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Institutions, Wages and Inequality: The Case of Europe and its Periphery (1500-1899)

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Abstract

This paper explores the long-run relationship between institutions and wage inequality in Europe and its periphery using a two-sector model. When institutions improve, wages rise across the board, thus reducing the costs of rural-urban migration and skills acquisition relative to the expected urban wage. The subsequent increase in the supply of urban craftsmen can in turn lead to a narrowing of the relative gap between skilled and unskilled wages. These predictions are borne out by the historical data. Cities with stronger institutions experienced: (i) higher skilled and unskilled real wages, and (ii) lower levels of urban inequality, as measured by the skilled-unskilled wage ratio.

Keywords: Institutions; Wage inequality; Rural-urban migration; European cities.

JEL Classification: J31, N13, N33, O10, O15

1 Introduction

In recent years, the research agenda on the sources of economic growth has seen a renewed focus on the role of institutions. Following North (1981) and North and Weingast (1989), much of this literature has dwelt on one particular form of institutions, namely the provision of a secure system of property rights protection. According to this view, a strong set of institutions is needed to prevent the executive or other elite groups from expropriating private wealth, in order to foster an environment conducive to the accumulation of capital, and hence growth. Several empirical studies have lent weight to this hypothesis. Knack and Keefer (1995), Barro (1997) and others identified a positive relationship across countries between aggregate growth and measures of institutions such as the rule of law. More recently, Acemoglu, Johnson and Robinson (2001) have used settler mortality rates as an instrumental variable for the risk of capital expropriation, in order to isolate the causal effect of institutions on income levels.

Within the context of Europe, institutions have also received much attention as a leading explanation for why some states grew faster than others over the course of history.¹ Treating states with absolutist rulers as synonymous with a high risk of expropriation, DeLong and Shleifer (1993) showed that European states run by such “princes” experienced slower growth in city populations than those governed as free republics. Acemoglu, Johnson and Robinson (2002) pursued this research program further by constructing improved indices on the strength of capitalist institutions stretching back to 1000A.D. Using these measures, the authors argued that access to Atlantic trade was crucial for growth, and that it was precisely those cities in Atlantic states with strong institutions where commerce expanded in full force. In particular, England (after the “Glorious Revolution” of 1688A.D.) and the Netherlands (after shaking off Spanish rule)

¹The large variance in the timing and pace of growth across Europe is well-documented: DeLong and Shleifer (1993) identified a northward shift in the growth of cities away from the Mediterranean Basin between 1000-1500A.D. Separately, Van Zanden (1999) and Allen (2001) highlighted that a divergence in wage levels had set in between Western and Eastern Europe well before the Industrial Revolution.

both succeeded in establishing political systems where the executive was made accountable to a parliament, thereby significantly reducing the risk of arbitrary expropriation and facilitating the rise of an urban merchant class.

This paper contributes to the growing literature on institutions, by examining how they affected wage outcomes in a broad sample of 17 cities in Europe and its periphery (including Istanbul).² The use of wage data confers two distinct advantages over existing studies. First, real wages are a more direct measure of the welfare of workers than aggregate incomes, since wages in principle capture the actual returns earned by labor.³ This paper thus tests for the effects of institutions on income levels directly, without having to use a proxy for the dependent variable such as urbanization or city population, as has been the practice to date. Second, the data from Allen (2001) and Özmucur and Pamuk (2002) includes wages for both skilled craftsmen and unskilled laborers, therefore allowing us to investigate the relationship between institutions and wage inequality. Historians have long documented how the urbanization of Europe was accompanied by a widening rich-poor divide, which on occasion boiled over into outbreaks of violence such as food riots (Hohenberg and Lees 1985, p. 131). Inequality was thus a very real problem, and it is relevant to ask how improvements in institutions might have affected the relative economic position of different groups of workers.⁴

Using the institutional indices coded by Acemoglu, Johnson and Robinson (2002), I find that cities which provided more stringent protection from expropriation witnessed higher subsequent real wage levels for both skilled craftsmen and unskilled laborers. More interestingly, stronger initial institutions were also associated with lower levels of urban wage inequality, as measured by the skilled-unskilled wage ratio. Last but

²For a full list of the cities, please refer to Appendix B on data sources.

³This argument in favor of factor returns data over aggregate figures as a measure of welfare is well-articulated in the economic history literature. See, for example, Williamson (1995).

⁴In a modern context, De Soto (2000) has proposed that the establishment of legal institutions to recognize the ownership of property occupied by the poor could help to improve their economic condition significantly, as this would allow the use of their property as collateral to gain access to credit.

not least, in a reduced sample for which data on land rents was available, there is some preliminary evidence that institutions also helped to improve the position of wage earners relative to landlords.

Importantly, these results are robust to the inclusion of other variables that have been proposed elsewhere to explain differences in income levels. For example, Lal (1998) and Engerman and Sokoloff (2002) have argued that factor endowments were a more fundamental determinant of growth, as the initial cards dealt by nature were often historically pivotal in shaping the types of political and economic institutions that societies set up to adapt to their environments. The empirical work in Section 3 thus includes controls for land and labor supplies. Nevertheless, the institutional variables were largely robust to the inclusion of these additional regressors, indicating that institutions were not entirely endogenous to factor endowments in the case of Europe and its periphery. Similarly, the effect of institutions was unchanged when a dummy for Atlantic trader states was included, which would appear to mitigate the over-riding importance placed on Atlantic commerce as a pre-condition for growth emphasized by Acemoglu, Johnson and Robinson (2002). Most recently, Glaeser et. al. (2004) have argued forcefully for the primacy of human capital over institutions as the source of long-run growth, by presenting evidence based on post-World War II data that initial human capital is a more robust positive correlate of growth than initial institutions. Hence, Section 3 also explores the use of adult literacy rates as an additional control. While this proxy for human capital often yielded a significant coefficient, it did not overturn the sign and significance of the institutions variables as a determinant of wage outcomes.

The above evidence linking stronger institutions to lower levels of inequality merits more explanation. Indeed, existing theoretical models have been more concerned with the reverse direction of causation, namely how an initial configuration of the income distribution can affect a society's choice of policies regarding state taxation (Alesina and Rodrik 1994) and the extent of property rights protection (Tokman 1999). These

models of endogenous institutions are however not particularly relevant to the context of pre-modern Europe, given the general absence of voting democracies then. On a similar note, Glaeser, Scheinkman and Shleifer (2003) have argued that in states with higher initial levels of inequality, legal institutions are more open to subversion by the rich or politically influential. Their model however takes a stylized investment game between individual players as a starting point, from which it is hard to relate the quality of institutions to labor market outcomes.

Instead, this paper develops a two-sector rural-urban model that treats changes to the institutional environment as exogenous, and explores the subsequent impact on wage inequality. The model features an explicit mechanism for rural-urban migration, which is similar in spirit to Harris and Todaro (1970). When the protection of property rights improves and wages rise for both skilled and unskilled workers, it becomes relatively less costly for rural laborers to migrate into the city and attempt to acquire a skilled craft. This induces a movement of workers into urban centers, raising the supply of skilled craftsmen. If this labor supply response is sufficiently large – specifically, if the probability of successfully acquiring skills remains sufficiently high – skilled wages could rise less relative to unskilled wages in the long-run equilibrium after labor supplies have adjusted, thus resulting in lower urban wage inequality. As long as unskilled wages are highly correlated across rural and urban sectors, overall wage inequality in the economy will also fall.⁵ Clearly, the intuition behind this model makes a point that extends to more general contexts: When the acquisition of skills is costly, an improvement in institutions can lead to an increase in the supply of skilled labor, since workers will be more willing to invest in education or training when they are more assured of reaping

⁵If one adopts the view that inequality is detrimental to growth, then improvements in institutions that reduce inequality could provide an impetus for further increases in income levels. Empirically, however, there is a lack of consensus on this link between inequality and growth, and hence this paper does not emphasize this potential feedback effect. For example, Alesina and Rodrik (1994) find that inequality appears to hurt growth, but Forbes (2000) estimates this relationship to be positive instead. Using non-parametric methods, Banerjee and Duflo (2003) argue for a non-linear trend, with both increases and decreases in inequality associated with lower subsequent growth.

the full rewards of their investment in human capital, instead of having their incomes expropriated away.

This paper proceeds as follows. In Section 2, I set up the model and trace out its main implications on how institutions affect real wages and wage inequality. Section 3 presents the empirical results, which show that the predictions of the model are largely consistent with the historical experience of Europe and its periphery. Section 4 concludes.

2 Institutions in a model of rural-urban migration

In this section, I present a two-sector general equilibrium model that provides a convenient framework for analyzing the impact of institutions on labor markets. Following De Vries (1984), I model the “post-medieval” European city as comprising a rural and an urban sector. The rural sector aggregates agriculture and rural proto-industrial activity, both of which are unskilled-labor intensive. Output in this sector is given by the Cobb-Douglas production function: $Q_R = A_R L_R^\alpha T^{1-\alpha}$, where L_R denotes unskilled labor employed in the rural sector, T is a fixed endowment of land, and A_R is a productivity term. On the other hand, the urban sector comprises manufacturing or crafts-based industries located in the city. Output here is given by: $Q_U = A_U L_U^\beta S^{1-\beta}$, where L_U denotes urban unskilled labor, S denotes skilled or artisan labor, and A_U is a technology parameter.⁶ Unskilled labor is a homogenous input to both sectors, whereas land and skilled labor are specific factors in the rural and urban sectors respectively.⁷

To introduce a role for property rights institutions, let $\mu \in [0, 1]$ denote the fraction

⁶Although there were good reasons for urban industry to relocate to rural areas to exploit the large supplies of labor there, the extent to which this actually happened was often limited by explicit barriers to the diffusion of knowledge (in Spain and Italy) and by monopoly controls against rural competition (in the case of the silk and textiles industries in Bologna). See Hohenberg and Lees (1985), p. 128.

⁷The model abstracts from physical capital, since most of the data for this paper comes from the pre-industrial period, during which the importance of physical capital as a factor of production was still relatively small. In Section 3, the empirical results are shown to be robust to the exclusion of London, the leading European industrial city for which physical capital would have mattered the most. All regressions were also run with city and period fixed effects, which should help to capture any systematic differences in the stock of physical capital in different cities or across time.

of its own marginal product actually received by factors of production. In other words, a fraction $1 - \mu$ is extracted by the “prince” or other state officials for their private consumption.⁸ Let p denote the price of urban sector output relative to that of the rural sector. Wages in this economy are then given by: $w_S = p\mu A_U(1 - \beta) \left(\frac{L_U}{S}\right)^\beta$ for skilled craftsmen, $w_L^U = p\mu A_U\beta \left(\frac{S}{L_U}\right)^{1-\beta}$ for unskilled urban workers, and $w_L^R = \mu A_R\alpha \left(\frac{T}{L_R}\right)^{1-\alpha}$ for unskilled rural workers.

The model allows for rural-urban migration, which played a significant role in the growth of Europe’s cities. Hohenberg and Lees (1985), for example, note that urban manufacturing would not have been able to expand without the additional labor provided by such migrants, since the rate of natural increase in cities fell far short of the actual growth rates in city populations (p. 127). Suppose then that any unskilled rural worker can move to the urban sector by paying a fixed cost Ψ_m , which includes the cost of transportation and the start-up costs of living in the city.⁹

Furthermore, an unskilled worker in the city can decide whether to pay an additional fixed cost Ψ_s to engage in training (such as an apprenticeship) to acquire a skilled craft. However, success at training is not guaranteed, and each apprentice becomes a skilled craftsman with probability $q(S) \in [0, 1]$, an outcome which they discover only after Ψ_s has been sunk. The analysis posits that $q'(S) < 0$, namely that as the pool of skilled workers increases, it becomes harder for a given unskilled worker to transition from L_U to S . This would be the case for example if there is *a priori* only a fixed fraction of the total labor force endowed with the necessary human capital for artisan work, so if S is already large, it becomes very difficult to find any more unskilled workers with the talent to acquire skills.¹⁰ In Appendix A.1, I show how this reasoning can be formalized

⁸Note that μ can just as easily be interpreted as the *expected* fraction of wealth that is not extracted, in order to introduce a stochastic element in the way expropriation is modelled.

⁹For simplicity, assume that return migration to the rural areas is costless, or at least that Ψ_m dwarves the costs of return migration.

¹⁰Conversely, one could also conceive of stories in which $q'(S) > 0$. For example, there might be learning-by-doing externalities in skilled crafts, so that the more skilled workers there are, the easier it becomes for outsiders to acquire skills. However, it seems likely that such spillovers are relatively

to justify this assumption that $q'(S) < 0$.¹¹

The equilibrium of the model is now determined as follows. Suppose that all workers receive an infinite stream of wages, but that each worker has a probability δ of dying at any instant. The present discounted value of lifetime earnings for say a skilled worker is thus: $\int_0^\infty e^{-\delta t} w_S dt = \frac{w_S}{\delta}$. Consider first the decision facing an unskilled worker in the urban sector. Skills acquisition will be worthwhile only if the expected returns from training exceed the returns from remaining unskilled, namely if: $q \frac{w_S}{\delta} + (1-q) \frac{w_L^U}{\delta} - \Psi_s \geq \frac{w_L^U}{\delta}$. Assuming that all urban unskilled workers have access to this option of training, this relationship must in fact hold with equality. A rearrangement of this equation yields: $w_S - w_L^U = \frac{\delta \Psi_s}{q}$, which pins down the gap between skilled and unskilled wages in the urban sector. Notice that this skill premium is increasing in the cost of skills acquisition Ψ_s and the instantaneous death rate δ , but falls with a higher success probability of training q . After substituting the relevant expressions for wages, this equation becomes:

$$\mu p A_U \left[(1-\beta) \left(\frac{L_U}{S} \right)^\beta - \beta \left(\frac{S}{L_U} \right)^{1-\beta} \right] = \frac{\delta \Psi_s}{q(S)} \quad (1)$$

Turning to the decision facing a rural unskilled worker, a similar reasoning shows that in equilibrium, the net returns from migration must be precisely equal to the returns from staying in the rural sector. This implies that: $\frac{w_L^U}{\delta} - \Psi_m = \frac{w_L^R}{\delta}$, or equivalently: $w_L^U - w_L^R = \delta \Psi_m$. Thus, the wage gap between urban and rural unskilled workers is increasing with the cost of migration Ψ_m and the death rate δ . Substituting in for the wages, this yields:¹²

$$\mu \left[p A_U \beta \left(\frac{S}{L_U} \right)^{1-\beta} - A_R \alpha \left(\frac{T}{L_R} \right)^{1-\alpha} \right] = \delta \Psi_M \quad (2)$$

limited, since one would otherwise see the pool of skilled craftsmen increase indefinitely.

¹¹There are some parallels here with the Harris-Todaro model, even though their canonical model only includes unskilled labor. Harris and Todaro (1970) look instead at the probability of a given migrant finding employment in the city, which they posit to be a decreasing function of the current urban unemployment rate.

¹²The expected returns from migration are given either by $\frac{w_L^U}{\delta}$ if the worker remains unskilled or by $q \frac{w_S}{\delta} + (1-q) \frac{w_L^U}{\delta} - \Psi_s$ if the worker becomes skilled. In equilibrium, however, these two quantities are equal by equation (1).

A third equation determines the allocation of the total workforce, N . To abstract away from issues of population growth, N can essentially be normalized to 1:¹³

$$L_R + L_U + S = N \quad (3)$$

To close the system, it remains to pin down p , the relative price of urban sector output. This is done by specifying a Cobb-Douglas aggregate utility function, $Q_R^\eta Q_U^{1-\eta}$, for the demand side of the economy, where $\eta \in [0, 1]$ is the share of aggregate income, Y , spent on rural sector output. Since $\eta Y = Q_R$ and $(1 - \eta)Y = pQ_U$, taking the ratio of these two equations gives:

$$p = \left(\frac{1 - \eta}{\eta} \right) \left(\frac{A_R}{A_U} \right) \left(\frac{L_R^\alpha T^{1-\alpha}}{L_U^\beta S^{1-\beta}} \right) \quad (4)$$

Equations (1)-(4) constitute a fully-specified general equilibrium system, with four equations in the four unknowns, L_R , L_U , S , and p . In what follows, I analyze how wage outcomes respond to changes in the institutional parameter, μ . Since these comparative statics require the movement of workers to restore the economy to equilibrium, it is appropriate to think of these as the long-run effects of institutions on the wage structure.

2.1 Effect of μ on urban wage inequality

Consider first how improvements in property rights institutions affect urban wage inequality, as measured by the relative wage:¹⁴

$$\frac{w_S}{w_L^U} = \left(\frac{1 - \beta}{\beta} \right) \left(\frac{L_U}{S} \right) \quad (5)$$

Clearly, any changes in the relative wage must be driven by how the relative labor supply term, $\frac{L_U}{S}$, responds to an increase in μ .

¹³ L_R , L_U and S would then be the shares of the total workforce allocated to these respective labor pools, while T would have to be re-interpreted as land per worker. This normalization does not affect the empirical estimation in Section 3 since the expressions for factor returns are all homogenous of degree 0 in the factor supplies.

¹⁴Although it would be interesting to look at the rural-urban wage gap as well, I focus only on urban inequality given the lack of data on rural wages, w_L^R .

The migration equation (2) and the skills acquisition equation (1) now play a central role in the reallocation of labor. When institutions improve, both migration and skills acquisition become cheaper relative to the raised level of expected wages in the urban sector. In other words, Ψ_m and $\frac{\Psi_s}{q}$ both fall relative to $qw_S + (1 - q)w_L^U$. This increases the incentives for rural workers to migrate to the city, and once there, to invest in an apprenticeship, thereby raising the population of cities and increasing the pool of skilled workers. If the probability of acquiring skills, $q(S)$, does not decline too quickly as S increases, then the supply of skilled workers relative to L_U will rise. This relative factor supply shift acts to reduce the skilled-unskilled wage ratio, and thus lowers wage inequality.¹⁵ This chain of logic is spelled out in the following proposition, which is proven formally in Appendix A.2.

Proposition: For the rural-urban economy defined by equations (1)-(4), the following comparative statics hold when income expropriation falls (μ rises):

- (i) The supply of skilled workers rises: $\frac{dS}{d\mu} > 0$.
- (ii) The response of L_U is indeterminate, ie the sign of $\frac{dL_U}{d\mu}$ is ambiguous.
- (iii) Wage inequality falls if and only if the following condition is satisfied:

$$\frac{d \ln(w_S/w_L^U)}{d\mu} < 0 \iff \left| \frac{Sq'(S)}{q(S)} \right| < \left(\frac{w_L^R}{\delta \Psi_m} \right) \left(\frac{N}{L_R} \right) \quad (6)$$

Parts (i) and (ii) of the proposition state that the influx of rural-urban migrants increases the pool of craftsmen S , but has an ambiguous effect on the supply of urban unskilled workers. This latter indeterminacy reflects the fact that there are two competing forces acting on L_U : On the one hand, workers from the rural sector are joining L_U ; but at the same time, workers can depart L_U after successfully acquiring skills.

¹⁵Note that μ is assumed to be equal across both sectors, to ensure that this analysis is not driven exclusively by exogenous differences in the extent of expropriation faced by the various types of workers. For example, if the urban sector witnessed a more extensive improvement in property rights than the rural sector, migration into the cities would become an even more attractive option, further reinforcing the decline in urban wage inequality.

The crux of the proposition is part (iii), which provides a necessary and sufficient condition for urban wage inequality to fall. This condition has a very intuitive explanation: When institutions improve, the skilled-unskilled wage ratio falls if and only if the relative supply of skilled workers increases, so that $\frac{L_U}{S}$ in equation (5) falls. This will be the case if the elasticity of q relative to S is small (in absolute value), so that the probability of acquiring skills does not decrease too quickly to constrict any further increases in S . Note that the lower the initial values of $\frac{\Psi_m}{w_L^R/\delta}$ or $\frac{L_R}{N}$, the easier it is for the inequality to be satisfied. This is because there is a larger potential pool of skilled workers when either the cost of migration relative to lifetime rural sector earnings is low, or when the initial fraction of the workforce located in the city is high, and this helps to relax the condition in (6) for wage inequality to fall.

It is useful to re-write (5) in a manner that will facilitate the regressions in Section 3. Define $U = L_U + S$ to be the total workforce in the urban sector, and $\lambda = \frac{S}{U}$ to be the share of skilled workers in the urban workforce. This change of variables is necessary, since data on the breakdown between L_U and S is not available, whereas U can be proxied by the population size of a given city. Equation (5) then becomes:

$$\frac{w_S}{w_L^U} = \left(\frac{1 - \beta}{\beta} \right) \left(\frac{1 - \lambda}{\lambda} \right) \quad (7)$$

Not surprisingly, one can show that $\frac{dU}{d\mu} > 0$, namely that the total urban workforce increases with the inflow of rural-urban migrants when institutions improve. Also, it is clear from (7) that wage inequality falls if and only if λ rises, so that the urban workforce becomes relatively more skilled. Hence, $\frac{d\lambda}{d\mu} > 0$ if and only if the condition in part (iii) of the proposition holds. (Please see Appendix A.2 for the detailed proof.)

It is worth noting here that the form of the condition in (6) is actually very general, as it is only changed slightly if we enrich the model in plausible ways. For example, suppose that as institutions improve, the cost of migration Ψ_m actually falls because transportation technologies improve with rising income levels. In Appendix A.3, I work

through this case where $\frac{d\Psi_m}{d\mu} < 0$, and show that the resulting condition for a fall in wage inequality is indeed similar to (6): The elasticity of q with respect to S must once again be sufficiently small, although the explicit expression for the upper-bound is slightly different.¹⁶ Other plausible extensions, say if Ψ_s should vary with μ , can also be accommodated in a similar way.

One can also consider the case in which a subsistence level of rural production is required to ensure that the migration of workers does not hollow out the rural sector to the point where agricultural output becomes insufficient. This can be done by specifying aggregate preferences to be of the Stone-Geary form, $(Q_R - \bar{Q}_R)^\eta Q_U^{1-\eta}$, where $\bar{Q}_R$ is the fixed minimum subsistence level of rural sector output. It turns out that the analysis of this case is mechanically identical to the proof of the Proposition, which continues to hold with only a change to the upper-bound of inequality (6). The details of this extension are fleshed out in Appendix A.4.

2.2 Effect of μ on urban real wages

What then about the relationship between institutions and the level of **real** wages? To analyze this, I introduce a convenient Dixit-Stiglitz price index for consumption, $(1)^\eta(p)^{1-\eta} = p^{1-\eta}$.¹⁷ In terms of this price index, real wages can be written as:

$$\frac{w_S}{p^{1-\eta}} = \left(\frac{1-\eta}{\eta}\right)^\eta \mu A_R^\eta A_U^{1-\eta} (1-\beta) \left(\frac{(1-\lambda)^{\beta(1-\eta)}}{\lambda^{1-(1-\beta)(1-\eta)}}\right) \left(\frac{L_R^\alpha T^{1-\alpha}}{U}\right)^\eta \quad (8)$$

$$\frac{w_L^U}{p^{1-\eta}} = \left(\frac{1-\eta}{\eta}\right)^\eta \mu A_R^\eta A_U^{1-\eta} \beta \left(\frac{\lambda^{(1-\beta)(1-\eta)}}{(1-\lambda)^{1-\beta(1-\eta)}}\right) \left(\frac{L_R^\alpha T^{1-\alpha}}{U}\right)^\eta \quad (9)$$

In what follows, I focus on the empirically relevant case where urban wage inequality falls when institutions improve, so that $\frac{d\lambda}{d\mu} > 0$. In both equations (8) and (9), an

¹⁶However, with this extension of the model, the sign of $\frac{dS}{d\mu}$ becomes ambiguous, so that we no longer have the sharp prediction that the supply of skilled workers rises.

¹⁷Specifically, this price index is proportional to the minimum expenditure function of a utility-maximizing representative consumer for this economy. It is thus equal (up to a multiplicative constant) to the minimum value of $Q_R + pQ_U$ subject to the constraint $Q_R^\eta Q_U^{1-\eta} = 1$.

increase in the supply of rural sector inputs, L_R and T , tends to increase urban real wages, as this generates a relative scarcity of inputs in the urban sector. Conversely, an increase in U that expands the urban workforce tends to depress real wages in the city. Conditional on the values of L_R , U and T , equation (9) states that real wages of urban unskilled workers will improve when institutions are strengthened, since the term $\mu \times \left(\frac{\lambda^{(1-\beta)(1-\eta)}}{(1-\lambda)^{1-\beta(1-\eta)}} \right)$ increases unambiguously. On the other hand, one cannot make a sharp prediction for real skilled wages. From (8), these will rise as long as the increase in μ more than offsets the decline in $\left(\frac{(1-\lambda)^{\beta(1-\eta)}}{\lambda^{1-(1-\beta)(1-\eta)}} \right)$.

It is easiest to convey the intuition behind these differing responses of real wages by looking at how institutions affect the relative price p . Using the change of variables introduced earlier, equation (4) can be re-written as:

$$p = \left(\frac{1-\eta}{\eta} \right) \left(\frac{A_R}{A_U} \right) \left(\frac{1}{(1-\lambda)^\beta \lambda^{1-\beta}} \right) \left(\frac{L_R^\alpha T^{1-\alpha}}{U} \right) \quad (10)$$

In Appendix A.5, I show that when $\frac{d\lambda}{d\mu} > 0$, the relative price of urban sector output necessarily falls, namely $\frac{dp}{d\mu} < 0$. When the share of skilled workers λ rises, this reduces the relative skilled wage, which in turn lowers the unit cost of production in the urban sector, and hence p as well. The price index $p^{1-\eta}$ thus falls when institutions improve. For unskilled workers, the net effect of an increase in μ will be to raise their real wages unambiguously, since the decline in the supply of unskilled labor raises their nominal wage w_L^U , while the cost-of-living falls at the same time. However, the outcome for skilled workers is less clear, since the lower relative price for urban sector output would place downward pressure on their nominal wages.¹⁸

¹⁸The use of a common price deflator for both skilled and unskilled wages requires some justification, given that Hoffman et. al. (2002) have pointed out that there were important differences in the consumption baskets of households in different income brackets prior to the 19th century. However, the Hoffman et. al. caveat applies more to the rich and the nobility, and less to wage earners including craftsmen, whose spending would likely still be heavily weighted towards necessities such as food. Also, when institutions improve and raise w_L^U , the rising cost of servants would lower the affordable standard of living of the rich and hence lower inequality even further.

2.3 Effect of μ on wage-rental ratios

As a final piece of this model, I briefly explore its implications for wage-rental ratios, since these serve as additional measures of inequality, specifically on the divide between workers and land owners. The returns to land ownership are determined in the rural sector by $r = \mu A_R(1 - \alpha) \left(\frac{L_R}{T}\right)^\alpha$, so that the wage-rental ratios are:

$$\frac{w_S}{r} = \left(\frac{1 - \eta}{\eta}\right) \left(\frac{1 - \beta}{1 - \alpha}\right) \left(\frac{T}{S}\right) = \left(\frac{1 - \eta}{\eta}\right) \left(\frac{1 - \beta}{1 - \alpha}\right) \left(\frac{1}{\lambda}\right) \left(\frac{T}{U}\right) \quad (11)$$

$$\frac{w_L^U}{r} = \left(\frac{1 - \eta}{\eta}\right) \left(\frac{\beta}{1 - \alpha}\right) \left(\frac{T}{L_U}\right) = \left(\frac{1 - \eta}{\eta}\right) \left(\frac{\beta}{1 - \alpha}\right) \left(\frac{1}{1 - \lambda}\right) \left(\frac{T}{U}\right) \quad (12)$$

Thus, with an increase in μ , $\frac{w_S}{r}$ falls as λ rises, and $\frac{w_L^U}{r}$ rises as $1 - \lambda$ falls (conditional on the levels of T and U). In particular, the model predicts that stronger institutions would help to narrow the income gap between unskilled workers and landlords; however, the gap between skilled workers and landlords would increase.

3 Empirical evidence from Europe and its periphery

The above theoretical discussion has highlighted a set of relationships between institutions and factor returns that can now be taken to the data. The empirical work below proceeds in steps to verify the predictions of the model, using a sample of 17 cities in Europe and its periphery. This sample contains a good geographic spread, including cities in Atlantic states such as London and Amsterdam, several cities in Eastern Europe such as Gdansk and Warsaw, and even Istanbul. First, I confirm that cities with stronger capitalist institutions typically had a larger population size relative to their surrounding rural population, suggesting that rural-urban migration did take place in larger volumes into such cities. Second, stronger institutions were also associated with higher real wages for both skilled and unskilled workers, even after controlling for land and population variables. Third, the impact on unskilled wages was larger than for

skilled wages, and as a result, cities with stronger property rights institutions had lower levels of wage inequality. Finally, I briefly explore the effects on wage-rental ratios.

The wage data used are from Allen (2001) for European cities and from Özmucur and Pamuk (2002) for Istanbul.¹⁹ The wages (in grams of silver per day) of building craftsmen are used as a measure of w_S , while the wages of construction laborers are used for w_L^U .²⁰ To obtain real wages, the silver wages are deflated by price indices constructed by the respective authors, based on consumption baskets comprising mainly foodstuffs, cloth and fuel items. Since we are interested primarily in the long-run equilibrium effects, I use 50-year averages of the data, for 1500-1549, 1550-1599, ..., 1850-1899, in order to net out short-term fluctuations. This provides up to 8 data points for each city.

For the key explanatory variable of interest μ , I use the institutional indices from Acemoglu, Johnson and Robinson (2002). These indices are coded on a scale of 1 to 7, with 1 being the weakest property rights regime and 7 the strongest. Two such indices are available: The “constraint” index measures the extent to which codified laws or regulations formally curbed the ability of the state to extract private wealth. On the other hand, the “protection” index looks at the degree of property rights protection actually extended to urban-based merchants. The correlation between these two indices is high (correlation coefficient: 0.88), and the results differ only slightly when either index is used. When coding the index for a given year, say 1500, the authors looked at a 40-year window and took an average over the years 1480, 1490, 1500, 1510 and 1520, in order to reduce the sensitivity of the coding to idiosyncracies in any one particular year. The indices can thus be seen as measures of initial institutions for each subsequent 50- or 100-year period, namely 1500-1549 or 1500-1599.

¹⁹Özmucur and Pamuk (2002) gathered and constructed their data in a manner similar to Allen (2001), in order to facilitate direct comparisons between the two sources.

²⁰Note that w_S and w_L^U need to be measures of wages after expropriation has taken place. This is plausible if wealth was extracted from employers and firms, so that the wages passed on to workers are net of these costs of expropriation.

3.1 Institutions and the expansion of city populations

One of the basic predictions of the model in Section 2 is that a reduction in income expropriation will prompt an increase in rural-urban migration, leading to an expansion of the urban workforce ($\frac{dU}{d\mu} > 0$). It is precisely this reallocation of labor that prompts changes in factor returns across the economy. I thus verify first that institutions have the predicted effect on the size of the urban workforce, using the city population as a proxy. Table 1 reports the results of this exercise, using estimates for the various population variables for the years 1500, 1550, through 1850, derived from Bairoch, Batou and Chèvre (1988), McEvedy and Jones (1978), and Chandler (1987). (For more details on data sources, please refer to Appendix B.)

Column 1 runs a regression of city population on initial institutions with city and period fixed effects, weighted by the total population of the state/country to which the city belongs. I use a lagged value of the institutional index (from 50 years before the city population data), in order to identify the subsequent effect on the size of cities. The results for both the “constraint” and “protection” index yield coefficients on the institutions variable that are highly significant at the 5% level. This finding that cities with better initial institutions had larger populations corroborates the central result of Acemoglu, Johnson and Robinson (2002) for this smaller sample of cities.

To see how institutions affected the location of workers across urban and rural sectors, I use the ratio of city to rural populations as the dependent variable in Columns 2-5. The institutions coefficients remain positive and significant, which suggests that there must have been more migration into cities with stronger institutions if the rate of natural increase was approximately equal in both urban and rural areas. This result does not change when the regression is run without observation weights (Column 3). Moreover, the findings are not driven single-handedly by London, the largest city in the sample by 1800; although deleting London in Column 4 reduces the institutions coefficients,

these are still significant at least at the 10% level.²¹ Column 5 examines what happens when the data for 1850 are dropped. The paradigm of rural-urban migration may well have been breaking down during this last time period, as migration across borders, particularly out of Europe into the New World, became a large-scale phenomenon (Taylor and Williamson 1997). The institutions coefficients are higher in Column 5 than in Column 2, suggesting that the model is indeed particularly relevant for the earlier years in the sample. Finally, Column 6 shows that the change in city population over each 50-year period was also positively and significantly correlated with initial institutions after controlling for the change in rural population over the same period.

Overall, there is strong evidence that cities which offered more stringent protection from expropriation supported higher levels of city relative to rural populations, suggesting that these cities must have drawn larger pools of migrants from the rural sector.

3.2 Institutions and real wages levels

How did the movement of labor into cities affect wage outcomes in the urban sector? Using equations (8) and (9) as a reference point, I test now for the effect of institutions on real wage levels.²² In what follows, I use estimates of the rural and urban population respectively as proxies for L_R and U . Although these figures are for the entire state/country and are not city-specific, they should be highly correlated to the actual numbers of rural unskilled workers and urban dwellers for a given major city. For T , I use an estimate of the agricultural land area from the Food and Agricultural Organization's (FAO) *Production Yearbook*. Given the lack of historical data on land utilization, the same figure for each country is used for every period, to serve as an estimate of

²¹The removal of the only city from Asia Minor, Istanbul, hardly changes the results from the benchmark specification in Column 2. The coefficient for the institutional variables is 0.0063 when using "constraint" and 0.0049 when using "protection", both significant at the 5% level.

²²Allen (2003) also conducts a regression analysis on the determinants of real wage levels using his own Allen (2001) data from European cities. Allen does not find a significant role for institutions, but this could be due to the fact that the variable he used was a cruder binary indicator variable for absolute rule by "princes" from DeLong and Shleifer (1993).

the potential land area that can be exploited.²³ To control for differences in η , β , A_R and A_U , all regressions are run with city and period dummies.²⁴ Insofar as the concentration of population in a city might be proportional to its ability to generate new urban technologies, a city population variable is also included to capture some of the effect of A_U . To address possible endogeneity issues, all the regressions use explanatory variables that are measures of initial conditions at the beginning of each 50-year period, so that the results reflect how a given configuration of factor supplies and protection from expropriation were associated with subsequent wage outcomes.

The empirical results for log real skilled wages come out strongly in favor of the conclusion that better initial institutions boosted the returns to skilled workers, even though equation (8) suggests that the direction of change could be ambiguous. A basic weighted regression of log real skilled wages with only institutions and city and year fixed effects on the right-hand side identifies a strong partial correlation:

$$\log \text{Real skilled wages} = \left\{ \begin{array}{ll} 0.0462^{**} & \times \text{“Constraint”} \\ (0.0182) & \\ 0.0599^{**} & \times \text{“Protection”} \\ (0.0181) & \end{array} \right\} + \text{City and period fixed effects} + \epsilon_{it}$$

These results are for the two separate regressions, using each institutional index in turn; ** indicates significance at the 5% level, while ϵ_{it} are i.i.d. normal error terms.

This effect is largely unchanged when controls for factor supplies, namely land and labor, are included in Table 2. In the benchmark regression in Column 1, the “constraint” coefficient is 0.0361, while that when using the “protection” variable is 0.0385. Since the effect of institutions works partly through the changes that it induces in L_R and U , it is not surprising that after controlling for these factor supplies, the coefficients on “constraint” and “protection” fall slightly in magnitude. Nevertheless, these variables

²³The regression results are largely unchanged when a measure of the total land area from the same FAO source is used instead.

²⁴For example, De Vries (1994) has argued that as economies grew and proto-industry expanded, workers in pre-modern Europe gradually diverted a larger share of their incomes towards purchasing these new consumer goods. This would suggest a decrease in η over time.

remain significant at the 5% level, indicating that institutions had an effect on real wages over and above their effect on factor supplies.

Reassuringly, the signs of the factor supply coefficients are broadly consistent with equation (8): Real skilled wages are positively (and significantly) correlated with the supply of rural workers, L_R , and negatively correlated to the size of the urban workforce, U . The “city population” coefficient is also positive and significant, as would be expected if this variable was correlated with A_U . The one result that is contrary to (8) is the sign of the land variable, which is negative and strongly significant in the regressions, although the equation suggests that it should be positive. This finding is consistent with the view that land-rich countries might have had fewer incentives to industrialize quickly, suggesting a feedback effect from a high T to the technology parameters A_R and A_U that is not captured by the model.

The next column of Table 1 examines some other possible forces that might have been driving income levels in Europe during this period. Column 2 addresses Acemoglu, Johnson and Robinson’s (2002) thesis on the primary importance of Atlantic trade in determining a city’s prospects for growth. The dummy variable for access to Atlantic trade (equal to 1 for England, the Netherlands, France and Spain) emerges with a positive coefficient, significant at the 10% level, but crucially, it does not alter the institutions coefficients by much. Although access to Atlantic trade did help to boost wages, the strength of property rights protection continued to be important. Column 2 also includes a control variable for the fraction of years of war, constructed from Kohn’s (1999) *Dictionary of Wars* by examining the same 40-year periods used in the coding of the institutional indices (namely 1480-1519 for the 1500 data point etc).²⁵ The results however suggest that this war variable does not add much explanatory power.²⁶

²⁵The results did not differ much when I used a 20-year window for coding instead, namely 1490-1509 for the 1500 data point etc.

²⁶When either the Atlantic dummy or the war variable is added individually to the regression, this has little effect on the size and significance of the institutions or factor supply coefficients in Column 1. This statement applies also to the results in Tables 3 and 5 for unskilled wages and the relative wage.

Figure 1 illustrates this relationship between institutions and real skilled wages. The vertical axis plots the residuals from the regression in Column 2 but excluding the “constraint” variable, while the horizontal axis plots the residuals from regressing “constraint” against the same set of covariates (weighted by the total population). The diagram clearly shows that real skilled wages were positively correlated with institutions after filtering out the effects of labor supplies, land and other controls. (The slope of the regression line in Figure 1 is 0.0352, significant at the 5% level.)

Moreover, the quantitative effect of an improvement in institutions was potentially large. A change in property rights regime from the worst possible to the best (from 1 to 7) would be associated with an average increase of $0.0361 \times 6 = 0.22$ log units in real skilled wages based on the “constraint” coefficient in Column 2. The mean log real skilled wage for the 59 data points in the sample with a “constraint” value of 1 was 1.77, and so this wholesale improvement in institutions would amount to a 12% increase on average in log real wages for such cities.

The rest of Table 1 performs some robustness checks. Column 3 shows that the results hold when the regression is run unweighted, although the significance level on “constraint” is lower. In Column 4, I remove a possible outlier, London, the only city in the sample that ever recorded a 7 on either institutional index. When this is done, both the “constraint” and “protection” coefficients actually increase, while remaining significant at the 5% level. It is thus unlikely that institutions are simply proxying for an omitted variable such as the level of industrialization, since the effect of institutions is in fact strengthened when the industrial leader, London, is deleted. Similarly, the removal of Istanbul does little to change the results from Column 2 (regression not shown). In Column 5, the coefficients more than double when the observations for 1850 are removed. The impact of institutions was thus larger when restricted to the pre-1850 period, prior to the massive outflow of labor to the New World.

Column 6 attempts to address the alternative view put forward by Glaeser et. al.

(2004) that human capital is a more fundamental determinant of long-run growth than institutions, by including measures of adult literacy rates from Allen (2003) as an additional control. This exercise should be taken as suggestive at best given the limitations of this data: Adult literacy in this context – defined as the ability to sign one’s own name – is clearly a very crude proxy for human capital, while a fair amount of estimation had to go into the construction of these figures. The results in Column 6 are nevertheless reassuring: Higher literacy rates were indeed associated with higher real skilled wages, but the institutions variables remained positive and significant, at least at the 10% level. Thus, there appears to be a role for both initial institutions and initial human capital in explaining these wage levels. Instead, it is the rural population and city population variables which dropped out of significance, suggesting that adult literacy may be accounting for some of the effect of these factor supply variables.²⁷

Next, I re-do the empirical exercise using real unskilled wages. The results in Table 3 confirm that stronger capitalist institutions were also associated with higher returns to unskilled labor, helping to validate equation (9) of the model. For starters, a simple weighted regression yields:

$$\log \text{Real unskilled wages} = \left\{ \begin{array}{ll} 0.0730^{**} & \times \text{“Constraint”} \\ (0.0175) & \\ 0.0767^{**} & \times \text{“Protection”} \\ (0.0178) & \end{array} \right\} + \text{City and period fixed effects} + \epsilon_{it}$$

When the labor and land variables are added to this regression in Column 1 of Table 3, the coefficients drop respectively to 0.0665 and 0.0627, but still maintain their statistical significance. Once again, the unskilled real wage is negatively correlated with urban population and positively correlated with the rural population, although there is still a negative sign on the land variable.

As before, these results are robust to various modifications. In particular, the institutions coefficients do not change with the inclusion of the Atlantic dummy or the war

²⁷The regression results in Column 6 are largely unchanged if either the war variable or the Atlantic dummy variable is dropped.

variable (Column 2). In fact, both the “constraint” and “protection” variables beat out the Atlantic dummy in a horse-race to see which is the more important determinant of real unskilled wages. Also, the results are not driven by the regression weights (Column 3), the London data (Column 4) or the Turkey data (regression not shown). As was the case in Table 2, the institutions coefficients increase when the 1850 data are excluded (Column 5), and remain statistically significant even with the inclusion of adult literacy (Column 6).

Importantly, the impact of institutions is much larger for unskilled than for skilled wages holding all else constant, with the institutions coefficients almost twice as large in the regressions for unskilled wages. This provides suggestive evidence that unskilled laborers gained relatively more than skilled craftsmen from better institutions. Based on Column 2, an increase from 1 to 7 on the “constraint” index would raise real unskilled wages by $0.0656 \times 6 = 0.39$ log points. Since the average log real unskilled wage for data points with a “constraint” value of 1 was 1.27, such a radical improvement in the property rights environment would have boosted log real unskilled wages dramatically by 31% on average. If one focuses on just the pre-1850 period (Column 5), this increase would have been even larger, since $0.1277 \times 6 = 0.77$, which is 61% of the mean log wage for cities where “constraint” equalled 1. Figure 2 illustrates that this positive relationship between the residuals of the log real unskilled wage and the residuals of the “constraint” index is clearly much steeper than in Figure 1. (The regression line in Figure 2 has a slope of 0.0642, significant at the 5% level.)

In summary, using either measure of the security of private property, I find that cities with stronger institutions experienced higher levels of real wage returns for both skilled and unskilled workers, after controlling for factor supplies and other city- and period-specific effects. Moreover, this effect is strongest for the pre-1850 period, during which the rural-urban migration story that drives the theoretical model was most likely to hold.

3.3 Institutions and wage inequality

The theory in Section 2, as well as the above regressions for skilled and unskilled wages, all provide good reason to believe that cities with better institutions also experienced lower levels of wage inequality. While the skilled-unskilled wage ratio is by no means a complete measure of inequality, it nevertheless serves as a reasonable indicator of inequality within labor markets. Table 4 presents the available data on this relative wage, highlighting the wide range of experiences that cities had with wage inequality. For example, Amsterdam, Antwerp, Naples and Leipzig all saw overall declines in the relative gap between 1500-1899, but Paris, Vienna, Gdansk, Warsaw and Istanbul all experienced rising wage inequality.

A simple weighted regression of the log relative wage on just the institutional variables and city and period fixed effects reveals a negative partial correlation:

$$\log \text{Relative wage} = \left\{ \begin{array}{ll} -0.0267^{**} & \times \text{"Constraint"} \\ (0.0124) & \\ -0.0169 & \times \text{"Protection"} \\ (0.0129) & \end{array} \right\} + \text{City and period fixed effects} + \epsilon_{it}$$

Although the “protection” coefficient is not significant at the 10% level, these results actually look stronger when one includes controls for factor supplies in Table 5. In Column 1 of this table, the magnitudes and significance levels improve for both “constraint” (−0.0304, significant at the 5% level) and “protection” (−0.0242, significant at the 10% level). In addition, I find significant coefficients on both “urban population” and “city population”, although these do not appear directly in equation (7). Note that neither the Atlantic dummy nor the war variable alters the negative relationship between institutions and wage inequality (Column 2).

The potential quantitative impact of a strengthening of institutions is large. Based on the “constraint” coefficient in Column 2, an increase in the institutional index from a score of 1 to 7 would result in an expected decrease of $0.0297 \times 6 = 0.18$ log points in the skilled-unskilled wage ratio. Given that the average log relative wage for cities with

a “constraint” value of 1 was 0.51, this represents an average decline in wage inequality of 35%. This negative correlation is illustrated distinctly in Figure 3 by the residual plot of the log relative wage on the “constraint” index. (The regression line has a slope of -0.0291 , significant at the 5% level.)

However, the robustness checks on the Column 2 specification raise some caveats. When the unweighted regression is run (Column 3), or when London is excluded (Column 4), the significance level of “constraint” drops to just 10% although its magnitude stays about the same, while the “protection” index drops out of significance. The removal of Istanbul changes the coefficients in Column 2 only slightly (-0.0292 with “constraint”, significant at the 5% level; -0.0233 with “protection”, marginally significant at the 10% level). Fortunately, Column 5 provides some room for comfort. When the 1850 data points are removed, the institutions coefficients in fact increase to -0.0431 and -0.0410 respectively and are both strongly significant at the 5% level. The inclusion of the adult literacy variable in Column 6 leads to a severe reduction in the significance levels of the factor supply terms, but interestingly, the institutions coefficients remain consistently significant at the 10% level, with a magnitude comparable to that in Column 5. That the literacy coefficients are imprecisely estimated suggests that institutions are the more robust determinant of wage inequality, although this could also reflect larger measurement error in the literacy variables.

Overall, I conclude that improvements in property rights institutions worked in the direction of reducing wage inequality, with unskilled workers receiving a larger boost in wage returns than skilled workers. This effect was especially strong pre-1850, before the age of mass migration to the New World.

3.4 Institutions and wage-rental ratios

This subsection considers a different measure of inequality in factor returns by looking at wage-rental ratios. The accumulation of wealth by landlords in pre-modern Europe

through the collection of rents is an important aspect of aggregate inequality that the regressions so far have not been able to address. To the extent that the rich and nobility were also landlords, wage-rental ratios may provide some sense of the gaps that existed between workers and the elite. The available data on land rents is admittedly sparse, and so I offer only preliminary results that might be indicative of how institutions affected the relative returns between labor and land.

In fact, annual rent series are available only for three countries, England, the Netherlands and France; these are matched with the wage data from London, Amsterdam and Paris respectively. The England data is taken from Clark (2002). Although there are several other sources on British land rents in the literature, Clark shows that his data correlates very well with almost all of these alternatives. The data for the remaining two countries are from O'Rourke and Williamson (2002). Since purchasing power parity conversion rates are not available, all annual rent and wage series are normalized setting 1700=1.0. This data will only be useful for the trends over time that it reveals for individual cities, but cannot be used to make valid comparisons across cities.

Table 6 summarizes the available data on wage-rental ratios. Notice first that the skilled wage-rental and unskilled wage-rental ratios move together very closely (correlation coefficient: 0.99), so the position of both skilled and unskilled workers appears very similar vis-a-vis landlords. What is particularly striking is the precipitous decline in wage-rental ratios in all three cities up to the middle of the eighteenth century, which was also documented in O'Rourke and Williamson (2002). Much of this likely coincided with the rise of enclosure, which significantly enriched the economic position of land-owners relative to laborers.

Given the relatively sparse data (just 24 data points), the regressions I run in Table 7 are very parsimonious. These use only the ratio of agricultural land area to urban population and the institutions variables as regressors, together with just city dummies. Despite this small sample, one does indeed obtain positive and significant coefficients on

$\frac{T}{U}$, as predicted by equations (11) and (12). A higher relative supply of land thus tends to reduce rents, and hence raise the wage-rental ratios. More interestingly, both the coefficients on “constraint” and “protection” are positive and significant at the 5% level in Columns 1 and 2, providing some evidence that improved protection from expropriation was associated with a narrowing in the gap between landlords and workers, both skilled and unskilled.

However, several caveats need to be raised. First, the empirical findings do not match up exactly with the theory. Equations (11) and (12) predict that a rise in μ should lead to a rise in $\frac{w_L^U}{r}$ and a fall in $\frac{w_S}{r}$. While the regressions match the prediction for $\frac{w_L^U}{r}$, they fail to do so for $\frac{w_S}{r}$, given that the two wage-rental series are so highly correlated. Second, the result is sensitive to the use of observation weights. In particular, the significance levels on the institutional variables drop out in the unweighted regressions in Columns 3 and 4. Finally, the inclusion of period fixed effects severely reduces the number of degrees of freedom. When time dummies are included in the regressions in Columns 1 and 2, the “protection” variable only remains significant at the 10% level, while the “constraint” variable is no longer statistically significant (regressions not shown).

In short, the above findings on the relationship between institutions and the wage-rental gap have to be taken as tentative. Part of the reason for the lack of robustness must lie in the fact that there are just 24 data points, and a firmer conclusion must await the availability of more data.

4 Conclusion

This paper has examined the relationships between institutions, real wages and wage inequality, within the context of the long-run growth experience of cities in Europe and its periphery. By focusing on wages instead of aggregate incomes, the study has been able to identify an additional channel of impact of improvements in property rights insti-

tutions, namely its effect on the relative wage gap between skilled and unskilled workers. Using a two-sector model with rural-urban migration, I showed how a lowering in the expropriation of factor incomes can lead to the expansion of the urban workforce, by reducing the effective barriers to migration and skills acquisition faced by rural workers. The subsequent movement of workers from the rural to the urban sector results in a narrowing of the relative skilled-unskilled wage gap, as long as the probability of an unskilled worker acquiring skills remains sufficiently high. The empirical evidence provides support for this model. Specifically, I find that cities which enforced more stringent property rights protection had higher real wages for both skilled craftsmen and unskilled laborers, a result which dovetails with earlier work on the positive impact of institutions on aggregate incomes (such as Acemoglu, Johnson and Robinson 2002). Furthermore, cities with better institutions also witnessed lower levels of wage inequality, as measured by the skilled-unskilled wage ratio.

More generally, I would like to argue that the insights of this study need not be restricted to this historical case-study of Europe. The comparative statics of the model in Section 2 rest upon the basic intuition that the incentives for workers to migrate and engage in skills training improve when institutions are strengthened, as workers would be more willing to undertake such costly investments when there is less expropriation of incomes. There is thus clearly scope for more work to be done to explore the link between institutions, growth and inequality, to see if similar relationships continue to hold in the more recent, post-World War II data.

Appendix A: Proofs of Results

Appendix A.1: Justifying $q'(S) < 0$

Why might the probability of a successful apprenticeship fall as the current supply of skilled craftsmen increases? Suppose that only a fixed fraction f of the total workforce, N , is endowed with the talent or ability to acquire skills. If the current supply of skilled workers is S , then there are only $fN - S$ potential craftsmen left among the pool of unskilled workers. Thus, the probability that an unskilled worker will successfully acquire skills conditional on the current level of S is:

$$q(S) = \begin{cases} \frac{fN-S}{L_U+L_R} = \frac{fN-S}{N-S} & \text{if } fN - S > 0 \\ 0 & \text{otherwise} \end{cases}$$

It is easy to check that $q'(S) = -\frac{(1-f)N}{(N-S)^2} < 0$ when $fN - S > 0$, namely when there are still potential entrants into the pool of skilled workers.

Appendix A.2: Proof of Proposition

Start first by substituting the expression for p from equation (4) into (1) and (2), and then rearrange these last two equations as follows:

$$\begin{aligned} \frac{1-\beta}{S} - \frac{\beta}{L_U} &= \frac{\delta\Psi_s}{q(S)} \frac{\eta}{1-\eta} \frac{1}{\mu A_R L_R^\alpha T^{1-\alpha}} \equiv C_s \\ \frac{1-\eta}{\eta} \frac{\beta}{L_U} - \frac{\alpha}{L_R} &= \frac{\delta\Psi_m}{\mu A_R L_R^\alpha T^{1-\alpha}} \equiv C_m \end{aligned}$$

Totally differentiating the above two equations in the variables L_R , L_U , S and μ , and using the fact from (3) that $dL_R = -dL_U - dS$, one obtains a pair of equations of the form $AdS + BdL_U = Cd\mu$ and $A'dS + B'dL_U = C'd\mu$, where:

$$\begin{aligned} A &= \frac{q'(S)C_s}{q(S)} - \frac{1-\beta}{S^2} - \frac{\alpha C_s}{L_R} & A' &= -\left(\frac{\alpha}{L_R^2} + \frac{\alpha C_m}{L_R}\right) \\ B &= \frac{\beta}{L_U^2} - \frac{\alpha C_s}{L_R} & B' &= -\left(\frac{1-\eta}{\eta} \frac{\beta}{L_U^2} + \frac{\alpha}{L_R^2} + \frac{\alpha C_m}{L_R}\right) \\ C &= -\frac{C_s}{\mu} & C' &= -\frac{C_m}{\mu} \end{aligned}$$

Cramer's rule then implies that $\frac{dS}{d\mu} = \frac{CB' - BC'}{AB' - BA'}$ and $\frac{dL_U}{d\mu} = \frac{AC' - CA'}{AB' - BA'}$. Using the expressions for A, A', B, B', C and C' , it is straight-forward to verify that $AB' - BA' > 0$ and $CB' - BC' > 0$. Hence, $\frac{dS}{d\mu} > 0$, establishing part (i) of the proposition. For part (ii), one can show that the sign of $AC' - CA'$, and hence of $\frac{dL_U}{d\mu}$, is ambiguous.

For part (iii), recall from (5) that $\frac{w_S}{w_L^U} = \left(\frac{1-\beta}{\beta}\right) \left(\frac{L_U}{S}\right)$. Thus:

$$\begin{aligned} \frac{d \ln(w_S/w_L^U)}{d\mu} &= \frac{1}{L_U} \frac{dL_U}{d\mu} - \frac{1}{S} \frac{dS}{d\mu} \\ &= -\frac{1}{AB' - BA'} \frac{C_s}{\mu L_U S} \left[\frac{\alpha(S + L_U)}{L_R^2} + C_m S \frac{q'(S)}{q(S)} + \frac{\alpha}{L_R} \right] \end{aligned}$$

where I have substituted in for $\frac{dS}{d\mu}$ and $\frac{dL_U}{d\mu}$, and then used the identities $\frac{1-\beta}{S} - \frac{\beta}{L_U} = C_s$ and $\frac{1-\eta}{\eta} \frac{\beta}{L_U} = C_m + \frac{\alpha}{L_R}$, to obtain the last line. It is then clear that $\frac{d \ln(w_S/w_L^U)}{d\mu} < 0$ if and only if the expression in square brackets is positive. This is equivalent after some algebra to the condition:

$$\left| \frac{Sq'(S)}{q(S)} \right| < \frac{\alpha N}{L_R^2 C_m} = \left(\frac{\mu \alpha A_R}{\delta \Psi_m} \right) \left(\frac{T}{L_R} \right)^{1-\alpha} \left(\frac{N}{L_R} \right) = \left(\frac{w_L^R}{\delta \Psi_m} \right) \left(\frac{N}{L_R} \right)$$

Thus, wage inequality falls in response to institutional improvements if and only if the elasticity of q with respect to S is sufficiently small (in absolute value).

In addition, $\lambda = \frac{S}{L_U + S}$ implies that $\frac{d\lambda}{d\mu} = -\frac{L_U S}{(L_U + S)^2} \left(\frac{1}{L_U} \frac{dL_U}{d\mu} - \frac{1}{S} \frac{dS}{d\mu} \right)$, so that $\frac{d\lambda}{d\mu} > 0$ if and only if $\frac{d \ln(w_S/w_L^U)}{d\mu} < 0$. Thus, wage inequality falls if and only if the share of skilled craftsmen in the urban workforce increases. Finally, $\frac{dU}{d\mu} = \frac{dL_U}{d\mu} + \frac{dS}{d\mu}$, which after some substitution can be shown to be positive. The total size of the urban workforce thus increases as institutions improve. QED.

Appendix A.3: Extension where $\frac{d\Psi_m}{d\mu} < 0$

The analysis of this extension proceeds in an identical fashion to Appendix A.2, except that since $\frac{d\Psi_m}{d\mu} < 0$, C' is now equal to $-\frac{C_m}{\mu} + \frac{C_m}{\Psi_m} \frac{d\Psi_m}{d\mu}$. Replacing C' throughout with this new expression, one finds that $\frac{d \ln(w_S/w_L^U)}{d\mu}$ is now given by:

$$-\frac{1}{AB' - BA'} \frac{C_s}{\mu L_U S} \left[C_m S \frac{q'(S)}{q(S)} (1 - \tilde{\Psi}) + \frac{C_m \tilde{\Psi}}{L_R} (L_R + \alpha S + \alpha L_U) + \frac{\alpha}{L_R^2} N \right]$$

where $\tilde{\Psi} = \frac{\mu}{\Psi_m} \frac{d\Psi_m}{d\mu}$. For wage inequality to fall, the expression in square brackets must be positive. Re-arranging this and using the fact that $1 - \tilde{\Psi} > 0$ (since $\frac{d\Psi_m}{d\mu} < 0$), this condition becomes:

$$\left| \frac{Sq'(S)}{q(S)} \right| < \frac{1}{1 - \tilde{\Psi}} \left[\left(\frac{w_L^R}{\delta\Psi_m} + \tilde{\Psi}\alpha \right) \frac{N}{L_R} + \tilde{\Psi}(1 - \alpha) \right]$$

The key point is that the necessary and sufficient condition for the relative wage to fall remains of the same form, namely that the absolute value of the elasticity of q with respect to S not be too large, although the expression for the upper-bound has changed. Note that this requires $\tilde{\Psi}$ to be of a relatively small magnitude, so that $\left[\left(\frac{w_L^R}{\delta\Psi_m} + \tilde{\Psi}\alpha \right) \frac{N}{L_R} + \tilde{\Psi}(1 - \alpha) \right]$ does not become negative. Also, the sign of $CB' - BC'$ is now ambiguous, so that we can no longer sign $\frac{dS}{d\mu}$ precisely. Extensions of the model in which Ψ_s changes with μ can be handled analogously, to show that the condition for wage inequality to fall remains of the general form $\left| \frac{Sq'(S)}{q(S)} \right| < K$, where K is some function of the model parameters.

Appendix A.4: Extension with Stone-Geary preferences

Now, suppose that aggregate preferences are given by $(Q_R - \bar{Q}_R)^\eta Q_U^{1-\eta}$ instead. Equations (1)-(3) of the original model continue to hold, but (4) is now replaced by:

$$p = \left(\frac{1 - \eta}{\eta} \right) \left(\frac{A_R}{A_U} \right) \left(\frac{L_R^\alpha T^{1-\alpha} - \tilde{Q}}{L_U^\beta S^{1-\beta}} \right)$$

where $\tilde{Q} = \frac{\bar{Q}_R}{A_R}$. We proceed as in Appendix A.2, substituting this expression for p into (1) and (2) to obtain:

$$\begin{aligned} \frac{1-\beta}{S} - \frac{\beta}{L_U} &= \frac{\delta\Psi_s}{q(S)} \frac{\eta}{1-\eta} \frac{1}{\mu A_R (L_R^\alpha T^{1-\alpha} - \tilde{Q})} \equiv \tilde{C}_s \\ \frac{1-\eta}{\eta} \frac{\beta}{L_U} - \frac{\alpha T^{1-\alpha}}{L_R T^{1-\alpha} - L_U^{1-\alpha} \tilde{Q}} &= \frac{\delta\Psi_m}{\mu A_R (L_R^\alpha T^{1-\alpha} - \tilde{Q})} \equiv \tilde{C}_m \end{aligned}$$

Totally differentiating this system of equations, one finds that A , A' , B , B' , C and C'

are now replaced by the following expressions:

$$\begin{aligned} A &= \frac{q'(S)\tilde{C}_s}{q(S)} - \frac{1-\beta}{S^2} - \frac{\alpha\tilde{C}_s}{L_R}\Delta_1 & A' &= -\left(\frac{\alpha}{L_R^2}\Delta_1(1+\Delta_2) + \frac{\alpha\tilde{C}_m}{L_R}\Delta_1\right) \\ B &= \frac{\beta}{L_U^2} - \frac{\alpha\tilde{C}_s}{L_R}\Delta_1 & B' &= -\left(\frac{1-\eta}{\eta}\frac{\beta}{L_U^2} + \frac{\alpha}{L_R^2}\Delta_1(1+\Delta_2) + \frac{\alpha\tilde{C}_m}{L_R}\Delta_1\right) \\ C &= -\frac{\tilde{C}_s}{\mu} & C' &= -\frac{\tilde{C}_m}{\mu} \end{aligned}$$

where $\Delta_1 = \frac{L_R^\alpha T^{1-\alpha}}{L_R^\alpha T^{1-\alpha} - \tilde{Q}} > 0$ and $\Delta_2 = \frac{\alpha\tilde{Q}}{L_R^\alpha T^{1-\alpha} - \tilde{Q}} > 0$. As in Appendix A.2, one can show that $AB' - BA' > 0$, $CB' - BC' > 0$, and that the sign of $AC' - BC' > 0$ is ambiguous. Hence, part (i) and part (ii) of the Proposition continue to hold in this extension, ie $\frac{dS}{d\mu} > 0$, but the sign of $\frac{dL_U}{d\mu}$ is uncertain. With some algebra, one can show that the condition for $\frac{d\ln(w_S/w_L^U)}{d\mu} < 0$ is now given by:

$$\left| \frac{Sq'(S)}{q(S)} \right| < \left(\frac{w_L^R}{\delta\Psi_m} \right) \left(\frac{N + \Delta_2(S + L_U)}{L_R} \right)$$

So once again, urban wage inequality falls if the elasticity of q is sufficiently small.

Appendix A.5: Proof that $\frac{dp}{d\mu} < 0$

The aim is to show that if $\frac{d\lambda}{d\mu} > 0$, then the relative price p falls when μ increases. Recall from equation (10) that $p = \left(\frac{1-\eta}{\eta}\right) \left(\frac{A_R}{A_U}\right) \left(\frac{1}{(1-\lambda)^\beta \lambda^{1-\beta}}\right) \left(\frac{L_R^\alpha T^{1-\alpha}}{U}\right)$. First, note that $\frac{dU}{d\mu} > 0$ implies $\frac{L_R}{d\mu} < 0$, so that $\frac{L_R^\alpha}{U}$ decreases when μ increases. It remains then to analyze how the term $\Lambda = \frac{1}{(1-\lambda)^\beta \lambda^{1-\beta}}$ changes in response to μ . Log-differentiating Λ with respect to μ , one obtains:

$$\frac{1}{\Lambda} \frac{d\Lambda}{d\mu} = \frac{\beta}{1-\lambda} \frac{d\lambda}{d\mu} - \frac{1-\beta}{\lambda} \frac{d\lambda}{d\mu} = \left[\frac{\beta + \lambda - 1}{\lambda(1-\lambda)} \right] \frac{d\lambda}{d\mu}$$

Using the fact that $1 - \beta = \frac{w_S S}{w_S S + w_L^U L_U} = \frac{w_S S}{w_S S + w_L^U (U - S)}$ and $\lambda = \frac{S}{U}$, one finds that $\beta + \lambda - 1 = \frac{S}{U} \left[\frac{(w_S - w_L^U)(S - U)}{(w_S - w_L^U)S + w_L^U U} \right] < 0$, where the last inequality arises because $S < U$ and $w_S > w_L^U$. It follows that $\frac{d\Lambda}{d\mu} < 0$, since $\frac{d\lambda}{d\mu} > 0$ by assumption. So p decreases when μ increases. QED.

Appendix B: Data Sources

Wages: From Allen (2001) for all cities, except Istanbul, which is from Özmucur and Pamuk (2002). The underlying annual series for Allen (2001) can be downloaded from: <http://www.econ.ox.ac.uk/Members/robert.allen/WagesPrices.htm>

Total population: From McEvedy and Jones (1978), with geometric interpolation where necessary. In units of millions of people. Cities are matched to states as follows: Antwerp (ANT) to Belgium; Amsterdam (ANS) to The Netherlands; London (LON) to England; Florence (FLO), Milan (MIL) & Naples (NAP) to Italy; Valencia (VAL) & Madrid (MAD) to Spain; Paris (PAR) & Strasbourg (STR) to France; Augsburg (AUG) & Leipzig (LEI) to Germany; Vienna (VIE) to Austria; Gdansk (GDA), Krakow (KRA) & Warsaw (WAR) to Poland; Istanbul (IST) to the Ottoman Empire.

City population: From Bairoch, Batou and Chèvre (1988) for all cities, except Istanbul, which is from Chandler (1987). In units of millions of people.

Urban and rural populations: Data on urbanization rates are from Acemoglu, Johnson and Robinson (2002). These in turn are computed from the data on city populations in Bairoch, Batou and Chèvre (1988) for Europe and in Chandler (1987) for the Ottoman Empire. There is some discrepancy in the definition of a city between these two sources. Bairoch et. al. records the populations of all cities that exceeded a size of 5,000 people at any time between 800-1800A.D., but Chandler provides only the populations of several major cities in the Ottoman Empire, implicitly using a more stringent criterion. However, this difference should have only a minor impact, since the bulk of the urban population should be captured in either case. Urbanization rates are converted to urban population estimates using McEvedy and Jones (1978). Data are obtained in this manner for 1500, 1600, 1700,

1750, 1800 and 1850, while the figures for 1550 and 1650 are filled in by geometric interpolation. The rural population is the total minus the urban population.

Agricultural land area: Sum of arable and pasture land areas in each country, as listed in the Food and Agricultural Organization’s (FAO) *Production Yearbook*, in tens of millions of hectares. This figure does not change much across different years of the *Production Yearbook*, and so a constant value is used for each country, taken from the earliest possible edition of the *Yearbook*.

Institutions: From Acemoglu, Johnson and Robinson (2002). The value of the index in a given year is the average over five observations centered on that year, for example, the 1500 data point is the average over 1480, 1490, 1500, 1510 and 1520. For earlier years, the index is only coded at century intervals. In particular, for 1550 and 1650, I use the index value from 1500 and 1600 respectively.

Dummy for Atlantic trade: Equals 1 for cities in England, the Netherlands, France, and Spain. Equals 0 otherwise.

Fraction of Years of War: Based on Kuhn (1999). Fraction of years in a given 40-year window (1480-1519 for the 1500 data point etc) during which a given country was at war or had a significant domestic rebellion.

Adult literacy: From Allen (2003), which is in turn gathered from sources listed in Table 2 of his paper. Estimates are of the proportion of the adult population that could sign their own name for each country, and are available for two points in time, 1500 and 1800. Figures for the years 1550, 1600, 1650, 1700 and 1750 were imputed by geometric interpolation, assuming a constant rate in the spread of literacy. Data is available for all states except the Ottoman Empire.

Land rents: For England, the series used is the England average in Table 8 of Clark (2002). Values are reported as period averages, so observations are attributed to

the midpoint of each period (for example, 1555 for the 1550-1559 period). The earliest data point available is for 1515 (for the period 1480-1549). For 1500-1515, I follow O'Rourke and Williamson (2002) in assuming that rents were constant. The full series is then obtained by geometric interpolation. For France and the Netherlands, data were obtained from Jeffrey Williamson via e-mail communication, based on O'Rourke and Williamson (2002). All data is normalized by setting 1700=1.0.

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Table 1
The Effect of Institutions on City Populations

Dependent variable:	City Pop. Weighted (1)	City/Rural Pop. Weighted (2)	City/Rural Pop. Unweighted (3)	City/Rural Pop. Weighted Excl London (4)	City/Rural Pop. Weighted Excl 1850 (5)	Δ City Pop. Weighted (6)
Initial Institutions ("Constraint" ; "Protection")	0.127** ; 0.103** (0.023) ; (0.022)	0.0063** ; 0.0051** (0.0013) ; (0.0013)	0.0104** ; 0.0103** (0.0021) ; (0.0022)	0.0020* ; 0.0021** (0.0011) ; (0.0010)	0.0082** ; 0.0087** (0.0018) ; (0.0019)	0.0729** ; 0.0650** (0.0151) ; (0.0149)
Δ Rural Pop.	—	—	—	—	—	0.0169 ; 0.0201 (0.0167) ; (0.0171)
Adjusted R^2	0.76 ; 0.73	0.86 ; 0.85	0.80 ; 0.79	0.88 ; 0.88	0.88 ; 0.88	0.51 ; 0.49
Number of observations	105	105	105	98	93	106

Notes: For each column, the first set of coefficients is from the regression using the constraint on executive index, while the second set uses the protection of capital index. Standard errors are in parentheses. ** denotes significance at the 5% level, and * at the 10% level. All regressions include a full set of city and period dummies. Weighted regressions use the total country population as weights.

Table 2
The Effect of Institutions on the Real Wages of Skilled Building Craftsmen

Dependent variable = ln(Real Skilled Wage)						
	Weighted (1)	Weighted (2)	Unweighted (3)	Weighted Excl London (4)	Weighted Excl 1850 (5)	Weighted (6)
Urban population	−0.216** ; −0.201** (0.039) ; (0.039)	−0.216** ; −0.202** (0.039) ; (0.039)	−0.175** ; −0.156** (0.049) ; (0.048)	−0.306** ; −0.269** (0.050) ; (0.048)	−0.185** ; −0.130** (0.063) ; (0.063)	−0.169** ; −0.133** (0.063) ; (0.064)
Rural population	0.057** ; 0.050** (0.016) ; (0.017)	0.057** ; 0.050** (0.016) ; (0.017)	0.053** ; 0.045** (0.018) ; (0.018)	0.081** ; 0.069** (0.018) ; (0.019)	0.035 ; 0.003 (0.024) ; (0.026)	0.037 ; 0.017 (0.024) ; (0.025)
City population	0.953** ; 0.908** (0.153) ; (0.156)	0.953** ; 0.908** (0.154) ; (0.157)	0.737** ; 0.674** (0.193) ; (0.196)	0.695** ; 0.690** (0.186) ; (0.190)	0.429 ; 0.345 (0.270) ; (0.274)	0.381 ; 0.318 (0.300) ; (0.302)
Agricultural land area	−0.219** ; −0.206** (0.043) ; (0.046)	−0.143** ; −0.134** (0.036) ; (0.036)	−0.129** ; −0.117** (0.031) ; (0.031)	−0.132** ; −0.124** (0.036) ; (0.037)	−0.183** ; −0.047 (0.053) ; (0.046)	0.100 ; 0.119* (0.062) ; (0.063)
Institutions variable ("Constraint" ; "Protection")	0.0361** ; 0.0385** (0.0172) ; (0.0177)	0.0359** ; 0.0382** (0.0175) ; (0.0180)	0.0364* ; 0.0451** (0.0207) ; (0.0217)	0.0572** ; 0.0485** (0.0120) ; (0.0194)	0.0846** ; 0.0887** (0.0234) ; (0.0236)	0.0486* ; 0.0548** (0.0257) ; (0.0259)
Dummy for Atlantic trader	—	0.422* ; 0.401* (0.220) ; (0.222)	0.386** ; 0.342** (0.123) ; (0.129)	0.417* ; 0.422* (0.220) ; (0.225)	−0.311 ; 0.205 (0.219) ; (0.230)	0.044 ; −0.025 (0.204) ; (0.206)
Fraction of years of war	—	0.004 ; 0.013 (0.090) ; (0.089)	0.007 ; 0.016 (0.085) ; (0.083)	0.0001 ; 0.018 (0.092) ; (0.092)	0.018 ; 0.057 (0.089) ; (0.088)	−0.048 ; −0.023 (0.092) ; (0.092)
Adult literacy	—	—	—	—	—	1.459** ; 1.390** (0.642) ; (0.639)
Adjusted R^2	0.75 ; 0.75	0.74 ; 0.75	0.69 ; 0.70	0.68 ; 0.67	0.65 ; 0.66	0.68 ; 0.69
Number of observations	118	118	118	110	106	99

Notes: For each column, the first set of coefficients is from the regression using the constraint on executive index, while the second set uses the protection of capital index. Standard errors are in parentheses. ** denotes significance at the 5% level, and * at the 10% level. All regressions include a full set of city and period dummies. Weighted regressions use the total country population as weights.

Table 3
The Effect of Institutions on the Real Wages of Unskilled Building Laborers

Dependent variable = ln(Real Unskilled Wage)						
	Weighted (1)	Weighted (2)	Unweighted (3)	Weighted Excl London (4)	Weighted Excl 1850 (5)	Weighted (6)
Urban population	−0.124** ; −0.098** (0.043) ; (0.043)	−0.124** ; −0.099** (0.043) ; (0.044)	−0.089* ; −0.058 (0.050) ; (0.050)	−0.198** ; −0.142** (0.056) ; (0.055)	−0.144** ; −0.064 (0.068) ; (0.068)	−0.126* ; −0.062 (0.071) ; (0.071)
Rural population	0.042** ; 0.033* (0.017) ; (0.018)	0.041** ; 0.032* (0.018) ; (0.019)	0.037** ; 0.024 (0.018) ; (0.018)	0.061** ; 0.043** (0.020) ; (0.021)	0.015 ; −0.031 (0.025) ; (0.028)	0.010 ; −0.024 (0.026) ; (0.028)
City population	0.502** ; 0.437** (0.167) ; (0.173)	0.503** ; 0.439** (0.168) ; (0.174)	0.348* ; 0.273 (0.197) ; (0.203)	0.290 ; 0.291 (0.208) ; (0.215)	0.094 ; −0.015 (0.288) ; (0.295)	−0.092 ; −0.195 (0.335) ; (0.337)
Agricultural land area	−0.190** ; −0.176** (0.047) ; (0.051)	−0.126** ; −0.113** (0.039) ; (0.040)	−0.117** ; −0.100** (0.031) ; (0.032)	−0.116** ; −0.105** (0.040) ; (0.042)	−0.118** ; −0.014 (0.056) ; (0.049)	0.048 ; 0.078 (0.069) ; (0.071)
Institutions variable ("Constraint" ; "Protection")	0.0665** ; 0.0627** (0.0188) ; (0.0197)	0.0656** ; 0.0615** (0.0192) ; (0.0199)	0.0625** ; 0.0625** (0.0211) ; (0.0224)	0.0861** ; 0.0692** (0.0223) ; (0.0220)	0.1277** ; 0.1296** (0.0250) ; (0.0253)	0.0910** ; 0.0951** (0.0286) ; (0.0290)
Dummy for Atlantic trader	—	0.355 ; 0.353 (0.241) ; (0.247)	0.326** ; 0.304** (0.126) ; (0.133)	0.339 ; 0.363 (0.246) ; (0.255)	−0.165 ; 0.089 (0.234) ; (0.247)	0.060 ; −0.057 (0.228) ; (0.231)
Fraction of years of war	—	0.028 ; 0.048 (0.099) ; (0.099)	0.036 ; 0.056 (0.087) ; (0.086)	0.032 ; 0.062 (0.103) ; (0.105)	0.064 ; 0.123 (0.095) ; (0.095)	0.021 ; 0.066 (0.103) ; (0.103)
Adult literacy	—	—	—	—	—	1.658** ; 1.618** (0.716) ; (0.715)
Adjusted R^2	0.70 ; 0.69	0.69 ; 0.69	0.69 ; 0.68	0.57 ; 0.55	0.62 ; 0.62	0.62 ; 0.63
Number of observations	118	118	118	110	106	99

Notes: For each column, the first set of coefficients is from the regression using the constraint on executive index, while the second set uses the protection of capital index. Standard errors are in parentheses. ** denotes significance at the 5% level, and * at the 10% level. All regressions include a full set of city and period dummies. Weighted regressions use the total country population as weights.

Table 4
The Skilled-Unskilled Wage Ratio in European Cities (1500-1899)

City	Ratio of skilled craftsmen wages to unskilled laborer wages							
	1500-1549	1550-1599	1600-1649	1650-1699	1700-1749	1750-1799	1800-1849	1850-1899
Antwerp	1.733	1.746	1.658	1.662	1.667	1.667	1.662	1.614
Amsterdam	1.452	1.489	1.444	1.400	1.315	1.293	1.315	1.313
London	1.563	1.500	1.592	1.495	1.400	1.548	1.633	1.548
Florence	1.828	1.974	2.255					1.472
Milan			1.780	1.951	1.906	1.862	2.000	1.808
Naples	2.061	1.571	1.472		1.229	1.500	1.737	
Valencia	1.548	1.288	1.193	1.493	1.509	1.490		
Madrid		1.984	2.513		2.275	2.019	2.063	1.979
Paris	1.571	1.636	1.606	1.594	1.608	1.788	1.657	1.607
Strasbourg	1.378	1.618	1.419	2.677	1.517	1.667	1.309	1.258
Augsburg	1.667	1.355	1.350	1.383	1.429	1.256		
Leipzig		1.737	1.943	1.795	1.676	1.613	1.523	1.520
Vienna	1.481	1.500	1.250	1.486	1.500	1.600	1.524	
Gdansk	1.333	2.238	1.684	1.791	1.763	1.405	1.667	
Krakow	2.000	1.793	1.235	1.414	1.500	1.310	2.167	2.239
Warsaw		1.440	1.750	1.593	2.789	2.176	2.224	2.209
Istanbul	1.754	1.711	1.618	1.517	1.557	1.862	1.830	1.995

Source: Allen (2001) and Özmucur and Pamuk (2002). Figures computed using the ratio of the averages of skilled wages to unskilled wages for each 50-year period.

Table 5
The Effect of Institutions on the Skilled-Unskilled Wage Ratio

Dependent variable = $\ln(\text{Real Skilled Wage} / \text{Real Unskilled Wage})$

	Weighted (1)	Weighted (2)	Unweighted (3)	Weighted Excl London (4)	Weighted Excl 1850 (5)	Weighted (6)
Urban population	−0.093** ; −0.103** (0.029) ; (0.030)	−0.092** ; −0.102** (0.029) ; (0.030)	−0.086** ; −0.099** (0.036) ; (0.035)	−0.108** ; −0.127** (0.039) ; (0.038)	−0.041 ; −0.067 (0.051) ; (0.052)	−0.043 ; −0.071 (0.054) ; (0.055)
Rural population	0.015 ; 0.018 (0.012) ; (0.013)	0.015 ; 0.018 (0.012) ; (0.013)	0.016 ; 0.021 (0.013) ; (0.013)	0.020 ; 0.026* (0.014) ; (0.015)	0.020 ; 0.034 (0.019) ; (0.021)	0.027 ; 0.041* (0.020) ; (0.022)
City population	0.451** ; 0.471** (0.115) ; (0.119)	0.450** ; 0.469** (0.116) ; (0.120)	0.389** ; 0.401** (0.140) ; (0.145)	0.405** ; 0.399** (0.145) ; (0.148)	0.335 ; 0.361 (0.217) ; (0.223)	0.473* ; 0.513* (0.258) ; (0.262)
Agricultural land area	−0.030 ; −0.030 (0.033) ; (0.035)	−0.017 ; −0.021 (0.027) ; (0.028)	−0.013 ; −0.017 (0.022) ; (0.023)	−0.016 ; −0.019 (0.028) ; (0.029)	−0.064 ; −0.032 (0.042) ; (0.037)	0.052 ; 0.041 (0.053) ; (0.055)
Institutions variable ("Constraint" ; "Protection")	−0.0304** ; −0.0242* (0.0129) ; (0.0135)	−0.0297** ; −0.0233* (0.0132) ; (0.0137)	−0.0260* ; −0.0174 (0.0150) ; (0.0160)	−0.0289* ; −0.0207 (0.0156) ; (0.0151)	−0.0431** ; −0.0410** (0.0188) ; (0.0191)	−0.0424* ; −0.0403* (0.0221) ; (0.0225)
Dummy for Atlantic trader	—	0.067 ; 0.048 (0.166) ; (0.170)	0.059 ; 0.038 (0.089) ; (0.095)	0.078 ; 0.059 (0.172) ; (0.175)	−0.147 ; 0.116 (0.176) ; (0.186)	−0.016 ; 0.033 (0.176) ; (0.179)
Fraction of years of war	—	−0.023 ; −0.035 (0.068) ; (0.068)	−0.029 ; −0.041 (0.062) ; (0.062)	−0.032 ; −0.044 (0.072) ; (0.072)	−0.046 ; −0.066 (0.072) ; (0.072)	−0.069 ; −0.089 (0.079) ; (0.080)
Adult literacy	—	—	—	—	—	−0.199 ; −0.229 (0.552) ; (0.555)
Adjusted R^2	0.41 ; 0.39	0.40 ; 0.39	0.39 ; 0.37	0.40 ; 0.39	0.33 ; 0.32	0.34 ; 0.34
Number of observations	118	118	118	110	106	99

Notes: For each column, the first set of coefficients is from the regression using the constraint on executive index, while the second set uses the protection of capital index. Standard errors are in parentheses. ** denotes significance at the 5% level, and * at the 10% level. All regressions include a full set of city and period dummies. Weighted regressions use the total country population as weights.

Table 6
Wage-Rental Ratios in European Cities (1500-1899)

City	Ratio of skilled craftsmen wages to land rental rates							
	1500-1549	1550-1599	1600-1649	1650-1699	1700-1749	1750-1799	1800-1849	1850-1899
Amsterdam	1.69	1.46	1.16	0.94	1.13	0.97	0.77	0.62
London	2.74	2.00	0.97	1.02	0.96	82.9	0.70	1.12
Paris	1.53	1.73	1.24	1.16	1.07	0.93	1.02	1.84

City	Ratio of unskilled laborer wages to land rental rates							
	1500-1549	1550-1599	1600-1649	1650-1699	1700-1749	1750-1799	1800-1849	1850-1899
Amsterdam	1.49	1.27	1.05	0.88	1.12	0.98	0.77	0.62
London	2.38	1.85	0.84	0.93	0.93	0.73	0.59	0.99
Paris	1.54	1.68	1.23	1.16	1.08	0.91	0.98	1.79

Source: Wage figures are from Allen (2001). Land rent figures for England are based on Clark (2002). Land rents for France and the Netherlands obtained via e-mail communication from Jeffrey Williamson, based on O'Rourke and Williamson (2002). Figures shown are 50-year averages, setting 1700=1.0; the levels of these ratios are thus not comparable across countries.

Table 7
The Effect of Institutions on Wage-Rental Ratios

Dependent variable:	ln(Skilled/Rent) Weighted (1)	ln(Unskilled/Rent) Weighted (2)	ln(Skilled/Rent) Unweighted (3)	ln(Unskilled/Rent) Unweighted (4)
(Agricultural land area / Urban population)	0.0028** ; 0.0026** (0.0006) ; (0.0007)	0.0027** ; 0.0025** (0.0006) ; (0.0007)	0.0023** ; 0.0023** (0.0007) ; (0.0007)	0.0023** ; 0.0023** (0.0006) ; (0.0007)
Institutions variable ("Constraint" ; "Protection")	0.1386** ; 0.1085** (0.0409) ; (0.0429)	0.1299** ; 0.0993** (0.0423) ; (0.0440)	0.0430 ; 0.0362 (0.0537) ; (0.0527)	0.0423 ; 0.0340 (0.0512) ; (0.0504)
Adjusted R^2	0.47 ; 0.37	0.50 ; 0.41	0.44 ; 0.44	0.49 ; 0.48
Number of observations	24	24	24	24

Notes: For each column, the first set of coefficients is from the regression using the constraint on executive index, while the second set uses the protection of capital index. Standard errors are in parentheses. ** denotes significance at the 5% level, and * at the 10% level. All regressions include city but not period dummies. Weighted regressions use the total country population as weights.

Figure 1
The Relationship between Log Real Skilled Wages and the “Constraint” Index

Notes: Slope of regression line is 0.0352, significant at the 5% level. Residuals are from a regression of log real skilled wages or the constraint index on urban population, rural population, city population, agricultural land area, the Atlantic dummy, fraction of years of war, and city and period fixed effects, weighted by total population of the country.

The Relationship between Log Real Unskilled Wages and the “Constraint” Index

Notes: Slope of regression line is 0.0642, significant at the 5% level. Residuals are from a regression of log real unskilled wages or the constraint index on urban population, rural population, city population, agricultural land area, the Atlantic dummy, fraction of years of war, and city and period fixed effects, weighted by total population of the country.

Figure 3
The Relationship between the Log Relative Wage and the “Constraint” Index

Notes: Slope of regression line is -0.0291, significant at the 5% level. Residuals are from a regression of log skilled-unskilled wage ratio or the constraint index on urban population, rural population, city population, agricultural land area, the Atlantic dummy, fraction of years of war, and city and period fixed effects, weighted by total population of the country.