

In all cases EBRD provides subjective scores ranging from 1 to 4+. A score of 1 indicates very little institutional change relative to the typical situation in a centrally planned economy, a score of 4+ indicates that the creation of market-supporting institutions in this area is largely complete, as the standards of developed market economies have been reached. Table 1 reports the value of each of these indices for 1998 and for the first year in which they were reported by EBRD (1994 or 1995) revealing considerable differences in institutional reform in the mid-1990s, and significant variation in reform progress since. (Appendix 1 provides the definition of the five dimensions.)

4 Institutional Change: Initial Conditions, Demand and Supply Factors

In this section we review the main set of causal factors that will be used in the empirical analysis of institutional change. We group them into three main categories: initial conditions, demand factors and supply factors.

4.1 Initial conditions and path dependencies

Institutions are historically specific and for this reason it is necessary to be sensitive to historical context. The inclusion of initial conditions among the causal factors of institutional change in the transition economies reflects the possibility that history could matter a lot for the development trajectories followed by these countries since the early 1990s.

A priori a large number of variables could influence transition paths and the resulting patterns of institutional change. We focus here on three main aspects: geographical factors, cultural factors, and the institutional legacy of central planning. We do not consider structural or macroeconomic distortions at the start of transition among our initial conditions, as these variables would be expected to change during the course of transition. Instead we test directly whether changes in economic structure and macroeconomic conditions have an influence on institutional change. Our set of initial conditions is thus essentially fixed and exogenous to the transition process itself.

Among the geographical factors we consider distance to Brussels and natural resource endowments. Proximity to the modern democratic and business-oriented societies of the EU could help in the process of institutional adaptation through diffusion effects, learning and cultural familiarity. Distance of the capital city from Brussels is thus introduced as

[12] For a more extensive discussion of these issues see Raiser, DiTommaso, and Weeks (2000b)

an initial condition (LOCAT_EU). Natural resource endowments on the other hand may have slowed institutional reforms or distorted institutional adaptation, as natural resource rents diminished the perceived need for reform and vested interests gained control over the policy agenda. Natural resource wealth (RICH) is measured as a dummy variable and taken from De Melo, Denizer, Gelb, and Tenev (1997).

Cultural factors comprise religious affiliation and ethnic heterogeneity. It is remarkable that the first wave of accession countries to the EU will include exclusively countries with a predominantly Roman Catholic or Protestant orientation. We distinguish between "western Christianity" (REL1=1), "eastern or orthodox Christianity"(REL2=1) and other (mainly Muslim) affiliations (both REL1 and REL2 =0).Ethnic heterogeneity arguably makes reforms more difficult as it opens cleavages unlikely to be resolved through material redistribution alone. Compensation of losers can be far more difficult under these circumstances Elster, Offe, and Preuss (1998). Alesina and la Ferrara (2000) also draw a link from ethnic heterogeneity to lack of trust. We construct a dummy variable (ETHNIC) that measures the importance of the dominant ethnic group in the population.

The legacy of central planning may affect institutional change primarily by shaping individual behaviour and thus the appeal of reforms aimed at radical socio-economic modernisation. We include as measures of starting points in this regard the number of years spent under central planning (MARMEM), and a dummy for the degree of established national sovereignty (STATE). The latter indicator captures the fact that new nations would need to spend considerable resources on consolidation which would not be available for institutional reform. Finally we also include GDP per capita (GDP_PC) and the rate of urbanisation among the initial conditions as proxies for the level of development. The full list of variables and their definitions appears in Appendix 2.

Since many of them are potentially highly correlated we construct a synthetic index of initial conditions using principal components analysis. [13] This analysis yields a set of common factors, which may broadly represent the various dimensions of initial conditions. The factor loadings obtained can then be used to construct country scores for each of these common factors. The country factor loadings are presented in Table 2 and the country scores are in Table 3.

[13] In an earlier version of this paper we experimented with different sets of initial conditions extended the earlier work of De Melo et al. (1997). It turns out that most initial conditions are highly correlated and the resulting weighted indices all produce the same ranking across countries. The results in this paper are thus not sensitive to the special initial conditions index we use.

4.2 Demand factors: liberalisation, privatisation, integration and the political process

The dominant view among reformers and their advisors during the early transition period was that because institutions would necessarily take time to develop, it was best to focus first on liberalisation and privatisation. This advice was predicated on the expectation that the creation of markets would result in some endogenous adaptation of institutions and at least make institutional reform easier further down the road. In this paper we test this claim.

We measure the extent of liberalisation as the number of years that have elapsed since a country reached full price liberalisation (ysPL). Measuring liberalisation in this way allows us to get more directly at the idea that the impact of liberalisation of the demand for institutions may operate with significant lags. In addition, this variable is likely to be exogenous to institutional change, unlike the scores for liberalisation themselves.

Markets alone are not sufficient to generate the demand for supporting institutions. Economic actors will only have appropriate incentives to demand change, if their property rights are protected. However, where state assets have become concentrated in the hands of few powerful oligarchs the pressure for institutional reforms that would constrain their economic and political power will be less than in countries with a more equal distribution of wealth (Hellman (1998)). We therefore focus on small-scale privatisation, which is less prone to capture by vested interests. In accordance with the liberalisation measure, small-scale privatisation is measured as the number of years since it was largely completed (ysSCP).

Turning to the impact of structural change on the demand for institutions, we focus on changes in the direction of exports and changes in the structure of employment. Exports to other transition economies may not generate the same demand for institutional adaptation than exchange with established market economies. We measure trade orientation by the share of exports to non-transition countries (EXSHE). By the same token, where workers remain locked in inefficient industries institutional change that might harden budget constraints for instance, might be resisted. We include an time-varying indicator of economic distortions defined as difference between the share of employment in industry predicted by the level of per capita income and the actual share (EMSIN).

Finally, the demand for institutions by economic actors will to some extent be mediated by the political process. If the political institutions in a given country do not grant individuals political rights to express their views and the civil liberties to follow new opportunities in line with their preferences, the impact of market liberalisation and privatisation on the demand for institutions may be fundamentally altered. We construct an average of the ratings of political rights and civil liberties obtained from Freedom House (ave_POL). All the demand factors and their definitions are reported in Appendix 3.

4.3 Supply factors: state capacity and reform priorities

Institutional change would be impossible without a state that enacts and enforces new rules and regulations. The capacity of the state is thus of primary importance in accounting for institutional change during the transition. However, it is difficult to find variables that may be used to measure state capacity. Some reflection of it is included in the vector of initial conditions. The assumption of constant state capacity is, however, unsatisfactory. In many transition economies a deep fiscal crisis has eroded the morale of public servants and the resources of most public services.[14] We thus include the ratio of general government expenditures to GDP (PUBE_GDP) as a variable capturing time varying aspects of state capacity.

There is one further variable that might be used to measure developments in state capacity over time. For many transition economies, macroeconomic stabilisation was the primary task at the beginning of the transition and institutional reforms really started to be considered only once this had been successfully secured. To some extent it might thus be argued that the state's capacity for institutional reform is higher in countries that have successfully stabilised. Our measure of macroeconomic stability uses a cumulative indicator, counting the number of years since the beginning of the transition in which a country achieved inflation below 30 per cent and a budget deficit below 5 per cent of GDP (ysIBD).

The two supply variables are also reported and defined in Appendix 3.

5 A Structural Equations Approach to Modelling Institutional Change

In this section we develop the empirical model that links the measures of institutional change with the causal factors discussed in the preceding section. Our modelling strategy has been extensively used in psychometrics and more recently in econometrics, and is founded upon the specification of a system of equations which specify the relationship between a set of unobservable latent variables, $\mathbf{y}^*$, a set of observable endogenous indicators $\mathbf{y}$, and a set of observable exogenous variables $\mathbf{x}$. This approach builds upon the

[14] For example, a high court judge in Kazakhstan currently currently earns US$ 180 per month, around three times the daily rate of a driver for an international organisation.

early work of Joreskog and Goldberger (1975) and Zellner (1970), and has been formalised in the LISREL[15] model of a set of linear structural equations. Excellent reviews of this literature are to be found in Bentler and Weeks (1980) and Aigner, Hsiao, Kapteyn, and Wansbeek (1984).

Below we provide a brief overview of the characteristics of our approach.

5.1 Model Specification

The structure of our model of institutional change is as follows. Each y_i $(i = 1, ..., m)$ represents an independent indicator of institutional change, denoted y^*, such that we may write

$$y_j = \Lambda_j^y y^* + \varepsilon_j,\ j = 1, \ldots, m, \tag{2}$$

where $\mathbf{\Lambda}^y = \{\Lambda_1^y, \Lambda_2^y, ..., \Lambda_j^y\}'$ denotes a $m \times 1$ parameter vector representing the expected change in the respective indicators following a one unit change in the latent variable. $\mathbf{\Theta}_\varepsilon$ denotes the covariance matrix of $\boldsymbol{\varepsilon}$, and initially we assume that $cov(\varepsilon_j, \varepsilon_i) = 0\ \forall i \neq j$ such that any correlation across the indicators is driven by the common factor y^*. Given these covariance assumptions we may think of (2) as a factor analysis model for the observable indicators $\mathbf{y} = (\mathbf{y}_1, y_2, ..., y_m)'$. In addition it is worth emphasising that the unique factor is actually composed of two components: a specific factor, say $\mathbf{q}$ and a pure random measurement error $\mathbf{e}$. With the exception of studies providing multiple records for each observation (i.e. panel data), it is not possible to separately identify the components of $\boldsymbol{\varepsilon}$. Finally, we let $\boldsymbol{\tau}$ denote the $m \times 1$ vector of diagonal elements of $\mathbf{\Theta}_\varepsilon$. We also posit that the institutional change is linearly determined by a vector of observable exogenous variables $\mathbf{x} = (x_1, x_2, .., x_s)'$, a latent time invariant component η_0 reflecting initial conditions, and a stochastic error ζ, giving

$$y^* = \gamma_1 x_1 + \gamma_2 x_2 + ... + \gamma_s x_s + \eta_0 + \zeta. \tag{3}$$

We write the country specific component η_0 as

$$\eta_0 = \boldsymbol{\alpha}' \mathbf{w} \tag{4}$$

where $\mathbf{w}$ is a vector of country characteristics which in combination reflect the initial conditions prior to transition, and $\boldsymbol{\alpha}$ is a vector of weights.

Examining (2) and (3) we may think of the model as comprising two parts: (3) is the structural (or state) equation and (2) is the measurement equation reflecting that the observed measurements are imperfect indicators. (2) is a special case of a factor analysis

[15] LISREL is an acronym for *linear structural relationships.*

model with a set of m observable indicators determined linearly by a single unobserved (common) factor - institutional change. The measurement equation specifies how the observed endogenous variables are determined by the unobservable construct institutional change together with a vector of predetermined variables $\mathbf{z}_j$. Any correlation between the elements of $\mathbf{y}$ results from the mutual association with y^*. The assumption that the partial covariance between all i, j indicator pairs should be zero is enforced by setting the off-diagonal elements of Θ_ε equals to zero. The structural equation specifies the causal relationship between the observed exogenous causes and institutional change[16].

Since y^* is unobserved it is not possible to recover direct estimates of the structural parameters $\boldsymbol{\gamma}$. However if we combine these two equations and solve for the reduced form representation, we may write

$$\mathbf{y} = \boldsymbol{\pi}\mathbf{x} + \mathbf{v}, \tag{5}$$

where $\boldsymbol{\pi} = \boldsymbol{\Lambda}^y\boldsymbol{\gamma}'$ is the $m \times s$ reduced form coefficient matrix and $\mathbf{v} = \boldsymbol{\Lambda}^y\zeta + \boldsymbol{\varepsilon}$ is the reduced form disturbance with covariance matrix

$$\boldsymbol{\Theta}_v = E[(\boldsymbol{\Lambda}^y\zeta + \boldsymbol{\varepsilon})(\boldsymbol{\Lambda}^y\zeta + \boldsymbol{\varepsilon})'] = \sigma_\zeta^2\boldsymbol{\Lambda}^y\boldsymbol{\Lambda}^{y\prime} + \boldsymbol{\Theta}_\varepsilon. \tag{6}$$

σ_ζ^2 is the variance of the structural stochastic error ζ. Note that the structure of the reduced form covariance matrix $\boldsymbol{\Theta}_v$ is characteristic of factor analysis models where the correlations between the observed variables (here indicators) are accounted for by the unobserved (common) latent variable. In this instance the common factor is ζ, $\boldsymbol{\varepsilon}$ denotes the vector of unique factors, and $\boldsymbol{\Lambda}^y$ the vector of factor loadings.

As presented, equations (5) and (6) are indeterminate since the reduced form parameters are invariant to a transformation given by, for example, $\boldsymbol{\Lambda}^y\kappa$, $\boldsymbol{\gamma}/\kappa$ and σ_ζ^2/κ, where κ is a scalar. This follows directly from the fact that institutional change is not directly observable. In order to be able to interpret parameter estimates it is necessary to define the origin and unit of measurement. We may do this in one of two ways. An approach which is used in factor analysis is to standardise the latent variables to have unit variance. Alternately, we may fix a non-zero coefficient in Λ^y, such that the unit of measurement for y^* is defined relative to one of the observed indicator variables. In this application we choose the former approach.

We also note the existence of a maintained assumption, namely that the set of indicators are valid proxies for institutional change. At this juncture the comments of Cliff (1983) in terms of the possible gap between observed indicators and the latent variable are pertinent. This problem has been referred to as the *nominalistic fallacy,* and concerns

[16] More general forms of (2) and (3) are possible including models which specify a measurement equation for $\mathbf{x}$ and models which allow elements of $\mathbf{y}^*$ to appear on the RHS of (3).

the problem of interpreting correlations and model output derived from the latent variable model as if they correspond directly to relations of the unobservable theoretical construct. In this context the recognition that all models are no more than approximations to an unknown truth is particularly relevant when the endogenous variable is not observed.

Based upon equations (5) and (6) there are two sets of restrictions on the reduced form. First, the $m \times s$ coefficient matrix $\boldsymbol{\pi}$ has rank 1, since the ms elements of $\boldsymbol{\pi}$ are expressed in terms of the $m+s$ elements of $\boldsymbol{\Lambda}^y$ and $\boldsymbol{\gamma}$. Second, the $m \times m$ covariance matrix $\boldsymbol{\Theta}_v$ represents the sum of a rank one matrix and a diagonal matrix, $\boldsymbol{\Theta}_\varepsilon$. The $m(m+1)/2$ unique elements of $\boldsymbol{\Theta}_v$ are expressed in terms of the $1+2m$ elements of $\boldsymbol{\Lambda}^y$, σ_ζ^2, and $\boldsymbol{\tau}$.

The question of identification can be addressed by examining the covariance matrix of the observed data, namely $\mathbf{z} = (\mathbf{y}'\mathbf{x}')$.[17] For the model given by (3) and (5) the $(m+s) \times (m+s)$ covariance matrix may be written

$$\boldsymbol{\Sigma}(\boldsymbol{\omega}) = \left[\begin{array}{ll} \boldsymbol{\Lambda}^y(\boldsymbol{\gamma}'\boldsymbol{\Phi}_x\boldsymbol{\gamma} + \boldsymbol{\sigma}_\zeta^2)\boldsymbol{\Lambda}^{y'} + \boldsymbol{\Theta}_\varepsilon & \boldsymbol{\Lambda}^y\boldsymbol{\gamma}'\boldsymbol{\Phi}_x \\ \boldsymbol{\Phi}_x\boldsymbol{\gamma}'\boldsymbol{\Lambda}^y & \boldsymbol{\Phi}_x \end{array} \right], \tag{7}$$

where $\boldsymbol{\omega}$ represents the vector of independent parameters in $\boldsymbol{\Lambda}^y$, $\boldsymbol{\gamma}$, $\boldsymbol{\Theta}_\varepsilon$ and σ_ζ^2. The diagonal elements are given by:

$$E[\mathbf{y}\mathbf{y}'] = \boldsymbol{\Lambda}^y(\boldsymbol{\gamma}'\boldsymbol{\Phi}_x\boldsymbol{\gamma} + \boldsymbol{\sigma}_\zeta^2)\boldsymbol{\Lambda}^{y'} + \boldsymbol{\Theta}_\varepsilon \tag{8}$$

and $E(\mathbf{x}\mathbf{x}') = \Phi_x$, with off-diagonal elements

$$E[\mathbf{x}\mathbf{y}'] = \boldsymbol{\Lambda}^y\boldsymbol{\gamma}'E[\mathbf{x}\mathbf{x}']. \tag{9}$$

(9) may be trivially rewritten in terms of the reduced form parameters, namely $\boldsymbol{\pi} = \boldsymbol{\Lambda}^y\boldsymbol{\gamma}' = E[\mathbf{x}\mathbf{x}']^{-1}E[\mathbf{x}\mathbf{y}']$.

The issue of identification is based upon whether the information contained in $\boldsymbol{\Sigma}$ is sufficient to deliver a unique set of values in $\boldsymbol{\omega}$. Since equation (9) expresses the $q = ms$ observable moments in terms of the $p = m+s$ structural parameters, then if $q - p \geq 0$ the set of mean parameters will be identified. If this condition holds, then the remaining parameters in equation (8) will be identified. Combining these two conditions, a necessary condition for identification of all parameters, p is

$$p \leq \frac{1}{2}(m+s)(m+s+1). \tag{10}$$

For the model to be exactly identified then there exists one and only one combination of the independent parameters in $\boldsymbol{\Lambda}^y, \boldsymbol{\gamma}$, and Θ_ε which generates $\Sigma(\boldsymbol{\omega})$[18]. A sufficient condition for the MIMIC model to be identified is that $m \geq 3$ and $s \geq 1$.

[17] See Goldberger (1973) and Wiley (1973), both in Goldberger and Duncan (1973).
[18] See Robinson (1974) for a discussion of identification with *multiple* latent variables.

5.2 Non-continuous Measurements

In the discussion to date we have implicitly assumed that all elements of $\mathbf{z}$ are continuous variables. In this study we have a mixture of ordinal and continuous measurements.[19] The principal problem here is that ordinal variates do not have a origin or a unit of measurement, and as such means, variances and covariances have no real meaning. In Appendix 1 we list the indicators of institutional change which in each case represent data on ordered categories. In light of this (2) does not represent a valid measurement equation. To see why consider the case of a single ordinal indicator y^o. Assuming that the underlying and unobserved latent variable is continuous, then it is apparent that the relation $y^o = \Lambda^y y^* + \varepsilon$ will only hold by chance. Bollen (1989) notes two additional and related problems. First, if the covariance structure hypothesis $\mathbf{\Sigma}^s = \mathbf{\Sigma}(\boldsymbol{\omega})$, where Σ^s is the observed covariance matrix, holds for data $\mathbf{z}$ when all variables are continuous, then this will not be the case when some data are ordinal. If, for example, the data are distributed multivariate normal then all the information is contained in the first and second moments. Second, the impact of excessive kurtosis and skewness in ordinal counterparts of normal variates results in inordinately high chi-square values, thereby generating invalid inference. Further information is available from a number of Monte Carlo studies. For example, Wylie (1976) and Olsson, Drasgow, and Dorans (1981) suggests that Pearson correlation coefficients between categorical measures are less than those between continuous counterparts, with the extent of attenuation an inverse function of the number of categories. A simulation study by Johnson and Creech (1983) found that if no account was taken of the use of ordinal data, the parameter estimates from equations (2) and (3) will be biased downwards, although by a small magnitude.

In this study we construct a version of $\mathbf{\Sigma}(\boldsymbol{\omega})$, denoted $\mathbf{\Sigma}^o(\boldsymbol{\omega})$, which consists of the following elements. For two continuous variables we may use the standard Pearson correlation coefficient. In cases when one or both variates are ordinal we introduce a threshold observational rule to link the ordinal variable with an underlying continuous counterpart.[20] Letting y^o and y^c denote, respectively, the ordinal variable with k categories and the continuous counterpart, the observational rule may be written as

$$y^o = \mathbf{1}(y^c \leq a_1) + \mathbf{1}(a_2 < y^c \leq a_3)2 + \ldots + \mathbf{1}(a_{k-1} < y^c \leq a_k)k. \tag{11}$$

Thresholds a_j, $j = 1, \ldots, k$, can be estimated once we assume a distribution for y^c, which

[19]Note that since all the observed indicators are ordinal the most appropriate method to account for the fact that the unobserved construct institutional change (y^*) has no origin or unit of measurement, is to standardise y^*.

[20]See Muthen (1983) for an overview of specification and estimation issues when recorded data are comprised of both continuous and categorical data.

in the case of standard normal generates thresholds

$$a_j = \Phi^{-1}(\sum_{i=1}^{j} \frac{N_i}{N}),\ j = 1, \ldots, k,$$

where $\Phi^{-1}(.)$ is the inverse of the standard normal distribution function and N_i is the number of observations in category i. Once these thresholds (or scores) have been estimated for each ordinal variable, it is possible to calculate the appropriate (polychoric) correlation coefficients. In summary and following Muthen (1983), the estimation of model parameters based upon the combination of a measurement error model for the *latent* indicators y^c, a threshold observational rule as in (11), and the structural equation (3), generates the following stages: i) estimation of threshold values; ii) estimation of population latent correlations *given* estimated thresholds; and iii) estimation of model parameters conditional upon i) and ii).

5.3 Specification Issues

The specification of a model based upon (2) and (3) and (11) represents a number of difficulties. Beginning with the specification of a measurement error model for institutional change, it is instructive to view this equation as a special case of factor analysis - namely a variant in which the correlation of the m indicators is explained by a single common factor. In this respect we see that this particular use of a factor model is different in the sense that the number of common factors is fixed a priori. Thereafter confirmatory factor analysis can be employed to test hypotheses as to whether the factorial composition is consistent with the data.

Although we believe that the process of institutional change is dynamic the structural equation (3) is static. By including a lagged dependent variable it is possible to add a dynamic component. However, there are a number of well known problems associated with such a model, the most significant being that the autoregressive parameter is biased and inconsistent in the face of fixed effects. One way to circumvent this problem is to allow for lags in the set of causal factors and thereby allow a process of adjustment. Note that if we add y^*_{it-1} to the right hand side of (3) we incorporate both lagged $\mathbf{x}'s$ *and* lagged stochastic term ξ_{t-1}. In addition, although the artificial variable η_0 representing the set of initial conditions faced by each country at the beginning of transition is time invariant, it is likely that the *impact* of these conditions will be time dependent. To allow for the possibility of a decay profile, we examine whether a specification which incorporates a time varying parameter represents an improvement.

Given that our measures for institutional change and for liberalisation and privatisation both derive from subjective ratings by EBRD economists there exists the potential for endogeneity of these causal factors. Errrors of judgement in institutional reform might well be correlated with errors of judgement in liberalisation and privatisation, while upgrade would typically be undertaken by looking at all transition indicators at the same time. To circumvent this problem we define liberalisation and privatisation as the number of years since a specific threshold was crossed (see Appendix 3).

Finally, in the above discussion we have assumed that Θ_ε is a diagonal matrix insofar as the *partial* correlation of indicators of institutional change is zero. In this particular study we relax this assumption given that the EBRD is the primary source for all these indicators, and therefore there may be both contemporaneous and temporal correlation across the measurement errors.

5.4 Estimation

The fundamental hypotheses for structural equations models is that the covariance matrix of the observed variables, Σ^s, may be parameterised based upon a given model specification.

A general form of a measure of fit between Σ^s and $\Sigma(\boldsymbol{\omega})$, where $\boldsymbol{\omega}$ is a vector of free parameters, may be written as

$$\begin{aligned} F(\boldsymbol{\omega}) &= (\mathbf{s}-\boldsymbol{\sigma})'\mathbf{W}^{-1}(\mathbf{s}-\boldsymbol{\sigma}) \\ &= \sum_{g=1}^{k}\sum_{h=1}^{s}\sum_{i=1}^{k}\sum_{j=1}^{i} w^{gh,ij}(s_{gh}-\sigma_{gh})(s_{ij}-\sigma_{ij}), \end{aligned} \tag{12}$$

where $\mathbf{s} = (s_{11},\ s_{21}, ..., s_{kk})'$ is a vector of the elements of the lower half of Σ^s and $\boldsymbol{\sigma} = (\sigma_{11}, \sigma_{21}, \sigma_{22}, ..., \sigma_{kk})'$ is the corresponding vector of $\Sigma(\boldsymbol{\omega})$. $w^{gh,ij}$ is a typical element of a positive definite matrix $\mathbf{W}^{-1}$ of order $q \times q$ where $q = m + s$. In the context of weighted least squares $w_{gh,ij}$ is a consistent estimate of the asymptotic covariance between s_{gh} and s_{ij}. If the original $\mathbf{z}$ are distributed multivariate normal (i.e. Σ^s has a multivariate Wishart distribution) then the general form for the asymptotic covariance matrix of Σ^s may be written

$$\lim_{n\to\infty} Cov(s_{gh}, s_{ij}) = (1/N)(\sigma_{gh}\sigma_{hj} + \sigma_{gj}\sigma_{hi}). \tag{13}$$

However, departures from normality such as excess kurtosis, result in a more complex expression for (13). A GLS version of (13) which assumes the asymptotic covariance of Σ^s is given by $\Sigma(\boldsymbol{\omega})$ has elements $w_{gh,ij} = (1/N)(s_{gi}s_{hj} + s_{gj}s_{hi})$. Browne (1984) has provided a number of extensions to classical theory, proposing an asymptotic distribution free estimator with typical element

$$w_{gh,ij} = v_{ghij} - s_{gh}s_{ij}, \tag{14}$$

where v_{ghij} are fourth order central moments. As noted by Joreskog and Sorbom (1996) the principal limitation of this approach is one of dimensionality, given that an estimate of $\mathbf{W}$ based upon (14) requires computation of $q(q+1)/2$ fourth order moments, which require large samples.

Under normality, model parameters are estimated based upon minimising the function

$$F = \ln|\underbrace{\Sigma(\boldsymbol{\omega})| - \ln|\Sigma^s}_{A}| + \underbrace{tr(\Sigma^s\Sigma(\boldsymbol{\omega})^{-1}) - (m+s)}_{B}, \tag{15}$$

where m (s) are the number of endogenous (exogenous) variables, and $tr(.)$ denotes the trace operator. As the elements of the predicted and observed covariance matrices converge in probability ($\rightarrow$) then terms A and B will both approach zero, given that $tr(\Sigma^s\Sigma(\boldsymbol{\omega})^{-1}) \rightarrow m+s$.

In this paper we estimate two model variants based upon (15): model 1 using $\Sigma(\boldsymbol{\omega})$ and model 2 using $\Sigma(\boldsymbol{\omega})^0$, where the latter covariance matrix incorporates an adjustment (using the threshold observational rule) for the ordinal structure of the data.

6 Results

The data for this research is derived from a cross-country database assembled at the EBRD (see Di Tommaso and Weeks (1999) for details). The individual variables were discussed in preceding sections and a full list with definitions is provided in the Appendices. The data cover a total of 25 countries over the period 1989 to 98 but data is not available for every year for all variables. Moreover, for several countries records are incomplete, particularly during earlier reform years.

In Table 4 we present 4 correlation matrices (pooled over the five years 1995-1999), which provide a useful summary of the key relationship among our ordinal indicators and identified causes of institutional change. In Matrix A we present the lower diagonal elements of the polychoric correlation matrix for the five ordinal indicators of institutional change. Since we believe that these indicators share a common component, namely institutional change, then the high pairwise correlations are not surprising. We note a particularly high correlation between enterprise reform and restructuring (ER_ST) and banking reform and interest rate liberalisation (BR_I). Matrix B presents the matrix of polychoric correlations for the ordinal indicators of institutional change and our measures of small scale privatisastion (ysSCP), price liberalisation (ysPL) and macroeconomic stability (ysIBD). In all cases the correlations reveal a relatively strong association between these causes and the ordinal indicators, with a particularly strong association for macroeconomic stability.

The correlations between the institutional change indicators and the measures of structural change (direction of trade and composition of employment) are presented in Matrix C. It also presents the correlation with government expenditures as the key indicator of state capacity. The correlations with exports to non-transition economies and government expenditures are strong and positive as expected. The correlation with our measure of employment structure is however unexpectedly signed. The positive correlation suggests that countries, which maintained a higher share of employment in industry relative to a benchmark given by their per capita income have achieved more institutional change. This reflects the continuing high level of industrial employment in many central European countries, whereas industrial employment has declined as a share of the total in most of the CIS.

Finally in Matrix D we present the correlations between the first component extracted from our principal component analysis of initial conditions (IC), together with a variable $IN_t = IC \times D_t, t = 95, 96, 97$ and 98, where D_t denotes a time dummy. Note that the initial conditions index is negative for more favourable starting points (see 2) and as such the negative correlations are therefore expected. There is a tendency for the correlation between initial conditions and institutional change to decline over time, although in a number of cases the pattern is non-monotonic.

The parameters in the MIMIC model of institutional change given by (2) and (3) are identified given that the identification condition (10) is satisfied[21]. Parameter estimates are given in Table 5. The Table has two columns representing the structural equation (left column) and the measurement equation respectively. Results in the left column can be interpreted as standard regression coefficients. Results in the right column contain the standardised the elements of $\mathbf{\Lambda}^y$ and may be interpreted as standardised factor loadings. For the measurement equation we also present the squared multiple correlation statistic, which provides a measure of the reliability of each indicator as a measure of unobservable institutional change.

As can be seen from Table 5, there is not much difference in the factor loadings across the five dimensions of institutional change, and the goodness of fit is also quite similar. This finding confirms the attention that EBRD places on all five dimensions of institutional change.[22] The largest weight is given to the measure of legal transition, confirming the centrality of legal and judicial change to the transition process. Although we start from an initial assumption of zero contemporaneous correlation across our five indica-

[21] The number of unique elements in $\sum^s$, $\frac{1}{2}(m+s)(m+s+1)$, is 78. The number of estimated paramters, p, is 12, and so the model is overidentified by 66 degrees of freedom.

[22] In a future paper we intend to test the MIMIC model against its OLS (using an unweighted average over the five indicators) and SUR alternatives.

tors (i.e. all off diagonal elements of Θ_ε were set to zero), given that all indicators were scored by EBRD staff, we checked our model for systematic cross-indicator errors. Under the heading *Covariance Parameters* we report the only significant covariance parameter, namely that between the banking reform and interest-rate indicator (BR_I) and the indicator of governance and enterprise restructuring (ER_ST).

The parameters of the structural equation (γ) are consistent with prior expectations, with the exception of the measure of employment distortions (see discussion further above). Market liberalisation and small scale privatisation have a positive impact on institutional change, although the impact is relatively small. A country that had not achieved a single year of price liberalisation (or small scale privatisation) since the start of transition a decade ago would, *ceteris paribus*, have a score for institutional change 0.8 (1.0) points below a country that fully liberalised (or privatised) in year one. This difference is roughly equal to the difference in 1999 between Hungary and Russia.

Exports to non-transition economies help institutional reform but the impact is very small. An increase in the share of exports to non-transition economies from 20-50 per cent for instance would explain only around 0.03 points difference in the institutional reform ratings. The effect of employment distortions is wrongly signed but significant. The difference between the most and the least distorted country in the sample would account for some 0.02 points difference in institutional reform.

Political liberalisation is also significantly associated with institutional change with a relatively large impact. The difference between no and full political rights would help to account for a full 1.3 points difference in institutional reform. The impact of macroeconomic stabilisation is also sizeable - around 1.5 times as large as that of liberalisation.[23] In other words, stabilising early has considerable pay-offs in terms of promoting subsequent institutional reforms. Maintaining a higher share of public expenditures in GDP also eases reforms (if internal balance is not compromised as a result). Every 10 percentage points increase in public expenditure over GDP would yield a 0.02 point increase in institutional reform - again a very modest impact. It seems on inspection of these results that economic reforms and political liberalisation are far more powerful determinants of institutional change than changes in the structure of the economy or in state capacity. We will check this result further in future research drawing on the full 1989-2000 EBRD dataset.

Although the set of initial conditions are fixed by construction, by interacting the first principle component with year dummies it is possible to allow for the possibility of

[23]Remember that both variables are defined the same way as the number of years with a certain threshold achieved.

a decay profile. In doing this we consider two competing models: Model A imposes both a fixed set of initial conditions and a fixed *effect* over the course of transition, and model B imposes a fixed set of initial conditions but allows for a decay profile in the effect. By comparing the χ^2 measure of fit for these two models,[24] we find that Model B represents a substantial and statistically significant improvement over Model A.

Of particular interest are the parameter estimates IN_t, $t = 95, 96, 97$, and 98 which represent the (lagged) impact of initial conditions relative to their impact in 1999 (see Table 5). As expected the largest effect of differences in starting points is felt early on, although we observe a non-monotonic profile. The difference between the country with the most favourable initial conditions (the Czech Republic) and the country with the worst starting point (Tajikistan) in the fixed set is sufficient to explain around one full point difference in institutional reforms in 1995, but only a third point difference in 1996, before the difference widens again to around two thirds of a point by 1998. With the new full 1989-2000 dataset of EBRD transition indicators it will be possible to examine more rigorously whether there is convergence or divergence in institutional performance across the transition economies over time.

The results in Table 5 are based on a model correcting for the ordinal nature of the institutional change indicators. In Table 6 we present summary measures of fit for model 1 (without correction) and model 2 (with correction). In interpreting these statistics we first consider an important caveat. Based upon the χ^2 statistic reported in row 1, it is possible to conduct a test of the fundamental covariance hypothesis, namely that $\Sigma^s = \Sigma(\boldsymbol{\omega})$. However, as Joreskog and Sorbom (1996) emphasise, the use of this statistic for such a test will in many circumstances not be valid[25]. A more appropriate use is to simply use this statistic as a measure of fit, with, as demonstrated by equation (15), large values, relative to degrees of freedom, indicating a poor fit. The overall fit of the model, $\chi^2 = 32.1$ with 62 degrees of freedom and a probability value of 0.953, provides clear evidence that model 2 based upon including a threshold observational rule to link the ordinal observed indicators to their continuous counterparts, is consistent with the data. In comparing model 2 with model 1, where model 1 takes no account of the ordinal data, we note that in both absolute and relative terms, model 1 represents a poor fit. We also perform a test of the null model given by the restriction $\boldsymbol{\gamma} = \mathbf{0}$. This restriction gives us the one factor measurement error model and implies the absence of causal links driving

[24] Note that we do not report the full set of parameter estimates for Model A. These are available from the authors upon request.

[25] The χ^2 measure reported in Table 5.6 is particularly sensitive to departures from multivariate normality. However, in model 2 by introducing a threshold obervational rule, we have accounted for these problems introduced by ordinal data. We also note that such departures tend to inflate the χ^2 statistic, and thus in this case there does not seem to be a problem.

institutional change. For both model 1 and model 2 this hypothesis is rejected.

Finally we note that although in the case of model 2 the reported χ^2 value corresponds to a good fit, given the existence of a large class of equivalent models which may fit the data equally well, we should not interpret this as evidence that it represents the best approximation to the *true* model.

7 Conclusion

This paper has provided a first empirical analysis of the factors driving institutional change in the transition economies. To this end, we have developed a MIMIC model of institutional change taking into account potential measurement error in our institutional change variables. Moreover, this approach has allowed us to focus on institutional change as an unobserved multi-dimensional process and has yielded an empirical approximation of this process as a weighted average of subjective progress indicators. We intend to explore this approach further in future research by testing it rigorously against alternatives, such as simple OLS or Seemingly Unrelated Regression. The weights generated by the MIMIC model on the five dimensions of institutional change suggest that a simple average may not be a bad approximation, but this needs to be tested further.

With regard to the determinants of institutional change, we find strong evidence that economic reforms and political liberalisation are more powerful forces than changes in economic structures induced by such reforms. In other words, we do find some support for the view in the early transition literature suggesting a positive spill-over from "early reforms" such as liberalisation and privatisation into institutional reforms. However, this effect operates with a significant lag. After 10 years of transition, a country with no progress in either liberalisation, privatisation or macroeconomic stabilisation would on aggregate have an institutional rating of 3 full points below a country that had introduced all of these reforms in year one. In other words, the current laggards in transition would be expected to take around a decade in order catch up with the present leaders.

The results in this paper remain preliminary. With a longer time series of economic reform ratings by EBRD now available, the conclusions in this paper need to be reviewed. It is put forward as a first methodological exploration into the kinds of issues faced by students studying a process, where exact measurement will remain elusive.

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Table 1: EBRD Transition Indicators: 1995 and 1999

Countries	Governance and enterprise restructuring		Competition policy		Banking reform and interest rate liberalisation		Securities markets and non-bank financial institutions		Overall Legal Effectiveness and Extensiveness	
	1995	1999	1995	1999	1995	1999	1995	1999	1995	1999
Albania	2	2	1	2	2	2	1	2	2	2+
Armenia	2	2	1	2	2	2+	1	2	2	3
Azerbaijan	2	2	1	1	2	2	1	2-	1	2-
Belarus	2	1	2	2	2	1	2	2	2	1+
Bulgaria	2	2+	2	2	2	3-	2	2	3	3-
Croatia	2	3-	1	2	3	3	2	2+	3	3-
Czech Republic	3	3	3	3	3	3+	3	3	4	3
Estonia	3	3	3	3-	3	4-	2	3	3	4-
FYR Macedonia	2	2	1	1	3	3	1	2-	2	2+
Georgia	2	2	1	2	2	2+	1	1	2	1
Hungary	3	3+	3	3	3	4	3	3+	4	4
Kazakhstan	1	2	2	2	2	2+	2	2	2	3-
Kygyzstan	2	2	2	2	2	2+	2	2	2	2+
Latvia	2	3-	2	3-	3	3	2	2+	2	3-
Lithuania	2	3-	2	2+	3	3	2	3-	2	3-
Moldova	2	2	2	2	2	2+	2	2	2	3
Poland	3	3	3	3	3	3+	3	3+	4	4
Romania	2	2	1	2	3	3-	2	2	2	3-
Russian Fed.	2	2-	2	2+	2	2-	2	2-	2	3-
Slovak Republic	3	3	3	3	3	3-	3	2+	3	3+
Slovenia	3	3-	2	2	3	3+	3	3	3	3+
Tajikistan	1	2-	1	1	1	1	1	1	1	na
Turkmenistan	1	2-	1	1	1	1	1	1	1	na
Ukraine	2	2	2	2	2	2	2	2	2	2
Uzbekistan	2	2	2	2	2	2-	2	2	2	2-

Table 2: Initial Conditions Analysis: Factor Loadings

	Eigenvectors	
	Fixed Set	
	Prin1	Prin2
GDP_PC	-0.361	-0.007
STATE	-0.283	0.312
RICH	0.276	0.206
URBAN	-0.251	0.278
LOCAT	0.455	-0.238
MARMEN	0.418	-0.019
ETHNIC	0.318	-0.000
REL1	-0.401	-0.411
REL2	0.068	0.746
Proportion of variance	0.43	0.17

Table 3: Initial Conditions: Country Scores

		Fixed set
Albania	CEE	-0.538
Armenia	CIS	-0.716
Azerbaijan	CIS	2.038
Belarus	CIS	0.227
Bulgaria	CEE	-1.223
Croatia	CEE	-1.794
Czech Republic	CEE	-2.928
Estonia	CEE	-1.023
FYR Macedonia	CEE	-0.08
Georgia	CIS	1.808
Hungary	CEE	-3.046
Kazakhstan	CIS	2.529
Kyrgyzstan	CIS	2.685
Latvia	CEE	-1.025
Lithunia	CEE	-1.511
Moldova	CIS	0.919
Poland	CEE	-1.94
Romania	CEE	-0.455
Russia	CIS	0.469
Slovak Republic	CEE	-1.934
Slovenia	CEE	-2.625
Tajikistan	CIS	2.918
Turkmenistan	CIS	2.906
Ukraine	CIS	1.310
Uzbekistan	CIS	3.027

Table 4:

(A)

	ER_ST	COMP_POL	BR_I	O_LE	SNB
ER_ST	1.000				
COMP_POL	0.854	1.000			
BR_I	0.966	0.685	1.000		
O_LE	0.723	0.690	0.648	1.000	
SNB	0.759	0.776	0.681	0.772	1.000

(B)

	ysSCP	ysPL	ysIBD
ER_ST	0.357	0.526	0.581
COMP_POL	0.453	0.444	0.514
BR_I	0.382	0.500	0.616
O_LE	0.479	0.575	0.572
SNB	0.509	0.451	0.601

(C)

	EXSHE_Lg	EMSIN_Lg	PUBE_GDP
ER_ST	0.631	0.625	0.729
COMP_POL	0.417	0.566	0.664
BR_I	0.536	0.425	0.565
O_LE	0.739	0.598	0.724
SNB	0.652	0.671	0.822

(D)

	IC	IN95	IN96	IN97	IN98
ER_ST	−0.774	−0.665	−0.463	−0.352	−0.300
COMP_POL	−0.588	−0.190	−0.302	−0.338	−0.352
BR_I	−0.582	−0.500	−0.332	−0.380	−0.281
O_LE	−0.757	−0.500	−0.332	−0.439	−0.414
SNB	−0.764	−0.500	−0.329	−0.423	−0.418

See Appendices 1 and 3 for variable definitions.

Table 5: Mimic Model of Institutional Change: Parameter Estimates

	Structural Equation γ	Indicators	Measurement Equation R^2
No. years with inflation<30% and budget deficit <5% of GDP	0.130 (0.039)	Governance and Enterprise Restructuring	0.395
Years with small scale privatisation	0.102 (0.037)	Competition Policy	0.291
Years with price liberalisation	0.081 (0.038)	Banking Reform and Interest Rate Liberalisation	0.302
Average score for political factors	-0.185 (0.043)	Securities Markets and Non-bank Financial Institutions	0.373
Lag of exports to EU as a share of total exports	0.082 (0.038)	Legal Transition Indicators	0.416
		Covariance Parameters	
Lag of the share of employment in industry respect to a market economy benchmark	0.078 (0.039)	Θ_ε(Gov.Ent.Rec,Bank.Ref.)	0.267 (0.096
Initial conditions 95	-0.139 (0.038)		
Initial conditions 96	-0.058 (0.037)		
Initial conditions 97	-0.080 (0.036)		
Initial conditions 98	-0.079 (0.039)		
Government expenditure to GDP	0.183 (0.041)		

$R^2 = 0.814$
Numbers in parenthesis denote standard errors

Table 6: Measures of Fit

	Model 1	Model 2
χ^2 FIT	126.03	32.07 (0.953)
χ^2 (independence)	1173.99	842.60
Goodness of Fit Index (GFI)	0.891	0.984
Adjusted Goodness of Fit Index (AGFI)[1]	0.629	0.955
Root Mean Square Residual (RMS)	0.034	0.043
Null model: $\boldsymbol{\gamma} = \mathbf{0}$	367.06	457.88

Notes:
The test statisitic χ^2 FIT is equal to $NT \times F_0$, where F_0 is the mimimum value of F in (10) and NT is sample size
$\chi^2_{62,0.95}$ *critical value* $= 82$

GFI[2] $= 1 - [tr(S{-}\mathrm{I}_{m+s})^2/tr(\Sigma(w)^{-1}S)^2]$

(I_{m+s} is an identity matrix of dimension m+s)

AGFI $= \frac{1-(m+s)(m+s+1)}{2d}(1 - GFI)$
(d = degrees of freedom)

RMS = square root of the squared discrepancies betweeen elements of $\Sigma(w)$ and Σ^s

[1] The adjustment is based upon using mean squares in the numerator and denominator.

[2] Note that under the fundamental covariance hypothesis $\Sigma(\omega) \rightarrow \Sigma^s$ $tr(\Sigma^{-1}S) \rightarrow tr(\mathrm{I}_{m+s})$, and therefore GFI goes to 1.