TABLE 1
Model Estimates

	log(Output) Est.	log(Output) SE	log(Output) Est.	log(Output) SE	log(Output) Est.	log(Output) SE
	A. Production Function, log(Output)					
log(Labor) β^{l}	.794	.034	.699	.327	.661	.041
log(Materials) β^{m}	.275	.028	.222	.138	.237	.080
log(Capital) β^{k}	−.008	.140	.153	.075	.102	.088
Serial correlation TFP ρ			.866	.198	.853	.157
Method	OLS		GMM		GMM	
RTS	Free		Free		Fixed at 1.05	
R^2	.941		.938		.826	
Hansen J-test			2.34		2.72	
Hansen J-test p-value			.126		.255	
Number of firms	166		159		159	
Observations	4,480		4,005		4,005	
	B. Markdowns and Markups					
Median markdown	1.541	.193	1.680	.450	1.486	.330
Average markdown	1.676	.224	1.828	.491	1.616	.361
Median markup	.884	.112	.714	.494	.763	.287
Average markup	.946	.120	.764	.535	.816	.315
Method	OLS		GMM		GMM	
RTS	Free		Free		Fixed at 1.05	
	C. Labor Supply					
	log(Wage) Est.	log(Wage) SE	log(Wage) Est.	log(Wage) SE		
log(Employment)	.066	.006	1.009	.265		
Method	OLS		IV			
First-stage F-statistic			462			
Hansen J-test			5.92			
Hansen J-test p-value			.014			
Observations	1,990		1,990			
Firm-level elasticity	155.56		10.172			

Note.—Panels A and B are estimated at the firm-year level, and panel C is estimated at the market-year level. Standard errors (SEs) in panels A and B are block-bootstrapped with 200 iterations. Standard errors in panel C are estimated using the Driscoll and Kraay (1998) correction to allow for both cross-sectional (i.e., intratemporal) and intertemporal dependence, using the STATA command ivreg2,draay(2).

average of the firm wages, weighted by their employment shares within each market. As mentioned above, there is barely any within-municipality wage variation. Moreover, 90% of the workers did not commute more than 10 km from their home, as shown in figure D.7. This shows that most workers were employed within the boundaries of the village where they lived.

To identify the labor supply curve, we need labor demand shifters, as firms choose employment levels with knowledge of the latent market-level labor supply shifters $\nu_{\ell t}$. We rely on two labor demand shifters. First, we construct an indicator variable for the coal demand shock between 1871 and 1875 due to the aftermath of the Franco-Prussian War of 1870, which coincided with a peak in the international coal price as shown in figure 2. After the Franco-Prussian War, the French coal basin in Lorraine was annexed by Germany, which resulted in a sharp increase in the international coal price and hence in the demand for coal in the Liège and Namur coal basin. This "coal famine" of the early 1870s was exacerbated by cold winters and other reasons for rapid increases in consumption (Murray and Silvestre 2020, 688). This instrument is measured as a dummy indicating the years between 1871 up to and including 1875. Second, we include cartel membership during the cartel period as a demand shifter, given that the cartel decreased coal supply and hence labor demand for the cartel participants. This is measured as the interaction term of the cartel dummy with the postcartel period. We control for cartel membership and for the time dummy indicating the postcartel period. Conditioning on these instruments and on log employment and wages to be observed, the market-level sample size drops from 2,624 to 1,990 observations.

The estimates are in panel C of table 1. The market-level inverse elasticity of labor supply Ψ^l is estimated at 1.009. This implies that at a monopsonistic firm the marginal revenue product of labor is twice the wage, whereas it would be 10% above the wage at a firm with a labor market share of 10%. Converting this market-level inverse elasticity to a firm-level labor supply elasticity, as explained in appendix section A.4, implies an average firm-level elasticity of 10.172. This is of a similar order of magnitude as the average labor supply elasticity in current-day studies as surveyed in Sokolova and Sorensen (2021). Based on the Hansen *J*-test, we can reject overidentifying restrictions.

Again, we perform a wide range of robustness checks. In appendix section C.2.4, we allow for time-changing labor supply coefficients and also include a linear time trend. In appendix section C.2.5, we reestimate the model using different labor market definitions, as well as assess the potentially confounding effects of the expansion of the railroad network throughout the nineteenth century. In appendix section C.2.6, we change the time window over which the coal price shock instrument is defined to 1871–74 and 1871–76. In appendix section C.2.7, we compare our results against two separate model specifications that rely on only the price surge instrument and the cartel membership instrument, respectively. We prefer to keep both instruments as the main specification because this gives both intertemporal and cross-sectional variation in the instrument.

B. *Markdowns and Markups*

1. Wage Markdowns and Price Markups

Using the production function coefficients, we can now estimate coal price markups μ_{ft} and wage markdowns μ^l_{ft} following equation (13) and the right-hand side of equation (14), respectively. The log markdowns are observed for 4,702 observations. The estimated moments are in panel B of table 1. Taking our preferred specification that does not impose an RTS parameter, we obtain a median wage markdown of 1.680, which implies a markdown wedge of miner wages δ^{l}_{ft} of 40% below the marginal revenue product of labor. The average wage markdown is 1.828. Although both the median and average wage markdowns were not significantly different from 1 over the entire time period, there is an important fraction of firms and time periods for which wage markdowns are significantly above 1. We will assess drivers of this wage markdown heterogeneity across firms and time further below. Imposing an RTS parameter of 1.05 results in slightly lower markdown estimates of 1.486 and 1.616 at the median and on average, respectively. The standard errors on the markdown estimates reduce, even when normalizing to the median and average markdown levels, because the production function is estimated more precisely when imposing this RTS restriction.

In contrast to the wage markdown, the coal price markup was much lower. Using the full model, the price markup at the median firm was 0.714; the average was 0.764. Hence, coal prices are below marginal costs. This does not mean that firms were loss making, given that the total profit margin is the combined wage markdown and price markup. The joint markup $\mu^{\mathrm{tot}} \equiv 1 + (\mu_{ft} - 1) + (\mu^{l}_{ft} - 1)$, which is the sum of the coal price markup and the wage markdown, is 1.44 at the median firm and 1.58 on average, which implies that these firms were making profits despite negative markups. The joint markup is negative for 13% of observations only.

Our low markup estimates suggest that coal mines had little market power downstream. This is no surprise, given that the relevant coal market size was much larger than Liège and Namur. Figure 2 shows that the coal price in Liège and Namur followed the international coal price up to 1897, which indicates that the firms in our dataset were price takers on the coal market. This is in line with recent historical research that has highlighted the increasing integration of the European coal market through the nineteenth century (Murray and Silvestre 2020). If the coal firms in the dataset were price takers on the coal market, this would imply that markup $\mu_{ft} = 1$, which cannot be rejected from our markup estimates. Our result of prices below marginal cost $\mu_{ft} < 1$, even if this finding is not significant, could be explained by monopsony power of coal buyers, such as large steel plants or railroad companies. If these industrial buyers have monopsony power over the coal mines, it is conceivable that they would use this power to push down coal prices to grasp the profit margins generated by monopsony power of the coal mines on the labor market.[29] Normally, monopsonistic buyers would not push prices below marginal

[29] This was also discussed in Rubens (2023b) in the context of Chinese tobacco markets.

costs because their suppliers would then exit the market. However, in our setting, coal firms do not exit the market when coal prices fall below marginal costs, because there is still the markdown wedge between marginal costs and input prices as an additional source of profits.

Taken together, the markdown and markup estimates above imply that coal firms mainly derived profits from market power on their labor markets, rather than on the coal market.[30] Still, equilibrium markdowns above 1 do not necessarily imply collusion: they could be due to noncollusive oligopsony power. In what follows, we will unpack the effects of collusion on the wage markdown, starting with a correlational analysis in the next subsection. As mentioned earlier, all estimates are derived from our preferred specification without assuming RTS, unless otherwise noted.

2. Evolution and Drivers of the Wage Markdown

Figure 3 plots the evolution of the wage markdown $\hat{\mu}^{l}_{ft}$ in all coal mines in Namur and Liège provinces between 1845 and 1913. Up to the 1870s, the median firm had a wage markdown ratio of around 1.5, which implies that wages were around a third below the marginal revenue product of labor. This ratio was relatively stable throughout the 1840s, 1850s, and 1860s. The average wage markdown, weighted by employment shares, was higher, around 1.75 on average.[31] During the late 1870s and 1880s, a long period of recession, median wage markdowns grew moderately to around 1.7. Despite short-run fluctuations, the wage markdown usually reverted to its long-term mean within 4–5 years.

Around 1900, there was a sharp increase in the wage markdown, both on average and at the median firm. The average wage markdown after 1897 was around 2.2, meaning that workers received less than 50% of their marginal revenue product. This wage markdown increase was persistent: there was no reversion to the pre-1897 steady-state level. The estimates in table 2 show that the increase in the wage markdown after 1897 was statistically significant. The wage markdown increase after 1897 does not reflect reallocation between firms but was the result of within-firm markdown growth. Figure D.11 compares the unweighted average wage markdown to the weighted average wage markdown, by employment usage. The unweighted average wage markdown grew by even more after 1897, which indicates that there was some reallocation away from the highest-markdown firms after 1897.

[30] Nonetheless, in app. sec. D.4, we find moderate positive effects of the 1897 cartel on the markups of its participants.

[31] Figure D.11 compares the unweighted and weighted markdown series, which up to the cartel period are very similar. Appendix sec. C.3.2 compares different weighting methods to construct aggregate markdowns.

FIG. 3.—Evolution of average and median wage markdowns, 1845–1913. This graph shows the evolution of the weighted average (by employment) markdown and median wage markdown in Liège and Namur coal mines from 1845 to 1913.

TABLE 2
MARKDOWNS: CORRELATIONS AND EVOLUTION

	log(Markdown)		log(Markdown)	
	Est.	SE	Est.	SE
	A. Markdown Correlations			
1(Employers' association)	.112	.052		
1(Cartel)	.080	.041		
1(1855 ≤ year < 1865)			−.021	.039
1(1865 ≤ year < 1875)			−.020	.038
1(1875 ≤ year < 1885)			.059	.045
1(1885 ≤ year < 1895)			.108	.047
1(1895 ≤ year < 1905)			.196	.045
1(1905 ≤ year < 1915)			.422	.054
Year fixed effects	Yes		No	
R^2	.094		.076	
Observations	4,432		4,705	
	B. Employers' Association, Pre- vs. Postcartel			
1(Employers' association)	.132	.042	−.058	.091
Period	1845–97		1898–1913	
R^2	.094		.130	
Observations	3,737		695	

NOTE.—Panels A and B are both estimated at the firm-year level. The reference category for the time dummies in panel A is the period between 1845–1859. Block-bootstrapped standard errors (SEs) are computed using 200 iterations.

What could explain the variation in wage markdowns across firms? The historical discussion in section II.D highlighted two key drivers. First, there was the pervasive nature of employers' associations throughout the nineteenth century. Based on internal communication by the union, we created a time-invariant variable indicating the union membership of each firm. A second big shift in the competitive environment of both coal and labor markets happened in 1897, when the coal cartel Syndicat des Charbonnages Liégeois was set up. The cartel statutes reveal which firms were part of said cartel.[32]

In column 1 of table 2, we compare markdowns across employers' association and cartel membership. Having to observe these membership statuses reduces the sample from 4,702, the sample on which markdowns are observed, to 4,429. We find that wage markdowns were 11.2% higher among employers' association members. This confirms anecdotal evidence of wage fixing through these employers' associations. Wage markdowns were also 8.0% higher for members of the coal cartel, but, given that the membership of the cartel and the employers' associations overlap, there is a possibility of multicollinearity here. Also, comparing wage markdowns at cartel and noncartel firms does not reveal the true effect of cartel membership on wage markdowns, as this variation could be due to a variety of markdown drivers. This again highlights the need for a more solid identification approach toward collusion.

In panel B of table 2, we compare the correlation between wage markdowns and employers' association membership between two time periods: the pre- and the postcartel period. The difference in wage markdowns between employers' association members and nonmembers that existed prior to 1897 entirely disappears after the introduction of the cartel in 1897. This suggests that the informal wage collusion that took place in employers' associations, which was not legally binding, was replaced as a driver of wage markdowns by the formal collusion through the coal cartel.

3. Markdown Heterogeneity

The homogeneous employers Cournot model has strong empirical implications for wage and markdown variation, which can be tested using our data and estimates. First, the Cournot model implies within-market markdown variation, whereas wages should be homogeneous. Moreover, in the absence of full collusion, wage markdowns should be higher for firms with high labor market shares, given that they face more inelastic firm-specific (residual) labor supply curves. Under full collusion, wage markdowns should be equalized within markets, and wage markdowns should no longer be increasing in firm size.

[32] For more information on the firm-level membership data, we refer to app. sec. B.2.

We test these implications using the markdown estimates from the production model. We regress the log wage markdown on the log labor market share in three specifications: one without any fixed effects, one including market fixed effects, and one including market × year fixed effects. The results can be found in the three sets of estimates in table 3, respectively. Panel A reports these correlations for all firms, panel B only for firms that are not part of the cartel, and panel C for the cartel firms. For the noncartel firms, there is quite some markdown variation within a given year and market: market × year fixed effects explain 55% of markdown variation. Moreover, there is a positive relationship between firm size, as measured by the labor market share, and markdowns, both with and without market × year fixed effects. However, when conditioning on the cartel members, we find that there is no longer a positive relationship between labor market shares and markdowns as soon as we control for market fixed effects. Although there is still some variation in wage markdowns within market-year cells for collusion firms, there is much less markdown heterogeneity than for noncartel firms. The latter is in line with the theory. The finding that markdowns are not exactly identical for cartel members could be due to imperfect discipline among the cartel members.

TABLE 3
Size-Markdown Correlations

	log(Markdown)					
	Est.	SE	Est.	SE	Est.	SE
	A. All Firms					
log(Labor market share)	.044	.001	.055	.003	.051	.004
Fixed effects	None		Market		Market × year	
R^2	.067		.192		.550	
Observations	4,671		4,671		4,671	
	B. Noncartel Firms					
log(Labor market share)	.037	.000	.053	.005	.065	.005
Fixed effects	None		Market		Market × year	
R^2	.046		.180		.561	
Observations	3,183		3,183		3,183	
	C. Cartel Firms					
log(Labor market share)	.043	.001	.008	.002	−.004	.002
Fixed effects	None		Market		Market × year	
R^2	.063		.188		.793	
Observations	1,472		1,472		1,472	

Note.—We regress log markdowns on the log labor market employment share at the firm-year level for all firms (panel A), firms outside the cartel (panel B), and firms in the cartel (panel C). We control for a linear time trend and either no, market, or market × year fixed effects. The sample sizes of panels B and C add up to 4,655 because the cartel information is unobserved for 16 observations. Standard errors (SEs) are block-bootstrapped with 200 iterations.

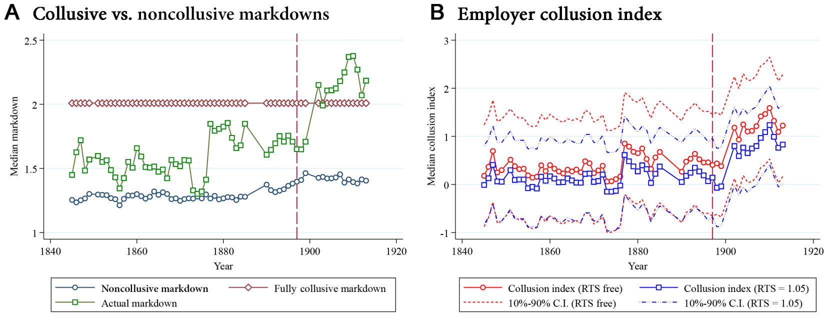

FIG. 4.—Employer collusion estimates, 1845–1913. *A*, The median markdown over time, along with the median of the lower and upper markdown bounds under no and full collusion. *B*, The median collusion index together with block-bootstrapped confidence intervals between 1845 and 1913. Two hundred bootstrap iterations are used.

C. *Employer Collusion*

1. Markdown Decomposition

We now decompose the estimated wage markdowns into a collusive and a noncollusive component and estimate the collusion index from equation (15). Figure 4*A* plots the evolution of actual wage markdowns and the collusive and noncollusive markdown bounds as defined in section III. The circles show the annual median of the lower markdown bound in the absence of collusion $\underline{\mu}^1$, the diamonds show the upper bound of markdowns under full wage collusion $\bar{\mu}^1$, and the squares show the estimated median markdowns, $\hat{\mu}^1_{ft}$, as estimated using the left-hand side of equation (14). Prior to the introduction of the cartel in 1897, the actual markdown lies above the noncollusive lower bound. This difference could be due to imperfect wage collusion devices such as the employers' associations.[33] After the introduction of the cartel in 1897, the estimated markdown level moves up to the fully collusive upper bound.

From 1870 to 1900, there was an increase in the median markdown level, but there was equally an increase in the noncollusive lower markdown bound. The moderate growth in markdowns prior to 1900 hence seems not to be related to wage collusion. However, around 1900, markdowns jump to the fully collusive upper bound for the wage markdown. Given that the noncollusive markdown does not grow after 1900, the vast increase in markdowns after the introduction of the coal cartel appears to have been entirely driven by wage collusion.

[33] This difference could also be due to any other deviation from the baseline Cournot model, such as search or adjustment frictions, firm differentiation, or dynamic labor supply. We examine input adjustment costs in sec. V.A and firm differentiation in app. sec. C.2.3.

2. Testing for Employer Collusion

We can now tackle the question of whether we are able to detect employer collusion during the cartel era without ex ante knowledge of said cartel. Figure 4*B* plots the evolution of median collusion by year for both our preferred specification that does not impose an RTS parameter (circles) and for the production function estimates that calibrate RTS at 1.05 (squares), along with 10%–90% confidence intervals. We find that the median markdown fluctuated around 0%–50% of the collusive markdown level up to 1900, but we cannot reject the null hypothesis of no wage collusion for any year up to 1900. From 1901, we can reject the null of no collusion for every year except 1903 at the 10% confidence level in the full model and from 1908 in the model that restricts RTS at 1.05. When imposing a 5% confidence level, which we do in figure D.12, we can equally reject the absence of collusion for 1905 in the full model and from 1908 in both the full and calibrated RTS versions of the model. The price data in figure 2 suggest that the collusive behavior within the cartel took off from 1904 onward, as this is the year in which Liège coal mine prices start moving toward the international coal price. Hence, the collusion estimates seem to be able to detect collusion due to the cartel, without requiring any a priori information about the cartel.[34]

D. Consequences of Employer Collusion

1. Counterfactual Setup

To assess the effects of the cartel on wages and employment, we need to close the model and solve for joint labor and product market equilibrium. Moving from a cartel to Cournot competition does not just change the wage markdown but also the marginal revenue product of labor. To solve for equilibrium, we assume symmetry within each labor market, meaning that in a labor market i with N_{it} firms, each firm has a labor market share of $1/N_{it}$. Although this symmetry assumption is clearly rejected by the data, we find that it provides a very close approximation to the truth when examining the market-level aggregate implications of collusion, as is the goal of the counterfactual. We show this in appendix section A.3. We also assume that all firms in a labor market have the same level of labor collusion and rely on the conduct parameter $\tilde{\lambda}_{it}$ as it was defined in equation (12). Using the symmetric firms assumption, the market-level parameter $\tilde{\lambda}_{it}$ can be written in function of the collusion index $\hat{\lambda}$ and of market structure:

[34] Admittedly, we did rely on cartel information as a demand shifter to estimate labor supply, but this is not strictly necessary. With the availability of demand shifters, one could identify collusion using our approach without requiring information about which firms are colluding or when.

$$\tilde{\lambda}_{it} = \frac{1}{N_{it}} + \hat{\lambda}_{it}\left(1 - \frac{1}{N_{it}}\right).$$

When member of a cartel, firms set $\tilde{\lambda}_{it} = 1$, as this implies full collusion. We examine two counterfactual scenarios to assess the effects of the cartel. First, we set $\tilde{\lambda}_{it} = 1/N_{it}$, which corresponds to the Cournot model. This is a world with a complete absence of employer collusion. Second, we set the conduct parameter to $\tilde{\lambda}_{it} = \bar{\lambda}$, with $\bar{\lambda}$ being the average collusion index in 1897, just before the cartel started. This counterfactual scenario assumes that the cartel did not happen but that firms continued to collude imperfectly, to the same extent as they did prior to the cartel.

To solve for equilibrium wages and employment, we also need to take a stance on the extent to which coal markets are competitive. We rely on two different models, which provide bounds for the cartel effects. In a first model, we assume that coal prices are exogenous to individual firms. This is equivalent to assuming that the coal market was transnational and that individual Belgian coal firms were all atomistic on this coal market. This assumption provides a lower bound on the wage and employment responses. In a second model, we impose Cournot competition on the coal market, which moves to perfect collusion as soon as the cartel enters. This second assumption implies that coal markets are the same as labor markets. As we discuss below, this provides an upper bound (in absolute value) for the wage and employment effects of the cartel. Given that the median markup estimate is not significantly above 1 and that the coal market was transnational, rather than local, we believe that the true effects of the cartel are closer to the lower bound than to the upper bound, at least for the median firm.

2. Model with Exogenous Coal Prices

We start with the model specification that assumes exogenous coal prices to each individual coal firm. Under this assumption, we do not need to impose and estimate a coal demand model. As derived in appendix A, solving the labor demand function derived from the production function (1) and the labor supply curve (3) delivers the following equilibrium wages and employment levels in each market i:

$$\begin{cases} W_{it}^{l} = \left(\dfrac{\beta^{l} R_{it} \nu_{it}^{(1/\Psi^{l})}}{1 + \Psi^{l}\tilde{\lambda}_{it}}\right)^{\Psi^{l}/(1+\Psi^{l})}; \\[2ex] L_{it} = \left(\dfrac{\beta^{l} R_{it}}{(1 + \Psi^{l}\tilde{\lambda}_{it})\nu_{it}}\right)^{1/(1+\Psi^{l})}. \end{cases}$$

Using these equilibrium expressions, we compute the counterfactual wage and employment levels under Nash-Cournot competition and under pre-1897 conduct. The cartel effects are summarized in panel A of table 4. Compared to a baseline model of Cournot competition on the labor market, the cartel decreased both wages and employment by around 10%. However, the collusion estimates from the previous section suggest that labor market competition was not Cournot prior to the cartel. If we compare the cartel to a baseline model in which labor market conduct remained constant at its 1897 average, we find that employment and wages decreased by 6%.

If coal prices were endogenous to individual coal firms, the counterfactual employment and wage effects of collusion would be larger, as collusion leads firms to internalize both the market-level labor supply and the market-level product demand curve. Hence, as we show in section IV.D.3, the exogenous coal price counterfactual constitutes a lower bound to the employment and wage effects of the cartel.

3. Model with Endogenous Coal Prices

Next, we extend the model to allow for endogenous coal prices. Now, we need to formulate and estimate a coal demand model as well. We impose equation (18) as the market-level coal demand curve, with a market-level inverse demand elasticity η and a market-level coal demand shifter ξ_{it}:

$$P_{it} = Q_{it}^{\eta}\xi_{it}. \tag{18}$$

We identify joint equilibrium on the labor and product market by solving the system of equations given by the labor supply curve (3), the production

TABLE 4
Effects of Employer Collusion

	Comparison: Cartel vs.	
	Cournot	Pre-1898 Conduct
	A. Exogenous Price	
Relative wage change	−.103	−.059
Relative employment change	−.102	−.059
	B. Endogenous Price	
Relative wage change	−.251	−.167
Relative employment change	−.249	−.166
Relative price change	.174	.100
Relative output change	−.283	−.195

Note.—Panel A summarizes the wage and employment effects of moving from the fully collusive coal cartel to either Cournot competition or to the estimated level of collusion prior of the cartel introduction, assuming exogenous coal prices. Panel B does the same for the model that allows for endogenous coal prices.