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Studies in Finance, Trade and Politics

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RISK MATTERS:

STUDIES IN FINANCE, TRADE AND POLITICS

Jonas Vlachos

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to Karin and Elias

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Stockholm, November 2001
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INTRODUCTION AND SUMMARY

This thesis consists of four self-contained empirical essays. The simple idea underlying three of the four papers is that risk and uncertainty affect political choice. Indeed, most people would have to be hard pressed to argue otherwise. This idea does, however, have some interesting implications; in particular, it means that the existence, or the development, of different types of insurance should affect the political agenda. For example, we rarely hear political demands for government intervention in the home-insurance industry – most likely because home-insurance markets are relatively efficient. Unemployment insurance, on the other hand, is usually high on the electorate's list of political priorities. Although this can partly be explained by a preference for redistribution, the severe moral hazard and adverse selection problems affecting private insurance for human capital, naturally raise demand for public insurance of this kind.

How far this point can be taken naturally depends on what types of markets and insurance we are analyzing. Since the focus of the essays is political choice at the macro level (openness to trade, regional integration, and the size of the public sector), two of the largest and most important markets in a modern economy, namely the financial sector and the labor market, are considered. The everyday importance of the labor market is self-evident, as is the fact that labor markets differ in the level of risk to which they are subject. The financial sector specializes in the handling of financial risks and is a large contributor to GDP in most developed countries.[1] If an important effect of the insurance characteristics of these markets on policy cannot be documented empirically, the starting point of this thesis is likely to be just a theoretical curiosity.

Although the questions and hypotheses investigated here only constitute a small subset of all possible links from markets handling risk to policy, the results indicate that these links can be important. Broadly speaking, this gives support for theories such as Persson and Tabellini (1996) who discuss the risk-sharing aspects of regional integration, and Fernandez and Rodrik (1991) who demonstrate how individual-

[1] Demirgüc-Kunt and Levine (1996) present estimates of the size of the financial sector. The contribution to GDP in 1990 was around five percent in the USA, and around nine percent in Japan.

specific uncertainty can create a bias away from welfare enhancing economic reforms. At a more general level, it can be argued that political choices affecting the functioning of markets are likely to change the tradeoffs faced by political actors, thereby also influencing the political agenda. Similar mechanisms have been considered in the literature on endogenous choice of policy instruments. Rodrik (1986), for example, demonstrates that although a production subsidy is more efficient than a tariff, the electorate would, if possible, pre-commit to tariff protection only. The reason is that lobbying for tariffs generates more free-riding between interest groups and hence, smaller distortions. Although the point that policy can influence the political game itself is easily made, it can severely complicate the analysis of economic and political reforms. Therefore, plenty of empirical work needs to be undertaken to better understand the strength and scope of these mechanisms.

FINANCIAL DEVELOPMENT

Whenever economic research leaves the world of theory, there is an attempt to put numbers on theoretical phenomena. This is always tricky, and especially so when dealing with macro-level indicators of various kinds. In two of the papers included in this volume, numbers are used to indicate the degree of development in the financial sector. These numbers are supposed to measure how well the financial sector mobilizes savings, allocates credit across time and space, exerts corporate control, as well as how efficient are the hedging, pooling, and pricing of risks. The fact that indicators similar to the ones used here are commonly used in the literature (starting with King and Levine, 1993), does not reduce the problems involved in attaching a single number to these complex institutions and markets.[2]

This problem is dealt with by using a variety of indicators of financial development. When the results are consistent across indicators, we put greater faith in them. To further complicate matters, however, financial systems have developed along different paths in different countries. The various indicators of financial development therefore supposedly also measure different aspects of the financial system. Traditionally, researchers have classified financial systems as either bank based (e.g. in Japan and

[2] This criticism naturally extends to other indicators such as "democratic development", "law enforcement", "corruption", and "social capital", commonly used in economic research.

Germany) or market based (e.g. in the United States and the United Kingdom). This division partly emerged due to the different regulatory systems created as responses to financial crises during the early phases of industrialization (see the account in Allen and Gale, 2000). Since there is considerable variation in the way the financial sector is regulated and in company law even within these two subgroups, there is a continuing debate on to what extent this grouping adds to the understanding of different financial systems. Moreover, recent research has failed to document that the classification can contribute to the understanding of economic performance at the aggregate (Levine, 2001) or industry (Beck and Levine, 2001) level. For these reasons, results for individual indicators should be interpreted with care.

Rather than focusing on the differences between bank- and market-based financial systems, there is a growing consensus that the legal environment – for example, protection of minority shareholders and creditors, and rules for information dissemination – is what matters for financial development (see e.g. La Porta et al., 1998). According to this view, financial contracts are upheld by more or less efficient law enforcement. Other mechanisms can substitute for weak regulation in one area, and what ultimately counts is the overall performance of the legal system. An important insight from this literature is that deeply rooted legal traditions play a central role in shaping a financial system. In particular, common-law legal systems seem to offer better protection of investors than civil-law systems. Since the legal tradition of a society is not easily changed, one possible implication of this theory is that financial development cannot be generated without encompassing legal reforms.

The relative ease with which countries can be divided into different legal traditions can explain the wide spread use of this approach. The approach is not unproblematic, however. Countries in the same legal tradition do, for example, find themselves at very different levels of financial development. Further, the development of the financial sector has not been monotonic over time. Rather, both expansions and reversals have characterized the pattern of development during the last century. These facts bring into question theories stressing the importance of the legal tradition, which is time-invariant by nature. In contrast, Rajan and Zingales (2001) argue that poorly developed financial markets can benefit industry incumbents by restricting the availability of financial capital to new entrants. If competition in goods and/or

financial markets increases for exogenous reasons (for example through a trade reform), the relative benefit of holding back financial development is decreased. Coffee (2001) is another dissenting voice, questioning the mechanism through which the legal traditions have affected the financial systems. He argues that strong securities markets created a constituency demanding a stricter legal protection of investors. The legal tradition played a role through the early separation of the private sector from central government control in several common-law countries. This separation generated self-regulatory mechanisms of securities markets that, in turn, facilitated the development of dispersed ownership. According to Coffee's view, the important difference between civil- and common law countries does not lie so much in the formal rules as in the structure of law-making.

These objections highlight some of the difficulties involved in the methodology employed in this dissertation. In order to fully comprehend the mechanisms at work, both case studies and time series research are needed.

CHAPTER 1

In "Markets for Risk and Openness to Trade: How are They Related?" (with Helena Svaleryd), we ask if there is an empirical relationship between financial development and openness to trade. Numerous theoretical papers have noted that trade policies can be used as an insurance against shocks from international markets. It follows that the development of markets for risk should reduce the incentives to rely on trade policy for insurance purposes. Feeney and Hillman (2001) explicitly demonstrate how asset-market incompleteness can affect trade policy in a model where trade policy is determined by the lobbying of interest groups. If risk can be fully diversified, special-interest groups have no incentive to lobby for protection, and free trade will prevail. Likewise, trade liberalization might increase the demand for financial services, thereby spurring the development of financial markets.

Using several indicators of both openness to trade and financial development, we find an economically significant relation between the two. In particular, the relation holds when using the well known, although criticized (Rodriguez and Rodrik 1999), Sachs-Warner index, and structurally adjusted trade, as indicators of openness. For tariff

levels and non-tariff barriers, the results hold only for relatively rich countries. Causality seems to be running both from openness to financial development and the other way around, depending on which indicator and methodology are used.

CHAPTER 2

Due to underlying technological differences, industries differ in their need for external financing (Rajan and Zingales, 1998). Since services provided by the financial sector are largely immobile across countries (Pagano et al., 2001), the pattern of specialization should be influenced by the degree of financial development. In the second essay, "Financial Markets, the Pattern of Specialization, and Comparative Advantage: Evidence from OECD Countries" (with Helena Svaleryd), we find this effect to be strong. In fact, the financial sector has an even greater impact on the pattern of specialization among OECD countries than differences in human- and physical capital. Further, the financial sector gives rise to comparative advantage in a way consistent with the Hecksher-Ohlin-Vanek model. Large and active stock markets, as well as the degree of concentration in the banking sector, produce the strongest and most consistent effects. The results also support the view that the quality accounting standards and the legal protection of creditors affect the pattern of industry specialization, while the depth of the financial system (measured by the amount of liquidity in an economy) is a source of comparative advantage.

CHAPTER 3

The third essay, "Who Wants Political Integration? Evidence from the Swedish EU-Membership Referendum" looks directly at the determinants of political attitudes towards regional integration and separation. More precisely, the regional voting pattern of the 1994 Swedish EU-membership referendum is analyzed. To explain this variation, an empirical investigation based on the extensive theoretical literature analyzing the determinants of regional economic and political integration is undertaken. Since enhanced possibilities of inter-regional risk sharing is one of the main gains from integration discussed in the literature (e.g Persson and Tabellini, 1996), special attention is given to this issue.

The empirical results show that individuals living in labor markets exposed to a high degree of risk were more negative towards EU-membership than those living in safe ones. It is also shown that inhabitants of high-income labor markets, with a high level of schooling and small receipts of central government transfers were relatively positive towards the EU-membership. Given the restrictive regulations limiting discretionary policies within the EU, these results suggest that inhabitants of safe and rich regions voted in favor of secession from the Swedish transfer system, rather than in favor of European integration.

CHAPTER 4

In the final essay, "Does Labor Market Risk Increase the Size of the Public Sector? Evidence From Swedish Municipalities", I study if a high degree of private labor-market risk is related to a larger public sector in Swedish municipalities. The theoretical hypothesis is based on Rodrik (1998), who argues (and shows empirically) that countries exposed to a high degree of external risk also tend to have larger governments. The safe public sector is expanded at the expense of risky sectors and hence provides insurance against income volatility. Several problems related to data availability and comparability that apply to cross-country studies are circumvented by using data on Swedish municipalities. Further, there is no need to aggregate the public sector across different levels of governance: local risk is directly related to the size of the local public sector.

The paper is not a complete parallel to Rodrik's study, however. Several alternative insurance mechanisms that do not exist between countries are available between municipalities. For example, the central government provides insurance against individual-specific risk such as unemployment and illness, private capital markets are better integrated within than between countries, and the central government can hand out grants to municipalities. Despite these mitigating factors, local labor-market risk is found to have a substantial impact on municipal public employment. It is also found that shocks increasing the size of the public sector across all municipalities tend to generate a larger increase in risky locations. For municipal public spending and taxation the results are, however, much weaker. Hence, labor-market risk affects the labor intensity of the municipal public sector, rather than its size.

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CHAPTER 1

MARKETS FOR RISK AND OPENNESS TO TRADE: HOW ARE THEY RELATED?*

Helena Svaleryd[+] and Jonas Vlachos[++]

Abstract

If protectionist trade policies aim to insure domestic industries against swings in world market prices, the development of financial markets could lead to trade liberalization. Likewise, trade liberalization could lead to the development of financial markets that help agents diversify the added risks. In this paper, we empirically address the hypothesis that there is a positive interdependence between financial development and liberal trade policies. We find a positive and economically significant relationship between the two, with causation running in both directions. The results are, however, somewhat dependent on the measure of trade policy being used.

JEL Classification: F13; G20

Keywords: Financial markets; Trade policy; Panel data

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[+] Department of Economics, Stockholm University, 106 91 Stockholm, Sweden. Email: hs@ne.su.se.

[++] Corresponding author: Department of Economics, Stockholm School of Economics, Box 6501, 113 83 Stockholm, Sweden. Email: jonas.vlachos@hhs.se.

1. Introduction

It has long been argued that trade restrictions can be motivated by insurance considerations in the absence of full risk diversification. Examples are Corden (1974), Hillman (1977), Cassing (1980), Newbery and Stiglitz (1984), Eaton and Grossman (1985), Cassing et al. (1986).[2] It follows that the development of institutions for risk diversification, e.g. financial markets, might reduce barriers to trade. Given the abundance of theoretical models, it is surprising that no empirical work has brought this hypothesis to the data. In this paper, we address the issue empirically and show that there exists a positive relation between openness to trade and the degree of financial sector development.

International trade brings about substantial changes in competition, technology, prices of intermediary and final goods, and in the long run even in factor endowments and the institutional features of a society. The exact outcome of trade liberalization for different individuals is therefore uncertain. Rodrik (1998) provides evidence that openness to trade also increases the permanent degree of income volatility in an economy.[3] The theoretical papers mentioned above argue that trade barriers can be welfare enhancing if private markets fail to pool such risks. Feeney and Hillman (2001a) explicitly demonstrate how asset market incompleteness can affect trade policy in a positive theory of trade liberalization. In their model, the degree of portfolio diversification determines the protectionist lobbying effort conducted by owners of sector specific capital. If risk can be fully diversified, special interest groups have no incentive to lobby for protection and free trade will prevail. This model suggests a causal effect from financial development to trade liberalization. Another possibility is that the demand for insurance increases after liberalization, thus promoting the development of the financial sector.

In the light of this literature, we ask the question whether institutions allowing for better insurance possibilities and risk diversification within a country are positively

[2] Dixit (1987, 1989ab) is a dissenting voice in this literature. When explicitly modeling the reasons behind the absence of insurance markets, he finds the scope for government intervention to be limited.
[3] Traca (2000) shows theoretically that we have reasons to expect trade to increase income volatility. Further empirical evidence for this view is given in Gottschalk and Moffit (1994), and Ghosh and Wolf (1997).

related to a liberal trade regime. In particular, we investigate whether the development of domestic financial markets is systematically related to trade policy. Moreover, since openness to trade increases aggregate income volatility, we expect international financial integration to reduce the demand for trade protection. This hypothesis is also brought to the data. Due to the difficulties involved in measuring trade restrictions, we employ a number of different measures. The expected positive relation between financial development, both domestic and international, and openness to trade appears clearly for some indicators. For other measures, the results are weaker and only seem to apply to relatively rich countries. Causality is an important issue, albeit a difficult one to fully resolve using cross-national evidence. We find some support for a causal link from financial development to openness for trade, and some support for the opposite. These results are conditional on the indicator we use.

This paper is organized as follows. Section 2 provides an extensive theoretical and empirical motivation for the study. Section 3 outlines the empirical methodology, while Section 4 describes the data. Special attention is given to the measurement of trade policy and financial sector development. Section 5 presents the results and Section 6 concludes.

2. Theoretical and empirical motivation

As mentioned in the introduction, a number of papers deal with trade policy as an insurance device within a social planner framework. Feeney and Hillman (2001a), however, provide a positive theory of trade policy. Specifically, they model a two-sector economy with independent productivity shocks that determine which sector will be exporting and import competing. Ex post the import competing sector can choose to lobby for protection and policy makers respond by implementing a tariff. The tariff increases the price for the import-competing good, but it also induces a consumption distortion in the economy thereby lowering aggregate welfare. In the standard case, when no portfolio diversification is possible, the equilibrium tariff will always be positive since the income gain from lobbying is larger than the consumption distortion for the import competing sector.

Next, domestic asset markets are introduced into this framework. Suppose that before the uncertainty regarding productivity is revealed, specific factor owners can trade in the asset markets. In the case when asset markets work without friction, the incentive for lobbying disappears since specific factor owners will optimally hold a fully diversified portfolio with specific capital from both sectors, and therefore only care about aggregate welfare. Suppose instead that the agents can only trade with a subset of capital. They may then be unable to reach a perfectly pooled equilibrium. The extent of lobbying and consequently tariffs will be determined by the difference between the income gain and the consumption distortion. Compared to the model without any trade in sector-specific capital, even limited access to capital markets reduces the payoff from protectionist policies. Thus, the degree of asset market incompleteness affects the lobby pressure for the imposition of tariffs and consequently how liberal a country's trade policies will be. Regardless of how susceptible the political sector is to private demand for protection, this effect will always be present. Given this setup, the empirical prediction would be a causal effect from financial development to trade liberalization. Another possibility is that the demand for financial services increases when the volatility of income goes up. In this case causality would run from openness to financial development.

The main focus of the Feeney-Hillman model is on domestic diversifiable risk and the functioning of domestic financial markets. The productivity shocks that hit the two sectors are independent implying that the risk can be domestically diversified.[4] The assumption that a significant share of the risk facing the agents can be domestically diversified receives support in studies of output shocks and volatility. Ghosh and Wolf (1997) use US data to show that shocks to output growth in a particular industry in a particular state is mainly driven by shocks to the sector, and that these shocks are only slightly correlated across sectors. Hence there is scope for risk diversification between industries within a country using domestic financial markets. Using international data, Clark and Shin (2000) provide further support for this view by showing that the main source of variation in output and employment for an industry in a country is due to shocks to that industry in that country (as opposed to shocks common to the whole

[4] In some cases, access to international asset markets can reduce lobbying pressure in the Feeney-Hillman model. For this to happen, an asymmetry of assets that can be traded must be present between

country or industry). By showing that shocks to the traded goods sector are larger than shocks to the non-traded goods sector, Ghosh and Wolf (1997) also provide indirect evidence that trade increase the volatility of output, i.e. support for the underlying assumption in the models mentioned in the beginning of this section.

The higher volatility of the tradable sector, compared to the non-tradable sector, is given a theoretical explanation in Traca (2000).[5] In his model, productivity shocks hit both sectors. The tradable sector is also subject to price shocks, uncorrelated to the productivity shocks. When shocks hit the non-traded goods sector, prices move to offset the volatility of aggregate income. Since world market prices are given, no such offsetting mechanism is at work in the traded goods sector. Hence volatility in this sector is higher.

Although the main focus of this paper is the impact of domestic financial development on trade policy, it is obvious that an international dimension exists as well. If trade raises aggregate risk, as Rodrik (1998) argues, it is not possible to diversify this risk in purely domestic financial markets. Therefore, the amount of international risk sharing should also have a positive impact on openness to trade. In addition, Feeney and Hillman (2001b) observe that internationally open financial markets eliminate, or reduce, the interest in strategic trade policy. This being said, the literature on international risk sharing indicates that this effect is likely to be small. When summarizing the evidence, Lewis (1995) shows that the amount of consumption smoothing that takes place internationally is limited and that this amount is quite persistent over time.[6] Moreover, there is a strong 'home-bias' in equity holdings, suggesting that portfolios are not optimally diversified.[7] One reason for this could be the blurred distinction between international and domestic financial markets. Due to the presence of internationally active corporations and the cross-listing of companies, it is possible that international diversification can be achieved within the domestic

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sectors. If we interpret the share of tradable capital as the degree of financial development, the assumption of asymmetries between sectors is quite odd.

[5] Traca cites empirical evidence by Gottschalk and Moffit (1994) to motivate his model.

[6] The last point is important because if international risk sharing is constant over time, this effect will be captured by the country specific fixed effects when running panel regressions.

[7] Stulz (1999) finds more recent evidence that globalization has so far had quite a limited impact on the cost of capital to firms. Kraay et al. (2000) show that countries' foreign asset positions have been very persistent over time and have mainly taken the form of loans rather than equity during the 1966-1997 period. Both these papers indicate that international capital markets are not yet well integrated.

market. It is clear that whatever measures we use to capture the degree of domestic financial development, they will also capture this effect.

Another issue regarding international financial integration is concerned with the timing of liberalization events. The Feeney-Hillman model suggests that financial integration should precede trade liberalization. Generally, however, trade liberalization seems to precede or be simultaneous with international financial liberalization.[8] In practice, it is difficult to separate trade and financial liberalization from each other. As shown by Tamirisa (1999), capital controls can effectively work as an impediment to trade. Thus, measures of financial openness, rather than explaining trade policy, may be part of what we wish to explain. Since trade and financial liberalization can be part of the same policy, questions concerning the timing between the two types of events are hard to sort out.[9]

3. From Theory to Estimation

There are, of course, other determinants of trade policy besides the concern for insurance. The optimal tariff argument makes it clear that countries (economically) large enough to affect international goods prices can increase their welfare by the introduction of a tariff. In a related vein, Alesina and Wacziarg (1998) argue that the cost of self-sufficiency is lower for large than for small countries. Countries with large markets should therefore be less open to trade than countries with small domestic markets. As the demand for variety in the choice of goods is likely to increase with wealth, per capita GDP is another probable determinant of trade policy.[10] This leaves us with the following basic trade policy equation to estimate:

$$\textit{Trade Policy} = f(\textit{Market Size, GDP, Financial Development})$$

[8] We thank an anonymous referee for making this point.

[9] There are of course other factors that can explain financial liberalization. One is that the gains from the removal of capital restrictions can increase after an increase in trade since the volume of international transactions has gone up. Another reason (not related to trade) is that governments running deficits might want to free capital movements in order to get access to international credit more cheaply.

[10] The inclusion of per capita GDP in the trade policy equation can be motivated by other arguments as well: Industrial specialization and dependence on imported intermediate goods are just some factors likely to increase with GDP.

Since the institutional environment is roughly the same for all sectors within a country, and risk diversification is essentially an inter-sector activity, the natural level of comparison is between countries. It is of course possible that the need and the opportunities for risk diversification differ between industries. However, since it is unclear in what way industries differ, we argue that the most suitable approach is to use country level data. Thus, we analyze how aggregate measures of financial development affect the aggregate trade policy choices made in a country.

Due to the level of aggregation, it is possible neither to discriminate between different sources of uncertainty, nor to be explicit about the mechanism that causes financial development to affect openness to trade. What is possible, however, is to control both for aggregate risk caused by openness and for aggregate income uncertainty. This is of importance since domestic asset markets cannot diversify aggregate risk.

A standard cross-section approach has the disadvantage of being a static approach to the essentially dynamic problem of financial development and trade policy. To allow for a time dimension, we make extensive use of panel data. Panel data has a number of advantages compared to both cross-section and time-series analysis, the most obvious being the ability to control for time and country specific fixed effects. In addition, the panel approach allows us to undertake causality tests, which are not possible in a cross-section setting.

4. Measurement issues and data

We now turn to the problem of how to measure trade policy and financial sectors with the available data.

4.1. Measuring trade policy

There is a huge literature discussing the pros and cons of different aggregate measures of the restrictiveness of trade policy (see e.g. Harrison 1996). The conclusion to be drawn from these studies is that no fully satisfactory measure is available. For our purpose, the measures should be objectively comparable across countries and time.

A popular direct measure of trade policy is the Sachs-Warner (1995) index. In their study of the period between 1950-94, a country is judged as open when it does *not* fulfill any one of the following criteria: (i) average tariffs are higher than 40%, (ii) non-tariff trade barriers cover more than 40% of imports, (iii) the economic system is considered socialist, (iv) major exports are monopolized by the state, and (v) the black market exchange rate premium exceeded 20%. The fraction of years between 1950-94 when the country is judged as open is then used to construct the index. A problem with the Sachs-Warner indicator is that it only considers the discrete nature of trade policy and not the degree of restrictiveness. Moreover, the different criteria might not be equivalent when evaluating the protectionist impact of trade policy. A more serious criticism of the Sachs-Warner index is put forward by Rodriguez and Rodrik (1999). They argue that the index serves as proxy for a wide range of policy and institutional differences and not only of trade policy. Despite these criticisms we will make use of the index, since it attempts to handle the problem of aggregating and combining different aspects of trade policy coherently across countries and time.

The second indicator of trade policy to be used is openness, measured as the ratio of the sum of imports and exports to GDP. Openness itself is not a measure of trade *policy* since trade is determined by other factors than policy. Lee (1993) constructs a simple measure of free trade openness, which controls for distance to the world's major trading economies and land area. The main advantage compared to the incidence measures is that all relevant trade restrictions are captured in a single, aggregate measure. The most obvious shortcoming here is the hypothetical counterfactual under free trade, making the measure sensitive to misspecifications of the trade equation. Despite these limitations, we will follow Lee and control for structural features of the economy and use trade share to GDP as a measure of trade policy (OPEN in this paper).

In addition to the measures described above we make some use of other indicators of trade policy. First, the ratio of import duty revenue to the value of total imports (IMPDUT) is used as a proxy for effective import tariffs. The main problem with the import duty ratio is that it does not have a linear relation to the effective degree of protection. If tariffs are prohibitive, they will not show up in this measure. Further, especially in poor countries, trade taxes are an important source of government

revenue. Finally, a drawback with all measures based on tariffs is that many restrictions on trade are not tariff-based. The second additional measure of trade policy, the pre Uruguay round non-tariff trade barriers (NTB), tries to remedy this problem. This is not unproblematic since it disregards the use of tariffs as impediments to trade. Moreover, the cover ratio is just an indicator of the share of all traded goods that are subject to quotas and similar restrictions. This means that nothing can be said about whether these restrictions are binding or not.

4.2. Measuring financial development

The purpose of the paper is to investigate whether the financial system, in its role as an insurance mechanism, is correlated with trade openness. Thus, we need a measure describing the financial systems ability to hedge, diversify, and pool risks.

The possible proxies for financial development can be divided into three different categories: the size of financial sector, the financial systems ability to allocate credit, and the real interest rate. Since the real interest rate is largely affected by macroeconomic factors, it will not be used in this study. A general problem with size-based measures is that the size of the financial sector does not necessarily measure its capacity to diversify risk.

The most popular measure for size of financial sector is the ratio of liquid liabilities to GDP, labeled LLY in this paper. A potential problem with this measure is that it can be too high in countries with undeveloped financial markets, since no other value-keeping asset than money exists. To allocate credit is a major function of the financial sector, especially the banking industry. Proxies focusing of the financial system's ability to allocate credits have been developed by for example King and Levine (1993). Since we are interested in the financial system's ability to diversify private sector risk, we will use credit issued to private enterprises divided by GDP, and label it DC.

Another measure focuses on the stock market. Levine and Zervos (1998) measure stock market capitalization by the value of listed companies on the stock market as share of GDP in a given year (here labeled MCAP). Although large markets do not

necessarily function effectively, many researchers use capitalization as an indicator of stock market development. Compared to the other measures, MCAP is intuitively better related to the underlying idea of portfolio diversification than LLY and DC. Unfortunately, this variable is not available prior to 1975, and the number of countries for which MCAP is available is also more limited than for LLY and DC. The correlations between LLY, DC, MCAP, and GDP (among other variables) are presented in Table A2.[11]

4.3. Measuring the possibilities of international risk-sharing

In their report *Exchange Arrangements and Exchange Restrictions,* the IMF annually summarizes the restrictions on international capital markets that each country imposes. These indicators are given the number one if a certain restriction is imposed, and zero otherwise. Although the binary nature of these indicators makes them less than ideal, they are the only available measure for a wide range of countries over time (we have data from between 1967-1993). Moreover, they have been found to have significant explanatory power on international consumption risk sharing (Lewis 1996), thus making them suitable for our purposes. In order to account for the differences in the *degree* of restrictiveness between countries, the simple annual average of four indicators is calculated and labeled CAPCONT.[12] These annual averages are then converted into five-year averages in order to fit the rest of the data.

4.4. Scope of the data set

The panel in this study is constructed for the years 1960-1994. To smooth short-term fluctuations, and to fill gaps in the series, all time varying variables are averages over five year periods. The selection of countries is based on the widely used Barro-Lee (1994) data set, which contains data for 138 countries. Data availability restricts the sample for some regressions to around 80 countries (for more details on the data, see

[11] In Table A2, the correlations in the 1990-94 cross section are presented. The results are essentially unchanged when considering the full sample of observations.

[12] The indicators are 'Bilateral payments', 'Restrictions for payments on capital transactions', 'Restrictions for payments on current account transactions', and 'Proscribed currency/payment arrears'. See Lewis (1996) for a thorough discussion of the data. Using alternate versions of this index does not affect the results to any significant degree. However, when using the widest measure (capital

Table A1 in the Appendix). When running panel regressions, we first remove countries for which one or more variables are available for one (or no) time periods. This procedure allows us to make better comparisons between fixed and random effects estimations. We also check for outliers and in the panel study remove Hong Kong and Singapore in the regressions with OPEN as dependent variable. This is done because these countries display an extreme degree of trade, which largely consists of transit trade.

5. Results

Now we are ready to formulate the specifications we would like to estimate. Theory predicts that there will be a positive relation between well developed financial markets and openness to trade, even when controlling for other determinants of trade policy.

5.1. The Sachs-Warner index

We begin by looking at the Sachs-Warner index. Given the discussion in Section 2, we use land area and population as proxies for country size, and GDP per capita as a measure of wealth. According to theory, the size proxies should have a negative effect and GDP a positive effect on openness. Moreover, we include dummies for geographical region (OECD, East Asia, Latin America and Sub-Saharan Africa) and proxies for financial development in the regression. In line with our hypothesis, we expect the proxies for financial development to enter with a positive sign. Since the Sachs-Warner index is an aggregate index based on the fraction of years a country has been open since 1950, we can not take the time variation of the variables into account. Since it is not obvious which time period to use in OLS regressions, we use the average of the explanatory variables to estimate the following equation:

$$Sachs\text{-}Warner_i = \alpha + \beta_1 AREA_i + \beta_2 POP_i + \beta_3 GDP_i + \beta_4 FD_i + \beta_5 Region_i + \varepsilon_i$$

transactions) by itself, this variable loses its significance. Since most countries have *some* restrictions on international transactions, this should be of no surprise.

where FD_i is one of our measures of financial development. The results from the estimations are presented in Table 1. The coefficients on LLY and DC are positive and significantly different from zero, as expected. MCAP, however, is not significant on conventional levels. One possible explanation is that the MCAP sample mainly includes relatively rich countries, which score high on the Sachs-Warner index. Per capita GDP is positive, but not statistically significant. Population enters, as expected, negatively into the regressions although not always statistically significant. The other proxy for size – area is never significant. Due to the construction of the index, the interpretation of the coefficient values of 0.52 and 0.37 is not fully clear. In the full sample of observations, the medians of LLY and DC are around 29 and 22, with standard deviations of 24 and 26, respectively. The median of the Sachs-Warner index is 17. An increase in LLY or DC by approximately one standard deviation (25 percentage points) would yield an increase in the index by about 10 points. For a country around the median of the Sachs-Warner scale (Philippines), an increase of 10 points means getting ahead of 12 countries in the openness ranking. Thus, the results indicate a strong positive relationship between openness and financial development.

(Table 1 here)

As discussed in Section 2, we would like to control for international financial openness. The measures at hand, however, partly overlap with the Sachs-Warner index and should thus not be included in the regressions. However, if we nevertheless include the measure CAPCONT, the results remain virtually unchanged.

As mentioned in Section 2, causality may run either from financial development to trade liberalization or the other way around if the demand for insurance increases after the switch of trade regime. Since the Sachs-Warner measure is constructed as an average over the period 1950-1994, it is not obvious how to investigate the question of causality between financial development and trade liberalization using this index. In Table 2, we show the percentage change in financial development for the five-year periods preceding and succeeding the actual date of liberalization (most countries just liberalize once and then remain open).[13] The first two rows indicate that after an economy moved from closed to open according to the Sachs-Warner index, financial

development accelerated. From the bottom two rows, however, we read that the countries that are open during the whole period have a higher initial level of financial development than closed countries. Investigating in this way shows that the direction of causality seems to run from trade liberalization to financial development. This indicates that the demand for the services provided by the financial sector increases after trade liberalization. Another possibility is that Rodriguez and Rodrik (1999) are correct when criticizing the Sachs-Warner index for reflecting a general switch to market friendly practices.[14]

(Table 2 here)

5.2. Openness: cross-sectional results

The other approach chosen to investigate whether financial markets affect trade policy is to use the direct measure of trade, i.e. OPEN. Since actual trade is not a measure of trade policy, we control for structural factors such as population and area. Not only does country size affect trade policy, it also affects the country's propensity to trade. This assumption is based on gravity models, which show that, everything else being equal, large countries will tend to trade less than smaller ones. GDP per capita is included given the reasons presented in Section 3. We use the measure of aggregate distance as a proxy for transportation costs, since high transportation costs are likely to decrease trade by making it less profitable. Finally, regional dummies are included. This gives us the following baseline cross-section specification:

$$OPEN_i = \alpha + \beta_1 AREA_i + \beta_2 POP_i + \beta_3 DIST_i + \beta_4 GDP_i + \beta_5 FD_i + \beta_6 Region_i + \varepsilon_i$$

The results from cross-sectional regressions on 1990-94 data are presented in Table 3. We estimate the baseline regression both with and without the inclusion of measures of financial development. In the presented regressions, all variables are in logs. We can motivate a log specification on theoretical grounds: it is reasonable to assume that the risk reducing effect is more important when starting from a low degree of

[13] The results are qualitatively the same when using a 10-year period rather than a five-year period.

financial development.[15] The results, however, are not contingent on the log specification. In the baseline openness equation (column 2), population and distance indeed come out with the expected negative sign. Moreover, the proxy for financial markets, LLY, is statistically significant and positive as predicted by theory. In columns 3 and 4, we show that this relationship holds for both DC and MCAP as well.

Comparing columns 2 and 3, we see that the adjusted R^2 increases by 8 percentage points when adding LLY to the openness equation, implying that multicollinearity between GDP and LLY is not what is driving the results.[16] Further, we see that the point estimates vary quite a bit between the proxies, making it difficult to judge the effect of the development of financial markets on trade. Moreover, it must be remembered that we are dealing with *proxies,* making the interpretation of the slope coefficients slightly unclear. However, if we consider the point estimate of 0.412,[17] then an increase in LLY by 10 percent of GDP would imply an increase in the trade to GDP ratio by 4 percentage points. To get some further intuition about the size of this effect, note that the median values of LLY and OPEN in the period 1990-94 are around 40 and 60, with a standard deviation of 27 and 41, respectively. Increasing LLY by one standard deviation would then be associated with an increase in the trade to GDP ratio by 11 percentage points. For the country with median openness, this is an increase in the trade share of GDP by 18 percent. Repeating this exercise for DC and MCAP shows that an increase by one standard deviation in the respective variable would increase OPEN by 7 and 5 percentage points respectively. All in all, these estimates indicate that an increase by one standard deviation in financial development increases openness to trade by between 8-18 percent for a country with median openness.

(Table 3 here)

[14] It is also possible that investments increase after liberalization (Wacziarg 1998), and that the proxies of financial development are related to investments. However, controlling for the investment share of GDP does not affect the results.

[15] This argument does not apply to the special case of CARA utility.

[16] For the log-specifications, the adjusted R^2 without the proxies for financial development are 0.55, 0.55, and 0.74, for the respective sample. This means that the increase in explanatory power from adding financial proxies is between 4 and 8 percentage points.

[17] The point estimate is higher in 1990 than any other time period. The coefficients on DC and MCAP are more stable over time.

The relationship between LLY and OPEN holds for all cross-sections except the 1960-64 period. For DC, the result is somewhat weaker, although still pretty strong. In 1985-89, the effect is borderline significant (p-value 0.115) and it fails to hold in 1960-64 and 1965-69. MCAP is strongly significant, both in the 1980-84, 1985-89 and 1990-94 periods. In the first period of availability, 1975-79, MCAP is not significant. In this period, the very small sample of countries is likely to be part of the explanation. Since MCAP is the proxy closest to our idea of portfolio diversification, it is encouraging for the hypothesis that it remains significant at the 1% level in the three last time periods.

Aggregate vs. diversifiable risk

In Section 2, the difference between domestically diversifiable risk and aggregate risk was discussed. Domestic insurance markets cannot diversify aggregate risk – access to international insurance markets is necessary for that purpose. In Table 4, we show the result of some tests for LLY and MCAP that consider these issues.[18] In the first column we add the index of international financial openness. The sign of this variable is negative (although not significant) as we would expect it to be if access to international financial markets helps facilitate an open trade policy. Including the variable CAPCONT does not affect the estimates of LLY and MCAP. In the next two columns, we control for variables likely to increase aggregate external risk. The first of these variables is an index on the product concentration of exports, CI,[19] the second the share of primary exports of all exports, PRIMSH. It is reasonable to assume that a high value on either of these variables will increase aggregate income volatility caused by price movements on the international markets. Since domestic risk markets cannot diversify aggregate risk, controlling for external risk is the equivalent to controlling for non-diversifiable risk. The inclusion of these variables does not affect the results for LLY or MCAP. In the last column we include the share of foreign direct investments to GDP, FDI, in order to account for the possibility of a spurious

[18] The results are essentially the same for DC. The reason why both LLY and MCAP are presented is that the sample differs quite a bit between the variables.

[19] The Gini-Hirschman index of concentration over 239 three-digit SITC export categories, calculated by UNCTAD.

correlation between financial development and trade.[20] It is plausible that FDI is positively related to both trade and financial development. This inclusion does not affect the basic results.

(Table 4 here)

5.3. Openness: panel results

By using a panel of data, we can go beyond the simple cross-section approach, controlling also for time and country specific effects, as well as bringing time into the analysis. The baseline panel specification is:

$$OPEN_{it} = \alpha + \beta_1 AREA_i + \beta_2 POP_{it} + \beta_3 DIST_i + \beta_4 GDP_{it} + \beta_5 FD_{it} + \beta_6 Region_i + \lambda_t + \nu_i + \varepsilon_{it}$$

Where λ_t is a time-specific effect, constant over countries, ν_i is a country-specific effect, constant over time, and ε_{it} is the usual residual. AREA and DIST will not be included in fixed effects estimations since these variables are time-invariant and can hence not be distinguished from the country specific effects. We introduce further variables later on.

The baseline panel results are presented in Table 5. The results are very similar to the ones in the cross-country setting, although the point estimates on our proxies for financial development are somewhat smaller. We have also run random effects regressions, but since these results are similar to the fixed effect estimations, we do not present them.[21] More generally, the fixed effects approach is an attempt to account for the changes in openness which have occurred between 1960-94. To account for these changes, we would need more (time varying) explanatory variables than the

[20] We have also controlled for the investment share to GDP, human capital, and population density but this does not affect the results. This is further evidence that the measures of financial development are not proxies for investments (compare note 14).

[21] The great exception in the difference in explanatory power – the 'within' R^2 that applies to fixed effect estimations and the 'overall' R^2 applying to random effects. The overall R^2 is between 0.7-0.8 depending on sample and specification. Since some of our control variables for trade are time invariant (DIST, AREA, regional dummies), it should be of no surprise that the within R^2 is much lower than the overall R^2.

ones we have included in the baseline regressions. However, the result that financial development is positively related to openness, even after controlling for both fixed country effects and time-specific events, is encouraging for our basic hypothesis.

(Table 5 here)

Aggregate risk, once again

To check the robustness of the panel results, especially with respect to aggregate and domestically non-diversifiable risk, we continue our study by including additional control variables. The results from these regressions are reported in Table 6.[22] The first variable we include is CAPCONT, the index capital controls. As we would suspect, the sign of the coefficient is negative, although the statistical significance varies between specifications. The negative sign gives support for the hypothesis that international risk sharing is an important determinant of trade policy. In Section 2 we mentioned some caveats with this measure. Most importantly, capital controls can effectively work as trade restrictions, thus belonging on the left-hand side of the regression. However, regardless of the exact mechanism involved, the sign should be negative. Domestic financial markets still have a positive and significant impact on trade in all specifications. Hence we can conclude that the basic result is not due to a correlation between international financial restrictions and financial development. Both the development of domestic asset markets and the integration on international financial markets have independent effects on openness to trade.

(Table 6 here)

Next, we include TOT, terms of trade shocks. TOT is defined as the growth rate of export prices minus the growth rate of import prices. We would expect the sign of this coefficient to be positive: a country is likely to trade more if export prices are rising and import prices are falling (or growing at a slower rate). This prediction is supported by the data and the inclusion of TOT does not affect the point estimates of

[22] All specifications have been estimated for DC as well. These regressions are virtually identical to the LLY-regressions so we do not present the results.

LLY and MCAP (TOT is not included in the baseline specification since it is not available for the full time period).

In the cross-section regressions we had to rely on export concentration, CI, and the share of primary resources in exports, PRIMSH, as proxies for aggregate external risk. The panel setting allows for more direct ways of approaching this problem. In columns 5 and 6 of Table 6, we control for *total* aggregate risk as measured by the standard deviation of per capita GDP during the period 1960-92. The variable is not significant and does not affect the coefficients on LLY or MCAP. In the last two columns, we control for aggregate *external* risk, measured by the standard deviation of terms of trade multiplied by average openness. Given its construction, it is of no surprise that this variable is positively related to openness. Since the key estimates are not affected by these inclusions, we conclude that the effect of financial markets on openness is not caused by a correlation with external risk. Rather, the stability of the coefficient even after the inclusion of aggregate risk measures indicates that domestic risk sharing is what matters for openness. As our controls for aggregate risk are time-invariant, we must use the random effects estimator for the last four regressions.[23] For these estimations, we perform the Hausman specification test. If the empirical model is correctly specified, and the Hausman test returns a significant result, this can be interpreted as evidence that the individual specific effects and the regressors are correlated and hence that fixed effects estimation should be used. As can be seen in the last row of Table 6, the Hausman test indicates that the random effects estimator is appropriate for the LLY regressions, but not for the MCAP regressions. Despite this limitation, we conclude that domestic asset markets have an independent positive relation with openness to trade, and that access to international asset seems to have a positive impact on trade, although this result is somewhat weaker.

5.4. Alternative measures of openness

As discussed above, there are many possible, but no ideal, measures of openness that have been used in the literature. Before moving on to issues regarding causality and

[23] It should also be noted that the effect of aggregate risk, measured in these ways, is captured in the country specific fixed effects when running fixed effects estimations.

simultaneity, we take a look at the measures IMPDUT and NTB's.[24] Although direct indicators of trade policy, these measures suffer from drawbacks mentioned in Section 4.1. For these reasons, and for the lack of very interesting results, this section is kept brief.

There are good reasons to expect the use of these barriers to vary between rich and poor countries. Therefore, we split the sample according to average per capita GDP (for the relevant time period). Since no significant effects are found for the full sample of countries for any of the trade policy indicators, Table 7 only includes the sub-sample regressions. When using the proxy LLY for financial development, no significant results are found in any specification (the same is true when using DC). When using MCAP, our favored measure, we see from columns one and three that the variable enters with the predicted sign in the sub-sample of relatively rich countries. The absence of results for poor countries could reflect their reliance on trade taxes as a source of government revenue. Since trade-related risk may also be diversified on international markets, we also include the measure of international financial integration, CAPCONT, to the regressions (not presented). NTB-cover ratios are not affected by this inclusion. We do find support for the hypothesis in rich countries, when import duties is used as a measure of trade policy. However, when both CAPCONT and MCAP are included in the regressions, the latter loses its significance.

It must be kept in mind that these results only apply for a sub-section of the sample and that they are quite weak. Despite this, they do point in the direction suggested by theory. The weak concordance with the results for OPEN and the Sachs-Warner index partly reflects the difficulties in putting a single number on a country's aggregate trade policy. However, it also suggests that financial development has a stronger link to trade than to trade policy.

(Table 7)

[24] A combined measure of import- and export duty ratios is also constructed: TARIFFS = (1+IMPDUT)(1+EXPDUT)-1), as suggested in Rodriguez and Rodrik (1999). This measure is highly correlated with IMPDUT and hence, the results are essentially equivalent.

5.5. Causality

So far, we have established a strong relationship between our proxies for financial development and openness to trade, but have not mentioned issues of causality and simultaneity in relation to the variable OPEN. Finding a causal link from financial development to trade would lend specific support for the Feeney and Hillman (2001a) hypothesis. The reverse causality, from trade to financial development, would not, however, contradict the underlying theoretical reasoning. Increased exposure to the fluctuations of the international market could increase the demand for portfolio diversification. Finally, one can easily imagine political decisions affecting both variables simultaneously.

In order to explore the causality aspects, we first turn to the concept of Granger causality tests. The Granger test amounts to checking whether the lagged independent variables are jointly significant in a regression of the dependent variable on its own lagged values, i.e. a regression of the following kind:

$$y_{it} = \alpha_0 + \sum_{r=1}^{n} \alpha_r y_{it-r} + \sum_{r=1}^{n} \beta_r X_{it-r} + v_i + \varepsilon_{it}$$

The inclusion of the lagged dependent variable on the RHS creates a dynamic panel data problem: the lagged dependent variable is correlated with the fixed effect. To eliminate the bias caused by the presence of fixed effects, the equation is estimated in first differences. Since first the differentiation induces MA(1) residuals, the lagged difference of the dependent variable has to be instrumented for. This will be done using the twice-lagged difference and the twice-lagged level of the dependent variable as instruments.[25]

The estimates presented in Table 8 include only one lag, since 5-year averages make the time series rather short. Columns 1 and 2 show that the causality runs both ways. LLY and DC do Granger-cause openness, but there is also an effect of openness on the two proxies for financial development. MCAP, however, does not Granger-cause OPEN instead the causality clearly runs in the opposite direction. Thus, which

[25] The dynamic panel data problem and its solutions are discussed in Baltagi (1995), chapter 8.

conclusion to draw depends on which proxy is used. Judging by the result for DC and LLY, there is an effect of financial markets on trade, while MCAP shows that there is no such effect. It should be kept in mind that we have very few time-periods available when testing for causality using MCAP.

(Table 8)

5.6. Simultaneity

In order to take into account the possible simultaneity, we use instruments for financial development to see if the exogenous component of each of the proxies is significant. Finding instruments for financial development that are not correlated with trade policy is not an easy task. LaPorta et al. (1997) have, however, come up with a number variables that could possibly serve our purposes. Unfortunately, these variables are all time-invariant, making panel regressions impossible.[26] Instead we limit our focus to the 1990-94 cross-section and attempt to instrument for our three proxies of financial development. The instruments we use are: (i) an index of minority shareholder protection that takes a value between 0-6 with a higher value indicating stronger minority rights, (ii) a 'rule of law' index constructed by the risk rating agency *International Country Risk* (ICR) that assesses the law and order tradition of a country and ranges between 0-10, with lower scores for less tradition of law and order, and (iii) the number of assassinations per million inhabitants. We use these instruments in order to capture the ideas that the protection of shareholders, an effective legal system, and a relatively safe environment are crucial elements for the development of financial markets. Why these instruments should have a relation to trade policy is less clear, and below we test for an independent effect of the instruments on openness to trade. Using these instruments limits the number of observations, and we end up with 36-38 countries in the regressions. Even though the sample is limited, we see in Table 9 that LLY, DC, and MCAP are all significant with the expected sign. The level of significance in a bit lower than in the OLS-estimations, but this should be of no surprise since we lose efficiency by using instrumental variables. Compared to the

[26] As explained above, all time-invariant variables are captured in the country specific fixed effects.

OLS-estimates in Table 3, all point estimates are much larger, suggesting elasticities between 0.2-0.6.[27]

(Table 9 here)

In order to account for the validity of the instruments, we also report the test statistics from the Hansen over-identification test in Table 9.[28] Since the 10% critical value equals 4.61, all three IV-regressions pass the test by a wide margin.

Using the instrumental variables of financial development on the Sachs-Warner index does not yield any significant results. One reason could be that the instruments are from the late 1980's or the early 1990's, while the Sachs-Warner index is an average between 1950-1994. Thus, even if the results had been significant, we would have had reasons to question the exogeneity of the instruments.

These results are broadly consistent with the results on Granger causality. LLY and DC Granger cause OPEN, and have an exogenous effect on OPEN when using instrumental variables. MCAP seems to have an exogenous effect on openness, even though the Granger tests tell a different story.

Finally, we consider the possibility that both openness and financial development are simultaneously caused by the same underlying variable, not yet included in the regressions. One obvious possibility is that market friendly policies in general could affect both financial development and trade policies. To control for this, an index of "regulatory burden" described in Kaufmann et.al. (1999) is added to the regressions. This index is a combination of measures of market unfriendly practices. Further, dummy variables of a country's legal origin as described in LaPorta et.al. (1997) are included on the right hand side of the regressions. The inclusion of these variables does not affect the results and the regressions are not presented.

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[27] In the IV-samples, the OLS-estimates for LLY, DC, and MCAP are 0.23, 0.15, and 0.21 respectively.

[28] This test assesses if the instruments have an independent effect of openness to trade beyond their ability to explain cross-country variation in financial development. The test statistic is obtained by running the residuals from the second stage regression on the instruments and multiplying the R^2 from this regression with the number of observations. Under the null-hypothesis that the instruments are not correlated with the error term, the test is distributed χ^2 with $(j-k)$ degrees of freedom, where j is the number of instruments, and k the number of variables instrumented for.

6. Conclusion

Previous work on the relation between trade policy and financial markets has been purely theoretical. This paper is a first attempt to empirically investigate the interdependence between financial development and openness to trade. Our major finding is that there does exist an economically significant relationship between these variables. In addition, the degree of integration on international financial markets has an independent effect on openness to trade. However, the results are somewhat dependent on the measure of trade policy that is used.

The positive relation is shown to hold both when the Sachs-Warner index and when structurally adjusted trade are used as measures of openness. The results are robust for a variety of specifications and econometric techniques. For proxies of tariff levels and non-tariff trade barriers the results are weaker and only hold for relatively rich countries. Evidence of simultaneity between trade and financial development is found, and the direction of causality seems to be running both from financial development to the volume of trade and in the opposite direction. Considering the possible simultaneity, we find support for an exogenous effect of financial markets on openness to trade using instrumental variables. However, when looking at trade liberalization according to the Sachs-Warner index, we find that financial development is stronger after, rather than before, the liberalization event. Despite our efforts to establish causality the results are inconclusive and hence the exact mechanism behind the effect is not revealed in this study. This is perhaps not surprising since issues concerning causality are notoriously difficult to settle using cross-national panel data.

This paper can be seen as a complement to the recent literature suggesting that risk reducing policies and trade policy is interdependent. In an oft-quoted paper, Rodrik (1998) presents theoretical and empirical arguments that countries more open to trade show greater fluctuations in income, and that the government sector is expanded in order to reduce these fluctuations. Since risk is aggregate in his model, private risk diversification within a country is not possible. Rodrik argues informally and shows empirically, however, that policies that reduce risk *between* groups within a country are also an important response to openness to trade. Agell (1999) follows the same

line of thought when showing that open countries are more prone to provide insurance through labor market regulations than closed ones. The common theme in these two papers is that risk-increasing policies can be optimally combined with policies reducing risk. Our paper makes the point that there may be interdependence between opportunities for private risk diversification and risk-increasing policies. This point should apply to other areas than trade policy and is of interest when discussing, among other issues, the timing and political feasibility of policy reform.

Our results qualify some conclusions drawn in the recent literature. Both Agell (1999) and Rodrik (1997) warn that the current trend towards globalization of economic activity leads to a greater exposure to risk, and thus greater demand for risk-reducing reforms, while simultaneously reducing the scope for government interventions especially through taxation and labor-market regulations. However, globalization also encompasses financial markets. If domestic asset markets and the integration of international asset markets facilitate liberal trade policies, as our results suggest, better domestic and international financial markets may well alleviate the negative effects of globalization pointed out by Agell and Rodrik.

Table A1. Descriptive statistics

Variable (years available)	Description	Source	Mean	Standard deviation	N obs
Assassina-tions	Number of assassinations per million inhabitants	World Bank1	0.261	0.474	70
AREA (1960-94)	Land area in km^2	WDI	761052	1693138	134
CAPCONT (1967-92)	Restrictions index on inter-national capital movements	IMF	0.503	0.301	757
CI (1990)	Export concentration index, defined in text	UNCTAD	0.401	0.247	122
DC (1960-94)	Financial resources to private sector, % of GDP	IMF-IFS	30.809	26.148	797
DIST (1960-94)	Distance to 20 major trading economies	Barro-Lee	5.953	2.311	90
FDI (1990-94)	Foreign direct investment, net inflow, % of GDP	WDI	1.798	2.997	126
IMPDUT (1975-94)	% tariff revenue to value of imports.	WDI	12.52	9.435	459
Log(GDP) (1960-92)	(Log of) real per capita GDP	PWT	7.685	1.027	853
LLY (1960-94)	Liquid liabilities, % of GDP	IMF-IFS	36.534	24.461	783
MCAP (1987-94)	Value of listed companies, % of GDP	World Bank1	26.975	40.976	217
Minority (1990)	Index of minority share holder protection	World Bank1	2.441	1.231	41
NTB (1985)	Non-tariff trade barrier cover ratio	Wacziarg (1998)	13.269	13.094	113
OPEN (1960-94)	Exports + imports, % of GDP	WDI	63.960	37.196	811
POP (1960-94)	Thousands of inhabitants	WDI	28941	101986	880
PRIMSH (1990)	Share of primary exports in total exports	WDI	0.657	0.313	116
Regburd (1990)	Index of market unfriendly practices	World Bank2	0.137	0.795	124
Rule-of-law (1990)	ICR index of law and order tradition	World Bank1	7.072	2.618	41
Sachs-Warner (1950-94)	Index of openness, Defined in text	SW	36.928	41.373	110
TOT (1965-92)	Growth of merchandise export prices minus growth of import prices	Wacziarg (1998)	-0.0067	0.066	803

PWT stands for Penn World Tables 5.6; WDI for World Bank, *World Development Indicators*; World Bank1 for the *Financial Structure and Economic Development Database*; World Bank2 for the World Bank *Governance Indicators* described in Kaufmann et.al. (1999); UNCTAD for *Handbook of International Trade and Development Statistics of UNCTAD*; IFC for International Finance Corporation, *Emerging Stock Market Factbook*; IMF for the IMF's *Exchange arrangement and Exchange Restrictions*; IMF-IFS for IMF's *International Finance Statistics;* Barro-Lee for Barro and Lee (1994); SW for Sachs and Warner (1995). SW, PWT and Barro-Lee are available at http://www.nuff.ox.ac.uk/Economics/Growth/datasets.htm. The *Financial Structure and Economic Development Database* and the *Governance Indicators* are available at http://www.worldbank.org/research.

Table A2. Correlation between main variables of interest (1990-94 cross section)[a)]

	OPEN	AREA	DIST	POP	SW	LLY	DC	MCAP	IMPD.	NTB
GDP	0.164*	0.190**	-0.516***	-0.043	0.750***	0.577***	0.752***	0.394***	-0.619***	-0.090
	(107)	(107)	(80)	(109)	(106)	(104)	(106)	(69)	(81)	(113)
OPEN		-0.291***	-0.038	-0.209***	0.342***	0.459***	0.221***	0.504***	-0.064	-0.005
		(130)	(90)	(130)	(120)	(125)	(127)	(76)	(96)	(107)
AREA			0.099	0.542***	0.018	0.003	0.167**	0.003	-0.110	-0.032
			(91)	(136)	(122)	(128)	(130)	(76)	(97)	(111)
DIST				0.104	-0.409***	-0.302***	-0.297***	-0.054	0.430***	0.158
				(92)	(91)	(86)	(88)	(61)	(68)	(86)
POP					-0.095	0.116	0.143*	-0.068	0.028	0.128
					(124)	(128)	(130)	(76)	(97)	(109)
Sachs-Warner						0.572***	0.641***	0.436***	-0.609***	-0.086
						(116)	(118)	(75)	(80)	(105)
LLY							0.737***	0.512***	-0.306***	0.083
							(128)	(73)	(95)	(102)
DC								0.622***	-0.539***	-0.025
								(75)	(97)	(104)
MCAP									-0.289**	-0.205
									(67)	(52)
IMPDUT										0.192*
										(89)

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. Figures in parentheses are number of observations. Note that the above correlations are virtually the same for the full sample of observations. [a)] The correlations between NTB and the other variables are for the 1985-89 period.

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Table 1. Sachs-Warner index

	SW	SW	SW
Constant	0.529	20.784*	14.197
	(0.053)	(1.926)	(1.463)
LLY	0.522***		
	(2.655)		
MCAP		0.241	
		(1.245)	
DC			0.376**
			(2.351)
GDP	0.001	0.002	0.001
	(0.637)	(0.689)	(0.637)
AREA	-0.004	-0.015	-0.010
	(-0.306)	(-1.262)	(-0.829)
POP	-0.326	-0.238	-0.386*
	(-1.627)	(-1.149)	(-1.720)
OECD	49.708***	50.733***	44.112***
	(3.672)	(2.661)	(2.929)
LAAM	10.692	7.436	2.195
	(1.174)	(0.620)	(0.219)
SAFRICA	-3.167	-1.344	-11.689
	(-0..346)	(-0.081)	(-1.158)
EASIA	38.483***	42.415***	34.026**
	(2.558)	(2.441)	(2.169)
N Countries	105	68	107
Adj. R^2	0.612	0.522	0.607

*** Indicate significance at 1%-level, ** at 5%-level,
* at 10%-level. t-statistics based on robust standard errors in parentheses.

Table 2. Before and after liberalization

	LLY	DC	MCAP
% Change	9.5	21.7	75.6
before lib.	(37)	(35)	(11)
% Change after	16.3	38.2	306.1
lib.	(23)	(21)	(8)
Initial level:[a)]	49.4	37.9	23.1
always open	(19)	(16)	(19)
Initial level:[a)]	13.6	11.1	3.6
always closed	(17)	(13)	(3)

Number of observations in parenthesis. [a)] 1960 for LLY and DC, 1980 for MCAP.

Table 3. Baseline cross section 1990-94

	Log(OPEN)	Log(OPEN)	Log(OPEN)	Log(OPEN)
Log(LLY)		0.412***		
		(3.988)		
Log(DC)			0.197***	
			(2.789)	
Log(MCAP)				0.138***
				(4.784)
Log(AREA)	-0.030	-0.017	-0.025	-0.017
	(-0.845)	(-0.551)	(-0.774)	(-0.709)
Log(DIST)	-0.289***	-0.343***	-0.335***	-0.486***
	(-2.801)	(-3.150)	(-0.335)	(-5.161)
Log(POP)	-0.223***	-0.208***	-0.223***	-0.252***
	(-4.683)	(-5.315)	(-5.608)	(-7.790)
Log(GDP)	0.065	-0.052	-0.052	-0.176**
	(0.759)	(-0.658)	(-0.593)	(-2.180)
OECD	-0.352**	-0.265	-0.357**	-0.273*
	(-2.178)	(-1.558)	(-2.037)	(-1.650)
LAAM	-0.298**	-0.010	-0.229**	-0.205*
	(-2.355)	(-0.064)	(-1.666)	(-1.677)
SAFRICA	-0.99	0.220	0.009	0.183
	(0.560)	(1.289)	(0.055)	(1.170)
EASIA	0.525***	0.610***	0.458***	0.533***
	(3.274)	(3.737)	(2.684)	(4.079)
Constant	6.602***	5.628***	6.851***	8.669***
	(7.252)	(7.427)	(9.127)	(11.250)
N obs.	77	75	77	55
Adj. R^2	0.552	0.635	0.592	0.793

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level.
t-statistics based on robust standard errors in parentheses.

Table 4. Additional controls. Dependent variable is Log(OPEN)

	Control: CAPCONT	Control: CI	Control: PRIMSH	Control: FDI
Log(LLY)	0.431***	0.419***	0.385***	0.436***
	(3.069)	(3.297)	(2.856)	(4.301)
Estimate of	-0.089	0.410*	0.196	0.101***
control	(-0.353)	(1.876)	(0.830)	(3.526)
N obs.	59	73	71	70
Adj. R^2	.631	0.656	0.695	0.744
Log(MCAP)	0.145***	0.147***	0.157***	0.131***
	(4.777)	(5.782)	(6.066)	(4.167)
Estimate of	-0.262	0.600***	0.587***	0.068**
control	(-1.050)	(4.513)	(3.522)	(2.327)
N obs.	38	54	54	53
Adj. R^2	0.832	0.827	0.820	0.812

***Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses. CAPCONT is an index of openness to international financial markets; TOT is the change in terms-of trade; CI is the export concentration index; PRIMSH is the share of primary resources in exports; FDI is foreign direct investments. All regressions include regional dummies, Log(AREA), Log(POP), Log(DIST), Log(GDP) and a constant.

Table 5. Panel, LLY and DC 1960-94, MCAP 1975-94.

	Log(OPEN) FE	Log(OPEN) FE	Log(OPEN) FE
Log(LLY)	0.185***		
	(5.410)		
Log(DC)		0.099***	
		(4.269)	
Log(MCAP)			0.060***
			(2.701)
Log(POP)	-0.202**	-0.088	0.352*
	(-2.244)	(-1.032)	(1.782)
Log(GDP)	0.191***	0.167***	-0.096
	(4.434)	(4.269)	(-0.977)
N obs	694	706	199
N countries	121	123	68
Within R^2	0.316	0.302	0.261

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses. All regressions include time-period dummies and country-specific fixed effects that are jointly significant.

Table 6. Additional tests. Dependent variable is Log(OPEN)

	Fixed Effects	Fixed Effects	Fixed Effects	Fixed Effects	Random Effects	Random Effects	Random Effects	Random Effects
Log(LLY)	0.167***		0.157***		0.139***		0.159***	
	(4.273)		(3.990)		(3.203)		(3.698)	
Log(MCAP)		0.051**		0.050**		0.061***		0.071***
		(1.996)		(1.969)		(3.019)		(3.753)
CAPCONT	-0.126**	-0.174	-0.150**	-0.166	-0.198***	-0.236**	-0.185***	-0.171
	(-1.917)	(-1.191)	(-2.241)	(-1.107)	(-2.808)	(-2.024)	(-2.664)	(1.600)
TOT			0.512***	0.134	0.594***	-0.037	0.624***	0.321
			(3.900)	(0.322)	(3.653)	(-0.093)	(3.753)	(0.812)
$\sigma(GDP)_{60\text{-}92}$					0.447	0.327		
					(1.395)	(0.938)		
$\sigma(TOT)_{71\text{-}90}$							0.022***	0.027***
$\times OPEN_{60\text{-}90}$							(3.920)	(3.583)
N obs.	596	175	579	173	431	154	420	151
N countries	116	62	115	61	81	52	79	52
Within R^2	0.316	0.254	0.320	0.255				
Overall R^2					0.716	0.848	0.760	0.873
Hausman[a)]					6.82	18.63	11.80	1786.87
[p-value]					[0.74]	[0.017]	[0.299]	[0.000]

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses. All regressions include Log(POP), Log(GDP) and time period dummies. The random effects estimates also include regional dummies, Log(AREA), and Log(DIST). $\sigma(GDP)_{60\text{-}92}$ is the standard deviation of GDP 1960-92, $\sigma(TOT)_{71\text{-}90}$ is the standard deviation of log differences in terms of trade 1971-90, $OPEN_{60\text{-}90}$ is average openness 1960-90. [a)] The Hausman test statistics is χ^2 distributed with the same degrees of freedom as the number of explanatory variables. The p-value indicates at which probability we can reject the hypothesis that the fixed effects and the regressors are correlated.

Table 7. Alternative measures of openness.

	IMPDUT > mean GDP/pc.	IMPDUT < mean GDP/pc.	NTB > mean GDP/pc.	NTB < mean GDP/pc.
Log(LLY)	-1.376	2.270	-3.437	6.337
	(-0.934)	(1.111)	(-0.753)	(1.233)
Log(MCAP)	-1.473**	1.508	-3.075*	0.850
	(-2.191)	(0.743)	(-1.790)	(0.428)

***Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses. In columns one and two, Log(POP), Log(GDP), time period and country specific fixed effects are also part of the regressions. In columns three and four, Log(AREA), Log(POP), Log(GDP), and continent dummies are also part of the regression.

Table 8. Granger causality between LLY/DC/MCAP and OPEN.

	Δ(OPEN)	Δ(OPEN)	Δ(OPEN)
$\Delta(OPEN)_{t-1}$	0.289	-0.082	1.231
	(0.699)	(-0.222)	(0.363)
$\Delta(LLY)_{t-1}$	0.277**		
	(2.521)		
$\Delta(DC)_{t-1}$		0.148**	
		(2.102)	
$\Delta(MCAP)_{t-1}$			-0.126
			(-0.406)
N obs	476	403	83
	Δ(LLY)	Δ(DC)	Δ(MCAP)
$\Delta(LLY)_{t-1}$	0.430**		
	(2.233)		
$\Delta(DC)_{t-1}$		0.462**	
		(2.193)	
$\Delta(MCAP)_{t-1}$			1.129**
			(2.119)
$\Delta(OPEN)_{t-1}$	0.115***	0.212***	1.565***
	(2.646)	(3.427)	(2.935)
N obs.	390	3.427	34

***Indicate significance at 1%-level, ** at 5%-level, * at 10%-level.t-statistics based on robust standard errors in parentheses. The lagged difference of the dependent variable is instrumented using the twice-lagged difference and the twice-lagged level of the dependent variable.

Table 9. Instrumental variables 1990-94

	Log(OPEN)	Log(OPEN)	Log(OPEN)
Log(LLY)	0.606**		
	(1.993)		
Log(DC)		0.606**	
		(2.011)	
Log(MCAP)			0.276*
			(1.866)
N obs.	36	38	37
R^2 of residuals on instruments	0.004	0.0001	0.053
OIR-test[a)]	0.144	0.004	1.961

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses. Regressions include constant, Log(GDP), Log(AREA), Log(DIST), Log(POP), and regional dummies. Instruments are an index of minority shareholder rights, the 'rule of law' index by *ICR*, and the number of assassinations per million inhabitants. [a)] The null hypothesis of the overidentification test is that the instruments are not correlated with the residuals from the second stage regression. Critical values for the test (2 d.f.): 10%=4.61, 5%=5.99.

CHAPTER 2

FINANCIAL MARKETS, THE PATTERN OF SPECIALIZATION AND COMPARATIVE ADVANTAGE. EVIDENCE FROM OECD COUNTRIES.[a]

Helena Svaleryd[*] and Jonas Vlachos[**]

Abstract

Due to underlying technological differences, industries differ in their need for external finance. Since services provided by the financial sector are largely immobile across countries, the pattern of specialization should be influenced by the degree of financial development. We find this effect to be strong. In fact, the financial sector has an even greater impact on the pattern of specialization among OECD countries than differences in human- and physical capital endowment. Further, it gives rise to comparative advantage in a way consistent with the Hecksher-Ohlin-Vanek model. Results on which aspects of financial systems that are of importance for specialization are also presented.

JEL classification: F14; G20; O16

Keywords: Financial intermediation; Financial systems; Specialization patterns; Comparative advantage

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[*] Department of Economics, Stockholm University. Email: hs@ne.su.se.

[**] *Corresponding author*. Department of Economics, Stockholm School of Economics, P.O. Box 6501, 11383 Stockholm, Sweden. Fax: +46-8-313207. Email: jonas.vlachos@hhs.se.

1. Introduction

In a modern economy, financial markets and financial intermediaries play an important role by mobilizing savings, allocating credit, and facilitating the hedging, pooling and pricing of risks.[1] That a well-functioning financial sector has strong, positive effects on a country's aggregate growth opportunities has been shown by, for example, Levine et al. (2000). Since the need for external financing through financial markets differs depending on in which type of activity firms are involved, it would be surprising if the growth effect was completely symmetric across sectors and firms. Resent research (Rajan and Zingales 1998, Demirgüc-Kunt and Maksimovic 1998, Beck and Levine 2001) has actually found evidence that firms and industries heavily dependent on external financing grow faster in countries with well-developed financial systems. Given these empirical results, it is only natural to expect trading and specialization patters to be influenced by the financial sector. This paper adds to earlier research, first by reporting that differences in financial development among OECD countries have an even greater impact on the pattern of specialization than differences in human or physical capital. Second, we find that well-developed financial intermediaries and markets have a positive effect on the content of external financing in net trade. In other words, the financial sector gives rise to comparative advantages in a way consistent with the Heckscher-Ohlin-Vanek (HOV) model.

When discussing financial development, it is important to keep in mind that the financial sector has developed along different lines in different countries. Consequently, there has been an intense debate on the relative merits of the different systems, traditionally divided into bank-based versus market-based systems. By using a variety of indicators of financial development, we attempt to assess the relative importance of different aspects of the financial system on specialization patterns and the strength of comparative advantage. Our findings indicate that large and active stock markets, and the degree of competition in the banking sector, have the strongest effect on both specialization patterns and the trade content of external financing. Further, there is support for the view that the quality of investor information and the

[1] The contribution of the financial sector to GDP is large. Demirgüc-Kunt and Levine (1996) present estimates varying from around 5% of GDP in the US to 9% in Japan during the years around 1990.

legal protection of creditors affect the pattern of industry specialization, while the depth of the financial system is a source of comparative advantage.

An obvious prediction of the standard Heckscher-Ohlin-Vanek (HOV) model is that a country well endowed with institutions of relatively high quality should tend to specialize in the production of goods relatively intense in the use of services provided by these institutions. This study treats financial markets and intermediaries as factors in the production of goods and services. A necessary condition for a production factor to give rise to comparative advantage is that it is immobile across countries. If financial intermediation were internationally mobile, however, we would not expect the strong growth effect of domestic financial development that is found in the empirical growth literature. Moreover, Jayrathne and Strahan (1996) show that the services provided by the financial sector are indeed highly immobile geographically, even within the USA.

This paper belongs to the small empirical literature investigating the effects of institutions on trade. In Svaleryd and Vlachos (2001), we find an economically significant relation between the degree of financial development and aggregate openness to trade. Anderson and Marcouiller (1999) find corruption and imperfect contract enforcement to be important determinants of aggregate bilateral trading volumes. To our knowledge, the present paper is the first to analyze empirically how financial markets affect industry specialization patterns and international competitiveness.[2] More broadly speaking, this is the first paper documenting that the institutional features of a society can give rise to comparative advantage. It also contributes to the literature on financial market and growth by focusing on absolute levels of production rather than growth rates. Finally, we provide new, indirect evidence to the debate on the relative merits of different financial systems to generate capital.

The paper is organized as follows. In Section 2, we start by discussing different aspects of the financial sector. Section 3 describes our measures and data of financial

[2] After the completion of this paper, the independent work by Beck (2001) was released. Beck addresses the same questions, using basically the same methods. He does not, however, control for a

intermediation and other variables. Sections 4 and 5 present the results for industry specialization and factor content of net trade, respectively, and Section 6 concludes.

2. The financial sector

2.1. The financial sector as an endowment

What do we mean by our claim that the financial sector effectively works as an internationally immobile factor endowment? The question is important since Wood (1994) has shown that the inclusion of internationally mobile production factors in studies of the factor content of trade can yield incorrect predictions. Especially, he argues that since capital mobility has (more or less) equalized real interest rates across countries, capital cannot be a source of comparative advantage. This line of reasoning abstracts from the well-known imperfections of financial markets arising from informational asymmetries and conflicting interests between creditors and debtors however (see, for example, Stiglitz and Weiss 1981).

These problems have given rise to financial intermediaries specializing in project evaluation and monitoring and information dissemination, thereby mitigating the negative effects of market imperfections. Two countries with the same real interest rate, but with financial sectors of differing quality, are thus, in practice, differently endowed with financial capital. Alternatively, the problem can be seen from the perspective of the firm or industry: Industries heavily involved in projects subject to especially strong informational problems stand to gain most from the development of financial intermediaries, even if this development does not affect market interest rates. There is a huge literature on the underlying causes of, and possible remedies to, these problems. The degree of project uncertainty (Huang and Xu 1999) and the share of investments in intangible assets (Myers and Majuf 1984) are just two of the factors that make financial intermediation more important. Hence, it should be clear that financial intermediaries do not just raise money for financing investments in physical capital. In fact, it is difficult to have a clear prior on the factor content of the investments made with financial capital, which makes us draw the conclusion that the

wider range of production factors. Hence, he cannot relate the size of the effect of financial markets on the pattern of specialization to the effect of other factors.

financial sector is best viewed as a type of human or organizational capital, specialized in overcoming market distortions in financing.

But are not the services provided by the financial sector internationally tradable, thereby erasing this source of comparative advantage? Several results in the empirical growth literature suggest otherwise. In a recent study, Levine et al. (2000) demonstrate that the *domestic* level of financial development is an important determinant of its economic growth. Following La Porta et al. (1997, 1998), they also show that a country's legal origin and the legal environment have strong effects on the development of the financial sector. Wurgler (2000) shows that more capital is allocated to growing industries, and less to declining industries, in countries with well-developed financial systems compared to other countries. This improved allocation of capital can explain why Beck et al. (2000) find that the development of financial intermediaries has a positive effect on total factor productivity growth.

Several micro-oriented studies also present results indicating the non-tradable character of financial services. Demirgüc-Kunt and Maksimovic (1998) show that firms highly dependent on external finance, located in countries with efficient financial and legal systems, tend to grow faster than similar firms in other countries. In an influential study, Rajan and Zingales (1998) show that the same result applies at the industry level. Giannetti (2000) presents evidence that the ease with which firms investing in intangible assets obtain loans depends on the legal system and the level of financial development. More direct evidence is found in Jayaratne and Strahan (1996), who show that financial services are difficult to trade geographically even within a country. In those US states that experienced relaxation in bank branch restrictions, the quality of bank loans improved and per capita income grew compared to those without banking deregulation. Finally, by investigating companies' cross-listing decisions, Pagano et al. (2001) conclude that geography is still of importance for financing.

All these studies demonstrate that the use of financial capital that, admittedly, is internationally mobile, to a large extent depends on the immobile institutional features of a society, summarized in measures of financial development. The problems pointed out by Wood (1994) when including measures of internationally traded physical

capital in HOV-studies, hence do not apply to the endowment of financial intermediaries. It is therefore reasonable to expect countries with well-functioning financial markets to have a comparative advantage in the production of financial services, and to specialize in industries highly dependent on external financing.

2.2. Views of the financial system

Although the above discussion is quite straightforward, it is also abstract. In reality, the financial system is not an entity that develops linearly along a single dimension; rather, there are intrinsic differences between different systems. Naturally, there is also a huge literature on the pros and cons of these different systems. Traditionally, the debate has been focused on bank-based versus market-based financial systems. Recently, however, new perspectives based on the overall efficiency of the financial sector, and its legal environment, have widened the debate.

Hardly anyone doubts that banks play an important role in the modern economy, by specializing in monitoring and screening firms, and by building long-term relationships with firms. Thereby, banks mitigate the informational problems between borrowers and lenders and allow them to provide financing for firms. Competitive equity markets could, however, potentially perform exactly the same functions as by banks, although Stiglitz (1985) argues that information disseminates quickly on well-developed markets. Hence, the individual investor has small incentives to acquire information in the market-based system, whereas the long-term relationships characterizing bank-based systems may mitigate this problem. Schleifer and Vishny (1986) present a similar argument i.e. that the easy by which the individual owner can sell their shares on a well-developed market reduces the incentives to exercise corporate control. All in all, it may well be that banks are better at assessing and controlling firms and managers, and hence at and providing financing.

Several objections have been raised to this negative view of markets. First of all, a well-developed stock market aggregates information about both firms and markets in a way not possible for an individual bank. Even though information spreads fast, there are large and quick gains to be made from acquiring superior information, possibly making markets better informed than banks. Second, corporate control may be

facilitated by stock markets through compensation schemes linked to stock market performance. Further, Hellwig (1991) argues that rent extraction by banks can reduce manager incentives for profitable investments due to their inside information. It is also likely that banks that issue debt have an incentive to be biased against high-risk projects, which can explain why Allen and Gale (2000) find that riskier industries attract more external funding in market-based economies. Another explanation for this might be that a well functioning stock market also expands the possibilities for risk diversification, thereby making high-risk projects more attractive for the individual investor.

Another possibility expressed by, for example, Huybens and Smith (1999) is that markets and banks are complements rather than substitutes. Then, it is the efficiency of the financial sector as a whole that is of importance, not whether the system is primarily based on markets or banks. Finally, as La Porta et al. (2000) have stressed, the legal system is a key determinant of the workings of the financial system. Especially, the legal system protects creditors and minority shareholders against expropriation by majority shareholders and managers. Legal investor protection is therefore associated with effective corporate governance and hence, constitutes a better staring point for cross-country comparisons of financial systems than the bank versus market framework.

Thus, there are four main views of the financial sector: the market based and the bank based views, the view that it is the overall size and efficiency that is of importance, and the view that it is the legal protection of creditors and shareholders that is of importance. In the next section, the different measures employed in this study are presented and related to these views.

3. Measurement issues and data

3.1. Financial dependence

The basic premise of this paper is that there are intrinsic technological reasons why industries differ in their dependence on external financing, and that these differences persist across countries. In the empirical trade literature, such assumptions are quite

standard regarding other production factors such as human and physical capital. It is even standard procedure to assume the inter-industry ranking of intensity in factor usage to be stable over time. Making this assumption for financial dependence might therefore be more empirically than conceptually difficult.

In an innovative paper, Rajan and Zingales (1998) tackle exactly the problem of how to measure industry differences in financial dependence.[3] This is achieved by noting that when financial markets work relatively without friction, the supply of external financing will be very elastic. Differences in the actual use of external financing in such an economy will hence mainly reflect differences in demand for this type of funding. By arguing that the U.S. financial markets are the most advanced in the world, Rajan and Zingales use data on the actual external financing pattern of U.S. firms to calculate their measure of financial dependence. More precisely, their measure is defined as capital expenditures minus cash flow from operations divided by capital expenditures. To smooth fluctuations, they use data on the firm's external financing and capital expenditure over a 10-year period. In order to prevent that excessive weight from being given to large firms, industry values for each of the industries in their study are calculated as medians rather than means. According to this indicator, drug and medicines (ISIC 3522) constitute the most financially dependent industry, while the tobacco industry (ISIC 314) is the least so.

3.2. Financial development

Ideally, a measure of how well developed the financial sector is should gauge how effectively financial intermediaries and markets manage to mobilize and allocate capital. Thus, the ideal measure of financial development should be related to the variety of intermediaries and markets available, the efficiency with which they evaluate and monitor firms, and the legal and regulatory framework assuring performance. Although there are no perfect measures available, the recently developed indicators in Beck et al. (1999) proxy for the different aspects of the financial system outlined in Section 2.2.

[3] Beck and Levine (2001) basically employ the same methodology as Rajan and Zingales, when asking if a bank-based or a market-based financial system is most conductive to the growth of financially dependent industries.

The first couple of proxies are related to the size and activity of the stock market, and are hence related to the market-based view of the financial system. We use the stock market capitalization to GDP ratio (MCAP) which equals the value of listed shares to GDP, as an indicator of the size of the stock market. Second, the total value of stock market trade to GDP (STRADE) is used to proxy for the activity of the stock market. Both these indicators suffer from the potential problem of capturing the forward-looking expectations of the economic agents, however if, for example, high growth and hence, high profits are anticipated, both MCAP and STRADE will increase. Although this could result in severe problems when considering the effect of these variables on growth as in Levine and Zervos (1998), it is not a problem since we here study within country and across industry differences. Another potential problem is that none of these measures reflect the amount of financing actually obtained by firms.

A commonly used proxy for the degree of overall financial development is the liquid liabilities to GDP ratio (LLY). This proxy is usually employed as an indicator of financial depth and has the advantage of being available for a wide range of countries. It is not, however, a direct measure of the financial sector's capacity to generate funds and may be most appropriate when other indicators are not available. A more direct aggregate indicator of the activity of financial intermediaries is the amount of credit given in an economy. More precisely, we use the ratio of private credit by deposit money banks and other financial institutions (DC) to GDP to proxy for this. One virtue of this measure is that it isolates credit issued to the private sector from the private sector. These two indicators are used to investigate the argument that the overall size and efficiency of the financial system is what is of importance for generating capital.

Next, we include indicators of the efficiency and market structure of commercial banks. A potential measure of the efficiency with which commercial banks channel funds from savers to investors is the net interest margin, i.e. the accounting value of a bank's net interest revenue as share of its total assets (MARGIN). This indicator serves as a proxy for the wedge between the prices faced by the parties on either side of a loan transaction. We define (CONC) as the ratio of the three largest banks' assets

to total banking sector assets, as an indicator of the market structure. A highly concentrated banking sector might be less competitive and hence less efficient than a competitive one.[4]

The next set of proxies is more related to regulatory efficiency and hence to the potential of raising funds, rather than the actual outcome. For this purpose, we (again) follow Rajan and Zingales and use the accounting standards for each country in 1990 (ACSTAN). International comparisons of accounting standards are made by the Center for International Financial Analysis and Research. This proxy is supposed to reflect the potential for obtaining financing by reducing information cost. Hence, it can be considered as an overall indicator of the quality of information available to investors. As a check for the consistency of this index, we also make use of the 1983 accounting standards (ACSTAN83).[5] Finally, we turn to indicators of the legal rights of creditors and minority shareholders. MINORITY is an index from zero to six of how well protected minority shareholders are. The higher is the value of this index, the better the legal protection against expropriation. CREDITOR is an index between zero and four, increasing in the legal rights of creditors relative to management and other stakeholders.

3.3. Data on other endowments and intensities

In order to measure the input requirements of human capital, we use the share of workers with post-secondary education in each industry, weighted by the relative size of the respective industry. The average number of years of secondary schooling in the population above 25 is used as a proxy for the national endowments of human capital.

Whether or not to include physical capital in the analysis is an open question. The answer is contingent on the mobility of physical capital; if it is a mobile resource, it should not be included. We choose to follow the convention and include physical capital, especially since we want to ensure that the indicators of financial dependence and endowments do not proxy for any other type of production factors. Physical

[4] As will be discussed later, competition in the banking sector can have both positive and negative effects for the generation of external financing to firms.
[5] The correlation between ACSTAN and ACSTAN83 is 0.70. ACSTAN83 is not available for Mexico.

capital intensities are calculated as the OECD-averaged capital formation to value added ratio, while the physical capital per workers measure capital endowments. [6] In order to capture the effect of natural resource endowments, the stock of agricultural- and forestland per worker is also employed. The intensity of the former is just a dummy for food production, whereas the latter is calculated using Swedish input-output data. For further details on all variables and sources, see the appendix.

3.4. Trade data

Production and trade data by three and four-digit ISIC industry codes for the OECD countries are obtained from the OECD/STAN database. Since we are forced to combine different data sources, our final data set includes data on 32 manufacturing industries in 20 countries. In other words, it must be kept in mind that trade in services and raw materials is not included in this study.

4. The pattern of specialization

We first approach the question of how different countries' factor endowments affect international trade by considering the pattern of industrial specialization. The hypothesis is that the international competitiveness of an industry in a certain country depends on the resource endowments of that country and the input requirements of the industry. Balassa pioneered this approach in a couple of influential papers (Balassa 1979, 1986).

One obvious candidate as an indicator of international competitiveness and industrial specialization is the ratio between production and consumption as suggested by Gustavsson et al (1999),

$$(4.1) \quad r_{ij} = \frac{Q_{ij}}{C_{ij}} = \frac{C_{ij} + X_{ij} - M_{ij}}{C_{ij}} = 1 + \frac{X_{ij} - M_{ij}}{C_{ij}},$$

[6] There are alternative ways of measuring both human- and physical capital intensities and endowments. We have employed several (see appendix) as checks on the robustness of the results.

where Q_{ij} is production, C_{ij} is consumption, M_{ij} is imports, and X_{ij} exports of good i in country j. It should be clear that when r_{ij} is greater than one, country j is a net exporter of good i, whereas a value lower than one indicates that the country is a net importer. In the analysis, r_{ij} is regressed on a set of variables constructed by interacting the input requirements of each industry i with the country characteristics of each country j. The larger the value of r_{ij}, the more specialized is country j in industry i.

In order to pick up fixed industry and country effects, a set of industry and country dummies is added to the regression. We take the logarithm of r_{ij} to ensure that the trade imbalances end up in the country fixed effects. To see this, consider the case of balanced trade. It must then be true that

$$(4.2) \quad \sum_i Q_{ij}^B = \sum_i C_{ij}^B .$$

For each country j there exists a parameter β_j such that

$$(4.3) \quad (1+\beta_j)\sum_i Q_{ij} = \sum_i C_{ij} .$$

By scaling each element in the production vector by $(1+\beta_j)$, a hypothetical value of production under balanced trade is derived. The relationship between the measure of specialization under balanced and unbalanced trade can then be expressed as

$$(4.4) \quad r_{ij}^B = (1+\beta_j)Q_{ij} / C_{ij} = (1+\beta_j)r_{ij} .$$

By taking the logarithms of (4.4), it should be clear that the country-fixed effects capture the trade imbalance parameter $(1+\beta_j)$.

An alternative measure of industry specialization would be the one used by Balassa (1986), namely

$$(4.5) \quad (X_{ij} - M_{ij})/(X_{ij} + M_{ij}).$$

The main difference between this measure and r_{ij} is that it can take on a negative value. Thus, it cannot be adjusted for trade imbalances by taking on logarithms.

Although the approach behind (4.1) and (4.5) is inspired by the HOV-theory, it should not be considered as a formal test of the HOV-theory. Leamer and Levinsohn (1995) raise theoretical objections to this type of studies when the number of goods is larger than the number of production factors. Bowen and Sveikauskas (1992) demonstrate, however, that these theoretical objections are of little practical importance in actual empirical analysis. The patterns of industry specialization are shown to be consistent with the net exports of factor services, especially for broad aggregates of production factors. What is important, though, is to adjust the dependent variable for trade imbalances. For this reason, we will mainly focus on r_{ij}, and keep the Balassa-measure for testing the robustness of the results.[7]

4.1. Estimation and Data

In order to estimate the impact of financial development on the pattern of industry specialization, we use data on industry factor input requirements and country-factor endowments. The expected sign of the interaction variables is usually positive, which means that a country well endowed with a certain factor will specialize in the industries with large input requirements of that factor. Exceptions are when net interest margin and bank industry concentration are used as proxies of efficiency and competition in the banking sector. Since higher values of these variables imply lower efficiency and competition, we expect the interaction between financial dependence and MARGIN and CONC to be negative.

This means that we estimate the following relationship:

$$(4.6) \quad \ln r_{ij} = \sum_{i=1}^{c} \beta_{1i} D_i + \sum_{j=1}^{n} \beta_{2j} D_j + \sum_{k=1}^{m} \beta_{3k} (\alpha_{ik} \times END_{jk}) + \varepsilon_{ij},$$

where i is the industry index, j is the country index, k is the factor index, D_i is a dummy for industry i, D_j is a dummy for country j, α_{ik} is the input requirement of

[7] The correlation between the two indicators of specialization is 0.69.

factor k in sector i, END_{jk} is the endowment of factor k in country j, and ε_{ij} is the error term.

4.2. Results

Table 1 shows the results from the estimation of (4.6). Seven of the ten interactions between financial dependence and financial development are statistically significant and all have the expected signs. Further, all other interaction variables are positive as expected, but the interaction of agricultural inputs is not significant. Given the highly regulated agricultural sector in most OECD-countries, it might not be surprising that natural advantage is not a key determinant of the pattern of agricultural production.

Establishing statistical significance is a first step, but is the effect of financial markets on the pattern of specialization of economic significance? In column one, we see that the coefficient on the interaction term between financial dependence and the market capitalization ratio takes the value of 0.194. In order to interpret the economic magnitude of this coefficient, the following experiment is helpful: Consider that the industry at the 75th percentile of financial dependence was located in the country at the 75th percentile of financial development, rather that in the country at the 25th percentile of financial development. Further, consider the same switch of locations for the industry at the 25th percentile of financial development. How much larger would the industry of high dependence be in the high development country compared to the low-dependence industry, given that all other variables take on their average values?[8] In specification (1), this exercise leads to an increase in $\ln(r_{ij})$ by 0.103. For all industries, the average value of $\ln(r_{ij})$ is –0.164. Hence, the switch of countries would lead to a 10.8 percent increase in r_{ij}, compared to the average value. For the other (statistically significant) proxies of financial development, the same number is 12.2, 5.4, 7.0, 6.1, 8.7, and 8.5 percent. In comparison, the same thought experiment with respect to human and physical capital gives an increase in r_{ij} by around 5 and 6 percent, respectively. The impact of the financial system on the pattern of specialization must thus be considered as very large.

[8] This thought experiment is from Rajan and Zingales (1998). Mathematically, this means the following calculation:
$\text{COEFF} \times \{\text{FINDEP}_{75} \times (\text{FINDEV}_{75}\text{-FINDEV}_{25}) - \text{FINDEP}_{25} \times (\text{FINDEV}_{75}\text{-FINDEV}_{25})\}$

When turning to the specific indicators of financial development, we see that both stock market indicators (MCAP, STRADE) are statistically significant. Moreover, the size-effect of these variables is the largest among all interaction terms. This shows that a well-developed stock market is the most important source of competitive advantage among financially dependent industries. In other words, we have indirect support for the view that a market-based financial system is the most efficient.

Neither of the aggregate indicators of financial-sector development, LLY and DC, seem to be of much importance for the pattern of specialization. The liquid liabilities ratio is not even close to statistical significance, while the credit ratio is weakly significant. The effect of DC is also among the smallest (although still large compared to the effect of human- and physical capital). One way of interpreting this is that the aggregate size of the financial sector is of less importance for raising funds, at least among OECD-countries. Thus, we have indirect evidence suggesting that the type of financial system is of importance.

Turning to the efficiency of the banking sector, the net interest margin (MARGIN) does not affect the pattern of specialization. The concentration index (CONC), which proxies for the degree of competition in the banking sector is, however, of importance. The result for banking concentration is interesting since it indirectly suggests that financially dependent industries have better access to credit when the banking industry is competitive. This contradicts Petersen and Rajan (1995) who show that competition in the credit market can be detrimental to the formation of firm-creditor relationships. The reason is that when creditors cannot hold equity claims, and the market is competitive, the creditor is forced to break even every period. For high-risk projects, this implies a very high interest rate that can distort the firms' incentives. In a monopolistic market, on the other hand, the creditor can cross-subsidy the firm over time – to the mutual benefit of both creditor and lender.[9] Rather, the result in column 5 constitutes indirect support of the view put forward by Rajan

[9] Petersen and Rajan (1995) also provide empirical evidence from the US, supporting this view. Using the same methodology as Rajan and Zingales (1998), Cetorelli and Gambera (2001) find that a concentrated banking sector supports the growth of financially dependent industries.

(1992). There, he suggests banks with market power extract rents and hence, reduce the firms' incentives to invest.

Accounting standards (ACSTAN, ACSTAN83), the indicators of the aggregate quality of information available to investors, are also significant – both statistically and economically. Given the severe informational problems in the financial markets, it should not be a surprise that good information has a positive effect on the generation of external financing.

Finally, the results concerning the view that the legal protection of outsiders against expropriation attempts by insiders are mixed. Minority shareholder protection (MINORITY) does not seem to affect the pattern of specialization, while the protection of creditors (CREDITOR) does.

[Table 1 here]

4.3. Sensitivity analysis

There are many different ways of measuring most variables in of regression 4.6. Since we want to ensure that the results presented above are not due to our choice of indicators, we perform a number of sensitivity tests. Each cell of Table 2 refers to an individual regression, and shows the estimates of the interaction terms between financial dependence and financial development.

In row 1, we replace Ln(r_{ij}) with the Balassa (1986) measure of industry specialization: $(X_{ij} - M_{ij})/(X_{ij} + M_{ij})$. The results are remarkably consistent with the ones in Table 1. Accounting standards for 1990 lose their significance, as does the index of creditor rights. The indicators of financial depth and domestic credit, on the other hand, now gain statistical significance. Making the same analysis for the size of the effect as for ln(r_{ij}), we obtain an increase in the dependent variable by 0.075, 0.090, 0.031, 0.062, 0.063, 0.052, and 0.044 for each of the significant interaction terms. For human capital, the size-effect is around 0.06 and for physical capital 0.025.

Thus, the large effect on the pattern of specialization previously found is not due to the choice of dependent variable.

In row 2, we instrument for financial development using each country's legal origin as instrument, as suggested by La Porta et al. (1998).[10] These instruments are a set of dummy variables taking the value one if a country is of British, Scandinavian, German, and French legal origin, respectively. To this set of instruments, we add the "rule of law" index produced by Business International Corporation. Although these instruments have successfully been used in other studies (e.g. Rajan and Zingales, 1998), we have some worries that there is too little variation in these variables since the analysis is limited to the OECD. However, the results from Table 1 are quite robust to the instrumentation, although the significance levels of the variables are generally somewhat lower. One exception is that CREDITOR gains both in statistical significance and size: the point estimate increases from 0.07 to 0.25. If we were to take this estimate seriously, an increase in the creditor rights' index from 1 to 3 would imply an increase in r_{ij} by 23 percent.

In row 3, the human capital indicator is now replaced by an interaction term where the number of scientists per worker in each country is used as country endowment of human capital.[11] In this specification, all interactions between financial dependence and financial development except the one based on financial depth (LLY) are statistically significant. The point estimates are very similar to the ones in Table 1. If we calculate the size-effect for this indicator of human capital, we get a value of around 6 percent, roughly the same as when secondary schooling is used.[12]

In row 4, physical capital intensities are replaced by the British industry level capital stock to value added ratio, which is done to verify that the results are not contingent

[10] The evolution and persistence of the French and British legal traditions are discussed theoretically in Glaeser and Shleifer (2001).

[11] The correlation between SECSCH and SCIENW is 0.53.

[12] Hanushek and Kimko (2000) measure labor force quality by using international mathematics and science test scores. They can thereby avoid the unrealistic assumption that schooling is of equal quality in different countries. Moreover, the use of test scores reduces the likelihood of proxying for general development effects rather than human capital. Using their indicator (HCQ1) rather than quantity based indicators such as SECSCH and SCIENW does not alter the results in this paper. The size effect of HCQ1 is 4.8 percent (results available upon request).

upon the flow-measure previously used. [13] The size-effect for this indicator is around 2 percent and hence, we can once again verify the results from Table 1.

In row 5, we exclude the US from the regression since the indicator of financial dependence is based on calculations on US firms.[14] This exclusion leaves the results unchanged.

Finally, we include interaction terms between industry intensity and country abundance of electricity and steel. Although these inputs are tradable, and hence should arguably be excluded from the regression, we include them to verify that the results for financial development are not spurious (similar production factors are also included by e.g. Ellison and Glaeser 1999, and Gustavsson et al. 1999). Both new variables are positive and significant but, as can be seen, the basic results are, if anything, strengthened by their inclusion.

[Table 2 here]

5. The factor content of trade

A different approach to the question of how financial development affects the pattern of trade between countries is to consider at the factor content of net trade. Basically, this approach amounts to investigating if the financial system can be a source of comparative advantage. Traditionally, the sources of comparative advantages have been analyzed within the framework of the Heckscher-Ohlin-Vanek (HOV) model. If we consider of financial intermediaries as immobile factor endowments, the prediction of the HOV-model is that a country endowed with well-developed financial markets will be a net exporter of external finance.

[13] The correlation between CVAI and CAPVA is –0.07 (not significant). That the two measures are not correlated is of course naturally a matter of concern. CVAI, however, is highly correlated with electric intensity, sometimes used as a proxy for capital intensity. This leads us to put more trust in the measure.

[14] It is by no means obvious why this should force us to exclude the US from the analysis. Rajan and Zingales do that in their paper, however, so we follow their example.

5.1. Estimation and data

The way to derive an empirical measure of the factor content of net trade, somewhat consistent with the HOV-theory though relaxing the assumption of balanced trade, has been shown in for example Leamer and Levinsohn (1995). In the present paper, we modify the Leamer-Levinsohn measure in the same way as Lundberg and Wikner (1997). More precisely, we calculate the following measure

$$(5.1) \quad Z_{jk} = \sum_i x_{ij} f_{ik} / \sum_i m_{ij} f_{ik} ,$$

where x_{ij} is the share of exports of sector i in country j, m_{ij} is the share of imports of sector i in country j, and f_{ik} is the input-requirement of factor k in sector i. Regardless of the trade balance, the ratio carries information about the relative factor content of exports to imports. Specifically, if $Z_{jk} > 1$ exports are more concentrated to k-intensive goods than imports (Lundberg and Wikner 1997).[15]

We use the Rajan and Zingales (1998) indicator of financial dependence as a proxy for the industry-requirements of external financing, as discussed previously (FINDEP). Likewise, we use the same indicators of financial development as before to proxy for the country endowment of financial intermediation. In this section, as well as before, we hope to be able to discriminate between what aspects of the financial system that are of importance for comparative advantage. This is done by studying at a whole series of indicators. All the measures are thoroughly described in Section 3.

The reason why we do not use exactly the same measure of factor content of net trade as suggested by Leamer and Levinsohn (1995) is that this would require data on world factor endowment of financial intermediation. The meaning of this is conceptually difficult to grasp. Rather than tackling these conceptual difficulties, we use the Z_{jk} of

[15]Equation (5.1) can thus be read as the factor content ratio under the restriction that balanced trade is achieved without a change in the composition of trade. The export (import) expansion needed to get rid of a trade deficit (surplus) is, in other words, assumed to be proportional across goods.

equation (5.1), which is very much in the spirit of the Leamer and Levinsohn measure.[16]

When constructing Z_{jk}, we have not taken the services of production factors in input goods into account. Thus, the net trade of external financing is calculated using only the direct and not the indirect input of services of financial markets.

5.2. Results

Japan has the largest net export of external financing according to definition (5.1). Other countries with high values in the Z_{fd} measure are Germany, Denmark and the U.K. At the bottom of the list, we find countries such as New Zealand, Australia and Greece.

When turning to regression analysis, indicators of human- and physical capital endowments, as well as the endowments of forest- and agricultural land, are included in the regressions. The reason is twofold. First, an industry's dependence on external financing may be a proxy for its human or physical capital intensity, for example, while a country's endowment of financial intermediation may be a proxy for its endowment in these production factors. To ensure that the results are not only an artifact of spurious correlation we control for the endowment of human and physical capital. Second, the exchange of external financing embodied in trade in services and raw material is not included, since the data covers manufacturing only. This may give a distorted picture of the factor content of trade for countries where raw material or services account for a large proportion of trade. Suppose that the external financing requirement in a sector not included in the data is very high (low). Then, the endowment figures will overstate (understate) the supply of external financing available for the manufacturing industry in countries where this particular sector is large. For this reason, it is necessary to include measures of the endowment of other production factors.

[16] The Leamer and Levinsohn measure, under the balanced trade restriction, would take the form: $\sigma_{jk} = \sum_i X_{ij}(\sum_i e_{ij} f_{ik} - \sum_i m_{ij} f_{ik}) / V_{jk} = 1 - s_j / a_{jk}$, where X_{ij} is the export of good i from country j, V_{jk} is country j's endowment of factor k and a_{jk} is country j's share of world endowments of factor k.

Table 3 presents results where the measure of factor content of net trade (Z_{fd}) is regressed on country endowment variables. There is definitely support for the hypothesis that the financial sector is a source of comparative advantage. The measures of stock market size (MCAP) and activity (STRADE) both positively enter the regression. The same is true for the proxy for the liquidity, or financial depth measure, of the financial sector (LLY), and the competition indicator of the banking sector (CONC). However, DC and the other proxy regarding the functioning of the bank sector (MARGIN) are not statistically significant. Moreover, there is no positive effect of a country's accounting standards or its legal framework on the net factor trade of external financing.

Thus, the effects of the endowment of financial intermediaries on a country's pattern of specialization and comparative advantage in trade are roughly the same. Notably, the size and activity of the stock market and the concentration of the banking system have a significant effect on both variables.

Admittedly, it is not easy to judge the economic effects of being endowed with well-developed financial intermediaries, since we deal with proxies of what we attempt to measure. To investigate the effect in the different proxies for financial intermediation, imagine an increase of, for example, STRADE by one standard deviation. This induces an increase in Z_{fd} by 23 percent from the mean. The impact of the other significant proxies is around 20 percent, or around 60 percent of one standard deviation. Another way of assessing the effect of the financial service endowment is to see what happens if it is excluded. Column 1 reveals that removing the proxy for financial intermediation reduces the adjusted R^2 from about 0.45 to 0.03. Thus, the statistically significant proxies have a remarkable effect on the fit of the regression.

[Table 3 here]

5.3. Sensitivity

As checks for robustness, we have replaced SECSCH with the test-based labor force quality indicator (HCQ1, results are not presented), but this has no effect on the

results presented in Table 3. We also include other control variables in the regressions. First, it may be the case that the indicators of financial development capture some aspect of economic development not accounted for by the other endowment variables. We therefore include GDP per capita, but the results remain unchanged. Second, the public sector is likely to be financed in other ways than through the private financial markets. Thus, for a country with a large public sector, the true endowment of financial intermediation available for private manufacturing may be larger than in a country with smaller public employment. Including the share of public employment does not, however, affect the results.

Finally, we run all specifications in Table 3 on an alternative measure of factor content of net trade. This measure is constructed as the ratio of factor content in net trade, corrected for trade imbalance, to factor content in consumption. Specifically,

$$(5.2) \quad W_{jk} = \frac{\sum_i f_{ik}(X_{ij} - M_{ij} - B_j \frac{Q_{iw}}{GDP_w})}{\sum_i f_{ik} C_{ij}},$$

where f_{ik} is the input-requirement of factor k in sector i, X_{ij} the exports of sector i in country j, M_{ij} the imports of sector i in country j, B_j country j's the trade imbalance and C_{ij} country j's consumption of good i. Q_{iw}/GDP_w is the share of world output of good i in world GDP. Once more this has little effect on the results presented in Table 3. All results remain qualitatively the same, except in the specification including LLY, where LLY is no longer statistically significant on conventional levels.

So far, the results show strong support for the hypothesis that the financial system can give rise to comparative advantage. Specifically, countries endowed with large and active financial intermediaries are more likely to have a larger net export of the services provided by the financial sector. To obtain a greater understanding of the impact, we derive a similar measure for net trade of human capital. The purpose of this paper is neither to test the validity of the HOV-model, nor to study the effect of a country's human capital endowment on the trade pattern. Nevertheless, it may be

enlightening to look at the equivalent measure of comparative advantage for human capital, mainly for two reasons. First, it might indicate whether the proxies used for human capital intensity and human capital endowment are valid. Second, it may make us more comfortable with our measure of net factor trade. Keep in mind, however, that the HOV-model has found limited support in the empirical trade literature.

A common procedure when studying the empirical support for the HOV-model is to conduct rank and sign tests (Leamer and Levinsohn 1995). According to the HOV-model, a country's ranking in net trade of a specific factor should correspond to its ranking in terms of endowment. We use the measure of net trade in factors as defined in equation 5.1. Table 4 shows the Kendall's rank test for the two production factors financial intermediation and human capital. We use the share of workers with post-secondary education, weighted by the relative size of that industry, as input requirements of human capital.

The results for the different proxies for financial intermediation are in line with the regression results presented in Table 3. All correlations, except for accounting standards, carry the expected sign and four are statistically significant. The correlations between our measure of the net trade of services of human capital and endowments of human capital are also positive (and significant in two cases out of three). In light of the fact that empirical research finds limited support for the HOV-model, we consider these results to be satisfactory.

[Table 4 here]

6. Conclusions

The main finding of this paper is that countries with well-functioning financial systems tend to specialize in industries highly dependent on external financing. Although this might not be surprising in itself, the size of the effect is. In fact, differences in financial systems are more important determinants of the pattern of specialization between OECD-countries than differences in human or physical capital. One plausible explanation for this phenomenon is that the differences in human- and physical capital within the OECD are fairly small. Hence, the relative size of the

effect might be smaller in a wider selection of countries. Whether this is true is a question left for future research. We also show that the financial system gives rise to comparative advantage in way consistent with the Heckscher-Ohlin-Vanek model.

Another way of studying this paper is as a robustness test of the Rajan and Zingales (1998) result that financially dependent industries grow faster in countries with well-developed financial markets. However, we approach this question by studying levels rather than growth rates. Given that Rajan and Zingales find strong signs of conditional convergence among industries (initially large industries tend to grow slower than initially small industries), it is by no means obvious that their result should carry over from growth rates to levels.

Especially strong results are found for indicators of stock market size and activity, as well as for competition in the banking sector. The latter result thus gives support to theories suggesting that banking concentration limit the amount of capital raised by firms (e.g. Rajan 1992).[17] The quality of a country's accounting standards and the legal protection given to creditors are also important determinants of the pattern of specialization. Financial depth and the aggregate amount of credit in an economy give rise to comparative advantage, but the results for the pattern of specialization for these indicators are mixed.

Since this is one of the first papers approaching the question at hand, we have aimed at simplicity and clarity in the empirical analysis. One extension of this study would be to allow for other amendments common in the empirical HOV-analysis. These amendments might be to allow for cross-country technological and demand differences as suggested by Trefler (1993), Davis and Weinstein (1996), and Harrigan (1997). Another interesting extension is to analyze the potential effects of the financial market on the choice of technology. Since financial markets are supposed (and shown) to solve information problems in the market place, they are likely to affect the choice of technology. Carlin and Mayer (1999) take a first step in this direction by showing that the financial system affects R&D. To get a better understanding of its effect on technology, we would also need a better grasp of why

[17] The finding that banking sector concentration is conductive for the growth of financially dependent industries reported in Cetorelli and Gambera (2001) also hinges on conditional convergence.

some industries are more dependent on external financing than others are. This might be a fruitful area for future research with implications for the literature on growth as well as international trade. Incorporating other institutional factors is another extension along these lines. As long as industries differ in their use of the services provided by these institutions (and the services are non-tradable), we would expect the pattern of specialization to be determined by institutional factors.[18]

Finally, there are at least two reasons for expecting the pattern in this paper to disappear over time. First, multinational corporations are supposedly insensitive to local financing conditions. To the extent that MNC:s continue to increase their share of international trade, local financial markets should exert a continuously smaller impact on the pattern of trade. The same applies if financial markets effectively become better integrated across countries over time.[19] Hence, extending the analysis along the time dimension might be fruitful.

[18] Naturally, measuring input requirements of institutional factors, and institutional quality is difficult. That wage-setting institutions compressing the wage distribution can affect the industrial composition is supported by evidence in Davis and Henrekson (2000).

[19] Petersen and Rajan (2001) document an increase in the physical distance between small firms and lenders in the US over time. This increase is correlated with higher bank productivity and hence, constitutes evidence that financial development reduces the need for proximity between borrowers and lenders.

Appendix

Table A1. Correlations and summary statistics: Endowments

	Mcap	Strade	Lly	Dc	Conc	Margin	Acstan	Acstan 83	Credi-tor	Mino-rity	Secsch	Scienw	HCQI	Kapwl	Agrilw	Rwood w	Elecw	Steelw
Mcap	1																	
Strade	0.80***	1																
Lly	0.72***	0.68***	1															
Dc	0.74***	0.75***	0.72***	1														
Conc	-0.36*	-0.50**	-0.47**	-0.18	1													
Margin	-0.46**	-0.50**	-0.43**	-0.68***	-0.11	1												
Acstan	0.39*	0.16	0.00	0.43*	0.24	-0.37*	1											
Acstan83	0.48**	0.35	0.01	0.50**	-0.02	-0.49**	0.70***	1										
Creditor	0.27	0.33	0.20	0.20	0.03	-0.36	0.08	-0.01	1									
Minority	0.49**	0.36	0.31	0.51**	-0.06	-0.25	0.45**	0.29	0.16	1								
Secsch	0.15	0.38*	0.17	0.45**	-0.11	-0.46**	0.31	0.31	0.30	0.45**	1							
Scienw	0.62***	0.81***	0.38**	0.71***	-0.23	-0.66***	0.42*	0.67***	0.30	0.34	0.53***	1						
HCQI	0.42*	0.53**	0.54***	0.47**	-0.19	-0.61***	0.03	0.05	0.61***	0.07	0.33	0.57***	1					
Kapwl	0.28	0.33	0.21	0.54**	-0.05	-0.68***	0.63***	0.56**	0.11	0.43*	0.63***	0.70***	0.34	1				
Agrilw	-0.02	-0.13	-0.12	-0.07	0.15	0.00	0.27	0.03	-0.39*	0.40*	0.11	-0.07	-0.45**	0.23	1			
Rwoodw	-0.15	-0.28	-0.20	0.07	0.49**	-0.24	0.50**	0.24	0-.16	0.28	0.27	0.00	-0.18	0.40*	0.20	1		
Elecw	0.01	-0.04	-0.17	0.20	0.38*	-0.26	0.59***	0.40*	-0.04	0.41*	0.37*	0.33	-0.06	0.63***	0.19	0.60***	1	
Steelw	-0.08	-0.12	-0.07	0.02	0.27	-0.30	0.25	0.12	-0.12	-0.09	-0.08	0.05	0.01	0.31	-0.03	0.59***	0.05	1
Mean	0.38	0.16	0.67	0.83	0.61	0.029	66	65	1.86	2.29	2.7	4.9	50,2	44.3	1.0	3752	18.7	1581
Stdev	0.27	0.15	0.30	0.42	0.23	0.013	10	11	1.06	1.42	1.1	2.4	5,8	12.5	1.4	4725	12.1	3292
75th perc.	0.48	0.21	0.75	1.09	0.46	0.018	74	73	3	4	3.1	6.5	54,2	53.0	-	-	-	-
25th perc.	0.15	0.04	0.50	0.49	0.87	0.044	61	61	1	1	1.9	2.7	44,6	39.9	-	-	-	-

*** indicates significance at the 1%-level, ** at the 5%-level, and * at the 10%-level. The last two rows display the values of the observations belonging to the 75th and the 25th percentile respectively.

Table A2. Correlations and summary statistics: Intensities

	Findep	Ahi	Cvai	Capva	Wood-int	Land-use	Elintl	Ironint
Findep	1							
Ahi	0.36**	1						
Cvai	0.14	0.10	1					
Capva	0.05	-0.13	-0.07	1				
Woodint	-0.06	-0.09	0.13	-0.11	1			
Landuse	-0.09	0.31*	-0.03	-0.13	-0.04	1		
Elintl	-0.19	-0.22	0.54***	-0.02	0.21	-0.05	1	
Ironint	-0.11	-0.06	0.20	-0.02	0.04	-0.03	0.14	1
Mean	0.34	0.07	0.15	0.41	1.7	0.03	185	0.03
Stdev	0.42	0.06	0.07	0.25	6.8	0.18	285	0.18
75th perc.	0.47	0.09	0.20	0.42	-	-	-	-
25th perc.	0.06	0.02	0.11	0.26	-	-	-	-

*** indicates significance at the 1%-level, ** at the 5%-level, and * at the 10%-level. The last two rows display the values of the observations belonging to the 75th and the 25th percentile respectively.

Table A3. Correlations and summary statistics: Dependent variables

	Mean	Stdev	Correlations	
Ln(r_{ij})	-0.164	0.482	1	-
Balassa	-0.210	0.431	0.69***	1
Z_{fd}	0.84	0.29	1	-
W_{fd}	-0.12	0.19	0.74***	1
Z_{hc}	0.98	0.52	-	-

*** indicates significance at the 1%-level, ** at the 5%-level, and * at the 10%-level.

Table A4. Industries and countries included

ISIC	*Sectors*	*Countries*
3110	Food	Australia
3130	Beverages	Austria
3140	Tobacco	Belgium
3210	Wearing Apparel	Canada
3220	Textiles, Apparel &	Denmark
3230	Leather & Products	Finland
3240	Footwear	France
3310	Wood Products	Germany
3320	Furnitures & Fixtures	Greece
3410	Paper & Products	Italy
3420	Printing & Publishing	Japan
3520	Other Chemicals	Mexico[1]
3522	Drugs & Medicines	Netherlands
3530	Petroleum Refineries	New Zealand
3540	Petroleum & Coal Pr	Norway
3550	Rubber Products	Portugal
3560	Plastic Products, n	Spain
3610	Pottery, China etc	Sweden
3620	Glass & Products	UK
3690	Non-Metallic Products	USA
3710	Iron & Steel	
3720	Non-Ferrous Metals	
3810	Metal Products	
3820	Non-Electrical Mach	
3825	Office & Computing	
3830	Electrical Machinery	
3832	Radio, TV & Communication	
3840	Transport Equipment	
3841	Shipbuilding & Repair	
3843	Motor Vehicles	
3850	Professional Goods	
3900	Other Manufacturing	

[1] Accounting standards from 1983 are missing for Mexico.

Data description

Dependent variables

$r_{ij} = Q_{ij}/C_{ij} = Q_{ij}/(Q_{ij} + M_{ij} - X_{ij})$, where Q_{ij} is the production, C_{ij} is the consumption, M_{ij} are the imports from the rest of the world, X_{ij} are the exports to the rest of the world, of industry i in country j. Average values 1989-91. Source: STAN. The definition is from Gustavsson et al. (1999).

Balassa = $(X_{ij} - M_{ij}) / (X_{ij} + M_{ij})$. Source: STAN.

$Z_{ik} = \sum_i x_{ij} f_{ik} / \sum_i m_{ij} f_{ik}$, where x_{ij} and m_{ij} are the shares of exports (imports) of sector i from (to) country j, and f_{ik} is the input requirement of factor k in sector i. The measure is called Z_{fd} when using external financing requirements and Z_{hc} when using human capital requirements. Source: x_{ij} and m_i are constructed from STAN average values in 1989-91, f_{ik} for external financing is FINDEP (Rajan and Zingales (1998)) and human capital intensity is AHI. The definition is from Lundberg and Wikner (1997).

$$W_{jk} = \frac{\sum_i f_{ik}(X_{ij} - M_{ij} - B_j \frac{Q_{iw}}{GDP_w})}{\sum_i f_{ik} C_{ij}}$$

, where f_{ik} is the input-requirement of factor k in sector i, X_{ij} the exports of sector i in country j, M_{ij} the imports of sector i in country j, B_j is country j's trade imbalance and C_{ij} country j's consumption of good i. Q_{iw}/GDP_w is the share of world output of good i in world GDP. The measure is called W_{fd} when using external finance requirements. Source: X_{ij}, M_{ij} and C_{ij} from STAN average values 1989-91. f_{ik} for external finance is FINDEP. Q_{iw} the sum of production over the 22 countries included in the study and GDP_w is the sum of GDP for the countries in the study. B_j = total export-total import of goods and services for 1990 from World Development Indicators.

Financial variables

External finance dependence:

FINDEP. Capital expenditure minus cash flows from operations divided by capital expenditures. Data source: Rajan and Zingales (1998).

Financial sector development:

MCAP: Stock market capitalization to GDP, average 1989-91. Source: Beck et al (1999).

STRADE: Stock market total value traded to GDP, average 1989-91. Source: Beck et al (1999).

LLY: Liquid liabilities to GDP, average 1989-91. Source: Beck et al (1999).

DC: Private credit by deposit money banks and other financial institutions to GDP, average 1989-91. Source: Beck et al (1999).

CONC: Market share of the three largest banks. Source: Beck et al (1999).

MARGIN: Net interest margin to total assets. Source: Beck et al (1999).

ACSTAN: Accounting standards 1990. Source: Rajan and Zingales (1998).

ACSTAN83: Accounting standards 1983. Source: Rajan and Zingales (1998).

Legal variables:

MINORITY: Index of minority share holder rights, range 0-6. Source: La Porta et.al. (1998).

CREDITOR: Index of creditor rights. Range 0-4. Source: La Porta et.al. (1998).

RULELAW: *International Country Risk (ICR)* index of law and order tradition. Source: La Porta et.al. (1998).

SCAND, GERMAN, FRENCH, ENGL: Dummies of legal origin. Source: La Porta et.al. (1998).

Human capital

Human capital intensities:

POSTSEC: The share of post-secondary schooling in total employment, 1990, Swedish industries. Source: SCB Regional Labor Statistics, unpublished.

AHI: $\sum_j \{(EMPLOYMENT_{ij} / WORKERS_j) \times POSTSEC_i\} /$ *number of countries*. Source: EMPLOYMENT from STAN average 1989-91, WORKERS from Penn World Tables 5.6.

Human capital endowment:

SECSCH: Average years of secondary schooling in the population over 25. Average 1985-90. Source: Barro and Lee (2000).

SCIENW: Number of scientists and engineers per worker. Year 1990 or the closest available (1988-93). Source: United Nations Statistical Yearbook.

HCQ1: Indicator of labor force quality, based on international mathematics and science test scores. Based on fixed world average test score. Source: Hanushek and Kimko (2000)

Physical capital

Physical capital intensities:

CVAI = $\sum_j capitalformation_{ij} / \sum_j value\ added_{ij}$, Average 1989-91. Source: STAN.

CAPVA = Capital stock/ Value added in UK. Average 1993-95. Source: OECD Statistical Compendium, Industry, Science and Technology, Industrial Structure Statistics – Industrial Surveys.

Physical capital endowment:

KAPW1: KSTOCK/WORKERS. Capital per worker, thousands of dollars. Average 1988-90. Source: KSTOCK: Real net capital stock in millions of US dollars. This is the accumulated, depreciated, and deflated series (15 years, 13.33% depreciation rate) of gross fixed capital formation in each country. Investment deflators were taken from Summers and Heston. Average 1988-90. From the Factor endowments database, (FEDB) compiled by Maskus and Poterba.

Natural resources

Natural resource intensities:

LANDUSE: Agricultural land intensities: Dummy for food production (ISIC 311/2).

WOODINT: Definition: Millions of SEK worth of input of forestry products divided by millions of SEK worth of production (times 100). Source: SCB (1992).

IRONINT: Use of iron ore. Dummy for iron & steel production (ISIC 3710).

ELINT1: Definition: Total amount of purchased electrical energy in megawatt hours divided by total number of thousands of hours worked. Average value 1990/1991. Source: SCB Industristatistik.

Natural resources endowment:

FORLANDW: FORLAND/WORKERS. FORLAND: Area of forests and woodland measured in thousands of hectares. Source: The Production Yearbook of the FAO. Average 1988-90. FEDB.

RWOODW: RWOOD/WORKERS. RWOOD: Round wood production, cubic meters. Average 1989-91. Source: United Nations Statistical Yearbook.

AGRILW: AGRILAND/WORKERS. AGRILAND: Area of arable land and land under permanent crops or permanent pasture in thousands of hectares. Source: The Production Yearbook of the FAO. Average 1988-90. FEDB.

ELECW: ELEC/WORKERS. ELEC: Indigenous production of electricity (Gwh). Average 1989-91. Source: OECD Basic Energy Statistics, various issues.

STEELW: STEEL/WORKERS. STEEL: Crude steel and pig iron production in metric tons. Average 1989-91. Source: United Nations Statistical Yearbook.

General country factors

GDPPC: GDP per capita. Average 1988-90. Source: Penn World Tables 5.6.

GDPPW: GDP per worker. Average 1988-90. Source: Penn World Tables 5.6.

POP: Population in thousands. Average 1988-90. Source: Penn World Tables 5.6.

WORKERS: Workforce in thousands. Average 1988-90. Source: Own calculations GDPPC*POP/GDPPW.

GDP: Total GDP. Average 1988-90. Source: Own calculations GDPPC*POP.

TOTEXP: Total manufacturing export value in dollars. Average 1989-91. Source: STAN.

TOTIMP: Total manufacturing import value in dollars. Average 1989-91. Source: STAN.

GOVSH: Government share of employment. Defined as government employment/WORKERS. Average value 1989-91. Source: OECD Economic Outlook.

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Table 1. Dependent variable is Ln(r_{ij})

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
FINDEP×Ln(MCAP)	0.218***									
	(3.12)									
FINDEP×Ln(STRADE)		0.165***								
		(2.86)								
FINDEP× Ln(LLY)			-0.002							
			(-0.020)							
FINDEP×Ln(DC)				0.161*						
				(1.89)						
FINDEP×CONC					-0.407**					
					(-2.40)					
FINDEP×MARGIN						-5.592				
						(-1.46)				
FINDEP×ACSTAN							0.012***			
							(2.53)			
FINDEP×ACSTAN83								0.017***		
								(3.39)		
FINDEP×MINORITY									0.014	
									(0.51)	
FINDEP×CREDITOR										0.066**
										(2.00)
AHI×Ln(SECSCH)	1.505***	1.344***	1.782***	1.469***	1.744***	1.552***	1.494***	1.441***	1.729***	1.650***
	(3.15)	(2.73)	(3.82)	(2.93)	(3.76)	(3.13)	(3.17)	(2.50)	(3.69)	(3.54)
CVAI×Ln(KAPW1)	2.223***	2.241***	2.375***	2.231***	2.366***	2.266***	2.186***	3.319***	2.356***	2.350***
	(2.73)	(2.76)	(2.89)	(2.72)	(2.90)	(2.77)	(2.68)	(2.88)	(2.87)	(2.87)
LANDUSE×Ln(AGRIW)	0.024	0.022	0.031	0.026	0.027	0.026	0.035	0.027	0.032	0.024
	(0.44)	(0.40)	(0.53)	(0.46)	(0.50)	(0.45)	(0.61)	(0.47)	(0.55)	(0.42)
FOREST×Ln(RWOODW)	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***
	(8.22)	(8.26)	(8.12)	(8.27)	(8.35)	(8.24)	(8.33)	(8.40)	(8.31)	(8.25)
% increase in r_{ij}[1]	10.8	12.2	-	5.4	7.0	-	6.1	8.7	-	8.5
ADJ R^2	0.300	0.300	0.281	0.289	0.290	0.286	0.295	0.330	0.282	0.285
# OBS.	619	619	619	619	619	619	619	587	619	619

Robust t-values in parenthesis. *** indicates significance at the 1%-level, ** at the 5%-level, and * at the 10%-level. Regressions include industry and country fixed effects. [1] The interpretation of this value is given in the text.

Table 2. Sensitivity analysis

	FD2 *MCAP*	*FD3* *STRADE*	*FD4* *LLY*	*FD5* *DC*	*FD7* *CONC*	*FD8* *MARGIN*	*FD1* *ACSTAN*	*FD6* *ACSTAN83*	*FD12* *MINORITY*	*FD13* *CREDITOR*
1) Balassa	0.159***	0.129***	0.184**	0.191***	-0.379**	-4.67*	0.005	0.009***	-0.004	0.036
	(2.73)	(2.82)	(1.90)	(3.13)	(-2.24)	(-1.82)	(1.48)	(3.23)	(-0.15)	(1.28)
2) IV	0.155	0.112*	0.278	0.283**	0.480	-9.780*	0.009*	0.019**	0.044	0.246***
	(1.24)	(1.67)	(1.43)	(1.94)	(1.46)	(-1.76)	(1.64)	(1.90)	(1.21)	(2.48)
3) Scientists per	0.209***	0.164***	-0.026	0.154*	-0.396**	-5.447*	0.012***	0.017***	0.014	0.067**
worker	(2.83)	(2.80)	(-0.33)	(1.73)	(-2.33)	(-1.38)	(2.47)	(3.12)	(0.51)	(2.00)
4) British capital	0.224***	0.167***	-0.007	0.170**	-0.407**	-6.038*	0.012***	0.018***	0.016	0.067**
intensities	(3.26)	(2.90)	(0.95)	(2.05)	(-2.41)	(-1.60)	(2.74)	(3.60)	(0.59)	(2.04)
5) US excluded	0.212***	0.162***	-0.014	0.146*	-0.391**	-5.919	0.011***	0.017***	-0.002	0.075**
	(2.97)	(2.67)	(-0.18)	(1.68)	(-2.12)	(-1.49)	(2.47)	(3.34)	(-0.05)	(2.16)
6) Electricity and steel	0.235***	0.172***	0.014	0.194**	-0.370**	-7.119*	0.014***	0.019***	0.023	0.068**
added	(3.36)	(2.97)	(0.19)	(2.27)	(-2.16)	(-1.85)	(3.04)	(3.64)	(0.82)	(2.07)

Robust t-values in parenthesis. *** indicates significance at the 1%-level, ** at the 5%-level, and * at the 10%-level. Each cell refers to an individual regression and shows the point estimate of the interaction between financial dependence and financial development. All regressions include indicators of human- and physical capital, agricultural- and forestland, as well as industry and country fixed effects. In row 1, the dependent variable is $(X_{ij} - M_{ij}) / (X_{ij} + M_{ij})$, in rows 2-6 it is $\ln(r_{ij})$. In row 2, we instrument indicators of financial development with the "rule of law"-index and dummies of legal origin. In row 3, the number of scientists per worker (SCIENW), rather than secondary schooling (SECSCH), is used to measure human capital endowments. In row 4, British capital intensities (CAPVA), than the capital formation to value added ratio (CVAI), are used to measure capital intensities. In row 5, the US is excluded from the regressions. In row 6, ELINTI×ln(ELECW) and IRONINT×ln(STEELW) are added to the regressions.

Table 3. Factor content of net trade (defined as in equation 5.1 Z_{fd}). Basic regressions.

	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}	Z_{fd}
MCAP		0.620**								
		(2.59)								
STRADE			1.278***							
			(3.08)							
LLY				0.578***						
				(3.47)						
DC					0.352					
					(1.40)					
MARGIN						-4.091				
						(-0.57)				
CONC							-0.681**			
							(-2.22)			
ACSTAN								-0.859		
×1000								(-0.17)		
CREDITOR									-0.014	
									(-0.29)	
MINORITY										-0.002
										(-0.04)
KAPW1	0.005	-0.000	0.001	-0.000	-0.002	0.002	0.003	0.005	0.005	0.005
	(0.74)	(-0.02)	(0.27)	(-0.07)	(-0.39)	(0.36)	(0.75)	(0.62)	(0.72)	(0.75)
SECSCH	0.030	0.039	-0.023	0.034	0.013	0.028	0.015	0.029	0.035	0.031
	(0.73)	(1.05)	(-0.46)	(0.91)	(0.25)	(0.60)	(0.34)	(0.62)	(0.84)	(0.562)
FORLANDW	-12.032	-9.894	-2.735	-6.353	-5.719	-13.206	-0.346	-11.749	-12.313	-11.979
×1000	(-1.19)	(-1.18)	(-0.36)	(-0.96)	(-0.52)	(-1.13)	(-0.05)	(-1.20)	(-1.15)	(-1.17)
AGRILW	-0.050	-0.046	-0.054	-0.039	-0.043	-0.040	-0.065	-0.050	-0.053	-0.049
	(-1.20)	(-1.24)	(-1.63)	(-1.19)	(-1.13)	(-0.92)	(-1.53)	(-1.18)	(-1.41)	(-0.87)
Constant	0.667***	0.614***	0.751***	0.453*	0.695***	0.908*	1.185***	0.705***	0.684***	0.668***
	(4.58)	(4.37)	(6.72)	(1.88)	(4.86)	(2.14)	(5.53)	(3.70)	(4.80)	(4.45)
Adj R^2	0.037	0.448	0.467	0.463	0.221	-0.005	0.381	-0.031	-0.024	-0.032
N.obs	20	20	20	20	20	20	20	20	20	20

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level. t-statistics based on robust standard errors in parentheses.

Table 4. Rank test between the factor content of net trade (Z_{jk} defined as in equation 5.1) and endowment

Production factor	*Proxy*	*Kendalls rank test*
Financial intermediation	STRADE	0.305*
	MCAP	0.221
	DC	0.286*
	LLY	0.324**
	CONC	-0.484***
	MARGIN	-0.074
	ACSTAN	-0.058
	MINORITY	-0.116
	CREDITOR	0.226
Human capital	SECSCH	0.238
	HCQ1	0.408***
	SCIENW	0.295*

*** Indicate significance at 1%-level, ** at 5%-level, * at 10%-level

CHAPTER 3

WHO WANTS POLITICAL INTEGRATION?
EVIDENCE FROM THE SWEDISH EU-MEMBERSHIP REFERENDUM*

Jonas Vlachos†

Abstract

The political determinants of regional integration and separation have been analyzed in much theoretical, but little empirical work. From a local perspective, the Swedish 1994 EU-membership referendum can be described as the choice between two different unions – Sweden and the EU. In this particular case, membership implies less discretion to handle income shocks. Local industry diversity is shown to have been an important determinant of the voting pattern. Further, high-income regions were more prone to vote in favor of membership. A reasonable interpretation is that safe and rich regions voted in favor of secession from Sweden, rather than integration with Europe.

JEL Classification: F15, H11, H70
Keywords: Regional integration; Secession; Risk sharing; Redistribution; Spatial econometrics.

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† Department of Economics, Stockholm School of Economics, P.O Box 6501, 113 83 Stockholm, Sweden. Phone: +46-8-7369649, Fax: +46-8-313207, Email: Jonas.Vlachos@hhs.se.

1. Introduction

During the last decade, the political economy of regional integration and disintegration has been under theoretical scrutiny (for surveys, see Alesina et al. 1995, and Bolton et al. 1996). Although there is a huge empirical literature on various aspects of regional integration (e.g Hess and van Wincoop, 2000), little empirical work has directly analyzed political attitudes toward the integration process. In this paper, the regional voting pattern in the Swedish 1994 EU-membership referendum is used to gain insights on the regional characteristics determining political preferences regarding integration.

Although enhanced possibilities of inter-regional risk sharing constitute one of the main themes in the literature discussing the possible gains from regional integration (e.g. Persson and Tabellini 1996ab), the European Union has chosen a constitution limiting the national discretionary powers to handle economic shocks in several dimensions.[1] This constitution is the starting point of the simple framework guiding the empirical analysis of the paper: by joining the EU, national governments renounce the freedom to deal with local income volatility.[2] Add to this that migration between regions is costly. A lower degree of local industry diversity hence implies higher costs of joining the EU. Consistent with this hypothesis, it is found that individuals in risky locations – in a standard portfolio sense – were significantly more negative towards EU-membership than those in safe ones. It is also shown that inhabitants of high-income labor markets, with a high level of schooling and small receipts of central government transfers were more positive towards EU-membership. Taken together, these results suggest that inhabitants of safe and rich regions voted in favor of

[1] Although an interesting issue in itself, no attempt is made to explain why the EU is less willing to allow for insurance-type transfers than national governments. One possibility is that the distribution of industry aid is more difficult to supervise at a more aggregate level, another that there is less solidarity between people living in different countries. Both effects suggest that moral hazard problems would be more severe within the EU than in individual nations, making it rational to limit the degree of discretion when creating a union. Persson and Tabellini (1996b) discuss federal risk sharing under different constitutions when subject to local government moral hazard.

[2] Since 1994, national support to manufacturing industries in the Euro-12 area has continuously declined from around 40 to around 28 billion Euro (EU 2000). This suggests that the attempts by the EU to limit government intervention at the industry level are (at least to some extent) successful. The funds handed out at the EU-level are mainly directed to agriculture and regions suffering from permanent economic backwardness. Further, EU rules on monetary policy, fiscal policy, competition, and public procurement all restrict the degree of national discretion.

secession from the Swedish transfer system, rather than in favor of European integration.

There is relatively little empirical work with which to contrast the results, and what there is, is based on cross-country data. Panizza (1999) shows that country size, per capita income, ethnic fractionalization, and democracy are all negatively related to the degree of fiscal federalism (defined as the percentage of revenues or expenditure of the central government out of total revenues or expenditure).[3] Alesina et al. (2000) focus on the relation between political and economic integration. They find that openness to trade mediates the economic benefits of country size, and that open trade regimes tend to generate secessions. This paper has two main advantages compared to previous studies. First, it looks directly at voter preferences on the decision to integrate (or secede) and second, by exploiting regional data, it circumvents some of the well-known problems with cross-country studies. More observations are available, problems of data comparability are much less severe and, perhaps most important, the variation between unobservables (or hard-to-observables), such as institutional quality, is usually lower within a country than between countries.

Whereas most of the theoretical literature tends to identify, implicitly or explicitly, a "region" with a "nation", the present paper suggests that a local labor market perspective can be fruitful. This is due to the fact that the local labor market is usually a place of residence from which exit is costly, both for social and economic reasons. Since tax revenue and hence, public services, are partly determined locally, it is obvious that every inhabitant is affected by whatever happens to his/her local market. Likewise, housing values are largely driven by local demand characteristics. From the viewpoint of the individual, what is usually portrayed as the choice between unification and separation of regions – as well as fiscal centralization or decentralization – could therefore well be described as the choice between different *types* of unification and centralization.

[3] These indicators are subject to several problems, especially when exploiting cross-national data. (i) Different levels of local governments should be weighted in different ways. (ii) Local governments may collect taxes without having autonomy in deciding the tax amount, or how to allocate the expenditures. (iii) Intergovernmental grants are not captured in the data.

This paper is organized as follows: In section 2, the theoretical argument guiding this study is presented. In section 3, the data is described, with special focus on the construction of an index of labor market risk, and some results are shown. Section 4 describes the problem that mightd be encountered when using spatial data and the estimates when taking these problems into consideration are presented. Section 5 concludes.

2. The theoretical argument

Although economic integration could foster market-based risk sharing between regions (see e.g. Asdrubali et al. 1996, and Mélitz and Zumer 1999), this paper highlights the political aspects of risk reduction. Entering the European Union involves renouncing national discretion along several dimensions. The Maastricht treaty entails the complete loss of monetary autonomy and caps on government borrowing during temporary shocks. Further, joining the EU would also mean joining a customs union, leaving Sweden with no discretionary trade policy towards third countries. Finally, the inner market builds on strict rules regarding industrial subsidies, competition policy and public procurement, leaving national governments with little discretion in times of industrial crises.[4] That the attempts to limit the interventions by national governments in the economy have been successful can be verified by the 30 percent decline in national state aid to industries since 1994 in the Euro-12 area (EU 2000). Since Sweden joined the EU, Swedish employment aid has declined by almost 50 percent and regional aid by more than 20 percent (NUTEK 2000).[5]

Based on these stylized facts, the binary choice between two different types of unions is studied. Specifically, one alternative is to vote in favor of membership in the

[4] An example related to Sweden is the policy to differentiate electricity tax rates depending on industry electricity consumption. The policy, which has been used to shield electricity dependent industry from price shocks, is now under heavy attack from the European Commission since it is claimed to distort competition. The EU-rules regulating state aid are summarized in EU (1999).

[5] As a share of total value added in industry, state aid has been reduced by about 27 percent from 1995 to 1999 according to the NUTEK (2000) study. Calculating the full extent of different subsidies is a complicated task and the results largely depend on which definitions have been used. In Sweden, one estimate of the value of industrial support in 1994 was around 50 billion Swedish kronor (about 5 billion Euro). The same year, corporate income taxation aggregated to 20 billion kronor (2 billion

European Union – with the loss of national powers this implies. The other alternative is to vote for continued autonomy.[6] Within the EU, the national government is assumed to have no discretionary power to compensate for industry-specific shocks, whereas there are no restrictions on these policies when being a non-member. An obvious objection to this extreme setup is that the European Union itself transfers resources between regions and industries. Although this is true, these transfers are mainly directed to industries and regions with permanent, rather than temporary, problems. Moreover, except in the case of agriculture, the transferable funds available at the EU-level are minute compared to the national budgets. Thus, it is fair to say that the EU prevents national governments from using many types of insurance mechanisms, while at the same time offering little replacement.[7]

Mathematically, the argument can be described in the following way. In a local market, c, labor is employed in sectors s=1,2,...n. Every period, each sector is hit by a normally distributed productivity shock, $\varepsilon_s \sim (0, \sigma_s^2)$. The share of labor in each city-sector is $s_{sc}=L_{sc}/L_c$, where L_{sc} is the number of employed in sector s in labor market c, and L_c is total employment in this market. Leaving a local labor market is assumed to be prohibitively expensive,[8] and there are no opportunities for private insurance across individuals or regions. Each individual gets an equal share of the total revenue to all industries in a local market. If Sweden joins the EU, the revenue of each industry changes by m_s and the national government is not allowed to insure against industry shocks. This means that total revenue in c in case of membership is:

$$(2.1) \quad R_c = \sum_{s=1}^{n} s_{sc}(R_s + m_s + \varepsilon_s).$$

Euro) (Barkman and Fölster 1995). To the best of my knowledge, no estimates employing the same methodology have been produced at a later date.

[6] Naturally, it can be questioned how real the Swedish autonomy would be outside the EU. The EES-agreement regulating trade between the EU and Sweden did imply that many of the EU regulations would automatically apply to Sweden, being members or not. That autonomy would be greater outside of the EU is quite obvious, however.

[7] This point is also made by Sorensen and Yosha (1998) who show that only a small fraction of shocks to EU-countries are smoothed by international transfers. Likewise, private markets (still) offer limited insurance. Given the importance of the national government in consumption and income smoothing, they argue that the strict fiscal discipline imposed by the Maastricht treaty should be relaxed.

[8] This assumption is less restrictive than what appears at a first glance. One could easily imagine that households own an immobile asset (say a house) that has a value proportional to the aggregate income of a region. The basic argument would hold even for this case.

The expected value of joining the EU is then given by

$$(2.2) \quad E\{R_c^y\} = \sum_{s=1}^{n} s_{sc}(R_s + m_s) = R_c + m_c,$$

and the total variance of returns is

$$(2.3) \quad Var(R_c^y) = \sum_{s=1}^{n} \sum_{t=1}^{n} s_{sc} s_{tc} \sigma_{st},$$

where σ_{st} is the variance/covariance of shocks. The inhabitants of a local market do not only care about this monetary revenue, but also have preferences over the election outcome for other reasons. If Sweden joins the EU, an individual i in labor market c receives utility $u_i(R_c^y) + y_i$, where y_i captures individual preferences over the membership. Agents are risk averse and hence, the utility function satisfies: $u'(R_c) > 0$, and $u''(R_c) < 0$. Since the shocks are assumed to be normally distributed, a two-term Taylor's series of the expected utility can be approximated by

$$(2.4) \quad u_i(R_c^y) + 1/2\, u_i''(R_c^y) Var(R_c^y) + y_i.$$

It is obvious that the expected utility decreases with the volatility of income. If Sweden does not join the EU, the national government has full discretion to deal with the industry-specific shocks. Assuming that the government is maximizing total utility, this means that the marginal utility between cities will be equalized in every period. In this extreme setup, full insurance against idiosyncratic shocks, as well as full redistribution, is, in other words, provided. Total revenue will be equalized across locations:

$$(2.5) \quad R_c = R = \sum_{s=1}^{n} s_s(R_s + \varepsilon_s - t_s),$$

where t_s is the tax (subsidy) paid (received) by the respective industry. The volatility of income is equalized across cities to aggregate national volatility:

$$(2.6) \quad Var(R) = \sum_{s=1}^{n} \sum_{t=1}^{n} s_s s_t \sigma_{st} .$$

Given that an individual receives utility n_i if Sweden stays out, the utility of each individual can be approximated by:

$$(2.7) \quad u_i(R) + 1/2\, u_i''(R) Var(R) + n_i.$$

A voter will cast his or her vote in favor of membership whenever (2.4) is greater than (2.7):

$$(2.8) \quad u_i(R_c^y) - u_i(R) + 1/2 \{u_i''(R_c^y) Var(R_c^y) - u_i''(R) Var(R)\} \geq n_i - y_i.$$

It should be clear that those who stand the most to gain from joining the EU are the inhabitants of relatively rich and stable regions, since they are no longer forced to participate in the transfer schemes towards poor and unstable regions. Hence, a "yes"-vote could best be described as a vote in favor of secession from Sweden, rather than a vote in favor of membership itself.

At a first glance, this result appears to contrast with the results in Persson and Tabellini (1996a), and Bolton and Roland (1997). In these models, the risky and poor regions have greater incentives to join the larger union in order to exploit the greater opportunities of risk sharing and redistribution within the larger economic entity. The difference is superficial, however. Joining the EU is here claimed to imply *leaving* a system of regional transfers, whereas unification in the above papers implies the *creation* of a regional transfer scheme. Had the EU and Sweden shared the same policies regarding redistribution and risk sharing, the results in this paper would have been reversed.

3. Empirics

Most of the data in this paper has been collected from *Statistics Sweden*, although the unemployment statistics are from *AMS,* the central body of the Swedish labor market authorities. Except when otherwise indicated, data refers to the election year, i.e.

1994. The units of analysis are the 109 local labor markets (LLM) that exist in Sweden. An LLM is an analytical (and not administrative) unit created by *Statistics Sweden* based upon the observed commuting pattern.[9] Since most variables are not available at the LLM-level, data has been aggregated using information from Sweden's 288 municipalities. Although the LLM:s are constructed in accordance with commuting patterns, it is possible that the conditions on one LLM affect choices made in a neighboring LLM. In other words, there is a possibility that LLM:s cannot be treated as independent observation units. In Section 4, I will describe - and attempt to correct for – the problems these spatial aspects induce to the econometric analysis.

3.1. Measuring Labor Market Risk

To construct the indicator of local labor market risk, the portfolio-theoretic approach used by Conroy (1975) and Diamond and Simon (1990) is followed. They derive what can best be described as an augmented Herfindahl index. This concept is borrowed from the literature on financial economics: each local market is viewed as a portfolio of industries, each industry is subject to random fluctuations. The total degree of labor market risk of a labor market is the weighted average of the variances and covariances of these fluctuations. The total variance in city c equals

$$(3.1) \quad \mathrm{RISK}_c = \sum_s^n s_{sc}^2 \sigma_{ss}^2 + \sum_s^n \sum_u^n s_{sc} s_{uc} \sigma_{su} \quad , s \neq u, c = 1,...,109$$

where s_{sc} is the employment share of sector s in labor market c, σ_{ss} is the variance of employment in sector s, and σ_{su} the covariance of employment in industries s and u.[10] Specifically, the variance/covariance matrix is constructed as follows: The variance element, σ_{ss}, is the variance in the rate of national employment growth in sector s, while the covariance elements, σ_{su}, equal the covariance between growth rates. To capture the degree of unexpected volatility, the employment growth rate of each industry was de-trended by taking the first differences of employment on the national level. The logarithms of the first differences were then regressed on a time trend. For

[9] See *Statistics Sweden* (1998) for a closer description of the LLM subdivisions.
[10] In the model, volatility is due to revenue fluctuations, rather than employment fluctuations. Revenue, however, is likely to fluctuate much less than employment, due to the public unemployment insurance.

each regression, the residuals were used to construct the variance/covariance matrix.[11] In order to calculate the variance/covariance matrix, sectoral employment data at the national level for 43 industries between 1985-1997 was used.[12] This amounts to the assumption that industries are subject to the same fluctuations, regardless of their geographical location. Obviously, this assumption is not a perfect description of reality, but it is necessary in order to capture both the within and between industry fluctuations in a systematic manner.[13]

In Table 1, the ten least and most risky markets are listed. It is obvious that markets in the north of Sweden had a large representation of unstable industries, which is in line with popular beliefs on this issue. It is not surprising that Stockholm, the largest labor market by far, finds its place among the least risky ones. Some of the labor markets are less obvious members of the low risk group. It should be kept in mind, however, that the measure of risk is derived using *fluctuations* around a trend. Local markets with a large share of steadily shrinking (or expanding) industries will be regarded as "safe" in this study. This does not imply that the labor market situation is good in any other sense of the word. Hällefors (the eighth safest market), for example, had an unemployment rate of 18 percent at the time of the referendum.[14]

At this point, it is appropriate to ask why local unemployment fluctuations are not used rather than the derived measure of industry risk. There are two main reasons for this. First, the theoretical argument is that industries, rather than geographical areas, are hit by shocks. It is possible that safe local markets (in the portfolio sense) actually experienced large employment fluctuations in the relevant time period. This does not mean, however, that rational agents should update their beliefs about the future riskiness of their local market. An analogy is that *ex ante* safe financial portfolios may very well exhibit larger fluctuations, *ex post,* than risky ones. Second, even though moving is prohibitively expensive in the model, it is not so in the real world. If people

[11] It can be noted that under the very strong assumptions of independently and identically distributed growth rates, the portfolio diversification index is equivalent to the well known Herfindahl index.
[12] This period is the longest possible for which industry data is consistent at the municipal level. The sectoral division is based on *Statistics Sweden's* "limited level". Industries are described in the appendix.
[13] Clark (1998) shows that industry shocks constitute an important part of the total variation in regional output.
[14] This said, note that the correlation between RISK and unemployment is 0.38.

do move in response to local employment shocks, the unemployment statistics will misrepresent true labor market risk.

3.2. Empirical Specification

The dependent variable is the percentage of votes in favor of EU-membership, labeled YES. In addition to the RISK-variable, INCOME is added to the regression. Given the argument in Section 2, we would expect richer regions to be more prone to enter the EU (or rather, leave the Swedish transfer system).[15] Several control variables are also used. Average years of schooling, SCHOOL, are added to the regressions since the benefits of the EU-membership might possibly benefit the well educated more than the less educated. Investment in general human capital makes individuals less sensitive to industry-specific shocks, making the loss of national discretion less important for this group. In most studies discussing the referendum, a clear north-south voting pattern is described, but not explained (see e.g. Gilljam 1996). In this paper, the north-south dimension is accounted for by creating a "trend" variable numbering the local markets, depending on the distance to Copenhagen, i.e. the nearest major EU-city. This ordering is then squared to take into account that labor markets increase in geographical size (and hence the distance to EU is increased) when moving north in Sweden.[16] Alesina and Spolaore (1997) give one possible interpretation of the trend variable. In their model, agents living far from the center stand the most to lose from political integration. Although this interpretation might be too literal, the expected sign of the variable TREND is negative.[17]

Since the nationwide Swedish political parties played a major role in the campaign prior to the referendum, it is likely that voters were influenced by the opinion of "their" party. In order to account for this effect, I use the percentage share of votes in the 1991 municipal election for the Social Democratic party, SAP, and the

[15] INCOME is defined as average income per working family. Alternative measures would be disposable income per family or per capita income, but the choice of income measure does not affect the results to any significant degree.

[16] Using the squared ordering markedly improves the fit of the regressions compared to using a simple south-north ordering of labor markets.

[17] One possibility is that one of the major advantages of the EU membership is the increased possibility of cross-border trade (especially in highly taxed goods such as alcohol). Naturally, the benefits from cross-border trade decreases (more than proportionately) with the distance from the border.

Conservative party, MODERAT, respectively.[18] The reason for using voting data from the 1991 rather than the 1994 election is that the 1991 election did not concern the EU-question. This way, the underlying political preferences are captured, and we rule out the possibility that the voters' attitudes towards the EU affected the voting pattern in the county election rather than the other way around. Since both the Social Democratic and the Conservative parties were strongly in favor of membership in the European Union, the expected signs on both SAP and MODERAT are positive.

The voter turnout ratio, TURNOUT, defined as the share of voters casting their votes in the referendum, is also included in the regressions. Although it is difficult to have a clear theoretical prior on the sign of this variable, it is possible that the willingness to vote varied systematically with the attitudes towards EU-membership. In US elections, a high turnout implies that a larger share of poor people with low education vote. In the present context, this type of effect would lead to a negative sign.[19] However, if the hypothesis of this paper is correct, then it is likely that the willingness to vote also varies systematically with LLM-risk. In other words, some of the effect that should be attributed to local market risk will be captured by TURNOUT. To account for this problem, TURNOUT will be instrumented for by using the turnout ratio of the 1991 and 1994 parliamentary elections.[20]

The referendum has commonly been described as large vs. small cities. For this reason, population size is included in the regressions. The sector most strongly affected by EU membership is probably agriculture. Due to the huge agricultural subsidies handed out by the EU, we would expect the share of agricultural employees, AGRI, to be positively related to the share of yes votes.

What is not yet controlled for is the expected sectoral impact of the EU membership (other than agriculture for which there is a clear prior). In other words, how do we control for the permanent change in sectoral revenue due to the membership (i.e. m_c of equation 2.2)? One possibility would be to consider actual changes in aggregate

[18] An attempt is also made to use voting shares of other parties, but these results are not significant.
[19] The correlation between TURNOUT and INCOME is 0.50, and between TURNOUT and SCHOOL 0.35.
[20] The correlation between the turnout ratio of the referendum and the 1991 and 1994 elections turnout ratios is 0.65 and 0.61, respectively.

sectoral employment since 1994 which would mean that all changes in employment that have occurred since 1994 are explained by the EU membership. Although this is obviously incorrect, this variable could give us some indication of the magnitude of m_c. Formally, we control for

$$(3.2) \quad m_c = \sum_{2=1}^{n} s_{sc} \Delta \overline{L_s} \text{ where,}$$

$$(3.3) \quad \Delta \overline{L_s} = \left[\frac{(L_{s,1995} + L_{s,1996} + L_{s,1997})/3 - L_{s,1994})}{L_{s,1994}} \right].$$

$L_{s,t}$ is the national employment in sector s at time t. This variable is labeled AVGCH9497. All in all, we have the following baseline equation to estimate[21]

$$\begin{aligned} (3.4) \quad \text{YES}_c = {} & \text{CONST}_c + \beta_1 \text{RISK}_c + \beta_2 \text{RISK}_c^2 + \beta_3 \text{INCOME}_c \\ & + \beta_4 \text{SCHOOL}_c + \beta_5 \text{TREND}_c + \beta_6 \text{MODERAT}_c + \beta_7 \text{SAP}_c \\ & + \beta_8 \text{TURNOUT}_c + \beta_9 \text{POP}_c + \beta_{10} \text{AGRI}_c + \beta_{11} \text{AVGCH9497} + \varepsilon_c. \end{aligned}$$

3.3. Results

Several interesting results are revealed just by inspecting the correlation matrix in the appendix. First of all, the correlation between YES and RISK is –0.46. A scatter plot of the variables is shown in Figure 1. This strong correlation is an indication that increased labor market risk did indeed contribute to making people less prone to vote in favor of membership in the EU. Second, the correlation between RISK and MODERAT is –0.39, while the correlation between RISK and SAP is 0.29 which suggests that political sympathies vary systematically with local market risk: the larger the risk, the more government intervention is asked for. Finally, the idea that recipients of large central government grants worried about the future of these transfers is given some support by the strong negative (-0.82) correlation with the share of yes-votes.

[Figure 1 about here]

In Table 2, column 1, the baseline estimate of equation (3.4) is presented. The fit is very good, with an adjusted R^2 of 0.86. As can be seen, all significant variables have the expected sign. TREND and TURNOUT are negative, suggesting that the further to the north, and the larger the share of people who used their right to vote, the more negative people in the respective labor market were towards the EU. A higher average income, more average schooling, a larger share of workers in the agricultural sector, and a larger share of Social Democratic or conservative supporters all seem to result in a more positive attitude towards the EU. Neither population size, nor the measure of expected change of employment, is significant (AVGCH9497 even has the incorrect sign). The main variables of interest, RISK and $RISK^2$, also have the expected (and statistically significant) signs, thereby giving support to the hypothesis that living and working in a riskier labor market did indeed make people less prone to vote in favor of EU-membership. All regressions are estimated using White's heteroskedasticty consistent standard errors. A Jarque-Bera test of normality of this first specification (which has a χ^2 distribution with two degrees of freedom under the null hypothesis of normally distributed errors) has a value of 1.60, which amounts to a probability value of 0.45. Thus, the null hypothesis cannot be rejected.

In column 2, the turnout ratios from the 1991 and 1994 parliamentary elections are used as instruments for TURNOUT in a standard 2SLS-regression. This is done to account for the possibility that TURNOUT itself is affected by labor market risk. Despite the high correlations between the instruments and TURNOUT (0.65 and 0.60), and the fairly good fit of the first stage regression (0.57), the point estimate of TURNOUT drops sharply and loses its statistical significance. The point estimate of RISK is, on the other hand, almost doubled. The size of the point estimates and the statistical significance of the political variables MODERAT and SAP also increase. In columns 3 and 4, the insignificant variables are dropped from the specification but the results remain stable.

[21] The term $RISK^2$ is entered since the fit of the regressions is found to be better using this specification. Actually, there are no reasons to expect the relationship to be linear.

Although the normality test indicates that outliers do not drive the results, I experiment by excluding labor markets with more than 50 percent yes-votes. As can be seen in column 5 (only 2SLS estimates are reported for the sake of brevity), this exclusion – if anything – strengthens the results for RISK. AGRI, MODERAT, and SAP lose some of their significance in this specification. Since OLS-estimates can be sensitive to outliers, column 6 reports the results from a median regression (that estimates the median of the dependent variable). The estimates are similar to the previous ones. Upon close inspection of the explanatory variables, it can be found that four labor markets display such a large labor market risk so large that they can be considered as outliers in the box-plot sense of the word. In an unreported set of regressions, these labor markets are excluded from the sample. This exclusion results in an increase in the point estimates of RISK (-0.37 and -0.64 in the OLS- and 2SLS specifications, respectively), but otherwise the results are stable.

One possible problem with the model specification is that the dependent variable is bounded between zero and one. The problem is likely to be of minor importance, however, since the minimum value of YES is 17.7 and the maximum is 65.3. The standard procedure of dealing with bounded dependent variables converts it to the form Log(YES/100-YES). This conversion does not affect the results to any substantial degree, either statistically or economically (the point estimates for RISK and $RISK^2$ are both significant at the 5-percent level). In the last column of Table 2, Log(YES/100-YES) is used as the dependent variable while all regressors are entered in logarithms. Now, the instrumented point estimate for TURNOUT is even positive (although not significant). Moreover, the SAP coefficient is not significant.

Establishing statistical significance is a first step; the next question is whether the effect is economically significantt. The point estimate of RISK is around -0.27 and that of $RISK^2$ 0.0034 when using OLS, and around –0.42 and 0.0053 respectively when using 2SLS. Since the standard deviation of RISK is 13.3, an increase in RISK by one standard deviation implies a decrease in YES by 3 percentage points when considering the OLS-estimates and 4.6 percentage points when looking at the 2SLS-estimates. Here, it is also interesting to consider the first-stage regression (not shown). In this, the coefficient on RISK is 0.11 and on $RISK^2$ -0.0014 (both are significant at the 5%-level). An increase in RISK by one standard deviation would then imply an

increase in TURNOUT by 1.19. Since the coefficient on TURNOUT is –1.31 in the OLS-specification, an increase in TURNOUT by 1.19 implies a decrease in YES by 1.56 percentage points. Thus, the total effect of a one standard deviation increase in RISK is around -4.6 percentage points, in the 2SLS-specification as well as in the OLS-specification.[22] Given that the mean of YES across local labor markets is 41.8, the effect is indeed substantial.

These calculations raise a potentially serious concern. In the sample, RISK ranges from around 14 to 72. The coefficients on RISK and $RISK^2$ indicate that the point at which the quadratic function turns from negative to positive is 39.8 (OLS-specification). This would mean that among the LLM:s with above average risk, increased risk is associated with a more positive attitude towards membership. Although not implausible, this result is highly dependent on the precision with which the coefficient of $RISK^2$ is estimated. By moving towards the lower bound of the 95-percent confidence interval of the point estimate, the turning point moves to 3608. Moving to the upper bound results in a turning point at 20. Thus, even minor errors in the estimate easily push the turning point far outside the relevant interval. Since the logarithmic specification produces highly significant negative point estimates, it is reasonable to expect the turning point to actually be above 72.

For the other statistically significant variables, increasing the independent variables by one standard deviation gives the following impact on the share of yes-votes (columns 3 and 4, respectively). SCHOOL: 4 and 3.4 percentage points. INCOME: 3.6 and 2.9 percentage points. AGRI: 1.8 and 2.0 percentage points. MODERAT: 0.8 and 1.3 percentage points. SAP: 1.5 and 1.9 percentage points. TREND: 7.2 percentage points. Altogether, the results presented in this section point in the direction suggested by theory. The results concerning the political variables do not appear to be very robust, however.

In Table 3, some further robustness tests are performed in order to exclude the possibility that the results are driven by an omitted variable (only 2SLS-estimates are shown). In column 1, the average age of the population in each labor market is

[22] The first-stage regression is obviously not specified to capture all variation in TURNOUT. However,

entered. Column 2 contains per capita welfare spending and column 3 the unemployment rate. Column 4 shows the results when population density and column 5 when the share of public employment is entered, respectively. These inclusions add very little to our understanding of the voting pattern.[23] In column 6, however, per capita grants from the central government to municipalities are added to the regression. The coefficient of this variable is negative and highly significant, which is in line with the hypothesis that those receiving large redistributive grants are less willing to join the EU. The point estimates of RISK, $RISK^2$, and INCOME are somewhat reduced, but still statistically significant in this specification.[24]

4. The Spatial Dimension

As mentioned above, some special econometric problems might arise when using geographical data of the type used in this study. Basically, the problems are due to the fact that the units of analysis, LLM:s, are not completely closed off from each other. Not only do people commute across imaginary borders; income, schooling, etc. are more similar between nearby regions than between distant ones. If this is the case, the attitudes towards EU-membership can "spill over" between labor markets. Using a different terminology, we have a problem of spatial autocorrelation. Formally, this type of spillovers would result in a systematic correlation between the dependent variable across regions, i.e. the so-called *space lag* case

$$y_c = \gamma \Omega y_c + \beta x_c + \varepsilon_c , \tag{4.1}$$

where Ω is a weight matrix of dimension $c \times c$, where c is the number of geographical units used in the study. γ is the coefficient of Ω in the regression. The elements in the weight matrix take the value of one if two regions are neighbors and zero otherwise. If

since the fit of the regression is fairly good (adj.R^2=0.51), these calculations are quite informative.

[23] Ringdal and Valen (1998) find that population density in the area of residence is an important determinant of the individual voting pattern. In the present data set, the correlation between population density and RISK is –0.32. In a specification including both variables, RISK is significant while population density is not. Thus, the fact that diversified labor markets also tend to be densely populated might be what drives Ringdal and Valen's results.

[24] When all six new control variables are entered simultaneously, the point estimates and significance levels of RISK, $RISK^2$, and INCOME are roughly the same as in column six.

4.1 describes the true model, then corr[y_i , y_j] ≠ 0 which would yield biased estimates using OLS.

Another possibility is that the true model is described by equation 4.2 – the so-called *space error* case. In this case, the error terms are systematically correlated across regions, but it is not the voting pattern per se that is behind the correlation. Formally, this takes the form:

(4.2) $$y_c = \beta x_c + \lambda \Omega \varepsilon_c + \eta_c$$

In this case, corr[ε_i , ε_j] ≠ 0 but η is iid. and the OLS-estimates will be inconsistent. Just as the model 4.1, a weight matrix is necessary to properly estimate 4.2. Since the autoregressive coefficient λ is not known, a full maximum likelihood estimation must be carried out.

The first issue at hand when taking the spatial structure into consideration is the construction of the weight matrix Ω. Given the nature of the problem, different specifications can be called for. In this paper, the simplest possible weight matrix, consisting solely of zeroes and ones, will be used. Ω will then be a 109×109 matrix, where the elements take the value one if two LLM:s are neighbors, and zero if they are not. Given the size of each LLM, it is not likely that any significant amount of movement takes place between LLM:s that are not neighbors. Each row in the matrix is normalized so that it sums to one, which means that the assumed impact of each labor market on its neighbors is decreasing in the number of neighbors of this labor market.

The baseline specification, without the insignificant variables, is reproduced in Table 4, column 1 for ease of comparison. The only difference from Table 2 is that the Akaike information criterion (AIC) rather than the adjusted R^2 is used. Since the spatial regressions are estimated using maximum likelihood, a more general measure of fit than the adjusted R^2 must be used. The regression is re-estimated using Ω to check for the different types of spatial autocorrelation. First, when testing for a spatial lag structure, the Lagrange Multiplier test suggested by Anselin (1988) is used. The

test yields a χ^2-distributed test statistic equal to 14.407 (one degree of freedom) with a probability value of 0.00, which means that the null hypothesis of no spatial lags can be firmly rejected. Second, the LM-test for a spatial error structure suggested by Burridge (1980) is performed. The test statistic is χ^2 -distributed with one degree of freedom. This test statistic has the value 13.33, with a probability value of 0.00. Hence, the null hypothesis of no spatial errors is also rejected. Based on this information alone, we cannot conclude whether the spatial lag or the spatial error model is the correct one to use. That spatial correlation should be taken into account is obvious, however.

Now, this regression is re-estimated assuming a spatial lag structure (i.e. equation 4.1). In other words, ΩYES is entered on the right-hand side of the equation and γ (GAMMA in Table 4) is estimated. γ is highly significant and the drop in the AIC, from 637 to 625, suggests that the spatially lagged variable captures a considerable part of the previously unexplained variation in YES. The point estimate of 0.24 implies that an increase in the number of yes-votes in a neighboring labor market by one percentage point increases the number of yes-votes by 0.24 percentage points. When inspecting the coefficients and significance levels of the other explanatory variables, some interesting changes can be found. First of all, the significance levels of RISK and $RISK^2$ are somewhat reduced. The point estimates are stable, however. Second, the point estimate and the significance level of MODERAT are both dramatically reduced. Thus, the effect of this variable seems to have been an artifact of the spatial structure of the data. Finally, the LM-test for spatial errors indicates that some spatial structure in the error terms is still present. In column 3, the results from a spatially lagged IV-estimation are shown. Two main differences can be found: first the instrumented variable TURNOUT is now (weakly) significant with a point estimate of –0.62 and second, the test for spatial errors is now barely significant.

Next, the spatial error structure is taken into consideration by re-estimating the regression specified according to equation 4.2. The variable LAMBDA is the λ of equation 4.2. As can be seen, this variable is highly significant, so obviously the weight matrix captures some of the variation previously unaccounted for, or attributed to other variables. The AIC now takes a value of 618, i.e. even lower than in the

spatial lag case. When conducting the LM-test for a spatial lag structure on this specification, the test is not significant.[25] Thus, the spatial error specification seems to be the one best describing the data. Some further tests are performed on this model: a spatial Breusch-Pagan test indicates that heteroskedasticity is not a problem,[26] and a likelihood ratio test shows that it was indeed appropriate to correct for spatial errors.[27]

Compared to the regression in column 1, some interesting differences can be found. First of all, RISK and $RISK^2$ are stable to the inclusion, although some of their significance is lost (just as in the space lag case). INCOME and SCHOOL are also robust towards the change in the specification. The party variables MODERAT and SAP, however, lose all their significance and the point estimates are sharply reduced. The point estimates of the rest of the variables are somewhat changed, but the basic relation between them and YES is robust. All in all, the results in this section show that the *space error* model is the one best describing the data. The coefficient on LAMBDA of around 0.6 tells us that when the expected number of YES votes is exceeded by one percentage point in one LLM, then YES will also increase by 0.6 percentage points in the neighboring LLM.

5. Conclusions and discussion

The aim of this paper has been to give some empirical input to the largely theoretical debate on the political economy of regional integration and secession. Using the regional voting pattern from the Swedish 1994 EU-membership referendum, it is found that regions with a well diversified industrial structure, high average income- and educational levels, close to the EU-center (i.e. in the south of Sweden), and a large agricultural sector were more positive to EU-membership than other regions. Further, the inhabitants of municipalities receiving large central government grants were less willing to join the EU.

Superficially, these results appear to contradict both Persson and Tabellini (1996a) and Bolton and Roland (1997). The former show that regions subject to large

[25] H_0: No spatial lag dependence. 1 degree of freedom; value 0.96; probability 0.33.
[26] Spatial B-P: H_0: No heteroskedastiticy. 9 degrees of freedom; value 7.86; probability 0.55.
[27] LR-test: 1 degree of freedom; value 44.76; probability 0.00.

idiosyncratic shocks have stronger incentives to join a union than stable regions. The latter present a model where poor regions are more prone to integrate since they stand more to gain from inter-regional transfer schemes. Based on the strict regulations regarding discretionary policies within the European Union, however, a more reasonable interpretation is available. From the individuals' viewpoint, joining the EU could well be described as leaving a system of transfers at the national level, whereas unification in the above-mentioned models implies the *creation* of an inter-regional transfer scheme. The inhabitants of stable and rich regions receiving little central government assistance could, in other words, be said to have voted for secession from Sweden rather than for integration with the EU. This interpretation highlights the importance of defining what "regions" and "unions" actually are, especially in models based on individual political choice. Further, it demonstrates that it is the actual institutional arrangements that are of importance for political choice – which is exactly the point made by Persson and Tabellini, as well as Bolton and Roland. The suggested interpretation offers an interesting potential explanation of why the political leadership was so much more positive towards EU-membership than the population as a whole.[28] From a national – rather than regional – perspective, greater risk sharing between nations could very well be encouraged by membership, even though this comes at the expense of less risk sharing between regions. Once again, what is safe and what is risky all depends on the perspective

Fernandez and Rodrik (1991) offer an alternative explanation to (some of) the empirical results in this paper. They show that uncertainty of gains and losses from an economic reform creates a bias towards the status quo. Joining the EU can be seen as undertaking an encompassing economic and political reform. To the extent that the uncertainty about gainers and losers depends on the insurance characteristics of the local industry structure, we would expect safe regions to be more prone to vote in favor of membership.

The basic relation between local industry diversity and voting behavior is interesting for several reasons. That the local market can act as an insurance mechanism adds a new dimension both to future empirical investigations in the field of political

[28] More than 80 percent of the members of parliament, compared to 52 percent of the electorate, were

economy, as well as to policy analysis. Second, political preferences can be expected to be conditional upon labor market diversity. The basic pattern of correlations indicates that people living on diversified markets vote more "right-wing" than people on less diversified ones. This suggests some degree of substitutability between insurance on the market and social safety nets.[29] Another open question is the generality of the results. Obviously, the cost of moving can differ both between and within countries. The effect of local conditions on voting behavior can be expected to decrease if relocation costs are reduced. Another possible extension of this work would be to decompose the sources of local labor-market risk into their national, regional, and industry-specific components. Since the optimal local policy response depends on the sources of risk, the different components can then be related to the referendum results.

in favor of the membership.

[29] In Vlachos (2001) it is shown that LLM-risk has a positive impact on municipal public employment in Sweden.

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Table A1: Descriptive statistics:

	YES	*RISK*	*TREND*	*SCH-OOL*	*IN-COME*	*POP*	*TURN-OUT*	*MODER AT*	*SAP*	*AGRI*	*PUB-LIC*	*POP-DENS*	*GRANT*	*UEMP*
Mean	41.79	31.81	4015	10.66	156.02	80884	82.59	16.18	41.53	0.02	0.35	30.09	7339	13.07
Media	43.31	29.14	3025	10.62	156.70	27642	82.96	16.78	41.02	0.02	0.35	16.19	6455	12.86
Max	65.28	72.83	11881	11.70	184.50	1.8×10^6	86.90	32.79	60.54	0.09	0.56	169.32	17638	20.69
Min	17.70	14.24	1	10.16	138.50	3424	74.30	3.73	23.92	0.00	0.14	0.28	-1001	4.99
Std. Dev.	10.50	13.36	3589.02	0.32	8.83	199226	2.08	5.73	7.87	0.02	0.07	36.29	3738	2.80

Table A2: Correlation matrix:[1]

	YES	*RISK*	*TREND*	*SCH-OOL*	*IN-COME*	*POP*	*TURN-OUT*	*MOD-ERAT*	*SAP*	*AGRI*	*PUB-LIC*	*POP-DENS*	*UEMP*	*AGE*
RISK	-0.46													
	(0.00)													
TREND	-0.72	0.67												
	(0.00)	(0.00)												
SCH-	0.46	0.04	0.02											
OOL	(0.00)	(0.71)	(0.80)											
IN-	0.63	-0.13	-0.34	0.52										
COME	(0.00)	(0.17)	(0.00)	(0.00)										
POP	0.68	-0.26	-0.35	0.80	0.61									
	(0.00)	(0.01)	(0.00)	(0.00)	(0.00)									
TURN-	0.25	-0.18	-0.34	0.35	0.50	0.38								
OUT	(0.01)	(0.07)	(0.00)	(0.00)	(0.00)	(0.00)								
MOD-	0.68	-0.39	-0.59	0.41	0.46	0.60	0.29							
ERAT	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)							
SAP	-0.38	0.29	0.59	-0.07	-0.23	-0.22	-0.31	-0.56						
	(0.00)	(0.00)	(0.00)	(0.47)	(0.02)	(0.02)	(0.00)	(0.00)						
AGRI	0.16	-0.04	-0.25	-0.09	-0.17	-0.02	-0.01	0.14	-0.39					
	(0.08)	(0.65)	(0.01)	(0.35)	(0.07)	(0.84)	(0.92)	(0.14)	(0.00)					
PUB-	-0.26	0.40	0.50	0.37	-0.31	0.05	-0.19	-0.12	0.26	0.04				
LIC	(0.01)	(0.00)	(0.00)	(0.00)	(0.00)	(0.63)	(0.04)	(0.23)	(0.00)	(0.69)				
POP-	0.63	-0.32	-0.47	0.46	0.42	0.63	0.21	0.54	-0.24	0.09	-0.05			
DENS	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.03)	(0.00)	(0.01)	(0.34)	(0.58)			
UEMP	-0.53	0.38	0.60	-0.17	-0.56	-0.34	-0.49	-0.48	0.47	-0.09	0.48	-0.29		
	(0.00)	(0.00)	(0.00)	(0.07)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.35)	(0.00)	(0.00)		
AGE	-0.55	0.09	0.28	-0.60	-0.76	-0.63	-0.46	-0.51	0.42	-0-07	0.16	-0.37	0.48	
	(0.00)	(0.33)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.47)	(0.10)	(0.00)	(0.00)	
GRANT	-0.82	0.56	0.69	-0.41	-0.73	-0.33	-0.45	-0.60	0.34	0.05	0.38	0.47	0.58	0.61
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.57)	(0.00)	(0.00)	(0.00)	(0.00)

[1] Probability values in parenthesis.

Table A3: Industry classification according to Statistics Sweden's "limited level"

Industry number	*Description*	*Created from (NACE)*[1]	*Industry number*	*Description*	*Created from (NACE)*[1]
00	Unspecified	00, blank	22	Wholesale trade	51
01	Agriculture	01	23	Retail trade	52
02	Forestry	02	24	Transports	60-63
03	Fishing	05	25	Postal services and telephony	64
04	Mining and mineral extraction	10-14	26	Banks and credit institutions	65, 67.1
05	Foodstuff	15-16	27	Insurance	66, 67.2
06	Textile and clothing	17-19	28	Real estate services	70
07	Wood	20	29	Staffing and recruitment	71
08	Paper and pulp	21	30	Computer consultants and services	72
09	Publishing and graphical printing	22	31	Other corporate services	74
10	Chemicals	23-24	32	Education	80
11	Rubber and plastics	25	33	Research and development	73
12	Dirt and rocks	26	34	Health care	85.1, 85.2
13	Steel and metal plants	27	35	Child care	85.321, 85.322
14	Metals	28	36	Elderly care	85.311, 85.323
15	Machinery	29	37	Other care	85.312-85.315, 85.325
16	Electric and optical appliances	30-33	38	Hotels and restaurants	55
17	Transport equipment	34-35	39	Interest organizations	91
18	Other manufacturing	36-37	40	Recreation, culture, and sports	92
19	Energy, water and disposals	40-41, 90	41	Other services	93, 95
20	Construction	45	42	Public administration	75, 99
21	Motor vehicle sales and gas stations	50			

[1] Standard for Swedish industry classification.

Figure 1

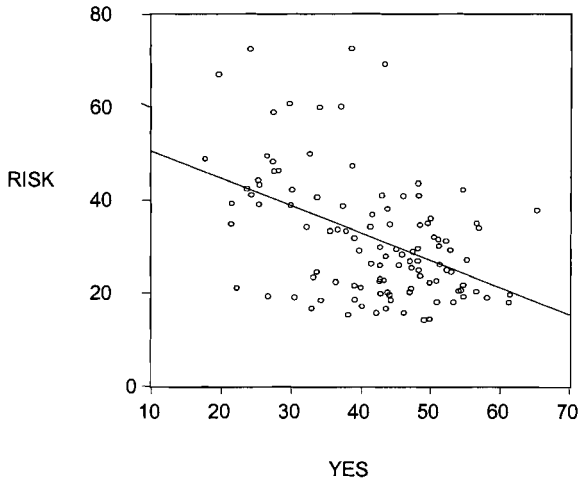

Table 1: The least and most risky local markets.

MARKET	*RISK*	*SOUTH-NORTH ORDERING*
Karlskoga	14.24	51
Arboga/Köping	14.36	57
Filipstad	15.30	60
Arvika/Säffle	15.74	41
Avesta/Säter	15.79	71
Mora	16.63	76
Västervik	16.64	36
Hällefors	17.17	62
Stockholm	17.98	56
Norrköping	18.07	44
Gällivare	49.59	103
Sollefteå	50.02	85
Storuman	59.02	96
Övertorneå	60.09	107
Kalix	60.22	105
Överkalix	60.88	104
Jokkmokk	67.16	102
Haparanda	69.35	106
Pajala	72.69	108
Boden	72.83	101

Table 2

	(1) OLS	*(2) 2SLS*	*(3) OLS*	*(4) 2SLS*	*(5) 2SLS*	*(6) 2SLS*	*(7) 2SLS*
RISK	-0.277	-0.417	-0.265	-0.414	-0.473	-0.451	-0.309
	(0.044)	(0.014)	(0.046)	(0.010)	(0.009)	(0.057)	(0.001)
$RISK^2$	0.003	0.005	0.003	0.005	0.006	0.006	
	(0.061)	(0.025)	(0.063)	(0.021)	(0.015)	(0.040)	
TURNOUT	-1.314	-0.204	-1.308	-0.203	-0.984	0.478	4.572
	(0.000)	(0.755)	(0.000)	(0.751)	(0.984)	(0.506)	(0.211)
SCHOOL	12.843	10.269	12.472	10.649	8.793	10.749	2.588
	(0.000)	(0.002)	(0.000)	(0.000)	(0.000)	(0.000)	(0.013)
INCOME	0.416	0.329	0.408	0.326	0.303	0.236	1.912
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.032)	(0.038)
POP	-0.209	0.066					
	(0.793)	(0.760)					
AVCH9497	-41.438	-36.808					
	(0.726)	(0.760)					
AGRI	89.028	96.360	91.401	98.868	52.895	84.77	0.115
	(0.002)	(0.003)	(0.001)	(0.002)	(0.158)	(0.094)	(0.009)
MOD	0.158	0.231	0.149	0.230	0.115	0.222	0.260
	(0.111)	(0.027)	(0.115)	(0.023)	(0.328)	(0.212)	(0.021)
SAP	0.192	0.236	0.194	0.242	0.168	0.329	0.317
	(0.014)	(0.006)	(0.008)	(0.004)	(0.063)	(0.006)	(0.130)
TREND	-0.002	-0.002	-0.002	-0.002	-0.002	-0.002	-0.151
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
CONST	-48.743	-103.99	-46.536	-107.33	-93.831	-153.30	-36.148
	(0.054)	(0.017)	(0.017)	(0.011)	(0.054)	(0.001)	(0.006)
ADJ. R^2	0.86	0.82	0.86	0.83	0.70	0.59	0.61
# OBS	109	109	109	109	84	109	109

White's standard errors have been used. P-values in parenthesis. Instruments for TURNOUT are the turnout ratios from the 1991 and the 1994 parliamentary elections. In columns (1)-(6), the dependent variable is YES. In column (5), labor markets with more than 50% yes-votes are dropped from the sample. In column (6), median regression rather than OLS is used, and in column (7), the dependent variable is Log(YES/100-YES) and all variables are entered in logarithms.

Table 3

CONTROL	*(1)* *2SLS* AGE	*(2)* *2SLS* WEL-FARE	*(3)* *2SLS* UEMP	*(4)* *2SLS* POP-DENS	*(5)* *2SLS* PUBLIC	*(6)* *2SLS* GRANT
RISK	-0.416	-0.408	-0.389	-0.408	-0.424	-0.347
	(0.009)	(0.009)	(0.020)	(0.011)	(0.015)	(0.029)
$RISK^2$	0.005	0.005	0.005	0.005	0.006	0.005
	(0.019)	(0.018)	(0.038)	(0.021)	(0.025)	(0.026)
INCOME	0.355	0.323	0.343	0.318	0.303	0.186
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.034)
EST. OF	0.232	1.665	0.106	0.021	-0.127	-0.788
CONTROL	(0.744)	(0.346)	(0.655)	(0.244)	(0.244)	(0.001)
ADJ. R^2	0.83	0.83	0.83	0.82	0.82	0.84

White's standard errors have been used. P-values in parenthesis. SCHOOL, TURNOUT, AGRI, MODERAT, SAP, TREND, and a constant are included in the regressions. Instruments for TURNOUT are the turnout ratios from the 1991 and the 1994 parliamentary elections.

Table 4

	OLS	SPACE LAG	SPACE LAG - IV -	SPACE ERROR	SPACE ERROR - IV -
RISK	-0.265	-0.256	-0.330	-0.253	-0.314
	(0.046)	(0.057)	(0.029)	(0.056)	(0.028)
$RISK^2$	0.003	0.003	0.004	0.004	0.004
	(0.063)	(0.094)	(0.045)	(0.018)	(0.019)
TURNOUT	-1.308	-1.083	-0.620	-1.130	0.049
	(0.000)	(0.000)	(0.097)	(0.000)	(0.919)
SCHOOL	12.472	11.142	10.350	11.366	9.197
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
INCOME	0.408	0.361	0.344	0.406	0.262
	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
AGRI	89.028	105.69	106.81	62.493	47.853
	(0.002)	(0.000)	(0.000)	(0.031)	(0.125)
MOD	0.158	0.083	0.126	-0.045	0.068
	(0.111)	(0.428)	(0.276)	(0.641)	(0.529)
SAP	0.192	0.125	0.156	0.028	0.070
	(0.014)	(0.059)	(0.036)	(0.669)	(0.350)
TREND	-0.002	-0.001	-0.001	-0.002	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
CONST.	-48.743	-52.667	-77.938	-41.653	-95.304
	(0.054)	(0.000)	(0.002)	(0.029)	(0.002)
GAMMA (γ)		0.244	0.186		
		(0.000)	(0.016)		
LAMBDA (λ)				0.637	0.686
				(0.000)	(0.000)
AIC[1]	637.59	624.80	na.	617.89	na.
LM – ERROR[2]	13.326	3.718	2.469		
	(0.000)	(0.054)	(0.116)		
LM – LAG[3]	14.407			0.959	0.086
	(0.000)			(0.327)	(0.770)

[1] The Akaike information criterion. [2] Test for spatial error structure. Chi-squared with 1 d.f. H_0: No spatial errors. [3] Test for spatial lag structure. Chi-squared with 1 d.f. H_0: No spatial lags. White's standard errors have been used. P-values in parenthesis. Instruments for TURNOUT are the turnout ratios from the 1991 and the 1994 parliamentary elections.

CHAPTER 4

DOES LABOR MARKET RISK INCREASE THE SIZE OF THE PUBLIC SECTOR? EVIDENCE FROM SWEDISH MUNICIPALITIES.[♣]

Jonas Vlachos[♦]

Abstract

It has been argued that the public sector is an insurance against otherwise uninsurable risks. If that is the case, it is reasonable to expect the public sector to be larger in regions where the private labor-market is risky. Using data from Swedish municipalities, this paper reports that labor-market risk has a substantial impact on public employment. The results for aggregate spending and taxation are, however, much weaker and labor-market risk thus affects the labor intensity of the municipal public sector.

JEL Classification: C23; H11; H40; J45

Keywords: Public sector size; Public employment; Labor market risk; Panel data

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[♦] Department of Economics, Stockholm School of Economics, P.O. Box 6501, 11383 Stockholm, Sweden. Email: jonas.vlachos@hhs.se.

1. Introduction

In a pair of influential papers, Rodrik (1998, 2000) shows that open economies tend to have larger public sectors. He also presents evidence in favor of a plausible explanation for this phenomenon: International trade raises the volatility of income, which can be decreased by the public sector. Optimally, an open trade regime therefore goes hand in hand with a larger government, when other insurance mechanisms are not available. Agell (1999) has subsequently argued that the globalization of economic activity increases labor market instability and hence, political demand for labor market rigidities. Although Alesina and Wacziarg (1998) have questioned the link from trade-induced risk to government size, the theoretical argument still has considerable appeal. Indeed, it would be surprising if the resources spent by the public sector were completely independent of the need for insurance. Here, the hypothesis that private sector risk increases the size of the public sector is addressed using data from Swedish municipalities. Several problems related to cross-country data availability and comparability are therefore circumvented.[1] The fact that there is no need to aggregate public spending across different levels of governance might be most important: local aggregate risk is directly related to the size of the local public sector.

The indicator of private-sector risk in local labor markets is found to have a substantial impact on municipal public employment. It is also found that shocks increasing the size of the public sector across all municipalities tend to generate larger increases in risky locations. While some of the results hold even for aggregate municipal spending and the rate of taxation, on balance the hypothesis can be rejected for these variables. Hence, labor-market risk affects the labor intensity of the municipal public sector, rather than its size.

By using municipal data, several well-known problems with cross-country data can be circumvented. First of all, data might not be comparable across countries. Second, institutional and cultural factors are not easily controlled for in a cross-country setting. Finally, there are plainly not enough comparable countries. Researchers are thus

forced to compare countries of very different degrees of development, hence aggravating the problems mentioned above. For these reasons, several researchers have successfully used municipal (or regional) data to test the predictions of different political-economy related models. The appeal of such an approach is obvious. Using municipal data in this study does not come without cost, however. There are three main reasons for expecting the effect of market-induced risk on government size to be smaller in Swedish municipalities than between countries. First of all, the Swedish central government provides insurance against individual-specific labor market risk through the unemployment insurance. No such mechanism exists between countries. Second, despite the international mobility of capital, private capital markets do not (yet) provide much risk sharing between countries (see e.g. Lewis 1995). These markets are therefore poor providers of insurance against aggregate country-specific shocks. Within a country, however, capital markets can certainly provide some insurance against municipal-specific risks. Sorensen and Yosha (2000) verify this by showing that interregional credit markets are highly integrated whereas the same is not true between OECD-countries. Finally, the Swedish central government has the means and power to transfer funds directly to the municipalities. Since we can find an economically significant effect of labor-market risk on local government size, despite these mitigating factors, this should be read as support for Rodrik's (1998, 2000) argument.[2]

A key issue in this paper is the indicator of labor-market risk, which is based on the industry structure of local labor markets; hence taking commuting patterns into explicit consideration. The behavioral consequences of local labor market risk have mainly been analyzed in the field of labor economics. Simon (1988) shows that frictional unemployment is higher in less diversified local labor markets and thus provides evidence that industrial shocks are less costly to people working in well diversified cities. Diamond and Simon (1990) probe deeper into the differences in unemployment risk between local labor markets. By building on a portfolio theoretic

[1] Swedish municipalities enjoy a constitutional right of self-governance, thereby making them appropriate units of analysis for questions regarding government size. Self-governance has given rise to considerable variation in the dependent variables used in this study.

[2] This said, countries have both monetary and fiscal policy instruments at their disposal to smooth temporary fluctuations. To some extent, Swedish municipalities can also use fiscal policy to this end (Dahlberg and Lindström, 1998). In fact, Pettersson-Lidbom (2001) even finds that municipalities use debt to strategically influence future policy makers.

view of labor markets pioneered by Conroy (1975), they show that cities with a more diversified industrial structure have both a lower probability of unemployment and lower wages. The reason is that a diversified labor market creates a portfolio effect for the individual worker, thereby making the aggregate risk lower compared to a specialized city. Thus, everything else equal, a worker is willing to accept lower mean wages on a diversified than on a specialized local labor market. All models used in local labor market analysis are based on the interplay between costly migration and local labor market conditions. Given the everyday importance of local markets, it should be of no surprise if political attitudes are also influenced by the insurance properties of different locations.

The paper is organized as follows. In Section 2, Rodrik's (1998) argument is presented. In Section 3, the data is described, with special attention given to the indicator of labor-market risk. Section 4 discusses the empirical specification, while the results are presented in Section 5. Section 6 concludes. In the appendix, all variables and data sources are described in detail.

2. Rodrik's model

Following Rodrik (1998), an economy with a fixed supply x of a manufactured good, not consumed within the economy, is considered. The manufactured good is traded for an import good at price π. In addition, there is a publicly and a privately produced good and the productivity of private production is enhanced by the import good. Since trade is continuously balanced, the economy can afford $x\pi$ of the import good. The private sector employs $(1-\lambda)$ of the total labor supply (which is normalized to one). Assuming that the production of the private good is linear, total production is $x\pi(1-\lambda)$. The supply of the publicly produced good is $h(\lambda)$, where $h'>0$ and $h''<0$. The size of the public sector, λ, is determined before the realization of π is known. Assuming that the public and the private good are perfect substitutes, the government's problem is to maximize $V(\lambda) = E[u(h(\lambda) + x\pi(1-\lambda))]$ with respect to λ. $u(\cdot)$ is the utility function of a representative household, displaying standard properties. If the mean of π equals π_m, and the variance equals σ^2, the Taylor expansion of $V(\lambda)$ gives the following expression (after taking expectations)

$$(2.1) \qquad V(\lambda) \approx u(h(\lambda) + x\pi(1-\lambda)) + \tfrac{1}{2} u''(h(\lambda) + x\pi_m(1-\lambda))(1-\lambda)^2 x^2 \sigma^2.$$

The first-order condition of this expression is then (if $z = x\sigma$),

$$(2.2) \qquad \{u'(\cdot) + \tfrac{1}{2} u'''(\cdot)(1-\lambda)^2 z^2\}[h'(\lambda) - x\pi_m] - \tfrac{1}{2} u''(\cdot)\, z^2 = 0.$$

As long as households display prudence ($u'''>0$), the first term is positive.[3] Then, the larger the aggregate risk z, the larger is the optimal level of λ. To see this, consider an economy, A, with zero aggregate external risk ($z_A = 0$). The first-order condition is then reduced to $h'(\lambda_A) = x\pi_m$. In another economy, B, risk is positive ($z_B > 0$) and hence $h'(\lambda_B) < x\pi_m$. Since $h''<0$, it must be that $\lambda_B > \lambda_A$. In other words, a relocation of resources to the safe public sector is the optimal response to a higher level of aggregate manufacturing risk.

Naturally, the providing a safety net would be the ideal way of insuring against aggregate risk. Swedish municipalities do not, however, have the power to directly redistribute income between households: this task is performed by the central government. Given that the local government wants to provide further insurance for its inhabitants, it must do this in less direct ways.

It should be noted that the issue here is not whether counter-cyclical policies can reduce income volatility, but whether the *size* of the public sector can do so.[4] The answer to this question is not obvious and ultimately, it is empirical. Fatás and Mihov (2001) document a strong negative correlation between government size output volatility across OECD countries, as well as across US states. Their study explicitly deals with several conceptual and econometric problems arising when investigating the proposed effect. In Table 1, the standardized variance of the pre-tax municipal per capita income for the period 1985-1999 is regressed on the 1985 values of municipal employment, spending, and tax rate, respectively (all variables are in logarithms). Per

[3] For a discussion of prudence in a macro context, see e.g. Romer (1996), chapter 7. In a micro context, see e.g. Menezes et al. (1980).

capita income and a dummy for municipalities performing tasks usually performed at county level (see footnote 5) are added as control variables in this regression. Admittedly, this is a very crude way of analyzing the relation between public sector size and income volatility, and the results should therefore be read as indicative rather than conclusive. As Rodrik (1998, p. 1023) notes, the coefficient on government size is downward biased in this regression if the hypothesis presented here is true. Despite this, a substantial negative effect on income volatility can be found for all three indicators of public sector size, however.

[Table 1]

3. Data

The constitutional right of self governance enjoyed by Swedish municipalities makes them appropriate units of analysis when studying the determinants of government size. In total, there are at present 289 Swedish municipalities. Nine of which are removed from the sample because they are either split or created in the time period 1983-1999.[5] The local tax rate, per capita municipal spending, and the share of inhabitants employed by the municipality will be used as dependent variables. There is considerable variation in these variables: The tax rate ranges from 9.7 to 23.6 percent, per capita spending from 20400 to 61500 SEK, and the share of municipal employment from 2.1 to 9.7 percent.[6]

When measuring the degree of labor market risk, commuting and other spillovers make the use of data on the municipal level unreasonable. Rather, data on the "local labor market"-level is used. Local labor markets are analytical, and not administrative, units of analysis whose boundaries are based on actual commuting patterns within and between municipalities. In total, there are 109 local labor markets in Sweden and data

[4] Sorensen et al. (2001) find that strongly procyclical revenue and weakly procyclical spending lead to procyclical budget surpluses in U.S. state and local governments. Dahlberg and Lindström (1998) find that Swedish local governments act to intertemporally smooth income.
[5] Three municipalities (Malmö, Gotland, and Göteborg) also perform tasks normally undertaken at the county (landsting) level. Whether these municipalities should be included in the sample is an open question. Here, they are included since they constitute a substantial share of the Swedish population. Further, the municipal specific fixed effects should capture the difference in responsibilities across municipalities. Excluding them has a miniscule effect on the results.
[6] These figures do not include municipalities with county-level responsibilities.

is available biannually between 1985-1999.[7] All in all, the data set covers 280 municipalities over eight time periods, giving a total of 2240 observations. The dependent variables from 1983 are included in order to reduce the loss of time periods when estimating dynamic specifications. All variables are entered in logarithms and the monetary variables are expressed in 1995 prices.

To construct the indicator of private aggregate local labor market risk, we follow the portfolio-theoretic approach used by Conroy (1975) and Diamond and Simon (1990) to derive what can best be described as an augmented Herfindahl index. The concept is borrowed from the literature on financial economics: each local market is viewed as a portfolio of industries, each of which is subject to random fluctuations. The total degree of labor market risk in a region is the weighted average of the variances and covariances of these fluctuations. The total variance in city c equals

$$\text{RISK}_c = \sum_s s_{sc}^2 \sigma_{ss}^2 + \sum_s \sum_u s_{sc} s_{uc} \sigma_{su} \quad , s \neq u, c = 1,...,109 \tag{3.1}$$

where s_{sc} is the employment share of sector s in labor market c, σ_{ss} is the variance of employment in sector s, and σ_{su} is the covariance of employment in industries s and u. Specifically, the variance/covariance matrix is constructed as follows: The variance element, σ_{ss}, is the variance in the rate of national employment growth in sector s at time t, while the covariance elements, σ_{su}, equal the covariance between growth rates. In order to capture the degree of unexpected volatility, the employment growth rate of each industry was de-trended by taking the first differences of employment on the national level. The logarithms of the first differences were then regressed on a time trend. For each regression, the residuals were used to construct the variance/covariance matrix and thus the risk index at each point in time provides an

[7] See *Statistics Sweden* (1998) for a closer description of the LLM subdivisions. That the indicator of risk is constructed at a different level of aggregation than the dependent variables potentially induces a downward bias in the standard errors (Moulton 1990). Here, there are, on average, only 2.6 municipalities per local labor market, which means that the problem is of little practical relevance. Allowing for possible cluster effects between municipalities in the same labor market yields an increase in the standard errors, but this effect is very small.

indication of the need for insurance against labor market induced risk. It can be noted that under the very strong assumptions of independently and identically distributed growth rates, the portfolio diversification index is equivalent to the well known Herfindahl index.

In order to calculate the variance/covariance matrix, sectoral employment data at the national level between 1985-1999 is used. This procedure amount to the assumptions that industries are subject to the same fluctuations, regardless of their geographical location, and that the pattern of fluctuations is stable over time. Obviously, these assumptions are not a perfect description of reality, but they are necessary to capture both the within and between industry fluctuations in a systematic manner. Another important question is which industries should be included in the study. To mainly capture the effects of economic fluctuations from outside the local market, the sample consists of all manufacturing and a few business-cycle sensitive service industries, making a total of 21 industries. [8]

4. Specification of the empirical model

The shares of population between 0-6, 7-9, 10-19, 65-79, and 80+, respectively are used as control variables. This division of groups is based on the cost differences between age groups.[9] Population density is also included to account for cost differences between municipalities. Per capita income is included in all specifications, while the average wage in the municipal sector is included when municipal employment is the dependent variable. As further controls, a dummy variable for left-wing majority and a dummy for undefined political majority are sometimes included as checks of robustness. It must be kept in mind that these political variables are likely to be affected by the degree of labor market risk (given that labor market risk is an important determinant of policy). The central government redistributes resources both

[8] The industry classification is based on *Statistics Sweden's* "limited level". A description of the industries is available in the appendix. The 21 industries constitute 35 percent of total employment and 52 percent of total private employment. Minor changes in the sample of industries do not change the results.

directly to, as well as between, municipalities. However, the system of redistribution has changed during the relevant time period, making it difficult to compare the impact of government grants on the size of the local government over time (see Bergström et al. 1998). This problem is further accentuated because of changes in grants-variable definitions over time. Another problem with central government grants is that for part of the period, they were mainly matching grants.[10] Hence, they are likely to be endogenous to the fiscal decisions taken in the municipalities. For these reasons, per capita grants will only be included to check for robustness. Since grants can be negative, they cannot be entered as a separate regressor in a logarithmic specification. Rather, they are added to per capita income.[11] In addition to these variables, I have also experimented by using the share of non-Swedish citizens, the share of women, the share of welfare recipients, and changes in aggregate employment at the municipal level. None of these variables has significant explanatory power, and their inclusion does not change the other estimates in any substantial way.

The general specification of the empirical model is then

$$(4.1) \qquad y_{it} = \beta z_{it} + \gamma \boldsymbol{X}_{it} + \lambda_t + v_i + \varepsilon_{it},$$

where y_{it} denotes a particular measure of government size in a municipality in a specific year, z_{it} is the indicator of labor market risk, and $\boldsymbol{X}_{it}$ the vector of control variables mentioned above. As concerns the remaining variables, λ_t is a time-specific effect, constant over municipalities, v_i is a municipal-specific effect, constant over time, and ε_{it} is the usual residual.

The advantage of using fixed municipal- and time effects is that they account for a large number of uncontrollable factors affecting the size of government. The time-specific effects, for example, control for business-cycle movements that could affect public spending. Since the fixed effects may be falsely attributed to changes in government size that should really be accounted for by other variables, this advantage

[9] Swedish Government Official Reports (1993).
[10] Aronsson and Wikström (1996) and Bergström et al. (1998) provide descriptions of the Swedish grants system.
[11] In fact, a different effect on spending from grants than from tax revenue is not supported by any utility function (Dahlberg and Jacob 2000).

comes at a cost, however. An attempt to account for these problems will be made in Sections 5.3 and 5.4.

5. Results

5.1. Fixed effects estimates

Table 2 shows the basic results for both the base and the full conditioning set of control variables.[12] The explanatory power is very good for the employment and taxation specifications (adjusted $R^2 = 0.93$), and quite good for spending (0.69). The indicator of labor market risk is highly significant when municipal employment is used as the dependent variable. In the base specification for per capita spending, labor market risk is weakly significant (at the ten percent level), but not in the full specification. The municipal tax rate, on the other hand, seems to be independent of labor market risk. A high per capita income increases the size of the public sector while population density has a negative effect. Municipalities where the public wage rate is high tend to have a lower number of employees and most of the demographic variables display a high level of statistical significance.

[Table 2]

The point estimates are not very large: 0.09 and 0.08 for the employment equation. There is, however, considerable variation in labor market risk. One standard deviation increase from the mean of this variable implies a change in the dependent variable by 4 percent. As a comparison, one standard deviation change in employment equals a 28 percent deviation from the mean of this variable. Some calculations show this effect to be somewhat larger than the effect of having a left-wing rather than a right-wing political majority in the municipality (see Pettersson-Lidbom 2000 for a thorough discussion of this topic).

[12] The full set includes central government grants (added to per capita income), a dummy for left-wing political majority, and a dummy for undefined political majority.

5.2. Dynamic specification. GMM.

Several studies (e.g. Bergström et.al 1998 and Pettersson-Lidbom 2000) of Swedish local governments have found that dynamics are important. Further, in a study of the demand for local public services, Dahlberg and Jacob (2000) have shown that the explanatory variables used in this paper are not strictly exogenous.[13] Since the fixed effects estimator is biased and inconsistent in a dynamic set-up, the Arellano and Bond (1991) GMM-estimator is used when introducing a lagged dependent variable to the specifications. This estimator is based on the use of lagged levels of the dependent and endogenous variables, while strictly exogenous regressors are estimated in differences. In order to check the model specification and the validity of the instruments, the Sargan over identification test is used. Since this test tends to over reject in the presence of heteroskedaticity, two-step estimates are used to judge the model specification. The reported coefficients are, however, based on (robust) first-step estimates, since Arellano and Bond show that the two-step standard errors tend to be downward biased. In order to save space, only the specifications where all regressors are treated as endogenous are presented.[14] The Arellano-Bond method is only valid if second-order autocorrelation is not present. For this reason, the test statistics of an AR(2) test (null hypothesis of no autocorrelation) are also presented.

[Table 3]

As shown in Table 3, the lagged dependent variables are highly significant in all specifications for all three dependent variables. Allowing for dynamics and endogeneity results in an increase of the point estimates of the risk variable by almost 100 percent (0.14-0.18). A 44 percent increase in labor market risk then implies an increase in municipal employment by 6-8 percent, which is indeed substantial. This is just the short-run effect, however. The long-run elasticity is calculated by dividing the coefficient on labor market risk by one minus the coefficient on the lagged dependent variable, which yields an elasticity of 0.52-0.72, i.e. very large indeed. Column 3 shows that the point estimate in the spending equation is now statistically significant

[13] The reason can be Tiebout-migration, causing both income and the demographic structure to be endogenous.
[14] The Sargan test rejects all specifications where the regressors are treated as strictly exogenous.

and large (short-run elasticity 0.2, long-run 0.33). This result is somewhat sensitive to the model specification, and when using the full conditioning set, statistical significance falls just short of the 10-percent level. Interestingly, labor-market risk has a small and marginally significant *negative* effect on the tax rate in these dynamic specifications. The reason could be that there is relatively little variation over time in this variable, and that tax increases were illegal during the first half of the 1990's.

5.3. Response to unobservable shocks. NLS.

One problem with the estimates presented in Sections 5.1 and 5.2 is that there is less variation over time in the degree of labor market risk within labor markets compared to the variation between labor markets. More precisely, the range of the risk variable within labor markets is 0.15-0.42, while the range between labor markets is 0.17-1.61. Since both the fixed effect- and the GMM-estimators mainly exploit the short-run time variation in the data, it is possible that some of the effects that should be attributed to labor market diversity are actually captured by the fixed effects. In particular, it is plausible that municipalities respond differently to similar shocks, depending on how well diversified their labor markets are. In all regressions, the time-effects are highly significant, indicating that some set of common shocks has influenced (i.e. increased) the size of the municipal sector in Sweden. These shocks may, for example, include business cycle effects, changes in preferences, and changes in the municipalities' responsibilities. Since we do not exactly know what shocks are captured by the time-effects, it is not obvious how to estimate their interrelation with labor-market risk. Blanchard and Wolfers (2000), however, suggest a method for dealing with this type of problem. In a study of the effect of labor market institutions on unemployment, they interact measures of country-specific institutions with shocks common to all countries in their study. Persson and Tabellini (2001) employ the same methodology when investigating the effects of political institutions on economic policy.

Following Blanchard and Wolfers, equation 4.1 is re-specified as

$$(5.1) \qquad y_{it} = \alpha + \beta z_{it} + (1+\eta (z_{it} - z_t))\, \varphi_t \lambda_t + \gamma \boldsymbol{X}_{it} + \nu_i + \varepsilon_{it}.$$

As above, y_{it} is the measure of government size, z_{it} is the indicator of labor market risk, $\boldsymbol{X}_{it}$ is the same vector of control variables as in the previous sections and v_i is the municipality specific effect. As before, the set of year dummies is used to estimates the effect of common events. Now, the effect of common events λ_t is proportional to $\eta(z_{it} - z_t)$, where z_t is the year average across municipalities. In other words, I measure whether the same unobservable shock has a different effect in municipalities deviating from average labor market risk. Because of the form of the specification, the variable η must be estimated using non-linear least square. An additional problem is that the lagged dependent variable cannot be included in these regressions since that would yield biased and inconsistent estimates.[15]

[Table 4]

In Table 4, results are presented that indeed support this hypothesis. The time effects on the size of the public sector and the interaction term η are highly significant in the employment and taxation equations, although no such effect can be found on spending. This means that any undefined shock increasing the size of the public sector is associated with a larger effect in risky labor markets than in safe ones.

To understand the size of this effect, the cumulative time-effect is calculated by subtracting the coefficient of the 1999 time-period dummy from the 1985 dummy (φ_{99} - φ_{85}). The time effect is then multiplied by the interaction term η, and by 44 percent (a one standard deviation increase from the mean in the risk variable), resulting in a cumulative effect of around 4 percent. As seen in the top row of Table 4, the point estimate of labor market risk (β) is around 0.1 in this specification. Altogether, these results suggest that two municipalities with the same per capita number of employees in 1985, going through the same sequence of shocks with a 44 percent difference in labor market risk, display a 8.5 percent difference in per capita employment in 1999.

[15] When only entering the lagged dependent variable in a fixed effects setting, the point estimates of labor market risk are insignificant for all specifications. As has been seen, however, when using the GMM-estimator to properly address the dynamic econometric problems, labor market risk does indeed

5.4. Response to observable shocks

The above results show that the response to common shocks indeed differs with labor market risk. Naturally, some curiosity arises as to what types of shocks are considered. We can, for example, expect the public sector to expand more during a temporary economic downturn when the private labor market is risky than when it is well diversified. To investigate this hypothesis, the GDP growth rate is interacted with the indicator of labor market risk. The interaction term is expected to be negative since risky labor markets have a greater incentive than safe ones to expand the public sector when the growth rate is low. The same reasoning applies when using the interaction between the total change in employment (at the local labor market level) and labor market risk. In other words: a negative (positive) point estimate indicates that the public sector is expanded to a relatively small (large) extent in risky labor markets when employment growth is high.[16]

[Table 5]

These predictions do find some support in Table 5, although the results are somewhat dependent on the type of shock. Interpreting the point estimates is quite difficult, but they carry the general message that labor market risk is one determinant as to how growth and employment shocks are handled at the municipal level.

5.5 Some additional tests

On average, staff costs constitute about 50 percent of total spending. In the full sample this ratio varies from 26 to 68 percent. So far, the results indicate that municipalities subject to large labor market risk tend to increase the number of employees rather than the size of the public sector per se, which implies that the public sector is relatively labor intensive in risky locations. A direct way of testing this hypothesis is to regress the ratio between total municipal labor costs and total

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come out significant. When including the lagged dependent variable in the NLS-specifications, the results are considerably weakened.

[16] Entering employment change as an independent variable does not change any results, and its point estimate is not statistically significant. The high correlation between changes in GDP and changes in local employment (0.86) makes it likely that the time-period dummies capture any direct effects.

expenditures on the indicator of labor market risk. In the first two columns of Table 6, the results from fixed effect estimates, as well as NLS-estimates using an interaction term between unobservable shocks and the indicator of labor market risk (as in Equation 5.1) are presented. As can be seen, there is indeed support for this hypothesis. The point estimates are similar to those in Table 2 and 4, indicating that the increase in labor intensity is roughly similar to the increase in public employment. One might also worry about the possibility that safe municipalities have a greater propensity to have services provided by private entrepreneurs. If that were the case, municipal spending would be constant, while municipal employment would be lower in these municipalities. The results in Table 6 are, however, a strong indication against this possibility.

[Table 6]

As an additional check of robustness, the ratio of public employees to the working age population (ages 20-64) is used as the dependent variable, rather than the per capita ratio used before. As can be seen in the last two columns of Table 6, the results are similar to the previous ones.

6. Conclusions

Consider two municipalities that only differ in the degree of aggregate private labor market risk. More precisely, one labor market is one standard deviation above the mean (44 percent) riskier than the other. These municipalities go through the same sequence of shocks to the size of the public sector between 1985-1999. The findings in this paper indicate that per capita public employment is between 8 (lower bound) and 30 (upper bound) percent higher in the risky relative to the safe municipality at the end of this period. Surprisingly, given these results, labor market risk does not seem to be an important determinant of aggregate municipal spending or taxation. Rather, the public sector in municipalities with a substantial representation of risky private industries is relatively labor intensive.

A reasonable interpretation of these results is that municipalities attempt to provide an insurance against unemployment risk, rather than income insurance per se. Since

income insurance is largely provided for by the central government, this may very well be an optimal policy. Support for this interpretation is given by the reduction in point estimates on risk in the spending equations that occur when central government grants are controlled for. In other words, these transfers seem to mitigate the adverse effects of a risky private labor market.

In this paper, only the risk due to the local composition of private industries has been considered. Although Clark (1998) has shown that industry shocks constitute a substantial share of the total variation of regional output, it would be interesting to include other sources of variation in the analysis. One possible extension of this work would therefore be to decompose local labor market risk into its national, regional, and industry-specific components. Since the optimal local policy response depends on the sources of volatility, the different components can then be related to policy choices both at the local and central levels. The greatest challenge of such an undertaking would be to construct indicators that capture risk, rather than fluctuations *per se*. So far, most of the research on regional output and income volatility has only dealt with the latter issue (see e.g. Clark and Shin, 2000).

Appendix. Data description and variables

Split (Borås (1490), Nyköping (480), Södertälje (181), Örebro (1880)) and new (Bollebygd (1443), Gnesta (461), Lekeberg (1814), Nykvarn (140), Trosa (488)) municipalities are removed from the sample.

The industry classification is based on *Statistics Sweden's* "limited level" which, in turn, is constructed using the MIS standard of Swedish industry classification. The following industries are included in the sample: fishery, transport equipment, agriculture, metals, rubber and plastics, textiles and clothing, wood, other manufacturing, mining and minerals, forestry, dirt and rocks, electric and optical appliances, steel and metal products, construction, machinery, telephone and postal services, publishing and graphical printing, paper and pulp, foodstuff, insurance, chemicals. The following industries are not included: Energy- water- and garbage disposals, motor vehicle sales and gas stations, wholesale trade, retail trade, transports and storage, real estate services, recruitment and staffing services, computer consultants and services, other corporate services, education, research and development, health care, child care, elderly care, other care, hotels and restaurants, interest groups, recreation and sports, other services, public administration, unspecified activities. The included industries are listed according to their variance (variance measure described in text), the most volatile come first.

The ten least risky (definition in text) labor markets in 1991 were: Sundsvall, Timrå, Karlstad, Stockholm, Hylte, Ånge, Gävle, Sandviken, Avesta, Borlänge. The ten most risky labor markets in 1991 were: Olofström, Trollhättan, Arboga, Tidaholm, Dorotea, Gnosjö, Gislaved, Oskarshamn, Skövde, Gullspång.

The data is biannual, and all monetary variables are expressed in 1995 prices after being deflated by the GDP-deflator.

RISK: Labor market risk measure calculated according to the portfolio measure described in the text. Source: *Statistics Sweden* "Registerbaserad arbetsmarknadsstatistik". Mean: 0.289. Std: 0.129.

INCOME: Total taxable personal income, lagged one period (due to delay in payments), divided by total population (at the end of the year). Sources: *Statistics Sweden* "Årsbok för Sveriges kommuner" and "Vad kostar verksamheten i din kommun?". Mean: 76870. Std: 15445.

INCOME (grants adjusted): Total taxable personal income, lagged one period (due to delay in payments), plus net grants and redistribution to the central government divided by total population (at the end of the year). Sources: *Statistics Sweden* "Årsbok för Sveriges kommuner" and "Vad kostar verksamheten i din kommun?". Mean: 80637. Std: 16062.

POPULATION DENSITY: Total population (at the end of the year) divided by area. Source: *Statistics Sweden* "Vad kostar verksamheten i din kommun?". Mean: 111.04. Std: 376.58.

SHARE –6: Share of population between 0-6 years of age (×100). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 8.67. Std: 1.11

SHARE 7-9: Share of population between 7-9 years of age (×100). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 3.89. Std: 0.55.

SHARE 10-19: Share of population between 10-19 years of age (×100). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 13.91. Std: 1.68

SHARE 65-79: Share of population between 65-79 years of age (×100). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 13.91. Std: 2.95.

SHARE 80+: Share of population above 80 years of age (×100). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 4.59. Std: 1.35.

LEFT-WING MAJORITY: 1 if SAP + VP obtains more than 50% of the valid votes, 0 otherwise. Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 0.396. Std: 0.489.

RIGHT-WING MAJORITY: 1 if MOD + KD + FP + CP obtains than 50 % of the valid votes, 0 otherwise. Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 0.360. Std: 0.480.

UNDEFINED MAJORITY: 1 if LEFT-WING and RIGHT-WING MAJORITY both equal 0, 0 otherwise. Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 0.245. Std: 0.430.

WAGE: Implicit unit labor cost, derived by dividing the total municipal labor costs (wages, employer taxes and fees) with total municipal employment. Sources: *Kommunförbundet* "Kommunal Personal" and *Statistics Sweden* "Vad kostar verksamheten i din kommun?". A few missing observations for 1995 have been replaced by the average of 1994 and 1996. Mean: 307894. Std: 45330.

EMPLOYMENT: Municipal full-time equivalent employees as a share of the population (×100). Source: *Kommunförbundet* "Kommunal Personal". A few missing observations for 1995 have been replaced by the average of 1994 and 1996. Mean: 5.46. Std: 1.54.

EMPLOYMENT PER WORKING AGE POPULATION: Municipal full-time equivalent employees as a share of the population aged 20-64. Calculated as EMPLOYMENT/(1-[SHARE-6]-[SHARE7-9]-SHARE[10-19]-[SHARE65-79]-[SHARE80+]). Mean: 9.88. Std: 2.94.

SPENDING: Total municipal expenditure divided by total population. Source: *Statistics Sweden* "Vad kostar verksamheten i din kommun" och "Årsbok för Sveriges kommuner". Nine missing values for 1995 have been replaced by the average of 1994 and 1996. Mean: 32820. Std: 5984.

STAFF COST SHARE OF TOTAL EXPENDITURE: Total municipal labor costs (wages, employer taxes and fees) divided by total expenditures. Sources: *Kommunförbundet* "Kommunal Personal" and *Statistics Sweden* "Vad kostar verksamheten i din kommun?". A few missing observations for 1995 have been replaced by the average of 1994 and 1996. Mean: 0.50. Std: 0.05.

TAX RATE: The municipal income tax rate (in percent). Source: *Statistics Sweden* "Sveriges officiella statistik". Mean: 18.04. Std: 2.30.

EMPLOYMENT SHOCK: The percentage change in total employment at the local labor market level between t-1 and t. Source: *Statistics Sweden* "Registerbaserad arbetsmarknadsstatistik". Mean: –1.54. Std: 5.87.

GROWTH SHOCK: Growth rate of real Swedish GDP between t-1 and t. Source: *Statistics Sweden* "Nationalräkenskaperna". Mean: 4.07. Std: 3.63.

DEFLATOR: The ratio between nominal GDP and GDP in 1995 year prices. Source: *Statistics Sweden* "Nationalräkenskaperna".

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Table 1. The effect of public sector size on income volatility.

Dependent variable:	*Income volatility*[a] *1985-99*	*Income volatility*[a] *1985-99*	*Income volatility*[a] *1985-99*
Employment (1985)	-0.486**		
	(0.204)		
Spending (1985)		-0.539***	
		(0.158)	
Tax rate (1985)			-0.731*
			(0.395)
Income (1985)	-0.047	-0.144	-0.520*
	(0.382)	(0.314)	(0.279)
Adj. R^2	0.12	0.15	0.12
# obs.	280	280	280

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Robust standard errors in parentheses. Regressions also include a dummy for municipalities with county-level responsibilities (coefficient not reported).

a) Income volatility is the variance of per capita income divided by the mean of per capita income.

Table 2. Basic specification. Fixed effects estimates

Dependent variable:	*Employment*		*Spending*		*Tax rate*	
Specification:	Base	Full	Base	Full	Base	Full
Risk	0.091***	0.075***	0.061*	0.049	-0.012	-0.017
	(0.032)	(0.031)	(0.035)	(0.035)	(0.145)	(0.014)
Income	0.609***		0.571***		0.089***	
	(0.062)		(0.067)		(0.028)	
Income		0.729***		0.552***		0.089***
(grants adj.)		(0.058)		(0.064)		(0.026)
Wage	-0.790***	-0.779***				
	(0.030)	(0.030)				
Population	-0.860***	-0.818***	-0.377***	-0.353***	-0.043*	-0.049**
density	(0.051)	(0.050)	(0.055)	(0.056)	(0.023)	(0.023)
Share –6	0.034	0.048	0.226***	0.207***	-0.014	-0.015
	(0.046)	(0.044)	(0.050)	(0.049)	(0.021)	(0.020)
Share 7-9	0.117***	0.101***	0.104***	0.084***	0.004	0.000
	(0.029)	(0.028)	(0.031)	(0.031)	(0.013)	(0.013)
Share 10-19	0.297***	0.251***	0.266***	0.198***	0.006	0.004
	(0.056)	(0.053)	(0.061)	(0.059)	(0.025)	(0.024)
Share 65-79	-0.140***	-0.146***	-0.013	0.015	0.023	0.021
	(0.041)	(0.040)	(0.044)	(0.044)	(0.018)	(0.018)
Share 80+	-0.241***	-0.217***	-0.090***	-0.078**	-0.011	-0.004
	(0.030)	(0.030)	(0.033)	(0.033)	(0.013)	(0.013)
Left-wing		0.034***		0.026***		0.019***
majority		(0.009)		(0.010)		(0.004)
Undefined		0.018***		0.018***		0.015***
majority		(0.006)		(0.007)		(0.003)
Within R^2	0.93	0.93	0.69	0.69	0.93	0.93
# obs.	2240	2240	2240	2240	2240	2240

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Standard errors in parentheses. Time-period and municipal-specific fixed effects are included.

Table 3. Dynamic specification. GMM-estimates

Dependent variable:	*Employment*		*Spending*		*Tax rate*	
Specification:	Base	Full	Base	Full	Base	Full
Risk	0.180***	0.147**	0.204**	0.153	-0.069*	-0.070*
	(0.069)	(0.068)	(0.102)	(0.101)	(0.037)	(0.039)
Lagged dep.	0.751***	0.726***	0.405***	0.410***	0.625***	0.631***
Variable	(0.070)	(0.069)	(0.041)	(0.041)	(0.123)	(0.124)
Sargan test[a]	474.00	477.44	337.63	398.98	560.65	558.99
(one-step)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Sargan test[a]	256.74	255.50	222.37	224.71	246.13	248.88
(two-step)	(0.563)	(0.550)	(0.713)	(0.638)	(0.265)	(0.223)
AR(2) test[b]	0.16	0.74	1.38	-1.18	-0.83	-0.61
(one-step)	(0.876)	(0.459)	(0.168)	(0.238)	(0.401)	(0.541)
# obs.	1680	1680	1680	1680	1680	1680

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Standard errors in parentheses. All point estimates are (robust) two-step. The same sets of control variables as in Table 2 are used.

a) Sargan gives the test statistic and p-value of the Sargan test for over-identifying restrictions. Both the test statistics from the first and second step estimations are presented since the test statistic from the first-step has been shown to over-reject in the presence of heteroscedasticity. The null hypothesis is that the instruments are valid/the model is correctly specified.

b) AR2 is the test statistic and p-value for no second-order correlation. The null hypothesis is that there is no second-order serial correlation.

Table 4. Unobservable shocks. NLS-estimates.

Dependent variable:	*Employment*		*Spending*		*Tax rate*	
Specification:	Base	Full	Base	Full	Base	Full
Risk [β]	0.106***	0.093***	0.061	0.043	-0.012	-0.015
	(0.032)	(0.032)	(0.035)	(0.036)	(0.014)	(0.014)
Risk deviation from mean × time dummies [η]	0.341***	0.365***	0.000	-0.156	0.069***	0.064***
	(0.054)	(0.064)	(0.130)	(0.186)	(0.025)	(0.025)
Cumulative time effect 85-99 [φ_{99} - φ_{85}]	0.283	0.23	0.10	0.07	0.22	0.22
Cumulative time effect in municipality with labor market risk 44% above mean. $\eta[\varphi_{99} - \varphi_{85}] \times 44\%$[b]	4.2%	3.7%	0%[a]	-0.5%[a]	0.7%	0.6%
Adj. R^2	0.94	0.94	0.82	0.82	0.95	0.95
# obs.	2240	2240	2240	2240	2240	2240

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Standard errors in parenthesis. The same sets of control variables as in Table 2 are used.

a) Not significant.

b) One standard deviation is 44 percent above the mean of the risk variable.

Table 5. Observable shocks. Fixed effects estimates.

Dependent variable:	*Employment*		*Spending*		*Tax rate*	
Type of shock:	Growth	Employ.	Growth	Employ.	Growth	Employ.
Risk	0.105***	0.079**	0.070**	0.059	-0.003	0.002
	(0.033)	(0.035)	(0.036)	(0.038)	(0.015)	(0.016)
Interaction between	-0.003**	-0.001**	-0.002	-0.002***	-0.002***	-0.0004
shock and deviation	(0.001)	(0.0006)	(0.001)	(0.0006)	(0.0006)	(0.0002)
from mean risk.						
Within R^2	0.93	0.92	0.69	0.66	0.93	0.93
# obs.	2240	1960	2240	1960	2240	1960

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Standard errors in parenthesis. The base set of control variables as in Table 2 is also included.

Table 6. Additional robustness tests

Dependent variable:	*Staff cost share of total costs*		*Public employment per working age population*	
Estimation method:	FE	NLS	FE	NLS
Risk	0.093***	0.090***	0.097***	0.112***
	(0.036)	(0.036)	(0.033)	(0.032)
Risk deviation from mean × time dummies		0.384***		0.356***
		(0.084)		(0.055)
Cumulative time effect 85-99		0.18		0.29
Cumulative time effect in municipality with labor market risk 44% above mean.[a]		3.0%		4.4%
R^2	0.29 (within)	0.49 (adjusted)	0.93 (within)	0.94 (adjusted)
# obs.	2240	2240	2240	2240

*** Indicate statistical significance on the 1%-level, ** on the 5%-level and * on the 10%-level. Standard errors in parentheses. The municipal wage rate, the population shares of age groups 0-6, 7-9, 10-19, 65-79, 80+, time-period and municipal-specific fixed effects are included in columns 1 and 2. In columns 3 and 4, the same controls as in Table 2 are included. [a)] 44 percent is one standard deviation above the mean.

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