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Abstract

In this paper we re-examine empirically the Structure-Conduct-Performance relationship between concentration and profitability using new data on Russia that allow us to overcome the endogeneity problem of market structure and expand on the traditional analysis in several ways. The analysis yields several important results. We find strong evidence that national concentration does increase profitability, but only in geographically dispersed industries, suggesting that regional markets are an important source of market power for oligopolistic industries. We find preliminary evidence that the effects of market structure are persistent in the long run. And we find for Russia that capital intensity is negatively related to profitability, suggesting that capital-intensive industries face higher adjustment costs and thus are less profitable after recessionary shocks.

JEL classification: L1; P2.

Keywords: Market Structure; Concentration; Competition; Structure Conduct Performance; Transition

1. Introduction

For decades industrial economists following the Structure-Conduct-Performance paradigm have been asking the question whether market structure matters to industry profitability.[1] The empirical results have been inconsistent and inconclusive. In spite of inconclusive evidence on the effects of market structure, however, recent analysis in industrial organization has focused on the determinants of market structure and the evolution of industry more generally.[2] As long as we are interested in the determinants of market structure though, the question of how and why it matters remains important. In this paper, we analyze Russian data that allow us to overcome serious empirical problems that have plagued previous studies and thus shed new light on the relationship between market structure, or concentration, and profitability.

Recent empirical studies using data from different countries include Salinger (1990), Conyon and Machin (1991), Wizarat (1992), Lai (1994), Conyon (1995), and Bhattacharya and Bloch (1997). Most, but not all, of these find a positive and significant relationship between concentration and profitability. Two empirical problems continue to hinder the empirical analysis, though. First and most important, in market economies, market structure is endogenous to profitability making it difficult to test for a causal relationship. Schmalensee (1989, p. 254) concludes "derived market structure is clearly affected by market conduct in the long run," and thus "all variables that have been employed in such studies are logically endogenous". Machin and Conyon (1991) and Conyon (1995) try to avoid this by using instruments to estimate

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[1] See Schmalensee (1989) and Hay and Morris (1991) for excellent reviews of this literature.
[2] See, for example, Sutton [1991 and 1997].

their equations, but their instruments are simply lagged values of concentration measures.

Second, researchers have been confined by the statistics on market structure published by governments. These statistics are restrictive in at least three ways: they usually only offer one measure of industrial concentration (for example, the four-firm concentration ratio); they only measure concentration at the national level; and they use imperfect definitions of markets. In fact, very few economists have analyzed the impact of the geographic dispersion of firms on the relationship between market structure and profitability; almost everyone assumes that the relevant market for analysis is the whole country. Collins and Preston (1969) construct a dispersion statistic over four regions of the United States and find that dispersion alone has a positive effect on profitability. Kwoka (1979) finds a similar result. Others find no effect of dispersion. No recent studies have examined, or even controlled for, spatial dispersion. Hay and Morris (1991, p. 234) conclude: "It remains unclear, therefore, just how important geographical considerations are in determining profits."

The data on Russia allow us to overcome these empirical problems. When Russia became a market economy, it inherited from the Soviet period an industrial structure that was largely determined by central planners. Thus market structure in Russia early in transition is truly exogenous to profitability.[3] Indeed, we see that there are important differences between Russian industrial structure early in transition and that found in many market economies. We use an economy-wide database on Russian firms, rather than government data on industries, which allows us to calculate several different concentration statistics. Further, we have information on firms' products and on firms' locations so that we can analyze different market definitions, in particular we test whether the dispersion of production changes the impact of nationally-measured concentration on profitability.

Russia provides a fascinating subject for this study not only because the data are so powerful, but also because it is an economy in transition. Even when we find a short-run equilibrium relationship between market structure and profitability, we must realize that most of these industries are not in long-run equilibrium. Thus we can also test for the impact of inherited market structure on longer-term profitability. Salinger (1990), for example, finds that concentration has a long-term as well as short-term impact on profitability in the United States.

We present robust evidence that market structure does matter to industry profitability when market structure is exogenous, but only when the geographic dispersion of production is considered—nationally-measured concentration increases profitability when industries' firms are dispersed across the population. This finding suggests that the market power of oligopolistic industries arises from monopoly power in regional markets rather than from tacit collusion, which one might hypothesize would be facilitated by firm clustering. We also find evidence that market structure effects are persistent over time. We do not find evidence that the government's definitions of industries (and thus markets) impedes analysis of market structure. We find strong evidence that, in Russia during transition, there is a negative relationship between capital intensity and profitability supporting the general hypothesis that capital-intensive industries adjust more slowly (see Caves, 1990). Finally, we present

[3] The profitability of firms becomes market determined very quickly, as soon as prices are market determined, which happened in Russia with the rapid liberalization of prices in January 1992. But clearly it takes much longer for market structure to change significantly.

preliminary evidence that import penetration does reduce industry profitability in transition economies.

In the next section, we present and discuss the specifications that we estimate and, in the following section, discuss the analysis of industry coverage and geographic dispersion. We briefly describe the dataset and methodology in section four and our results in section five.

2. Specification

2.1 Oligopoly model

Conyon and Machin (1991) present a modified Cowling-Waterson (1976) general oligopoly model for the theory behind the hypothesized positive relationship between concentration and the price-cost margin. Take a firm with a profit function $\Pi_{ij} = (P_j - AC_{ij}(q_{ij}))q_{ij} - FC_{ij}$, where Π_{ij} is profits, $AC_{ij}(q_{ij})$ is average cost, q_{ij} is output, and FC_{ij} is fixed costs for firm i in industry j. P_j is the price of output in industry j, which is a negative function of industry output, $Q_j = \Sigma q_i$ for i in j. The first order condition for profit maximization is

$$\frac{\left(P_j - AC_{ij}\right)}{P_j} = \frac{S_{ij}}{\varepsilon_j}\left(1 + \lambda_{ij}\right) + \frac{\left(MC_{ij} - AC_{ij}\right)}{P_j} \tag{1}$$

where $S_{ij} = q_{ij}/Q_j$ is firm i's market share, ε_j is industry j's price elasticity of demand, and λ_{ij} is firm i's conjectural variation (that is, the expectation of the change in other firms' outputs in response to a change in own output).

We sum over all i in industry j and assume constant returns to scale ($MC = AC$) to get the equation for industry profitability:

$$PCM_j = \frac{H_j}{\varepsilon_j}\left(1+\Gamma_j\right) \qquad (2)$$

where PCM_j is the weighted sum of the price-average cost margins for firms in industry j, H_j is industry j's Herfindahl-Hirschman Index, and Γ_j is the weighted mean conjectural variation. The constant returns to scale assumption reduces the profitability equation to Cowling and Waterson's altered Lerner condition.[4] This condition predicts a positive relationship between concentration, as measured by the HHI, and industry profitability as long as $(1 + \Gamma_j) > 0$. Under Cournot conjectures, $\Gamma_j = 0$, but the condition predicts a positive relationship even under some degree of negative conjectures.

At this point we can take the log of equation (2) to get a linear equation, or as most authors do, simplify the condition to the linear estimating equation

$$PCM_j = \beta_0 + \beta_1 CONC_j + \mu \qquad (3)$$

where *CONC* is a measure of market structure, or concentration (see Salinger, 1990 or Conyon, 1995). The term for the industry price elasticity of demand is dropped. Estimates of elasticity often do not exist, and even when they do, they are usually found to perform poorly in the regressions (Salinger, 1990).

2.2 Basic specification

The major objection to using equation (3) to estimate the impact of concentration on profitability is the endogeneity problem, that is, that market structure and profitability are jointly determined. Clarke and Davies (1982) raise this objection and further point out that both the Herfindahl index and the price-cost margin may be written as a reduced form function of product market demand and costs. The industrial structure observed during early transition Russia, however, was inherited from the Soviet period and therefore determined by central planners rather than market forces. That is, there was no market conduct to determine market structure.

Because this assumption is key to our study, we discuss it in detail for readers who are less familiar with Soviet-type economies. Planners in Soviet Russia maximized a variety of objectives, including, but not limited to, cost minimization, when making decisions about building new firms. The objectives also varied over the various levels of decision-making power. Ministers, for example, often maximized the returns to the firms in their Ministries with little to no consideration for the overall efficiency of the economy.

Of course decisions about the number and size of enterprises were not totally arbitrary. Ickes and Ryterman (1997) propose a theory for the process of enterprise selection and industrial evolution under central planning. Their theory is based on two assumptions: first, that ministers did want to minimize costs, and second, that enterprises were not allowed to exit. The second assumption is non-controversial; the first less so. The model predicts that mature Soviet industries will comprise larger enterprises that are more efficient, and smaller (but not small) enterprises that are less efficient. Even if we accept this theory though, there is no reason to believe that the resulting market structure is endogenous to the demand and costs of a market economy. That is, the costs that ministers were minimizing were based on relative prices that were highly out of sync with market prices (and demands that were out of sync with consumer sovereignty). For example, transportation costs of goods and

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[4] Conyon and Machin (1991) do not assume constant returns to scale, and they construct a returns-to-scale measure in order to test the impact of this effect. Conyon (1995) retains the constant-returns-to-

energy costs were greatly undervalued during the Soviet period. Such a cost structure no doubt resulted in not only sizes of firms but also locations of firms that are inconsistent with market conditions.

Not only is it easy to understand theoretically that the inherited industrial structure is exogenous to market forces, we can also see empirically that Russia's inherited industrial structure is quite different from that found in industrial market economies. Brown, Ickes, and Ryterman (1994) use 1989 data on the population of Russian civilian industrial enterprises to compare the market structure of Soviet Russia with that of the United States in 1987.[5] They find several significant differences. First, they find an almost complete absence of small manufacturing firms in Russia, a marked contrast to the U.S. Not surprisingly given the absence of small enterprises, Brown, et. al. also find that many Soviet Russian industries have far fewer enterprises than U.S. industries have firms. In spite of that fact, Soviet Russian industries are not, for the most part, highly concentrated according to four-firm and eight-firm concentration ratios. Finally, the correlation between CR4's across U.S. and Soviet Russian four-digit industries is quite low at .06. In contrast, we know that the rank correlations between manufacturing industries' concentration levels in industrialized market economies is quite high (Schmalensee, 1989).

Another objection to the specification in equation (3) is the interpretation of β_I. The standard oligopoly model interpretation is that a positive and significant estimate indicates that firms in concentrated industries enjoy market power. Demsetz (1973) argues that a positive coefficient could also result from perfect competition. Very

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scale assumption.

[5] The Russian enterprises were assigned to industries according to U.S. SIC codes by PlanEcon in Washington D.C. so as to make the data comparable with data on U.S. industries, especially manufacturing industries.

large firms in an industry should have lower costs, and thus industries where such large firms account for a significant market share should show higher profitability. Authors have debated this efficiency rents interpretation (see Salinger, 1990 and Peltzman 1990). In the Russian case, however, the efficiency rents interpretation is hard to defend because, again, we have no evidence that the larger enterprises inherited from the Soviet period are characterized by lower market costs.[6]

One final problem with the estimation of equation (3) has been the measure of *CONC*. While the Lerner condition specifies the Herfindahl index as the measure of concentration, most studies, due to data availability, have relied on concentration ratios. Studies of the United States like Salinger's (1990) usually use the four-firm concentration ratios calculated by the Census Bureau. Others countries' statistical offices publish concentration ratios anywhere between three-firm and eight-firm. While the correlation between Herfindahl's and CR4's is quite high (Scherer and Ross, 1990), it is usually difficult to test whether they indeed yield similar estimates. Kwoka (1981) extrapolates one- through ten-firm concentration ratios and finds that although the statistics are highly correlated, the estimated coefficients and the significance of those estimates in standard regressions on profitability vary considerably. He finds that the two-firm concentration ratio gives the strongest results. In this study we are able to calculate any and all measures of concentration that we wish to test and ask the question whether the power of a given measure to explain profitability reveals

[6] In fact a quick check with the data raises serious doubts that large firms are systematically more efficient than small firms in Russia. When we compare the cost/output ratios in 1992 of the largest 25% of firms with the smallest 25% of firms (where size is defined by employment), we find that in 166 industries the large quartile has lower costs, but in the other 166 industries the small quartile has lower costs.

something about what in particular about concentration matters. In the end we focus on the CR2, CR4, and HHI.[7]

Some authors calculate "corrected" concentration measures that take into account import penetration (see Caves, 1985 for a survey). Imports are clearly endogenous to profitability though—industries with high margins attract more imports. Also, imports can respond much more quickly to market conditions than market structure can. In fact, many advisors to transition economies propose import competition as a solution to possible inherited monopolization before domestic market structure is able to adjust. Thus, because we want to test for causal effects of exogenous structure, we do not "correct" our measures of market structure to account for import penetration. Rather, at the end of the paper, we test the hypothesis that import penetration introduces competition and thus reduces industry profitability by including import penetration as a separate variable in the regressions.

2.3 Additional variables

At this point, economists traditionally add a variety of different variables to the regressions, both to control for certain effects and to test for certain effects. Given the sensitivity of estimates to different specifications as found by Bhattacharya and Bloch (1997) and many others before them, we carefully motivate the inclusion and exclusion of additional variables in our analysis.

Bain (1951, 1956) argues that barriers to entry form an important part of market structure and thus affect profitability. His concern with entry is not unrelated

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[7] For the results presented in the paper, we calculate concentration using output, but all the results are robust to calculating concentration using employment. Output, or more precisely, sales is the standard value used, but employment may be a more reliable statistic in these economies during transition when prices are changing frequently.

to the later theory of contestability introduced by Baumol, Panzer, and Willig (1982). Baumol, et. al. theorize that any firm, no matter the market structure, will set the competitive price as long as markets are contestable, that is, as long as there are no sunk costs to entry. Bain (1956) suggests three types of barriers to entry: absolute cost advantages due to incumbency (e.g. patents, supply networks), product differentiation leading to brand loyalty, and scale economies. Authors starting with Comanor and Wilson (1967) try to objectively measure the latter two barriers in order to include entry barrier variables in concentration-profitability regressions.

Product differentiation is a barrier to entry to the extent that incumbents enjoy brand loyalty, making it difficult for entrants to gain market share even with lower prices. This differentiation is usually measured by advertising expenditures, as advertising contributes to product differentiation and represents a sunk cost of entry. In Russia though, brand loyalty to incumbents is not an issue. Given the low quality of Soviet goods, buyers have no reason to maintain brand loyalty with former Soviet firms. As advertising becomes increasingly important to business in Russia, all firms—incumbents and entrants—need to advertise equally. Thus we do not include a variable for advertising.[8] Scale economies may very well be an important barrier though. As do most authors, we measure scale economies using minimum efficient scale (*MES*)—the minimum market share needed to be profitable in the industry.

We use the capital-output ratio (*CAPOUT*) to control for the bias in the price-cost margin measure of profitability. The price-cost margin does not adequately account for depreciation and does not include the cost of taking risks, both of which

[8] Another reason not to include an advertising variable is that theory also suggests that industries where consumers rely on advertising for information should actually attract more entry than industries where advertising is not important because entrants can buy advertising while they cannot buy reputation. Kessides [1986] finds that advertising has a positive effect on entry.

are proportional to capital intensity. We also include dummy variables for industry branches to control for sectoral demand shocks, although we do not report these estimates in the tables.

The basic specification so far is no different from what one might run for a market economy. While we want to test the standard relationship, we do need to consider whether there are particular transition factors that could introduce omitted variables bias into the standard regression. One factor that stands out is privatization. It is highly plausible that the speed and success of privatization varies systematically across industries, and it is also probable that there is a systematic relationship between privatization and profitability. We control for the degree of privatization using the share of industry output accounted for by firms that are private and mixed private and state ownership (*PRIVSHARE*).[9]

Thus, the basic specification we run is

$$PCM_j = \beta_0 + \beta_1 CONC_j + \gamma_1 MES_j + \gamma_2 CAPOUT_j + \gamma_3 PRIVSHARE_j + \mu \quad (4)$$

3. Market definition

As discussed above, the data for Russia help us to improve on the standard analysis by eliminating the endogeneity problem and by allowing us to calculate any desired concentration measure. In addition to these benefits, the data allow us to

[9]There are two factors that are sometimes suggested for inclusion, which do make sense theoretically for Russia, but which never once yield statistically significant coefficients estimates (or alter the other estimates appreciably), and so we leave them out of the final specifications. The first is absolute capital requirements, another barrier to entry. Bain (1956) posited that absolute capital requirements could be a barrier to entry because capital markets are imperfect. While people stopped assuming imperfect capital markets for the advanced market economies, such an assumption can certainly be made for Russia during transition. The second is risk, as measured by the variance of profitability. Transition is a period of great uncertainty, and it would seem that entrants would in fact avoid industries with variable outcomes. The variance of profitability among the incumbents may not be a good indicator of risk to entrants though, which could explain why the variable is unimportant econometrically.

tackle another major problem of the standard analysis—market definition. The definition of markets implied by the industrial classification codes used in each country can be wrong in two ways. First, the way firms are assigned to industries may not accurately represent the full set of firms that produce products in that industry. This problem is greater for a country like Russia where production is assigned to industries by company rather than establishment. Second, the concentration measure reported by governments assumes that the relevant geographic market for competition is the national market, that is, that each firm in the country faces competition from each other firm. Clearly, however, this is not the case in all industries, certainly not in a country as large as Russia, but likely not in most countries.[10]

In order to test whether the assignment of firms to industries affects the outcome of concentration-profitability analysis, we use the data on products to construct a measure of industry coverage, *COVER*, that closely mimics the coverage statistic calculated and used by the U.S. Census Bureau for evaluating U.S. Standard Industrial Classification codes. A higher value of *COVER* suggests that the definition of industry in the data is a closer approximation to the market. Interacting *COVER* with *CONC* allows us to control for the accuracy of the concentration measure. Thus, we test for the importance of industry market definition by estimating the following:

$$\begin{aligned} PCM_j &= \beta_0 + \beta_1 CONC_j + \theta_1 COVER * CONC + \gamma_1 MES_j + \gamma_2 CAPOUT_j \\ &+ \gamma_3 PRIVSHARE_j + \mu \end{aligned} \tag{5}$$

where the coefficient on concentration in a perfect coverage industry is $\beta_1 + \theta_1$.

[10] Of course a third problem of market definition is that the range products included in an industry definition may overstate or understate the range of products that actually compete with one another, but we cannot control for that here.

To test for the impact of regional markets, where we take the Russian *oblast*, or province, to be the level of disaggregation,[11] we construct a variable, similar to the dispersion variable of Collins and Preston (1969), that measures the geographic dispersion of production in each industry relative to the geographic dispersion of the population of Russia. The variable, *DISPER*, takes values from zero to one, where one indicates that each oblast accounts for the same percentage of production in that industry as it accounts for percentage of total population.

In theory, the spatial distribution of production can affect profitability in several ways. Generally, economists beginning with Eaton and Lipsey (1978) hypothesize that dispersion decreases competition and should lead to greater profitability. The Eaton and Lipsey theory is essentially that the combination of small markets and sunk costs inhibits entry. On the other hand, one could hypothesize that small markets restrict the ability of firms to fully exploit economies of scale, thus decreasing their profitability. A third way to think about the impact of dispersion is that if the markets relevant for competition are regional rather national, then the national measure of concentration has little explanatory power for industry profitability. Take, for example, a product produced and sold locally like bread. The size of the market alone may not affect profitability, but the size of the market determines how many firms are competitors. So the bread industry could be monopolistic if each region has only one bread producer, even though national concentration would still be low.

To test for these different effects, we estimate the following:

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[11] There are currently 89 Russian *oblasts* (which include republics and autonomous *okrugs*). In the RERLD there are only 78 *oblasts*. The 78 do account for the whole population, except Chechnya and Ingushetia; some of them are aggregations of the republics and autonomous *okrugs* that have become independent during the transition period.

$$PCM_j = \beta_0 + \beta_1 CONC_j + \beta_2 DISPER_j * CONC_j + \beta_3 DISPER_j + \gamma_1 MES_j + \gamma_2 CAPOUT_j + \gamma_3 PRIVSHARE_j + \mu \tag{6}$$

For industries with complete dispersion, the coefficient on national concentration is $\beta_1 + \beta_2$. When firms are all clustered in one area (roughly $DISPER_j = 0$), the coefficient is just β_1.

4. Data

For this analysis we use the Russian Enterprise Registry Longitudinal Database (RERLD), an economy-wide database on all medium and large industrial firms in Russia.[12] We created the RERLD from Russian government registry cross-sectional datasets on industrial firms for each year.[13] The RERLD has annual data from 1992 to 1995 for a set of basic variables including output, employment, profitability, capital stock, ownership, and others. See appendix 1 for a precise definition of each variable. To define industry we use the Russian industry code OKONH, which is a five-digit industry code. See appendix 2 for a discussion of the possible variations on industry definition over the years. We limit the sample to manufacturing industries and exclude any industries that cover miscellaneous goods. Table 1 lists the means of the variables that we use in the regression analysis. We see that there has been a noticeable decline in average profitability from 1992 to 1995.

[12] The database comprises the population of Russian enterprises with 200 or more workers and a large sample of smaller enterprises.

[13] We are extremely grateful to Michael Barz, Evgeniya Bessonova, Elizaveta Brazdnikova, Maria Gorban, Maria Lopukhova, Igor Tchekhov, and Staffan Westersträn for excellent research assistance in the construction of the database.

Table 1
Variable means

Variable	1992	1993	1994	1995
Profit/output	0.246	0.240	0.210	0.132
CR2 by output	0.438	0.435	0.449	0.450
CR4 by output	0.592	0.587	0.600	0.597
HHI by output	0.208	0.201	0.211	0.205
CR2 by employment	0.384	0.374	0.378	0.374
CR4 by employment	0.533	0.522	0.524	0.520
HHI by employment	0.177	0.168	0.172	0.163
Average industry share		0.710	0.754	0.765
Minimum efficient scale	0.102	0.091	0.093	0.098
Capital/output	1.079	0.195	1.858	2.399
Private share		0.465	0.701	0.774
Import penetration	0.052	0.257	0.222	0.442

We use the data on firms to calculate the concentration statistics. Our statistics are not entirely accurate because we do not have information on most of the very small firms. Small firms have a trivial influence on concentration measures though, and we can see in appendix table 1 that the size and output of the missing firms are both quite small. Table 2 reports the correlation coefficients between different concentration measures and concentration across years. The correlation between the CR4 and the HHI in Russia is lower than in a country such as the U.S. For example, the 1992 correlation in Russia is 0.78 while for the U.S. in 1982 it was 0.95.[14] The table shows that industrial concentration is beginning to change even during the early years of transition. The correlation between the Herfindahl in 1992 and the Herfindahl in 1995

[14] Scherer and Ross (1990), p. 72.

is 0.78. Thus 1995 market structure is not as exogenous as 1992 market structure, but it is still a long way from being market determined.[15, 16]

[15] The 1995 data contain foreign-owned firms, which were excluded by Goskomstat from earlier registries. So the 1995 data should depict a greater change in market structure than the previous year-to-year changes. The bias is not great though, as there were not many foreign-owned firms prior to 1995.

[16] For a detailed description of Russian industrial structure and discussion of changes in structure during transition, please see Brown and Brown [1998].

Table 2
Correlations between concentration ratios across measures and across years

	CR292	CR492	HHI92	CR293	CR493	HHI93	CR294	CR494	HHI94	CR295	CR495	HHI95
CR292	1.00	0.95	0.90	0.94	0.91	0.84	0.89	0.86	0.81	0.83	0.82	0.75
CR492	0.95	1.00	0.78	0.91	0.96	0.73	0.87	0.91	0.71	0.82	0.87	0.68
HHI92	0.90	0.78	1.00	0.86	0.75	0.93	0.81	0.74	0.88	0.75	0.68	0.78
CR293	0.94	0.91	0.86	1.00	0.95	0.90	0.92	0.90	0.83	0.87	0.86	0.77
CR493	0.91	0.96	0.75	0.95	1.00	0.78	0.89	0.94	0.73	0.85	0.91	0.70
HHI93	0.84	0.73	0.93	0.90	0.78	1.00	0.82	0.72	0.88	0.76	0.70	0.77
CR294	0.89	0.87	0.81	0.92	0.89	0.82	1.00	0.95	0.90	0.90	0.88	0.80
CR494	0.86	0.91	0.70	0.90	0.94	0.72	0.95	1.00	0.78	0.88	0.93	0.72
HHI94	0.81	0.71	0.88	0.83	0.73	0.88	0.90	0.78	1.00	0.79	0.71	0.81
CR295	0.83	0.82	0.75	0.87	0.85	0.76	0.90	0.88	0.79	1.00	0.96	0.90
CR495	0.82	0.87	0.68	0.86	0.91	0.70	0.88	0.93	0.71	0.96	1.00	0.80
HHI95	0.75	0.68	0.78	0.77	0.70	0.77	0.80	0.72	0.81	0.90	0.80	1.00

5. Results

We begin by analyzing just 1992—the first year when profitability is market determined, but while market structure is still entirely exogenous. Table 3 reports the estimates for 1992 for equations (4), (5), and (6) estimated by OLS using three different output concentration measures. These regressions do not include the *PRIVSHARE* variable because mass privatization did not begin until the end of 1992. The estimates on concentration in the basic regressions in the first three columns suggest that exogenous market structure does not have a causal effect on profitability. When we control for industry definition in the middle three columns, the estimates on concentration do not improve.[17]

The regressions in the last three columns test for the importance of geographic dispersion of industry in understanding the relationship between concentration and profitability. We see that the interaction with dispersion makes the estimated coefficients on all three measures of concentration positive and statistically significant. The F-statistics on these regressions are highly significant, and the R^2's generally improve. Thus when national concentration is high *and* firms are geographically dispersed, profitability is higher. For example, the estimated coefficient on the CR2 for perfectly dispersed industries ($\beta_1 + \beta_2$) is 0.66, meaning that a one percentage point

[17] Note that the number of observations in the coverage regressions is quite a bit lower than in the basic regressions. This reduction arises because many industries do not produce the most of any one product listed in the product data, and so have a missing value for the coverage variable. Either these industries really are not distinct or their primary products are missing from the product data. If the former is the case, then it is appropriate to exclude the industries from the regressions; if the latter is

increase in the CR2 is associated with a 0.66 percentage point increase in industry profitability.

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true, the exclusion could be introducing selection bias. Given that the estimates are not significant, the distinction is unimportant.

Table 3

OLS Regressions of 1992 profitability on exogenous market structure

Variable	Basic			Coverage			Dispersion		
CR2 92	-0.019			0.145			-0.166***		
	(0.046)			(0.126)			(0.055)		
CR2 92 *Cover				-.041					
				(.125)					
CR2 92 *Dispers 92							0.826***		
							(0.245)		
CR4 92		0.018			0.139*			-0.100*	
		(0.034)			(0.082)			(0.052)	
CR4 92 *Cover 93					-0.017				
					(0.081)				
CR4 92 *Dispers								0.621***	
								(0.179)	
HHI 92			-0.078			0.176			-0.237***
			(0.074)			(0.339)			(0.057)
HHI 92 *Cover 93						-0.103			
						(0.357)			
HHI 92 *Dispers 92									1.66***
									(0.630)
Dispers 92							-0.148***	-0.143**	-0.069
							(0.055)	(0.064)	(0.048)
MES	0.014	-0.032	0.061	-0.122	-0.153	-0.034	0.140**	0.087	0.110**
	(0.064)	(0.051)	(0.071)	(0.159)	(0.129)	(0.215)	(0.060)	(0.059)	(0.055)
Capout	-0.022**	-0.022**	-0.022**	-0.028**	-0.027**	-0.028**	-0.018**	-0.018*	-0.019**
	(0.011)	(0.011)	(0.011)	(0.014)	(0.013)	(0.014)	(0.009)	(0.010)	(0.009)
Constant	0.281	0.270	0.280	0.267	0.248	0.284	0.303	0.292	0.279
	(0.029)	(0.030)	(0.029)	(0.034)	(0.035)	(0.033)	(0.038)	(0.046)	(0.035)
Diagnostics									
Prob > F	0.023	0.031	0.013	0.031	0.000	0.162	0.000	0.000	0.000
R^2	0.08	0.09	0.09	0.12	0.14	0.10	0.14	0.13	0.14
N	241	241	241	152	152	152	241	241	241

Robust standard errors are reported below the coefficient estimates. We proxy 1992 coverage with 1993 coverage, as we do not have product data for 1992. Estimates on branch dummies are excluded from the table. *Significant at 10% in a two-tailed test. **Significant at 5% in a two-tailed test. ***Significant at 1% in a two-tailed test.

Controlling for dispersion actually leaves a slightly negative effect for concentration alone, so for example, the estimated coefficient on the CR2 for a "perfectly" geographically clustered industry is -0.166. This result suggests that even in concentrated industries, when firms are located in only one or a few places, they compete with each other on the national market and do not enjoy oligopoly rents. In fact, their rents are slightly negative, and we might hypothesize that this effect arises because geographic clusters of firms in the same industry are easy targets for government regulators or tax authorities. There is clearly no evidence that geographic clustering facilitates tacit or even explicit collusion, or rather, the evidence suggests that in Russia, even if collusion is facilitated, other intervention erodes those rents.

We see that the capital-output ratio has a negative and significant effect on profitability in all the regressions. Although these results are the opposite of what is usually hypothesized, Salinger (1990) finds similar results on capital-sales ratios in the late 70's and early 80's in the United States and Lai (1994) finds similar results for Taiwan. Caves (1990) suggests an explanation for Salinger's U.S. results: profitability is lower in capital-intensive industries because the duration of mistakes when capital is sunk is longer, so the negative findings for the U.S. could be related to recessionary conditions after the oil shock. This explanation suits the Russian case quite well. Much of the capital stock inherited from the Soviet period is poorly maintained and out-of-date. The need to update the capital stock thus causes the adjustment period for capital-intensive industries to be longer.[18] Generally the results for the U.S. and Russia suggest that the relationship between profitability and capital intensity varies in sign depending on macroeconomic conditions that necessitate adjustment.

Although we see in table 2 that industry structure does change slightly from year to year during the early 90's, there is still a very high correlation between concentration in 93-95 and that in 92. Thus, extending our assumption of exogeneity over all four years, we can better estimate the equations by pooling the data. In table 4 we report the results from the regressions on the pooled data. Here and in the subsequent tables, we only report results for the CR2. As table 3 shows, the three measures do not perform differently. We choose the CR2 because it returns the most stable estimates in all the reported and unreported regressions for the paper.

Columns one, four, and seven report the estimates from the pooled OLS regressions on all three models. The additional observations do not improve the explanatory power of concentration in the basic or coverage-corrected models, but confirm the negative effect of capital intensity. The pooled dispersion estimates support the conclusions from the 1992 regressions that industries that are concentrated but regionally dispersed do indeed have higher profitability, while national concentration in geographically clustered industries has a small, negative effect on profitability. One might wonder whether there are unobserved industry-specific effects that introduce omitted variables bias into the OLS estimation. In fact, although such a possibility seems highly likely in all the empirical studies of this kind, authors rarely address it. Here, the F-test on significance of group effects (Greene, 1997) is insignificant, suggesting there are no industry-specific effects. Later tests, however, provide evidence that industry-specific effects do exist, which would make the OLS results inconsistent. The insignificance of the F-test probably arises from year-to-year measurement errors which weaken the fixed-effects estimation.

We also employ both between-effects estimation and random-effects estimation and report the results in table 4. The between-effects estimator does not improve the ability of concentration to explain profitability in the basic or coverage-corrected regressions. In column eight though, we see again that nationally concentrated

[18] The actual coefficient estimates on this variable jump around quite a bit from year to year. This variance results from the variance in the valuation of capital from year to year, as can be seen in table

industries that are also geographically dispersed are indeed more profitable. The between-effects estimator is well-suited for testing the impact of inherited, or exogenous, market structure, because it abstracts from the changes after 1992 and focuses on the explanatory power of the differences between industries. Assuming that market structure is primarily exogenous over all four years, the between-effects estimation improves on any single year regression because it averages out any year-to-year noise in the data—missing data on any large firms in a single year could influence the values of the variables for those industries in that year. Indeed we see that the R^2's for the between-effects dispersion regression is higher than those for the 92 and pooled OLS dispersion regressions.

1.

Table 4
Regressions of profitability on concentration using pooled data

	Basic			Coverage			Dispersion		
	OLS	Between effects	Random effects	OLS	Between effects	Random effects	OLS	Between effects	Random effects
CR2	-0.024	0.009	-0.043	-0.015	-0.050	-0.053	-0.080**	-0.055	-0.094**
	(0.029)	(0.044)	(0.030)	(0.077)	(0.083)	(0.082)	(0.041)	(0.060)	(0.046)
CR2*Cover				-0.010	0.070	0.002			
				(0.081)	(0.086)	(0.087)			
CR2*Dispers							0.671***	1.049***	0.631***
							(0.163)	(0.241)	(0.192)
Dispers							-0.022	0.023	-0.029
							(0.040)	(0.060)	(0.051)
MES	0103	0.004	0.125**	0.360*	0.297**	0.398**	0.185***	0.165*	-0.189***
	(0.060)	(0.083)	(0.058)	(0.210)	(0.125)	(0.205)	(0.067)	(0.089)	(0.066)
Capout	-0.025	-0.009***	-0.025***	-0.029***	-0.030***	-0.031***	-0.0245***	-0.007**	-0.025***
	(0.009)	(0.003)	(0.008)	(0.005)	(0.007)	(0.006)	(0.009)	(0.003)	(0.009)
Private share	0.022	0.115	0.014	0.005	0.001	-0.012	0.028	0.141***	0.017
	(0.022)	(0.031)	(0.024)	(0.025)	(0.033)	(0.023)	(0.022)	(0.031)	(0.024)
Constant	0.247	0.165	0.257	0.263	0.272	0.284	0.213	0.071	0.229
	(0.028)	(0.033)	(0.033)	(0.026)	(0.034)	(0.032)	(0.039)	(0.047)	(0.045)
Diagnostics									
Prob > F	0.000	0.002		0.000	0.000		0.000	0.000	
R^2	0.126	0.118		0.196	0.252		0.145	0.191	
Prob > χ^2			0.000			0.000			0.000
Breusch-Pagan test									
Prob > χ^2			0.000			0.000			0.000
Hausman test									
Prob > χ^2			0.000			0.341			0.000
N	745	253	745		197	422	745	253	745

Standard errors (robust for OLS and random effects) are reported in parentheses. Estimates and standard errors for the random-effects model come from GLM estimation, while the diagnostics come from GLS estimation. The two methods are asymptotically equivalent. Estimates on branch dummies are excluded from the table. *Significant at 10% in a two-tailed test. **Significant at 5% in a two-tailed test. ***Significant at 1% in a two-tailed test.

Columns three, six, and nine in table 4 report the random-effects estimates on the pooled data. The estimates and t-statistics in all three models are concordant with the estimates from the pooled OLS and between-effects regressions. The Breusch-Pagan Lagrange multiplier test does indicate that random effects, and thus industry-specific effects, are present. The random-effects estimation partials some of these effects out, but nonetheless, the Hausman test is significant in the basic and dispersion models, meaning that there are still industry-specific effects correlated with the right-hand side variables. Thus the random-effects and between-effects estimates are also inconsistent.

The econometric solution here is normally to use the fixed-effects estimates since they do not depend on the assumption that industry-specific effects are uncorrelated with the right-hand side variables. The fixed-effects estimator, however, is solely a within-effects estimator, meaning that it only looks at how the changes in concentration over time affect the changes in profitability. Such changes in market structure are exactly what are endogenous to market conduct though, and thus exactly what we want to exclude. In addition, we believe that that there is non-trivial measurement error from year to year, to which fixed-effects estimation is highly sensitive. Thus fixed effects is not appropriate here.

In table 5 we present the results of the longer-run analysis. Here we test whether the exogenous market structure in 1992 has a persistent influence on profitability in later years. The first column presents the estimates of a regression of 1995 profitability on 1992 market structure controlling for changes in market structure over time. The estimated coefficient on the 1992 concentration-interaction term is significant at 1 percent. Interestingly, the sign on the absolute value of the change in national concentration is negative and significant. This result likely reflects the endogeneity of changes in market structure—less profitable industries experience greater changes in market structure. The second column presents the result of the random-effects estimation of the original dispersion model where 1992 concentration is entered as a constant. The 1992 interaction term produces a positive and statistically significant estimate at 1 percent. These results tentatively, since the time period is relatively short, support Salinger's (1990) finding for the U.S. that market power is persistent.

Table 5
Effect 1992 market structure in the longer run

Variable	OLS	Variable	Random effects
CR2 92	-0.004	CR2 92	-0.057
	(0.058)		(0.039)
ΔCR2	-0.058		
	(0.050)		
|ΔCR2|	-0.329***		
	(0.152)		
CR2 92*Dispers 92	0.787***	CR2 92* Dispers 92	0.901***
	(0.272)		(0.203)
Δ(CR2*Dispers)	-0.060		
	(0.201)		
|ΔCR2*Dispers|	3.516		
	(2.371)		
Dispers 95	0.040	Dispers 92	-0.001
	(0.050)		(0.049)
MES 95	0.112	MES	0.189***
	(0.090)		(0.053)
Capout 95	-0.007	Capout	-0.023***
	(0.007)		(0.003)
Private share 95	0.130***	Private share	0.010
	(0.051)		(0.018)
Constant	0.024	Constant	0.190
	(0.067)		(0.034)
Diagnostics			
Prob > F	0.000		
R^2	0.24		
Prob > χ^2			0.000
Breusch-Pagan test			
Prob > χ^2			0.000
Hausman test			
Prob > χ^2			0.000
N	248		737

Robust standard errors are reported in parentheses. Estimates and standard errors for the random-effects model come from GLM estimation, while the diagnostics come from GLS estimation. The two methods are asymptotically equivalent. Estimates on branch dummies are excluded from the table. *Significant at 10% in a two-tailed test. **Significant at 5% in a two-tailed test. ***Significant at 1% in a two-tailed test.

The evidence that market power does exist raises the question, especially relevant in the transition context, of whether import penetration introduces competition to profitable industries. In table 6, we present the results of the random-effects estimation of the dispersion model including a separate variable for import penetration and interacting import penetration with dispersion. In both regressions, the coefficient estimate on imports is negative and statistically significant, suggesting that imports do reduce profitability, and the inclusion does not appreciably change the signs and significances of the other variables. In the regression where import penetration is interacted with dispersion, the estimate simply reflects that there are more profits to be eroded by import competition in industries with regional markets. We consider this evidence preliminary though, as only the random-effects estimators on the pooled data yield statistically significant result. When both imports and the interaction term are included, the estimates are not statistically significant, likely due to multicollinearity.

Table 6
Random-effects regressions with imports

Variable		
CR2	-0.098**	-0.107
	(0.048)	(0.049)
CR2*Disperse	0.660***	0.693
	(0.195)	(0.199)
Disperse	-0.039	-0.008
	(0.053)	(0.054)
Import	-0.042***	
	(0.017)	
Import*Dispers		-0.149***
		(0.047)
MES	0.196***	0.191***
	(0.067)	(0.067)
Capout	-0.024***	-0.024***
	(0.009)	(0.008)
Private share	0.023	0.022
	(0.025)	(0.024)
Diagnostics		
Prob > χ^2	0.000	0.000
Breusch-Pagan test Prob > χ^2	0.000	0.000
Hausman test Prob > χ^2	0.000	0.000
N	745	745

Robust standard errors are reported below in parentheses. Estimates and standard errors for the random-effects model come from GLM estimation, while the diagnostics come from GLS estimation. The two methods are asymptotically equivalent. Estimates on branch dummies are excluded from the table. *Significant at 10% in a two-tailed test. **Significant at 5% in a two-tailed test. ***Significant at 1% in a two-tailed test.

Finally, tables 4 and 5 and appendix table 1.2 (which presents the results of the yearly regressions of profitability on the dispersion specification) reveal that the *PRIVSHARE* variable sometimes has positive and significant coefficient estimates. One might interpret this result as showing that privatization increases profitability, but note that this result never appears in the random-effects estimations where within effects are included. Thus, we can only conclude that there is a correlation without knowing the direction of causality.

6. Conclusion

The data on industrial firms in Russia from the RERLD allow us to improve and expand on the traditional concentration-profitability analysis in several ways. First, the market structure in early transition Russia is completely exogenous to market determination, eliminating the endogeneity problem from the regression analysis. Second, the firm-level data allow us to construct a variety of concentration statistics, control for the coverage of industry definitions, and test for the influence of geographic dispersion on the effects of market structure.

The analysis yields several important results. When industries' firms are geographically dispersed, we find strong evidence that exogenous national concentration does increase industry profitability. Geographically clustered industries, however, seem to face competition regardless of the size and number of firms in the industry, suggesting that clustering is not able to facilitate implicit or explicit collusion. We also find preliminary evidence for Russia that supports Salinger's (1990) findings for the U.S. that the effects of market structure are persistent in the long run. The use of different concentration measures does not have an appreciable effect on the analysis, even when the correlation between these measures is much lower than that found in the U.S. Controlling for industry coverage does not improve the ability of national concentration to explain profitability generally.

We find for Russia that capital intensity is negatively related to profitability during transition, supporting Caves' (1990) hypothesis that capital-intensive industries face higher adjustment costs and thus have lower profitability after recessionary shocks. Finally, we find preliminary evidence that import penetration does reduce industry profitability early in transition.

Appendix 1. Definition of variables

PCM_j is the sum over firms in industry *j* of the ruble value of profits (losses) divided by the sum over the same firms of the ruble value of output.

$CONC_j$ is the two-firm concentration ratio (value of output of the two largest firms in the industry as a share of the total value of output in the industry), four-firm concentration ratio, or the Herfindahl Hirschman Index (sum of the squared output shares of each firm in the industry). We performed the same analysis, not reported in the tables, with these three measures calculated using data on employment.

$COVER_j$ is calculated by first assigning products to industries according to which industry produces the largest physical volume of the product. Then for each industry we calculate the percentage of the production of each assigned product that is produced by firms in that industry, and we take the mean over these percentages to get a single statistic for industry coverage. So for example, a *COVER* value of 0.5 for industry X means that on average, the firms classified in industry X produce 50 percent of the products in industry X. Our measure differs from the U.S. measure in that U.S. the Census Bureau has value data on product production, so it can sum the total value of the products produced by establishments in the industry and take it as a percentage of the total value of the product production. Since we only have physical output data for products, we are constrained to take the average of the percentages for each product. See U.S. Department of Commerce (1993) for more details.

$DISPER_j$ is one minus ((the sum of the absolute values of the difference between the output share of the oblasts and the population share of the oblasts) divided by 2). This measure (except the one minus part) was first used by Collins and Preston (1968, 1969.

MES_j is the average output per plant among the 50% of the firms with lowest cost, as a share of total industry output.

$CAPOUT_j$ is the sum over firms in industry *j* of the average value of fixed assets used in industrial production (main activity of the enterprise) divided by the sum over the same firms of the value of output.

$PRIVSHARE_j$ is the share of output in *j* from firms with private or mixed private and state ownership.

$IMPORTS_j$ is the share of domestic sales in industry *j* accounted for by imports. The variable was constructed using data from the Russian Customs Committee on traded goods classified by United Nations product codes. These codes were assigned, as best possible, to Russian OKONH industries. Combining the RERLD data on output and the trade data, we created the variable *imports/(imports+output-exports)* for each industry. The variable is imprecise, however, not only because the industry assignment is imperfect, but also because the trade data include all sales, while the output data are biased by any missing firms in the database.

The branch dummies used are energy, metal, chemical, machine, forest, construction, light, and food. The excluded branch is other, which does include non-miscellaneous industries.

Appendix table 1.1 Comparison of RERLD and Goskomstat industry statistics

Russian Enterprise Registry Longitudinal Database						
Year	Number of Enterprises	Employment	Mean	Median	Minimum	Maximum
1992	22,073	17,486,919	792	248	1	94,550
1993	25,013	16,940,158	677	199	1	96,038
1994	26,064	14,834,595	569	173	1	98,873
1995	30,867	14,106,624	457	128	1	99,242

Goskomstat statistics on total industry						
Year	Number of Industrial Enterprises	Industrial Employment	Mean	# Enterprises Missing from RERLD	Employment Missing from RERLD	Mean Employment of Missing Enterprises
1992	61,075	20,020,000	328	39,002	2,533,081	65
1993	104,059	18,864,000	181	79,046	1,923,842	24
1994	137,999	17,440,000	126	111,935	2,605,405	23
1995	136,674	16,006,000	117	105,807	1,899,376	18

Sources: RERLD, Goskomstat of Russia

Appendix table 1.2
Year-by-year OLS regressions

Year	1993	1994	1995
CR2	-0.039	-0.132	-0.035
	(0.052)	(0.086)	(0.062)
CR2*Dispers	0.562***	1.322***	0.562**
	(0.200)	(0.364)	(0.257)
Dispers	-0.002	-0.069	0.057
	(0.052)	(0.079)	(0.055)
MES	0.115	0.280**	0.140*
	(0.122)	(0.131)	(0.079)
Capout	-0.098***	-0.020*	-0.008
	(0.039)	(0.012)	(0.008)
Private Share	0.052**	0.098**	0.134***
	(0.026)	(0.049)	(0.049)
Constant	0.228	0.161	0.028
	(0.041)	(0.093)	(0.066)
Diagnostics			
Prob > F	0.000	0.000	0.004
R^2	0.13	0.14	0.19
N	246	247	252

Standard errors are reported in parentheses. Estimates on branch dummies are excluded from the table. *Significant at 10% in a two-tailed test. **Significant at 5% in a two-tailed test. ***Significant at 1% in a two-tailed test.

Appendix 2. Industry definitions in the RERLD

As explained in the text, the RERLD is a panel database of Russian firms that we created from separate cross-sectional datasets—the Russian statistical office does not maintain the data as a panel. When linking the firms across years, we noticed some problems with the industry codes. First, there are several cases where all the firms in one industry in one year seem to "move" to another industry in the next year, that is, they all have a new industry code. We assume that these are code changes rather than market structure changes and correct the codes to reflect the new code in all years. We also find many cases where individual firms change codes across years. These changes are especially prevalent in 1995 and 1996.

Such code changes could reflect three activities: First, firms could be changing their product mix to the degree that they truly belong in a different industry. Casual analysis of the product data though, reveals that many firms with changing industry codes do not report significant changes in their product mix. Second, firms could be reporting different industry codes for reasons other than true structural change, for example, they may decide to report the code for the industry in which they hope to compete after restructuring or they may simply make mistakes when completing the government forms. Third, firms or Goskomstat could be correcting past classifications to better reflect the true state of firms. We have no way of determining the relative importance of each of these explanations. Each, however, suggests a different treatment of the industry codes in the data, and thus we conduct our analysis on three different industry datasets.

According to the first explanation, we create a set where firms are assigned to industries each year based on the industry code they report that year. This set over-represents true structural change. Consistent with the second explanation, we create a set where continuing firms are constrained to the industry they start in. This set still allows for structural change in the form of entry and exit, but under-represents true structural change. Finally, we create a set where firms are constrained to the industry they report in the most recent year. Again this set under-represents true structural change, and it biases the measurement of market structure in the direction of the present. Table 2 in the paper reports the correlations between the concentration measures using the "first year" definition. It turns out that the analysis is robust to all three constructions. The tables in the paper are based on the "first year" data.

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