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Liquidity measures, liquidity drivers and expected returns on an early call auction market

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July 2011

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Carsten Burhop[†] / Sergey Gelman[‡]

Abstract

We analyze the determinants of illiquidity and its impact on asset pricing for purely call-auction traded stocks on Berlin Stock Exchange using 22 years of daily data (1892-1913). We use the Lesmond et al. (1999) measure of transaction costs to proxy illiquidity. We show that transaction costs were low and comparable to today's costs. Liquidity was negatively correlated with active informed trading, particularly being low for small and distressed stocks and in crises times. Liquidity concerns were a major driver of asset pricing: we find significant illiquidity level and illiquidity risk premia as well as an explicit premium for informed trading.

JEL-Classification: G12, G14, N23

Keywords: Transaction Costs; Liquidity Premium; Informed Trading

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I. Introduction

In this paper, we empirically investigate theoretical predictions for the interplay of liquidity, informed trading and asset pricing using daily stock prices from the Berlin Stock Exchange from the period 1892-1913. The data are especially insightful for studying the link between information asymmetry and liquidity because the design of the market closely resembles the assumptions of sequential auction games (as in Kyle, 1985) with insiders observing the amount of noise trading (as in Rochet and Vila, 1994). More specifically, the Berlin Stock Exchange, which was then the major German stock exchange, was a call auction market with an official broker arranging one price fixing a day. Moreover, informed traders could submit their orders after having inspected the order book. This close alignment of actual market design and theoretical assumptions allows us to infer insider trading from stock price dynamics employing theoretical predictions: high positive serial correlation of stock returns indicate the presence of a strategically acting informed trader. Furthermore, the presence of informed trading justifies the application of Lesmond, Ogden and Trzcinka (1999, further LOT) indirect measure of transaction costs as liquidity measure.¹

One of the implications of theoretical market microstructure models is that a high information-to-noise ratio leads to low liquidity (e.g. Rochet and Vila, 1994, p. 145). We test this hypothesis indirectly, analyzing the impact of states with likely high information asymmetry on transaction costs. Such states are small and distressed companies as well as market crises.

Beyond measuring transaction costs and identifying the link between information asymmetry and liquidity, we address three theoretical propositions with respect to liquidity premia. Here we strongly benefit from nearly absence of microstructure noise in the data: There are no bid-ask spreads in the call auction market by design. This rules out the upward bias in liquidity premia, outlined in Asparouhova et al. (2009).

First, both informed and uninformed traders dislike illiquid stocks and require a premium for holding them (Acharya and Pedersen, 2005; Amihud and Mendelson, 1986). By and large, the empirical literature supports this view (Asparouhova et al., 2009; Eleswarapu, 1997; Chalmers and Kadlec, 1998). Second, insider trading deteriorates wealth of liquidity traders and should deter their participation in stock trading subject to heavy informed trading (Dow, 2004). However, if real investment is sufficiently information elastic, insider trading can be welfare enhancing (Bernhardt et al., 1995; Leland, 1992). Third, uninformed traders dislike stocks which are illiquid in bad times, as they can not be used to offset income flow shocks (Acharya and Pedersen, 2005). We assess these three hypotheses in an asset pricing framework, where LOT transaction costs proxy for illiquidity, first order autocorrelation of daily returns proxy for informed trading and a regression coefficient of individual transaction costs shocks on market returns proxy for co-movement of illiquidity and market downturns.

1 More generally, Pagano (1989) proves the close correlation of transaction costs and liquidity.

We show that transaction costs at the Berlin Stock Exchange a century ago were pretty low and of about the same size as they are in modern financial markets. The LOT measure indicates that the cost for a roundtrip transaction were about 0.97 percent of the share price. This compares to an estimate of 1.23 percent for the largest decile of firms listed at the New York Stock Exchange for the period 1963-1990 (Lesmond et al., 1999). We find support of the negative impact of the information-to-noise ratio on liquidity: transaction costs are higher for small stocks and after a year of negative returns; moreover, illiquidity increases in crises and decreases with turnover. Estimation of multifactor asset pricing models provides strong support of a liquidity premium: for one percentage point higher transaction costs, investors require about 3.6 percentage points higher expected annual return. We find some support of the informed trading premium: securities, which reveal symptoms of informed trading yield higher expected returns. The magnitude of this impact is strongly significant from an economic point of view. Moreover, we find a significant positive premium for liquidity risk: investors value securities, which transaction costs go up stronger in case of market downturns, significantly lower. Thereby usual asset pricing benchmarks – market risk and size – turn out to have no impact on the cross-section of stock prices.

Our finding of comparatively low transaction costs supports the theoretical superiority of call auction markets over the nowadays prevalent continuous trading or dealership markets (Pagano and Roell, 1996). In contrast to the theoretical prediction, empirical results using data from modern markets are not as clear cut. For example, data from the Tel Aviv stock exchange show that prices and liquidity increase when stocks move from a call auction to continuous trading (Amihud et al, 1997; Kalay et al., 2002). However, data from the Singapore and London stock exchange illustrate that the introduction of opening and closing call auctions decreases the extent of price manipulation and increases the extent of price discovery (Chang et al., 2008; Chelley-Steeley, 2008). Moreover, Pagano and Schwartz (2003) show that introduction of the closing auction on the Paris Bourse in 1996 led to the reduction of execution costs. In addition, experimental studies point out that call auction markets reduce asymmetric information between different groups of traders and lead to lower transaction costs, but reduce the speed of information processing (Schmitzlein, 1996; Theissen, 2000).

Beyond its contribution to financial economics, our paper also supports recent findings from economic history showing that Germany's historical stock market was quite efficient. Starting with the work of Weigt (2005), it has been shown that stock price differentials among German stock exchanges (Weigt, 2005: 199) and between the Berlin Stock Exchange and other major European stock exchanges were small (Baltzer, 2006), that stock prices reflected the risk and return characteristics of the shares quite well (Weigt, 2005: 224), and that the Berlin Stock Exchange was weakly information efficient (Gelman and Burhop, 2008). Furthermore, Gehrig and Fohlin (2006) estimate in a paper closely related to our work that the effective spreads of samples of Berlin traded shares during the benchmark years 1880, 1890, 1900, and 1910 were low and decreasing in firm size.

The remaining parts of the paper are organised as follows. In Section II we give a short description of price fixing at the Berlin Stock Exchange at the turn of the 20th century and describe our data sources. The LOT measure of round-trip transaction costs is illustrated in Section III, along with a brief description of implemented econometric techniques. The results are presented in Section IV, followed by robustness checks in Section V and conclusion in Section VI.

II. Market Structure and Data Description

Shares were traded at the Berlin Stock Exchange six days per week using a call auction mechanism. Prices were fixed once a day by official, government appointed brokers. The brokers' association allocated two official brokers to each stock listed at the exchange. They jointly fixed the official price of the share and they both had the duty to act as brokers for the stock, i.e. they could not decline to take orders. They started taking orders at noon and stopped taking orders not earlier than 1.30 p.m. and not later than 2 p.m. Orders were made orally by representatives of banks and other participants on the trading floor. The official broker orally repeated the order and his substitute recorded the order into the order book. The order book was arranged in four columns, one for unlimited buying orders, one for limited buying orders, one for unlimited selling orders, and one for limited selling orders. The official price had to reflect the true commerce at the stock exchange. At the official price, it had to be possible that all unlimited buy and sell orders as well as buy orders with a higher price limit and sell orders with a lower price limit were carried out. Whenever the official broker expected a major price change (i.e., a price change of more than one percent), he had to make a written announcement to the trading floor. Moreover, in this case, a state commissioner joined the two official brokers to monitor the price fixing. The first tentative price was prepared in public and all interested parties could attend this event. Moreover, it was still possible to place further orders or to cancel formerly made orders. Afterwards, the two official brokers went to the back office, where the official quotation was registered, signed by the state commissioner, and published in the official price list. Prices were quoted in percent of the face value of a share (typically 1,000 Mark) and the tick size, which could have relevance for price impact, was 0.05% of face value (Obst, 1921: 380, 386-392).

Turning to transaction costs, we can distinguish three types of observable costs: taxes, broker fees, and bank fees. Transactions at German stock exchanges were taxed from 1881 onwards. More specifically for the period under consideration here, the stock market turnover tax was 0.01 percent of the underlying transaction value between 1892 and April 1894. From May 1894 onwards, the tax was doubled to 0.02 percent; another increase to 0.03 percent followed in October 1900. In addition to turnover taxes, the fees for brokers influence transaction costs. The fee for official brokers was 0.05 percent of the underlying transaction value (Gelman and Burhop, 2008). Furthermore, fees for the banks or other intermediaries varied between 0.1 and 0.33 percent (Weigt, 2005: 192). In sum, broker fees, fees for intermediaries, and turnover

taxes added up to a total cost for a roundtrip transaction (i.e., buying and selling of a share) in the range of 0.252 to 0.82 percent.

To investigate the size of actual transaction costs and to evaluate whether they changed over time, we use daily stock prices for the period 31 December 1891 to 31 December 1913 collected from the *Berliner Börsenzeitung* – Germany’s leading financial daily of the pre-1913 period – for a sample of 27 continuously traded corporations from the Berlin stock exchange. The data were obtained from Gelman and Burhop (2008) who construct a daily stock market index for the period 1892-1913.² The sample contains 6,692 daily returns. Descriptive statistics of individual stocks are shown in Table I. The average daily return of an individual stock was slightly above 1 basis point and the average standard deviation with 94 basis points was about a half of the modern stock return volatility, but corresponds to the values reported by Gehrig and Fohlin (2006, p. 10, p. 12) for 1890, 1900 and 1910. Most of the stock returns are negatively skewed; all of them are leptokurtic, somewhat stronger than the modern day stock returns.

To make some statements in how far our sample is representative for the whole universe of stocks traded on Berlin stock exchange, we compare the size of the companies included into the sample to the full cross-section in 1900 (the only data available to us for all listed stocks). From 826 listed companies there is market capitalization data only for 764 companies.³ The aggregate market capitalization of our sample accounts for 16% of the total market capitalization. The average capitalization of all listed stocks with reported data was with 1.1 million Mark, about five times smaller than the average capitalization of the selected 27 companies, which amounted to 5.2 million in 1900. Both a simple t-test of the mean as well as a non-parametric Mann-Whitney test indicate that our sample is biased towards larger stock. In fact, the average size rank (in descending order) of our sample stocks is 170.1 compared to the expected 382.5; 5 out of the 10 largest companies listed on the exchange belong to the sample.

2 Starting point for the index construction was the collection of daily share prices from the *Berliner Börsenzeitung* for a sample of 39 continuously listed non-insurance corporations from the Berlin stock exchange. Insurance companies were excluded from the index since trading in them was heavily restricted. They only issued *vinkulierte Namensaktien*, registered shares with restricted transferability. Then securities with the portion of zero daily returns in the period under study of one third or higher were deleted from the index. 27 corporations remained.

3 Details available on demand

Table I. Distributional properties of stock returns of the Gelman-Burhop-index constituent companies

	Name	Mean (ann.)	Median (ann.)	Max.	Min.	Std. Dev.	Skew- ness	Kur- tosis	Proportion of zero returns	$\rho(1)$	Average Mar- ket Cap. (1,000 M)
1	AG für Anilinfabrikation	0.0678	0.0000	0.1257	-0.2270	0.0082	-3.85	126.25	0.1638	0.0008	2727
2	Allgemeine Elektrizitätsgesellschaft	0.0336	0.0000	0.0526	-0.0611	0.0065	-0.18	11.94	0.0807	0.0820*	14997
3	Berlin-Anhaltische Maschinenbau	0.0134	0.0000	0.1037	-0.0878	0.0078	-0.65	25.62	0.1540	-0.0474*	1179
4	Bochumer Bergwerk (Lit C)	0.2457	0.0000	4.6522	-0.3611	0.0603	68.59	5285.9	0.2192	0.0060	351
5	Deutsche Bank	0.0294	0.0000	0.0333	-0.0544	0.0042	-1.64	24.72	0.1001	-0.0119	32778
6	Dresdner Bank	0.0129	0.0000	0.0446	-0.0554	0.0048	-0.85	17.74	0.1062	0.0230	19931
7	Darmstädter Bank (BHI)	0.0006	0.0000	0.0642	-0.0846	0.0044	-1.21	40.81	0.1903	-0.0231	16689
8	Deutsche Jute Spinnerei und Weberei	0.0465	0.0000	0.0674	-0.1040	0.0079	-0.53	16.89	0.2061	0.0024	430
9	Deutsche Spiegelglas	0.0687	0.0000	0.0921	-0.0838	0.0080	-0.30	18.32	0.1877	0.0716*	643
10	Erdmannsdorfer Spinnerei Gelsenkirchener Bergwerksgesellschaft	0.0001	0.0000	0.1143	-0.0774	0.0106	0.62	13.70	0.2497	-0.0425*	286
11		0.0254	0.0000	0.0484	-0.0858	0.0071	-0.82	13.87	0.0555	0.0240	15580
12	Gerresheimer Glashütten	0.0480	0.0000	0.0739	-0.1208	0.0079	-1.61	33.11	0.2452	-0.0551*	1112
13	Hallesche Maschinenfabriken	0.0266	0.0000	0.1000	-0.2788	0.0093	-6.04	169.13	0.1666	-0.0287*	667
14	Harpener Bergbau AG	0.0212	0.0000	0.0668	-0.0682	0.0075	-0.28	11.77	0.0517	0.0282*	10400
15	Kattowitz AG für Bergbau und Eisen	0.0348	0.0000	0.0609	-0.0603	0.0068	-0.26	13.77	0.1313	0.0402*	4652
16	Maschinenfabrik Kappel	0.0711	0.0000	0.2014	-0.2138	0.0113	-0.61	48.92	0.1725	0.0721*	340
17	Norddeutsche Wollkämmerei	0.0192	0.0000	0.0738	-0.0838	0.0080	0.01	15.22	0.2025	0.0272*	1926
18	Schaffhausen'scher Bankverein	0.0000	0.0000	0.0454	-0.0409	0.0037	0.07	22.24	0.2688	0.1121*	14043
19	Oberschlesische Portland-Cement AG	0.0390	0.0000	0.1267	-0.0943	0.0098	0.72	23.25	0.1515	-0.0238	440
20	Rheinische Stahlwerke	0.0151	0.0000	0.1095	-0.1427	0.0085	-0.77	39.52	0.1230	-0.0337*	3745
21	Rositzer Zuckerfabrik	0.0435	0.0000	0.0833	-0.0826	0.0092	0.05	11.73	0.1413	0.0237	717
22	Chemische Fabrik vormals Schering	0.0162	0.0000	0.0652	-0.0657	0.0083	0.17	10.17	0.1630	0.0610*	1298
23	Schlesische Zinkhütten	0.0336	0.0000	0.1079	-0.0853	0.0066	-0.52	33.99	0.2360	-0.0582*	7947

Name	Mean (ann.)	Median (ann.)	Max.	Min.	Std. Dev.	Skew- ness	Kur- tosis	Proportion of zero returns	$\rho(1)$	Average Mar- ket Cap. (1,000 M)
24 Schlesische Leinen-Industrie	-0.0043	0.0000	0.0703	-0.0679	0.0063	-0.56	21.28	0.3194	-0.1558*	1015
25 Schultheiss Brauerei	0.0097	0.0000	0.0897	-0.0853	0.0057	-1.23	65.22	0.1714	-0.2016*	2522
26 Siemens Glas-Industrie	0.0287	0.0000	0.0438	-0.0576	0.0058	-1.12	19.17	0.1966	-0.0273*	2290
27 Stettiner Chamottewaren	-0.0096	0.0000	0.0804	-0.2538	0.0088	-3.85	126.25	0.1452	0.0281*	2640
Average	0.0347	0.0000	n/a	n/a	0.0094	1.61	231.87	0.1703	-0.0040	5976
Gelman-Burhop index	0.0687	0.0909	0.0296	-0.0562	0.0032	-1.68	30.78	n/a	0.165*	161344
Equally-weighted price index	0.0327	0.0600	0.1780	-0.0325	0.0036	17.80	913.90	n/a	0.167*	161344

Notes: Mean and median returns are presented on the annual basis (mean daily return x300) for illustrative purposes.

* denotes significance of the autocorrelation coefficient on the 5% level.

Although our sample is skewed towards larger companies it spans a wide range of stocks in terms of size, from the largest (Deutsche Bank) to 590th and 495th from the top (Bochumer Bergwerk and Erdmannsdorfer Spinnerei correspondingly). The difference in distribution of logged market capitalization (which is supposed to stay in linear relationship with liquidity, see Gehring and Fohlin (2006) and Section IV) is by far not that striking, with averages of 19.9 and 21.4 for 764 stocks and our sample respectively. The scope of zero returns seems also to be adequate: it ranges from 5% of trading days for Harpener Bergbau to 32% of trading days for Schlesische Leinen, with an average of 17% or approximately 51 days per year. This is slightly less than 22% reported by Gehrig and Fohlin (2006) for their sample of 114 stocks from the Berlin stock exchange in 1900. Hence, our sample can be regarded as representative of stocks traded on Berlin Stock Exchange at the turn of the 20th century except for a small bias towards larger and more liquid stocks.

Furthermore, the dynamic properties support our hypothesis on informed insider trading: for most of the stocks (15 out of 27) we report positive daily return autocorrelation. The latter makes also indirect effective spread measures, based on the bid-ask bounce, inapplicable.

Beyond the reported properties of stock prices we have also obtained dividend data from the *Berliner Börsenzeitung*. As can be seen from the comparison of the average return on Gelman-Burhop performance index to the one of the price index the dividend yield was on average about 3.5% annually.

Moreover, due to withholding of the turnover tax there is data on aggregate annual trade volume of all securities in Imperial Germany, which can serve as a proxy of the overall trading activity (obtained from Wetzel 1996). The aggregate trade volume time series behaves stationary with the approximately same value of stocks traded in 1913 as in 1892 (see Appendix 8).

III. Econometric technique

In an information-efficient stock market, prices of stocks should incorporate new information instantaneously. However on the real-world stock exchanges the presence of transaction costs induces some deviations from such behaviour. Uncovering these deviations and analyzing them allows tracking back full transaction costs.

This idea is exploited in a measure of transaction costs, proposed by Lesmond et al. (1999). The LOT measure reflects the total costs of a roundtrip transaction, which includes not only the difference between bid and ask prices, but also all further expenses carried by the trader, including the price change induced by the trade itself (so called price impact, see Lesmond 2005). The LOT measure is based on the idea that transactions will only occur if the deviation of the market price from the true value of a stock is larger than the costs of a transaction. Thus, there are upper and lower thresholds $-\tau_i^l$ and τ_i^h – such that the measured return r is non-zero only if the true return exceeds the threshold:

- (1a) $r_{i,t} = R_{it}^* - \tau_i^l$ if $r_{i,t}^* < \tau_i^l$
(1b) $r_{i,t} = 0$ if $\tau_i^l < r_{i,t}^* < \tau_i^h$
(1c) $r_{i,t} = R_{it}^* - \tau_i^h$ if $r_{i,t}^* > \tau_i^h$.

The true return depends on the market return $r_{m,t}$ in a linear way: $r_{i,t}^* = \beta_i r_{m,t} + e_{i,t}$.

The estimated difference between the upper and the lower threshold – i.e. τ_i^h less τ_i^l – is a measure of the roundtrip transaction costs.

We use the following maximum likelihood estimator, developed by Lesmond et al. (1999), to estimate the LOT measure:

$$(2) L(\tau_i^l, \tau_i^h, \beta_i, \sigma_i | r_{it}, r_{mt}) = \prod_1 \frac{1}{\sigma_i} \phi \left[\frac{r_{it} + \tau_i^l - \beta_i r_{mt}}{\sigma_i} \right] \times \prod_0 \left[\Phi \left(\frac{\tau_i^h - \beta_i r_{mt}}{\sigma_i} \right) - \Phi \left(\frac{\tau_i^l - \beta_i r_{mt}}{\sigma_i} \right) \right] \\ \times \prod_2 \frac{1}{\sigma_i} \phi \left[\frac{r_{it} + \tau_i^h - \beta_i r_{mt}}{\sigma_i} \right]$$

S.T. $\tau_i^l \leq 0, \tau_i^h \geq 0, \beta_i \geq 0, \sigma_i \geq 0,$

Where $\Phi()$ is the standard normal cumulative distribution function. Region 1 (indicated by “1” below the Π) corresponds to the negative expected latent variable when the observed is nonzero ($\hat{r}_{it}^* < 0$, or equivalently $r_{mt} < 0$ and $r_{it} \neq 0$), region 2 to the positive expected latent variable if the observed is nonzero ($r_{mt} > 0$ and $r_{it} \neq 0$) and region 0 corresponds to the observation with zero observed returns ($r_{it} = 0$). σ_i denotes the root out of the residual variance, measured over the non-zero returns region.

The LOT measure thus includes the bid-ask spread, fees, transaction taxes, costs of information acquirement and processing as well as price impact. Its size should be therefore larger than the regulated costs, i.e., the sum of broker fees, bank fees, and transaction taxes. We calculate this measure for each company and each year, and then provide also aggregated estimates across companies and years.

In this paper we also calculate standard errors and confidence intervals for the transaction cost estimates, which is novel to the literature. It allows assessing the credibility of the estimates and inferring the significance of cross-section and time-series differences. We obtain standard errors for each stock i and year t from the standard expression:

$$(3) \sigma(S_{it}) = \sqrt{\text{var}(\tau_{it}^h - \tau_{it}^l)} = \sqrt{(\text{var}(\tau_{it}^h) - 2 \text{cov}(\tau_{it}^h, \tau_{it}^l) + \text{var}(\tau_{it}^l))},$$

where $\text{var}(\tau_{it}^h)$, $\text{var}(\tau_{it}^l)$ and $\text{cov}(\tau_{it}^h, \tau_{it}^l)$ are the elements of the coefficient variance-covariance matrix, yielded by maximum likelihood estimation in (2). To obtain standard errors of annual averages we take into account possible cross-correlations of stock returns:

$$(4) \sigma(\bar{S}_t) = \sqrt{\frac{\Sigma' \cdot \Omega \cdot \Sigma}{27^2}},$$

where $\Sigma' = (\sigma(S_{1t}) \cdots \sigma(S_{27t}))$ is a row vector with standard errors for each stock obtained for the year from (3) and Ω is a 27x27 correlation matrix of residuals from the limited dependent variable regressions, estimated by (2). For the standard errors of company transaction costs averages (4) can be simplified, as we can assume independence of estimates across time:

$$(5) \sigma(\bar{S}_i) = \sqrt{\frac{\Sigma'_i \cdot \Sigma_i}{22^2}},$$

where $\Sigma'_i = (\sigma(S_{i1892}) \cdots \sigma(S_{i1913}))$ is a row vector with standard errors for the stock i obtained for years 1892-1913. Confidence intervals are then estimated in a standard way under the assumption of normality of estimates.

However, the precision of LOT transaction costs estimates relies on the explanatory power of the market model for stock returns. Thus, if further information sources or factors, such as returns on Fama-French (1993) small minus big (SMB) and high minus low (HML) factor portfolios have significant influence on individual stock returns, effective transaction costs may be substantially under- or overestimated.⁴ Yet, since the LOT-measure proved to be a good proxy for transaction costs in modern financial markets (see Goyenko et al. 2009; Lesmond 2005) we see it as justified to use it for the historical data in our study.

As we find considerable differences in transaction costs across companies, similar to Gehrig and Fohlin (2006), we run cross-section regressions of estimated average transaction costs on a set of explanatory variables:

$$(6) \bar{S}_i = \alpha + \beta'X_i + \varepsilon_i,$$

where X denotes a vector of explanatory variables and β a vector of corresponding coefficients. However, as we observe remarkable time variation of transaction cost estimates we also run a panel regression:

$$(7) S_{it} = \alpha + \beta'X_{it} + \mu_i + \lambda_t + v_{it},$$

where μ_i denote cross-sectional individual effects, λ_t denote year effects and v_{it} is an idiosyncratic error term.

We rely on the standard technique in the asset pricing literature, the Fama-MacBeth (1973) regression, when analyzing the impact of transaction costs on the cross-sectional variation of returns. It is based on the assumption, that expected returns of stocks are fully described by the linear combination of risk premia and factor loadings for all relevant factors:

$$(8) E[Z_i] = \lambda'B_i,$$

whereby $Z_{it} = r_{it} - r_{ft}$ denotes excess return, λ' is a transposed vector of risk-premia and B_i is a vector of factor loadings or risk characteristics of company i . Given the values of factor

4 We are grateful to Christian Julliard for this comment.

loadings for each stock in each period the risk premia are estimated running T cross-section regressions (one for each period) and averaging the estimates:

$$(9) Z_{it} = \lambda'_t B_{it}$$

$$(10) \bar{\lambda} = \frac{1}{T} \sum_{t=1}^T \hat{\lambda}_t$$

The corresponding standard errors for each k -th element of the risk-premia vector are calculated from the corrected time variance of the estimated premia:

$$(11) \text{var}[\lambda_{kt}] = \frac{1}{T} \sum_{t=1}^T (\hat{\lambda}_{kt} - \bar{\lambda}_k)^2$$

$$(12) \text{stderr}[\bar{\lambda}_k] = \sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right) / T},$$

where z_{mt} denotes the excess return of the market index. For the risk factor k to be priced the corresponding risk premium should be significantly different from zero.

To obtain the illiquidity risk factor loadings β_i^{ll} we calculate the sensitivity of unpredicted transaction costs to market movements using the following linear regression:

$$(13) \tilde{s}_{it} = \omega + \beta_i^{ll} \cdot r_{mt} + u_{it}.$$

Unpredicted illiquidity is defined as the residual from a second order panel vector-autoregression of transaction costs and annual stock returns (without dividends):

$$(14) X_{it} = A_0 + A_1 X_{it-1} + A_2 X_{it-2} + C z_{it-1} + U_{it},$$

where $X_{it} = \begin{pmatrix} r_{it} \\ s_{it} \end{pmatrix}$ is the vector of dependent variables and $U_{it} = \begin{pmatrix} \epsilon_{it}^r \\ \tilde{s}_{it} \end{pmatrix}$ the residual vector; z_{it}

denotes the fraction of the market capitalisation of the company in the aggregate market capitalisation. A_0 and C are vectors and A_1 and A_2 matrices of coefficients which are kept invariable across companies.

IV. Results

1. Estimated transaction costs

Table II presents the averages across all shares of the annual LOT measure of round trip transaction costs as well as the average for the full sample period 1892-1913.

The transaction costs at the Berlin Stock Exchange varied between 0.66 percent (in 1906) and 1.68 percent (in 1901). The transaction costs were positive for any randomly chosen yearly

period and they were always higher than the lower bound of the regulated fees. The average transaction costs amounted to 0.97 percent. Therefore, we broadly confirm the result presented by Gehrig and Fohlin (2006), who estimated an average LOT measure of 0.71 percent for the four benchmark years 1880, 1890, 1900, and 1910. Moreover, we find our transaction cost measures rather precisely estimated, with 95% confidence bounds being about ± 10 basis points for most of the years. Significant illiquidity increases are revealed in 1901, 1910, 1912 and 1913 relative to the respective previous years. Significant transaction cost decreases appear in 1894, 1902 and in 1911 as compared to the respective previous years.⁵

TABLE II: ANNUAL AVERAGE OF TRANSACTION COSTS

Year	LOT	Std. error	95% confidence interval	
			lower bound	upper bound
1892	1.454	0.062	1.333	1.575
1893	1.584	0.067	1.452	1.716
1894	1.072	0.050	0.975	1.169
1895	0.925	0.046	0.835	1.015
1896	0.805	0.039	0.729	0.881
1897	0.814	0.041	0.735	0.893
1898	0.908	0.044	0.821	0.995
1899	0.878	0.045	0.789	0.967
1900	1.029	0.057	0.917	1.141
1901	1.678	0.073	1.534	1.822
1902	0.977	0.224	0.537	1.417
1903	0.848	0.040	0.769	0.927
1904	0.825	0.041	0.744	0.906
1905	0.696	0.036	0.625	0.767
1906	0.658	0.034	0.591	0.725
1907	0.775	0.042	0.693	0.857
1908	0.846	0.045	0.757	0.935
1909	0.731	0.040	0.653	0.809
1910	1.039	0.046	0.949	1.129
1911	0.713	0.036	0.642	0.784
1912	0.883	0.044	0.797	0.969
1913	1.124	0.049	1.028	1.220
Average	0.966	0.012	0.942	0.990

Own calculations based on daily returns for 27 stocks for the period 1892-1913.

Expressed in percent of share price, equally weighted averages. Four outliers were dropped. Standard errors are calculated taking into account cross-correlations between stocks, see (3)-(4). Confidence interval is given by $\bar{S}_{LOT,t} \pm 1.96s.e.(\bar{S}_{LOT,t})$

5 The explosive increase in the standard deviation of the transaction cost estimates in 1902 is caused by the untypical behavior of Bochumer Bergwerk stock returns.

It may come as a surprise that transaction costs were rather stable at the German stock exchange over the last century. We find that the 27 companies under study at the turn of the twentieth century had, on average, lower transaction costs than the 2nd tier German blue chips at the turn of the twenty-first century: Applying the same technique to 47 MDAX companies for 1995-2009 yields an average LOT measure of 2.6 percent.⁶

Evidence for other modern stock markets supports the impression that transaction costs were quite low at the Berlin Stock Exchange a century ago. Goyenko et al. (2009) document LOT measures for the Dow Jones Industrial Average index of 0.6 percent for the mid 1970s and 1980s which is comparable to our results for the Berlin Stock Exchange index in mid 1900s. Very advantageous is the comparison to the modern emerging markets: Stocks in the Gelman-Burhop (2008) index have lower transaction costs according to the LOT measure than any of the 31 emerging markets in the 1990s, covered in the study of Lesmond (2005). Their average transaction costs range from 2.3 percent for Taiwan to 18 percent for Russia. This reported superiority of the Berlin Stock Exchange at the turn of the 20th century can be interpreted as evidence in favour of higher liquidity of call auctions compared to modern continuous trading (Pagano and Roell, 1996). Some portion of the observed higher liquidity could be also due to the broader sample of securities per country, analyzed in Lesmond (2005).

2. Explaining transaction costs

Transaction costs varied across companies (see Appendix 1). Whereas textile companies, such as Deutsche Jute Spinnerei und Weberei and Erdmannsdorfer Spinnerei report LOT measures of 1.1 percent and 1.7 percent – which could be found also for median modern Chinese stock (Lesmond, 2005) – the transaction costs of banking sector stocks like Deutsche Bank (0.38 percent) and Dresdner Bank (0.45 percent) is on the same level with Dow Jones companies in the 1980s and 1990s (Goyenko et al., 2009). These deviations however, cannot be attributed fully to industrial differences: companies included into the index stemming from the banking sector have a much higher market capitalization, e.g., the value of Deutsche Bank was on average 114 times the value of Erdmannsdorfer Spinnerei.

The explanation may rather have informational origins: assuming the same share of trading relative to market capitalization across companies, the volume of trade for large companies was higher, allowing market makers to incorporate new information faster. Furthermore, large companies had probably better newspaper and analyst coverage, providing more thorough information to investors, thus decreasing information asymmetry. In turn, lower information asymmetry lessened the proportion of informed trading and thus provided for lower transaction costs.

6 The results are available upon request from the authors.

Some evidence for the size hypothesis can be obtained from a simple cross section regression of average transaction costs on the log of the market capitalization. One should nevertheless be cautious as Amihud and Mendelson (1986) reveals the possibility of a reversed causal relationship: transaction costs can raise expected returns and thus reduce the market capitalization of a company. To avoid the endogeneity problem and to ensure the pre-determinacy we use the market capitalization of 1892 (which is measured at the beginning of the year) to explain company transaction costs averaged over the twenty-two year sample. For twenty-six companies we obtain (standard errors of estimates are in parenthesis):⁷

$$(15) \quad \overline{S_i^{LOT}} = 4.92 - 0.19 \cdot \ln(MC_i^{1892}) + \hat{e}_i, \\ R^2 = 0.64, \quad \hat{e}_i \approx (0, 0.20^2)$$

All estimates are highly significant and support the hypothesis that market capitalization decreases transaction costs: raising the market capitalization by 2.3 million Mark (what corresponds to a one unit change of log market capitalization at the mean of the variable) leads to 0.19 percentage points lower transaction costs.⁸ In addition, market capitalization explains almost two thirds of the inter-company transaction cost variation in our sample.⁹

Another issue possibly relevant for transaction costs is tick size, which was 0.05 percent of the nominal (face) value of a stock. Thus, our transaction costs measure expressed in percent of the price could be higher for stocks with lower value. However, including the (log) price level at the beginning of the sample does not significantly help to explain the cross-section of transaction costs: While other coefficients barely change, the coefficient in front of the log price level is of the expected sign, but insignificant.

$$(16) \quad \overline{S_i^{LOT}} = 5.15 - 0.19 \cdot \ln(MC_i^{1892}) - 0.05 \ln(P_i^{1892}) + \hat{e}_i, \\ R^2 = 0.64, \quad \hat{e}_i \approx (0, 0.20^2)$$

Given that prices fluctuate considerably during the 22 year period, using the first year market capitalization and price level could be insufficient to uncover the hypothesized relationship. Therefore, we run regressions of transaction costs on market capitalization and on both market capitalization and price level in a balanced panel set-up with individual effects. We as-

7 We exclude Bochumer Bergwerk henceforth from the analysis, as it has unusually high transaction costs due to several months long periods of non-trading.

8 Our estimation coincides with the one reported by Gehrig and Fohlin (2006) for the year 1900 for the log of the book value and is considerably close to their results for 1890 and 1910.

9 Using equation (15) we could address the size bias, in the previous sub-section: Since the average (log) market cap in our sample is about 1.5 units higher than the population average in 1900, the population average transaction costs can be expected to be about 29 basis points higher than reported in Table 2. Still they turn out to be lower than they are in modern emerging markets and for constituents of 2-tier developed market indices.

sume that trade volume is proportional to market capitalization not only across companies but also across time. If higher trading volume of larger firms is associated with lower transaction costs, then we should find the same relationship in the panel regression as in the cross section regressions (15) and (16). As market capitalization is clearly non-stationary over the 22 year sample, we use the fraction of the overall market capitalization contributed by each company. Furthermore we include time-effects to account for changes in the overall market capitalization. As an alternative measure we include the aggregate annual trade volume of all securities in Imperial Germany per year, which under our assumption of proportionality, should capture changes in the overall market capitalization. In order to treat the non-stationarity of log price levels we take first differences and obtain returns (neglecting dividends). To address the previously outlined possible problem of reversed causality we use lagged log price changes. Since market capitalization is reported for the beginning of each year, we do not face possible reverse causality with regard to this variable. We perform random effects and fixed effects estimations, while the result of the Hausman test allows using random effects. Since transaction costs are believed to be rather persistent (Bekaert et al. 2007, Amihud 2002), we use White period standard errors, which account for clustering by stocks. Furthermore, we directly address the issue of persistence including the lagged dependent variable as a regressor in our last specification. To avoid the problem of collinearity of the lagged dependent variable and the error term we transform all variables using forward orthogonal deviations, following Arellano and Bover (1995) and then apply general method of moments estimation for dynamic panel data (further DPD-GMM).¹⁰

Table III: Panel regression to explain the size of transaction costs

	(1) FE	(2) FE	(2b)FE	(3) FE	(4) FE	(5) RE	(6) GMM
Constant	1.00*** (0.06)	0.96*** (0.05)	0.97*** (0.07)	0.94*** (0.07)	1.79*** (0.19)	1.85*** (0.20)	
S_{it-1}							0.37*** (0.08)
$MC_{it}/\sum MC_{it}$	-3.12** (1.57)	-2.59** (1.19)	-2.56 (1.74)	-2.59 (1.91)	-2.58 (1.59)	-4.29*** (0.73)	-2.34*** (0.59)
$\Delta \ln P_{it-1}$		-0.25 (0.16)	-0.55*** (0.16)	-0.34** (0.16)	-0.45*** (0.16)	-0.44*** (0.16)	-0.36*** (0.10)
$\ln TV_t$					-0.20*** (0.05)	-0.20*** (0.05)	-0.07* (0.04)
t_{1901}				0.26*** (0.08)			
t_{1913}				0.25*** (0.07)			
Time effects	Y	Y	N	N	N	N	N
Firm effects	Y	Y	Y	Y	Y	N	Y
R^2	0.56	0.60	0.45	0.48	0.46	0.27	0.63

Estimates of LS individual effects models as well as DPD-GMM for the transaction costs (LOT measures) for the sample period from 1892 to 1913 for the panel of 26 companies of the type: $S_{it} = \alpha + \beta' X_{it} + \mu_i (+\lambda_t) + v_{it}$. White period standard errors are reported in parenthesis. Values marked with ***, ** and * are significant at 1%, 5% and 10% level respectively. R^2 is calculated as one minus the fraction of the residual variance to the variance of the dependent variable.

Columns 1-2 and 5-6 in Table III support the hypothesis: relative market capitalization has a significant negative impact. According to column 1, if the share of market capitalization in the

10 The lagged LOT measure is instrumented by its second lag, all other explanatory variables are instrumented by themselves.

index increases by one standard deviation (5 percentage points) transaction costs decrease by 0.16 percentage points. The statistical significance of the size variable depends on the specification: it is the highest in the random effects specification (column 5) but diminishes in fixed effects specifications without time effects (columns 2b-4). The reasons are probably twofold: First, market capitalization is less variable in time than in cross section: In fact, the between variance of the market capitalization variable constitutes about 96% of its overall variance. Second, as we omit time effects in fixed effects specifications (columns 2b-4) standard errors surge. Including the omitted lagged dependent variable in the DPD-GMM specification (column 6) makes the estimates more efficient and substantially reduces their variance, which makes them significant on the 5% level.¹¹

Previous year log price changes have a negative impact on transaction costs, which is significant in all specifications without time-effects (columns 2b-6 of Table III). This result supports the findings of Griffin et al. (2004) and Bekaert et al. (2007), who find that returns help predicting liquidity on modern financial markets. Furthermore, time effects play a substantial role: they explain about 15 percent of the transaction costs variation (compare explanatory power in columns 2 and 2b). About one-seventh of the explanatory power of time can be attributed to two years: 1901 and 1913 (see column 3) – which are known to have caused worsening information efficiency (Gelman and Burhop, 2008). In 1901 the bankruptcy of Leipziger Bank, one of Germany's largest banks, caused a stock exchange turmoil and possibly high degree of uncertainty about fair prices of shares which forced speculative traders to act more conservative, thus reducing liquidity. In 1913, the fear of a Balkan war led to similar effects on the financial market.

The log of the German securities trading volume during year t , TV_t , (see columns 4-6 of Table 3) as an indicator of the overall trading activity picks up only about 1% of the variance of transaction costs, but is highly significant. The lower R^2 of the random effects vs. the fixed effect specification suggest that about 16 percent of the variance is explained by cross-sectional individual firm effects.

Standard fixed and random effects regression residuals exhibit a strong and highly significant autocorrelation.¹² In fact, DPD-GMM in column 6 reveals a highly significant autoregressive coefficient for illiquidity, supporting earlier empirical evidence of the persistence of transaction costs (Bekaert et al. 2007, Amihud 2002). Our estimates of the impact of previous period returns, company size and turnover stay about the same compared to the corresponding static specification (column 4), but the standard errors decrease substantially for size and previous period returns. This implies that neglecting the dynamics does not substantially bias the results, but does worsen the efficiency of estimates.

11 The result of a negative correlation of transaction costs with size proves to be rather stable over time: An equivalent to column 1 panel regression of transaction costs of 47 MDAX stocks over 1999-2009 on the fraction of overall market capitalization yields a coefficient of -5.34, which is also significant on the 10% level and explains about 6% of the variation of transaction costs. Results are available on demand.

12 Details are available on demand

Hence, we find some support for increasing illiquidity with rising information asymmetry or a larger information-to-noise ratio: In particular, a decline in company size leads to higher illiquidity. The evidence is weaker for the relevance of corporate distress periods for illiquidity. Contemporaneous backdrops in trading activity and crises deteriorate liquidity significantly.

3. Transaction costs, liquidity, and asset prices

The large dispersion of transaction costs should be reflected in asset pricing. Here we test three hypotheses of the liquidity impact.

First, as Amihud and Mendelson (1986) noted, given the set of investment opportunities, investors should avoid assets which have lower liquidity yielding same returns. This should in the long run decrease the price of such securities and raise their return. Therefore, in the long run one should find a positive relation between transaction costs and expected returns in the cross-section.

Second, we hypothesize that noise traders (ordinary, non-informed investors) should in the long run avoid stocks with pronounced informed trading (proxied by positive autocorrelation), since in these stocks noise traders' orders are more likely to get matched if the information is unfavorable. Thus, there should be a premium for informed trading on top of the transaction costs.

Third, inline with Liquidity-adjusted CAPM of Acharya and Pedersen (2005), we expect a premium for liquidity risk. Acharya and Pedersen (2005) suggest three liquidity risk channels: the covariance of individual stock liquidity with market liquidity, the covariance of individual stock returns with market liquidity and the covariance of individual stock liquidity with market returns. As the authors report strong correlation between these measures and with the level of illiquidity, we decide to use only the channel with the strongest economic effect, namely the sensitivity of individual stock illiquidity to market return (Acharya and Pedersen 2005, p. 398).¹³

To perform the tests we analyze excess returns, calculated as total returns (price changes plus dividends) less the risk free rate. Including dividends is important as the companies may compensate investors with higher dividends for lower prices. In line with the asset pricing literature, we use monthly return data. We run Fama-MacBeth (1973) regressions with Shanken (1992) corrections for the traditional CAPM and several multifactor extensions, including

13 Gernandt et al. (2011) choose another risk channel – sensitivity of individual stock returns to market liquidity shocks – and find no significant impact on asset pricing on the contemporary (1901-1919) Swedish stock market.

transaction costs, serial daily return autocorrelation, liquidity risk beta and we control for size.¹⁴

Table IV. Results of cross-sectional asset pricing regressions

	(1)	(2)	(3)	(4)	(5)
Constant	.0018 (.0014)	-.0024 (.0021)	.0024 (.0107)	.0034 (.0108)	-.0034 (.0023)
Market beta $\bar{\lambda}_\beta$	-.0003 (.0019)	.0016 (.0020)	.0013 (.0021)	-.0001 (.0022)	.0024 (.0022)
Transaction cost lagged $\bar{\lambda}'_{TC}$.3266** (.1324)	.3068* (.1773)	.3055* (.1771)	.2490* (.1413)
Size $\bar{\lambda}_s$			-.0002 (.0004)	-.0002 (.0004)	
Autocorrelation $\bar{\lambda}_\rho$.0105* (.0058)	
Illiquidity risk premium $\bar{\lambda}_l$					-.0020*** (.0007)
Average R^2	0.07	0.12	0.16	0.20	0.17
# of stocks	26	26	26	26	26
# of cross-sections T	264	252	252	252	252

Estimates of the Fama-MacBeth (1973) regressions for the sample period from 1892 to 1913 for 26 companies. Reported coefficient values $\bar{\lambda}_k$ are averages of 264 (252 for columns (2)-(5)) regression estimates of the type: $Z_{it} = \alpha_i + \lambda'_i B_{it} + u_i$, where λ'_i denotes the transposed vector of risk premia and B_{it} denotes the vector of risk factor loadings, which serve as explanatory variables in each cross section. Standard errors are calculated as

$$\sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right)} / T$$

, according to the Fama-MacBeth (1973) procedure with Shanken (1992) correction, and are reported in parenthesis. Values marked with ***, ** and * are significant at 1%, 5% and 10% level respectively. Average R^2 is an arithmetic mean of R^2 for each cross-section.

The liquidity risk beta is calculated as a regression slope of unpredicted individual illiquidity shocks on market return shocks. Unpredicted illiquidity shocks are residuals of a panel VAR(2) of annual returns and illiquidity measures (analog to Bekaert et al., 2007). As the risk free rate proxy we use the money market rate obtained from the NBER (series: 13018). Size is the log of market capitalization and varies on the annual basis. Transaction costs are our LOT estimates, which also vary yearly. Market betas and the first-order autocorrelation coefficient of daily price percentage changes are constant for each company throughout the sample. We also include a constant as we do not demean the explanatory variables.

As expected, the premium for transaction costs is significant and positive in all specifications (see Table IV). A one percentage point higher transaction cost (which is equivalent to moving from the most liquid stocks to the bottom of our sample, see Appendix 1) raises expected

14 We are aware of possible within firm and within month error clustering, as outlined in Petersen (2009). Having a considerably greater time dimension than cross-section dimension makes the within month clustering the primary problem. However, as Petersen (2009) shows, Fama-MacBeth (1973) technique is able to address it adequately. Turning to within firm clustering, it could be a problem in our data at a first glance, as our right hand side variables are very persistent, since transaction costs change only yearly and betas and autocorrelation coefficient stay constant throughout the sample. But our dependent variable – return – is not persistent at all, thus yielding slightly negatively correlated residuals and thus nullifying the problem of underestimation of standard errors.

monthly return by 25 to 33 basis points or 3% to almost 4% annually, depending on the specification. This range covers the 3.5% annual premium obtained by Acharya and Pedersen (2005) for the US value-weighted portfolios in 1964-1999. The illiquidity premium estimates also suggest an average holding period of three to four months, which is required for returns net of transaction costs to become equal across different stocks.

Furthermore, our proxy for insider trading is positive with tendency to significance (see Table IV, column 4): stocks, which daily price percentage changes are positively autocorrelated have on average higher expected returns. This implies that informed trading rather deteriorates the value of a company and for sure does not increase wealth. The effect of investing in stock with the highest autocorrelation coefficient of 0.11 instead of in stock with the lowest one of -0.20 would lead to an increase in expected monthly returns by 33 basis point or about 4% annually.

The illiquidity risk premium is, as predicted by theory, negative. The sign is due to undesired negative sensitivity of illiquidity to market movements: negative market shocks increase illiquidity and vice versa. Thus, the expected return is higher for those stocks, which liquidity deteriorates during market downturns. Our result for the premium on the individual illiquidity sensitivity to market returns is statistically highly significant, inline with results obtained by Lee (2011) for a large battery of stocks from 54 countries and 1988-2007 sample period (Lee 2011, Tables 3-4), whereas Acharya and Pedersen (2005) for modern US data fail to find significant premium for this liquidity risk channel alone. The economic extent of the liquidity risk effect in our data is rather strong: if sensitivity to liquidity risk moves from 0 to -0.7 (about one standard deviation) the expected return increases by 14 basis points per month. The difference between maximum and minimum liquidity risk sensitivity is about four times as large and would lead to 55 basis points increase (see Appendix 7). Annualized the full range move in autocorrelation would yield a 6.6% return increase, which by far exceeds the overall liquidity risk effects of 1.1%, reported by Acharya and Pedersen (2005, p.398) for US and 1.5% reported by Lee (2011) for the global market, but is comparable to the economic effect of 5.6% for modern emerging markets (Lee 2011, p. 146).

Moreover, our results reveal that CAPM does not hold since the market risk premium is insignificant in all four specifications, which is consistent with empirical results of Gernandt et al. (2011) for the contemporary (1901-1919) Swedish stock market and with results of Acharya and Pedersen (2005) and Lee (2011) for modern US and global data. In addition, there is no significant size effect, which supports Lee (2011) results for modern stocks except emerging markets. Seemingly, size to a large extent proxies liquidity risk, which is much better captured here by transaction costs. However, the inclusion of the size variable increases, due to correlation with transaction costs, the standard error of the latter coefficient, which leads to some loss in significance of illiquidity.

The results of this subsection suggest, that liquidity solely drives asset pricing and causes expected return variation of the magnitude of 7% (liquidity level plus informed trading differ-

ences) to 9.6% (liquidity level plus liquidity risk differences) per year. It shows that investors value liquidity even more in a more efficient call auction market design than what has been reported for continuous trading (or combined) mechanisms for modern US or global markets (Acharya and Pedersen 2005, Lee 2011). Opposite to Bernhardt et al., (1995) and Leland (1992) we find that active informed trading seems to destroy company value by considerable amount (about 4% p.a.).

V. Robustness checks

The results of the previous section rely upon the assumption that the standard LOT measure is a good proxy of illiquidity. To address concerns that this is not the case, especially if the standard estimate of LOT may be distorted by a falsely specified function of latent returns (market model), we repeat the tests using an extended market model, which includes excess returns to the SMB portfolio as a risk factor.

To construct the SMB portfolio we form “small” and “big” portfolios, which are equally weighted portfolios of the smallest five and largest five companies respectively. Portfolios are rebalanced at the beginning of each year based on 1 January market capitalization. SMB factor return is calculated as a return of a portfolio with a unit long position in the “small” portfolio and a unit short position in the “big” portfolio. The list of the constituent companies is in the Appendix 5. In fact, the augmented market model does explain non-zero returns sufficiently better than a simple market model: average R-squared increases from 16% to 25%. The fit increases tremendously for small stocks (see Appendix 2 and 6). The transaction cost estimation is then performed maximizing

$$\begin{aligned}
(17) L(\tau_i^l, \tau_i^h, \beta_{M,i}, \beta_{SMB,i}, \sigma_i | r_{it}^*, r_{mt}, r_{SMB,t}) &= \prod_1 \frac{1}{\sigma_i} \phi \left[\frac{r_{it} + \tau_i^l - \beta_{M,i} r_{mt} - \beta_{SMB,i} r_{SMB,t}}{\sigma_i} \right] \\
&\times \prod_0 \left[\Phi \left(\frac{\tau_i^h - \beta_{M,i} r_{mt} - \beta_{SMB,i} r_{SMB,t}}{\sigma_i} \right) - \Phi \left(\frac{\tau_i^l - \beta_{M,i} r_{mt} - \beta_{SMB,i} r_{SMB,t}}{\sigma_i} \right) \right] \\
&\times \prod_2 \frac{1}{\sigma_i} \phi \left[\frac{r_{it} + \tau_i^h - \beta_{M,i} r_{mt} - \beta_{SMB,i} r_{SMB,t}}{\sigma_i} \right] \\
\text{S.T. } \tau_i^l &\leq 0, \tau_i^h \geq 0, \beta_{M,i} \geq 0, \sigma_i \geq 0.
\end{aligned}$$

Here region 1 (indicated by “1” below the Π) corresponds to the negative expected latent variable when the observed is nonzero ($\hat{r}_{it}^* < 0$, or equivalently $\beta_{M,i} r_{mt} + \beta_{SMB,i} r_{SMB,t} < 0$ and $r_{it} \neq 0$), region 2 – to the positive expected latent variable if observed is nonzero ($\beta_{M,i} r_{mt} + \beta_{SMB,i} r_{SMB,t} > 0$ and $r_{it} \neq 0$), and region 0 corresponds to the observation with zero observed returns ($r_{it} = 0$). Note that since we cannot impose any restrictions on sensitivity to

SMB risk, we have to pre-estimate the latent variable to define the regions and to solve the likelihood iteratively.

Obtained transaction costs are generally rather similar: the correlation with the LOT measure in a panel set-up is about 97%, the average for the period is very close to the LOT average with 0.94 percentage points (see Appendix 3 and 4). However, the Augmented LOT measure has higher variance in the cross-section: transaction costs for low-liquidity stocks tend to be higher and for the high-liquidity ones tend to be lower.

Qualitative findings on liquidity and transaction costs drivers remain the same: liquidity is lower for small and distressed stocks and declines in crises periods (Table V).

Table V: Panel regression to explain the size of transaction costs – Alternative LOT

	(1) FE	(2) FE	(2b) FE	(3) FE	(4) FE	(5) RE	(6) GMM
Constant	0.983*** (0.065)	0.938*** (0.059)	0.942*** (0.067)	0.916*** (0.066)	1.831*** (0.292)	1.901*** (0.289)	
S_{it-1}							0.355*** (0.107)
$MC_{it} / \sum MC_{it}$	-3.524** (1.631)	-2.927** (1.486)	-2.887* (1.693)	-2.930* (1.669)	-2.913* (1.679)	-4.670*** (0.783)	-2.703*** (0.636)
$\Delta \ln P_{it-1}$		-0.1631 (0.147)	-0.475*** (0.135)	-0.265* (0.143)	-0.362*** (0.139)	-0.358*** (1.39)	-0.310*** (0.104)
$\ln TV_t$					-0.216*** (0.069)	-0.217*** (0.069)	-0.077* (0.045)
t_{1901}				0.266*** (0.082)			
t_{1913}				0.230*** (0.080)			
Time effects	Y	Y	N	N	N	N	N
Firm effects	Y	Y	Y	Y	Y	N	Y
R^2	0.54	0.58	0.44	0.45	0.45	0.27	0.76

Estimates of LS individual effects models as well as GMM for the augmented LOT measure per year for the sample period from 1892 to 1913 for the panel of 26 companies of the type: $S_{it} = \alpha + \beta' X_{it} + \mu_i (+\lambda_t) + v_{it}$. White period standard errors are reported in parenthesis. Values marked with ***, ** and * are significant at 1%, 5% and 10% level respectively. R^2 is calculated as one minus the fraction of the residual variance to the variance of the dependent variable.

Asset pricing analysis with Augmented LOT supports our previous findings (see Table VI): there is a substantial liquidity premium of about the same magnitude, which is statistically significant at least on the 10% level if we do not include the size characteristics. The informed trading proxy tends to lead to higher expected returns, and the liquidity risk premium is of a similarly high economic and statistic significance. Thus, the choice of possibly incomplete model for the latent returns seems not to distort our finding on drivers of illiquidity as well as on the impact of illiquidity and informed trading on the prices of assets.

Table VI: Results of asset pricing regressions – Alternative LOT

	(1)	(2)	(3)	(4)	(5)
Constant	.0018 (.0014)	-.0013 (.0020)	.0053 (.0106)	.0059 (.0107)	-.0038 (.0024)
Market beta $\bar{\lambda}_\beta$	-.0003 (.0019)	.0010 (.0021)	.0011 (.0021)	-.0002 (.0023)	.0034 (.0024)
Augm. LOT lagged $\bar{\lambda}'_{TC}$		0.2927** (0.1277)	0.2498 (0.1747)	0.2567 (0.1737)	.2587* (.1350)
Size $\bar{\lambda}_S$			-.0003 (.0005)	-.0003 (.0005)	
Autocorrelation $\bar{\lambda}_\rho$.0099* (.0060)	
Illiquidity risk pre- mium $\bar{\lambda}_l$					-.0018*** (.0007)
Average R^2	0.07	0.12	0.16	0.19	0.17
# of stocks	26	26	26	26	26
# of cross-sections T	264	252	252	252	252

Estimates of the Fama-MacBeth (1973) regressions for the sample period from 1892 to 1913 for 26 companies. Reported coefficient values $\bar{\lambda}_k$ are averages of 264 (252 for columns (2)-(4)) regression estimates of the type:

$Z_{it} = \alpha_i + \lambda'_i B_{it} + u_i$, where λ'_i denotes the transposed vector of risk premia and B_{it} denotes the vector of risk factor loadings, which serve as explanatory variables in each cross section. Standard errors are calculated as

$$\sqrt{\text{var}[\lambda_{kt}] \cdot \left(1 + \frac{\bar{z}_{mt}^2}{\text{var}[z_{mt}]}\right)} / T$$

, according to the Fama-MacBeth (1973) procedure with Shanken (1992) correction, and are reported in parenthesis. Values marked with ***, ** and * are significant at 1%, 5% and 10% level respectively. Average R^2 is an arithmetic mean of R^2 for each cross-section.

VI. Conclusion

We find an early call auction market at the Berlin Stock Exchange about as liquid as modern stock exchanges with transaction costs averaging about one percent between 1892 and 1913 according to the measure proposed by Lesmond et al. (1999). Thus, transaction costs a century ago were quite similar to today's cost, possibly due to the efficient trading design. We find some indirect but robust evidence that the ratio of informed to uninformed investors drives liquidity: we find significantly higher transaction costs for cases, where this ratio is believed to be higher: for small and distressed stocks, for the periods of rapid and stark market downturns.

Even for this rather liquid call auction market we find that liquidity seems to be the main driver of asset pricing: we find economically and statistically significant liquidity level and liquidity risk premia, whereas market risk has no impact. Therefore, we support the literature on the modern markets with continuous trading and can conclude that stock liquidity seem to matter irrespective of market design and how liquid the market is on average. This has implications for financial authorities and market agents. The former should take measures of enhancing liquidity – either through a more efficient market design or by attracting trading activity – which could be rather costly but still welfare enhancing. The latter, e. g. portfolio investors, have to consider (relative) illiquidity level and risk even on highly liquid markets.

We find also some weak evidence of an informed trading premium, thus active insider trading seems to deteriorate value on top of general illiquidity. This would imply that large stock holders should abstain from frequent trading on insider information to maximize their end-of-the-holding-period wealth.

Hence the impact of informed trading on illiquidity and on asset pricing deserves further closer study. In this context future research could answer limitations of the current paper, in the first place constructing alternative proxies for informed trading from the data available for modern stock markets.

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Appendix 1: Average transaction costs of corporations

Number	Name	Average LOT mea- sure	Standard error	95% confidence interval	
				Lower	upper
1	AG für Anilinfabrikation	0.943	0.026	0.892	0.994
2	Allgemeine Elektrizitätsgesellschaft	0.520	0.020	0.482	0.558
3	Berlin-Anhaltinische Maschinenbau	0.902	0.025	0.854	0.950
4	Bochumer Bergwerk (Lit C)	3.164	0.269	2.637	3.691
5	Bank für Handel und Industrie	0.543	0.014	0.516	0.570
6	Deutsche Bank	0.384	0.016	0.353	0.415
7	Dresdner Bank	0.446	0.015	0.417	0.475
8	Deutsche Jute Spinnerei und Weberei	1.109	0.025	1.060	1.158
9	Deutsche Spiegelglas	1.097	0.027	1.045	1.149
10	Erdmannsdorfer Spinnerei	1.689	0.035	1.621	1.757
11	Gelsenkirchener Bergwerksgesellschaft	0.427	0.021	0.387	0.467
12	Gerresheimer Glashütten	1.284	0.029	1.228	1.340
13	Hallesche Maschinenfabriken	1.112	0.029	1.054	1.170
14	Harpener Bergbau AG	0.425	0.022	0.383	0.467
15	Kattowitzer AG für Bergbau und Eisen	0.667	0.020	0.627	0.707
16	Maschinenfabrik Kappel	1.239	0.033	1.174	1.304
17	Norddeutsche Wollkämmerei	1.135	0.028	1.081	1.189
18	Oberschlesische Portland-Cement AG	1.094	0.013	1.069	1.119
19	Rheinische Stahlwerke	0.781	0.030	0.723	0.839
20	Rositzer Zuckerfabrik	1.053	0.025	1.005	1.101
21	Schaaffhausen'scher Bankverein	0.572	0.028	0.518	0.626
22	Chemische Fabrik vormals Schering	1.001	0.025	0.952	1.050
23	Schlesische Zinkhütten	0.959	0.022	0.916	1.002
24	Schlesische Leinen-Industrie	1.183	0.022	1.139	1.227
25	Schultheiss Brauerei	0.684	0.018	0.650	0.718
26	Siemens Glas-Industrie	0.776	0.018	0.740	0.812
27	Stettiner Chamottewaren	0.905	0.027	0.852	0.958

Source: Gelman and Burhop (2008), own calculations. Standard errors are calculated assuming independence of transaction cost estimates across time.

Appendix 2: Explanatory power of the market model for non-zero returns

Number	Name	Average R-squared
1	AG für Anilinfabrikation	0.092322
2	Allgemeine Elektrizitätsgesellschaft	0.329350
3	Berlin-Anhaltinische Maschinenbau	0.116825
4	Bochumer Bergwerk (Lit C)	0.065104
5	Bank für Handel und Industrie	0.351353
6	Deutsche Bank	0.397444
7	Dresdner Bank	0.488626
8	Deutsche Jute Spinnerei und Weberei	0.073646
9	Deutsche Spiegelglas	0.051038
10	Erdmannsdorfer Spinnerei	0.030019
11	Gelsenkirchener Bergwerksgesellschaft	0.508946
12	Gerresheimer Glashütten	0.054653
13	Hallesche Maschinenfabriken	0.043489
14	Harpener Bergbau AG	0.462738
15	Kattowitzer AG für Bergbau und Eisen	0.192785
16	Maschinenfabrik Kappel	0.043751
17	Norddeutsche Wollkämmerei	0.057130
18	Oberschlesische Portland-Cement AG	0.096344
19	Rheinische Stahlwerke	0.264628
20	Rositzer Zuckerfabrik	0.065095
21	Schaaffhausen'scher Bankverein	0.269807
22	Chemische Fabrik vormals Schering	0.072648
23	Schlesische Zinkhütten	0.089340
24	Schlesische Leinen-Industrie	0.010887
25	Schultheiss Brauerei	0.043544
26	Siemens Glas-Industrie	0.054630
27	Stettiner Chamottewaren	0.105594
	Average	0.164138

Numbers in the third column represent for each stock averages of 22 R-squared values obtained from yearly market model regressions for non-zero return observations.

Appendix 3: Average alternative transaction costs

Year	ALOT	Std. error	95% confidence interval	
			lower bound	upper bound
1892	1.475	0.060	1.358	1.592
1893	1.601	0.062	1.480	1.722
1894	1.029	0.038	0.954	1.105
1895	0.903	0.031	0.842	0.964
1896	0.763	0.028	0.709	0.818
1897	0.797	0.029	0.740	0.853
1898	0.877	0.032	0.815	0.940
1899	0.834	0.035	0.765	0.903
1900	0.975	0.038	0.901	1.049
1901	1.785	0.102	1.585	1.985
1902	0.952	0.114	0.729	1.175
1903	0.841	0.028	0.787	0.896
1904	0.793	0.029	0.737	0.850
1905	0.657	0.025	0.608	0.705
1906	0.617	0.026	0.565	0.669
1907	0.702	0.029	0.645	0.760
1908	0.816	0.031	0.755	0.878
1909	0.666	0.027	0.612	0.719
1910	1.030	0.034	0.964	1.096
1911	0.689	0.026	0.638	0.741
1912	0.816	0.028	0.762	0.870
1913	1.062	0.041	0.981	1.143
Average	0.940	0.010	0.921	0.960

Own calculations based on daily returns for 27 stocks for the period 1892-1913.

Expressed in percent of share price, equally weighted averages. Four outliers were dropped. Standard errors are calculated taking into account cross-correlations between stocks. Confidence interval is given by

$$\bar{S}_{ALOT,t} \pm 1.96s.e.(\bar{S}_{ALOT,t})$$

Appendix 4: Average alternative transaction costs of corporations, included in the investigation

Number	Name	Average alternative transaction cost measure	Standard error	95% confidence interval	
				Lower	upper
1	AG für Anilinfabrikation	0.966	0.023	0.921	1.011
2	Allgemeine Elektrizitätsgesellschaft	0.451	0.015	0.421	0.480
3	Berlin-Anhaltinische Maschinenbau	0.893	0.022	0.851	0.935
4	Bochumer Bergwerk (Lit C)	3.334	0.171	2.998	3.669
5	Bank für Handel und Industrie	0.454	0.011	0.433	0.474
6	Deutsche Bank	0.342	0.008	0.325	0.358
7	Dresdner Bank	0.361	0.009	0.343	0.378
8	Deutsche Jute Spinnerei und Weberei	1.011	0.023	0.966	1.055
9	Deutsche Spiegelglas	1.024	0.025	0.974	1.073
10	Erdmannsdorfer Spinnerei	1.524	0.033	1.460	1.588
11	Gelsenkirchener Bergwerksgesellschaft	0.315	0.013	0.290	0.341
12	Gerresheimer Glashütten	1.382	0.032	1.319	1.446
13	Hallesche Maschinenfabriken	1.126	0.024	1.078	1.174
14	Harpener Bergbau AG	0.324	0.014	0.297	0.351
15	Kattowitzer AG für Bergbau und Eisen	0.623	0.016	0.591	0.655
16	Maschinenfabrik Kappel	1.129	0.028	1.074	1.183
17	Norddeutsche Wollkämmerei	1.205	0.040	1.127	1.284
18	Oberschlesische Portland-Cement AG	0.994	0.025	0.945	1.043
19	Rheinische Stahlwerke	0.706	0.020	0.668	0.744
20	Rositzer Zuckerfabrik	1.008	0.026	0.958	1.058
21	Schaaffhausen'scher Bankverein	0.530	0.011	0.509	0.552
22	Chemische Fabrik vormals Schering	1.022	0.024	0.974	1.070
23	Schlesische Zinkhütten	0.998	0.023	0.952	1.044
24	Schlesische Leinen-Industrie	1.249	0.026	1.198	1.300
25	Schultheiss Brauerei	0.709	0.015	0.679	0.739
26	Siemens Glas-Industrie	0.783	0.017	0.749	0.817
27	Stettiner Chamottewaren	0.946	0.024	0.900	0.993

Source: Gelman and Burhop (2008), own calculations. Standard errors are calculated assuming independence of transaction cost estimates across time.

Appendix 6: Explanatory power of the augmented market model for non-zero returns

Number	Name	Average R-squared
1	AG für Anilinfabrikation	0.107059
2	Allgemeine Elektrizitätsgesellschaft	0.348324
3	Berlin-Anhaltinische Maschinenbau	0.141805
4	Bochumer Bergwerk (Lit C)	0.074821
5	Bank für Handel und Industrie	0.369678
6	Deutsche Bank	0.411315
7	Dresdner Bank	0.508748
8	Deutsche Jute Spinnerei und Weberei	0.229672
9	Deutsche Spiegelglas	0.158531
10	Erdmannsdorfer Spinnerei	0.250039
11	Gelsenkirchener Bergwerksgesellschaft	0.520432
12	Gerresheimer Glashütten	0.065981
13	Hallesche Maschinenfabriken	0.097760
14	Harpener Bergbau AG	0.469876
15	Kattowitzer AG für Bergbau und Eisen	0.199731
16	Maschinenfabrik Kappel	0.304432
17	Norddeutsche Wollkämmerei	0.069061
18	Oberschlesische Portland-Cement AG	0.294411
19	Rheinische Stahlwerke	0.272544
20	Rositzer Zuckerfabrik	0.126384
21	Schaaffhausen'scher Bankverein	0.289018
22	Chemische Fabrik vormals Schering	0.082454
23	Schlesische Zinkhütten	0.105702
24	Schlesische Leinen-Industrie	0.016795
25	Schultheiss Brauerei	0.047726
26	Siemens Glas-Industrie	0.078423
27	Stettiner Chamottewaren	0.130465
	Average	0.245094

Numbers in the third column represent for each stock averages of 22 R-squared values obtained from yearly augmented market model regressions for non-zero return observations.

Appendix 7. Market and liquidity risk betas

Number	Name	Market beta	Liquidity risk beta
1	AG für Anilinfabrikation	0.91	-1.03
2	Allgemeine Elektrizitätsgesellschaft	1.07	0.03
3	Berlin-Anhaltinische Maschinenbau	0.84	-0.11
5	Bank für Handel und Industrie	0.89	-0.21
6	Deutsche Bank	0.77	-0.33
7	Dresdner Bank	1.07	-0.45
8	Deutsche Jute Spinnerei und Weberei	0.73	-1.62
9	Deutsche Spiegelglas	0.67	-1.10
10	Erdmannsdorfer Spinnerei	0.67	-0.49
11	Gelsenkirchener Bergwerksgesellschaft	1.31	0.02
12	Gerresheimer Glashütten	0.43	-1.28
13	Hallesche Maschinenfabriken	0.78	-1.29
14	Harpener Bergbau AG	1.36	-0.17
15	Kattowitzer AG für Bergbau und Eisen	0.88	-0.87
16	Maschinenfabrik Kappel	0.82	-1.92
17	Norddeutsche Wollkämmerei	0.89	0.12
18	Oberschlesische Portland-Cement AG	0.95	-0.72
19	Rheinische Stahlwerke	1.35	-0.79
20	Rositzer Zuckerfabrik	0.81	-1.07
21	Schaaffhausen'scher Bankverein	0.77	-0.14
22	Chemische Fabrik vormals Schering	0.74	0.79
23	Schlesische Zinkhütten	0.63	-1.60
24	Schlesische Leinen-Industrie	0.46	-0.32
25	Schultheiss Brauerei	0.39	-1.67
26	Siemens Glas-Industrie	0.59	-0.91
27	Stettiner Chamottewaren	0.86	0.83
	Average	0.92	-0.60

Market beta is the slope of regression of monthly stock excess returns on excess market returns. Liquidity risk beta is the slope of regression of yearly individual stock illiquidity shocks on excess market returns.

Appendix 8. Trade volume of stocks in Imperial Germany, 1892-1913

Year	Trade volume in bill. mark
1892	72.07
1893	45.78
1894	73.60
1895	79.17
1896	51.40
1897	63.62
1898	59.20
1899	71.89
1900	59.77
1901	47.84
1902	50.72
1903	53.78
1904	56.34
1905	80.28
1906	63.79
1907	38.84
1908	37.80
1909	73.80
1910	84.72
1911	87.36
1912	91.61
1913	60.64
Average	63.82

Source: Wetzel (1996)