Essays in Labor Economics

by

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B.Sc. Mathematics
Tel-Aviv University, 2000

Submitted to the Department of Economics
in partial fulfillment of the requirements for the degree of

Doctor of Philosophy in Economics

at the

Massachusetts Institute of Technology

June 2006

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Abstract

My dissertation is a collection of three essays that consider various aspects of income inequality and the demand for skill.

The first chapter uses the advent of the US Interstate Highway System to examine the effect of reducing trade barriers on the relative demand for skilled labor. The Interstate Highway System was designed to connect major cities, to serve national defense, and to connect the US to Canada and Mexico. As an unintended consequence, many rural counties were connected to the highway system. I find that these counties experienced an increase in trade-related activities, such as trucking and retail sales. By increasing trade, the highways raised the relative demand for skilled manufacturing workers in skill-abundant counties and reduced it elsewhere, consistent with the predictions of the Heckscher-Ohlin model.

The second chapter examines the effect of the division of labor on the demand for information processing. I find that manufacturing industries with a more complex division of labor employ relatively more clerks, who process information that is used to coordinate production. An early information technology (IT) revolution that took place around 1900 raised the relative demand for clerks in manufacturing, and significantly more so in industries with a complex division of labor. The increased demand for clerks likely contributed to the subsequent onset of the High School Movement. Interestingly, recent changes in IT have enabled firms to substitute computers for clerks, and I find evidence that this substitution occurred at a faster rate in more complex industries.

The third chapter, coauthored with Liz Ananat, examines the effect of marital breakup on the economic outcomes of women with children. We find that having a female firstborn child increases the probability that a woman's first marriage ends in divorce. Using this exogenous variation we find that divorce has little effect on a woman's average household income, but it does increase the probability that her household ends up in the lowest income quartile. While women partially offset the loss of spousal earnings by receiving more child support and welfare, combining households, and increasing their labor supply, divorce still increases the odds of household poverty.

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Acknowledgements

First and foremost, I thank my advisors, Daron Acemoglu and Joshua Angrist. Their helpful suggestions, comments, and constructive criticism have been invaluable. Their guidance has taught me important lessons for tackling difficult questions.

I also thank David Autor and Esther Duflo for helpful advice at various stages of my thesis writing. Manuel Trajtenberg and Yoram Weiss also deserve thanks for their suggestions and encouragement, especially in the early phases of my research.

Many of my fellow students at MIT have contributed in important ways to this project. I thank Todd Gormley and Marek Pycia for their friendship, encouragement, and helpful suggestions. I thank Liz Ananat, the co-author of the third chapter of my thesis, for making our joint work an enjoyable experience. I have also benefited from discussions with many other fellow students, especially Emek Basker and Gerard Padrò-i-Miquel.

This thesis has also benefited from the advice of many people outside the economics community. In particular, I thank Pamela Siska for helping me improve the clarity of my writing, and Fred Gey and Richard Weingroff for providing key data sources.

I thank my parents, Shulamit and Graham, for all their encouragement through the years. Finally and most importantly, my wife, Dahlia, and my daughter, Maya, shared this experience with me. For their support and understanding, I express my deepest gratitude.

Contents

Chapter 1

The Effect of Trade on the Demand for Skill – Evidence from the Interstate Highway System

1.1 Introduction

The effect of reducing global trade barriers on inequality has been the subject of intense debate (Freeman 2004). The Heckscher-Ohlin (H-O) framework emphasizes the role of factor endowment differences as determinants of trade patterns. In a two-factor H-O model with two economies, the removal of trade barriers favors high-skilled workers in the skill-abundant developed world and low-skilled workers in the less developed world. But recent empirical and theoretical work challenges the applicability of the H-O framework for the analysis of the labor market consequences of trade.[1]

The principal empirical challenge in assessing the general-equilibrium effect of international trade on labor demand is identification. Recent work estimates the effects of trade liberalization (Attanasio, Goldberg, and Pavcnik 2003) and exchange rate shocks (Verhoogen 2004) on labor demand in developing countries. While these case-studies are informative, they may be insufficient to determine the effect of removing trade barriers on the demand for skill. First, the consequences of trade liberalization depend on the distribution of industrial protection, while exchange rate shocks affect exporters and importers in opposite ways. Second, governments that liberalize trade or face rapid currency devaluation may affect labor

[1] The H-O model cannot fully explain recent changes in worldwide labor demand (Krugman 1995; and Berman, Bound, and Machin 1998). Against this backdrop, theoretical work suggests that reduced trade barriers may increase the demand for skilled labor even in economies with a low skill endowment (Acemoglu 2003; Kremer and Maskin 2003; Matsuyama 2004; and Antras, Garicano, and Rossi-Hansberg 2005).

markets directly. Finally, concurrent pervasive skill-biased technical change may also change the demand for skill. Taking a different approach, Borjas, Freeman, and Katz (1997) use factor-content analysis to estimate the effect of trade on wages. However, these calculations rely on a fairly restrictive set of assumptions. To better understand the effect of trade on the demand for skill, we require exogenous variation in trade barriers that affects a wide range of industries and allows us to control for other concurrent changes in the labor market equilibrium.

In this paper, I use the advent of the United States Interstate Highway System as an interesting policy experiment to estimate the effect of reducing trade barriers on the demand for skill. The construction of the Interstate Highway System began after funding was approved in 1956, and by 1975 the system was mostly complete, spanning over 40,000 miles. The highways were designed to address three policy goals. First, they were intended to improve the connection between major metropolitan areas in the United States. Second, they were planned to serve U.S. national defense. And finally, they were designed to connect with major routes in Canada and Mexico. As a consequence – but not an objective – many rural counties were also connected to the Interstate Highway System.[2] Rural counties crossed by the highways experienced an exogenous reduction in barriers to trade, providing an opportunity to examine how product market integration affects relative factor demand.[3]

I show that large trucks used the rural Interstate Highways much more intensively than other types of vehicles. As the highway construction was being completed, the trucking industry grew very rapidly and trucks became the primary mode for cross-county commerce. I find that highways increased trucking income and retail sales by about 7-10 percent per capita in rural counties they crossed, relative to other rural counties. This suggests that highway counties took advantage of the reduction in trade barriers to increase their trade

[2] An extensive literature, dating back to Fogel (1964) examines the effects of transportation infrastructure on growth. Of this literature, my approach is closest to Chandra and Thompson (2000), who estimate the effects of the Interstate Highway System on growth in rural counties.

[3] Horiba and Kirkpatrick (1981), Davis et al. (1997) and others use within-country variation in factor endowments to test various predictions of the Heckscher-Ohlin framework. But this literature has not identified exogenous variations in regional factor endowments or in trade barriers.

with other counties.

I interpret the changes in highway counties, relative to non-highway counties, as the mean effect of reduced trade barriers. In order to test the predictions of the two factor H-O framework, I interact this variation with pre-existing differences in human capital endowment between the rural counties, as proxied by the fraction of high school graduates among persons 25 years and older in 1950. I find that on average highways did not change the wage-bill of (high-skilled) non-production workers relative to the wage-bill of (low-skilled) production workers in manufacturing.[4] But in rural counties that had a highly educated workforce, highways increased the relative wage-bill of non-production workers, and where the workforce was less educated, highways decreased the relative wage-bill of non-production workers. These results are robust to the inclusion of time-varying controls for geographic location and land abundance. This finding is consistent with the H-O prediction that trade increases the demand for skill where it is relatively abundant and decreases it elsewhere.

Using my estimates I calculate the elasticity of the wage-bill of non-production workers relative to production workers in manufacturing with respect to the ratio of domestic trade to local GDP. In a county that exceeds the mean level of education by one standard deviation this elasticity is roughly equal to 1. This finding suggests that trade may contribute to changes in labor market inequality, but its effects are not very large.[5]

Another prediction of the H-O model is that trade shifts employment towards industries intensive in the relatively abundant factor. To test this prediction, I calculate a measure of skill intensity of the manufacturing workforce in each county using data on 2-digit Standard Industrial Classification (SIC) industries. I find that highways did not significantly shift employment to skill-intensive manufacturing industries in skill-abundant counties, nor did it shift employment to low-skill industries where skill was scarce.

One possible interpretation of the change in the demand for skill and the absence of a shift in the industry composition in response to reduced trade barriers is that compositional

[4] I focus on manufacturing due to the availability of wage and employment data.

[5] However, it is possible that removing trade barriers across countries may have other effects on inequality.

changes may have taken place within industries or product classes.[6] Another possibility is that trade may have also increased the demand for skilled workers in skill-abundant counties through other channels.

My interpretation that highways affect county-level outcomes by removing trade barriers faces several potential challenges. First, political agents may have changed highway routes in response to economic or demographic conditions in rural counties, contrary to the original planners' intent. In order to address this concern, I instrument for highway location using the original plan of routes proposed in 1944. I also construct a second instrument, based on the fact that an Interstate Highway is more likely to run through a rural county that lies to the north, south, east, or west of the nearest major city. Estimates using these instrumental variables (IV) are consistent with the ordinary least squares (OLS) estimates. In addition, I find that measures of trade and demand for skill do not differ significantly between highway and non-highway counties before highway construction was completed. Second, my empirical strategy assumes that counties approximate separate labor markets. This assumption is consistent with the finding of many recent studies that wage differences vary persistently across regional markets in the United States.[7] In the sample I use, about three-quarters of the workers are employed in their county of residence, suggesting that counties are plausible units for the analysis of local labor markets. My results also suggest that the effect of highways on the relative wage and relative employment of high skilled workers were the same in sign. This finding is consistent with a change in the relative demand for skill, as we would expect from opening to trade, so the effect on the relative wage bill is unlikely driven purely by migration. Finally, one might argue that highways could have affected patterns of commuting, changing the geographic skill distribution of employment. However, I find that highways had little effect on passenger car traffic, and that the fraction of workers who commute to work did not increase in highway counties relative to other counties.

[6]Schott (2004) shows evidence of specialization within product classes in international trade.

[7]See for example Acemoglu, Autor and Lyle (2004); Bernard, Redding and Schott (2005); and Card and DiNardo (2001).

Section II describes a simple theoretical framework, which considers the effects of trade on the relative demand for skilled workers. Section III presents a brief historical overview of the planning and construction of the Interstate Highway System. Section IV discusses the data and the samples I use. Section V discusses the effects of highways on trade. Section VI estimates the effect of highways on the relative demand for skilled workers, and Section VII reports estimates of their effect on the industrial composition of employment. Section VIII presents conclusions.

1.2 Theoretical Framework

To frame the key questions of this investigation, it is useful to first discuss the theoretical implications of reducing trade barriers for the wage distribution. The model I use is an extension of the Heckscher-Ohlin framework with a continuum of goods as in Dornbusch, Fischer, and Samuelson (1980). The model assumes that differences in factor endowments determine the patterns and consequences of trade. The analysis begins with a single closed economy, and then considers two economies that differ only in their endowments and trade with each other. The model predicts that trade increases demand for the relatively abundant factor and shifts employment towards industries intensive in that factor. These predictions persist even when factor prices are not equalized and when migration between the economies is possible.

Consider an economy with two factors of production: a continuum H of high-skilled workers and a continuum L of low-skilled workers. There is a continuum of goods z on the interval $[0, 1]$. The production function for each good is

$$Q(z) = F_z\left(H(z), L(z)\right), \tag{1}$$

where $H(z)$ and $L(z)$ are the employment of high- and low-skilled labor in industry z. I assume that the production functions are twice continuously differentiable, increase in each of the arguments (with diminishing marginal returns), and satisfy constant returns to scale and the Inada conditions. The goods are ranked in a strictly decreasing order of skill intensity in production and there are no factor intensity reversals.[8] I assume that all factor and product markets are perfectly competitive with profit-maximizing firms and free entry.

Each consumer is endowed with one unit of labor of her type. Consumers are assumed to have an identical Cobb-Douglas utility function:

$$U = \int_0^1 b(z) \ln d(z) dz, \tag{2}$$

where $d(z)$ is the quantity of good z consumed, and

$$\int_0^1 b(z) dz = 1. \tag{3}$$

The model thus assumes that income effects and differences in preferences play no role in determining the patterns of trade.

1.2.1 Closed Economy Equilibrium

I examine the existence and properties of a closed economy equilibrium, which is characterized by individual optimization, producer optimization and market-clearing. First, individuals maximize their utility subject to their budget constraint, so they spend a constant fraction $b(z)$ of their income on each good z at all prices and all levels of income.

Second, firms are competitive, so they maximize their profits

$$\pi(z) = P(z)Q(z) - w_H H(z) - w_L L(z), \tag{4}$$

where w_H and w_L are the wage rates for high- and low-skilled workers, and $P(z)$ is the price

[8] In other words, I assume that the ranking of industries in terms of their relative skill intensity is invariant to factor prices.

of good z. Free entry implies a zero profit condition:

$$P(z)Q(z) = w_H H(z) + w_L L(z). \tag{5}$$

Finally, the equality of supply and demand for every good implies:

$$P(z)Q(z) = b(z)\left(w_H H + w_L L\right). \tag{6}$$

Combining the last two expressions we get:

$$x(z) \equiv L(z)/L = \frac{b(z)(1+\omega h)}{1+\omega h(\omega; z)}, \tag{7}$$

where $x(z)$ the intensity of low-skilled labor in industry z relative to the economy as a whole, $\omega = w_H/w_L$ is the skill premium. The ratio of low- to high-skilled workers employed in the entire economy and in the production of good z are $h = H/L$ and $h(\omega; z) = H(\omega; z)/L(\omega; z)$. The market-clearing conditions for low- and high-skill labor are

$$\int_0^1 x(z)dz = 1 \text{ and } \int_0^1 x(z)h(\omega; z)dz = h. \tag{8}$$

Combining these expressions we get an equilibrium condition for the closed economy as a whole:

$$\phi(\omega; h) \equiv \int_0^1 \frac{b(z)(1+\omega h)}{1+\omega h(\omega; z)}\left[h(\omega; z) - h\right] dz = 0. \tag{9}$$

Since the production functions satisfy the Inada conditions, there are low values of ω such that $h(\omega; z)$ is higher than h for all z and hence $\phi(\omega; h)$ is positive at those values of ω; similarly, there high values of ω for which $\phi(\omega; h)$ is negative. Because the production functions are assumed to be neoclassical, $\phi()$ is continuous in ω. Thus there exists an equilibrium level of skill premium in autarky, $\omega = \omega^A$.

Next we note that $\phi(\omega; h)$ is strictly decreasing in ω:

$$\frac{\partial \phi(\omega; h)}{\partial \omega} < 0, \tag{10}$$

so the equilibrium skill premium is unique. The leftmost curve in Figure 1 shows the existence and uniqueness of the equilibrium skill premium in the closed home economy. Given the equilibrium skill premium, ω^A, relative price structure and the supply and demand of each good are uniquely determined.

Predictably, an increase in the skill endowment of the economy reduces the skill premium:

$$\frac{\partial \phi(\omega; h)}{\partial h}|_{\phi=0} = -1 \Rightarrow \frac{\partial \omega^A}{\partial h} = \left[\frac{\partial \phi(\omega; h)}{\partial \omega}\right]^{-1} < 0. \tag{11}$$

1.2.2 Open Economy Equilibrium

Consider opening the economy to trade with another such economy, which differs only in its factor endowments. The foreign economy has a high-skilled labor force of size H^* and a low-skilled labor force of size L^*. The foreign economy is assumed to have a lower fraction of skilled workers: $h^* \equiv H^*/L^* < h$. Figure 1 demonstrates that the equilibrium skill premium in the foreign economy, ω^{A*}, is higher than the skill premium in the home economy, as shown in (11).

First I analyze the equilibrium where the fraction of skilled workers does not differ greatly between the two economies, so trade equalizes factor prices (of course, if the fraction is identical trade has no effect). Appendix A shows that the equilibrium skill premium with trade and factor price equalization, ω^T_{FPE}, is characterized by the equation:

$$\phi(\omega^T_{FPE}; \widehat{h}) = 0, \tag{12}$$

where $\widehat{h}$ is the ratio of the stock of high-skilled workers to low-skilled workers in the two economies together, and $h^* < \widehat{h} < h$. Using (11) we conclude that the skill premium increases in the home economy (which has a high skill endowment) and decreases in the foreign economy (see Figure 1). Moreover, the effect of opening to trade on the skill premium increases with the difference in relative factor endowments. Since factor supply is assumed constant, the relative wage bill of skilled workers in the home economy, $S \equiv \omega h$, increases with trade. In the foreign economy, where skill is scarce, trade decreases the relative wage bill of skilled workers.

Because preferences for consumption goods are homothetic and identical, the skill composition of goods consumed in both economies is equal. The home economy employs more skill in production, so it must be a net exporter of skill. When trade equalizes factor prices all commodities can be produced at equal costs in both economies. The exact pattern of production (and trade) is thus indeterminate, except in one important respect: the skill-abundant economy will, on average, export skill-intensive goods. When trade leads to complete specialization, the pattern of trade is precisely determined, so the skill-abundant economy is also a net exporter of skill. We therefore conclude that opening to trade shifts production towards skill-intensive goods in the skill-abundant economy; the opposite is true for the skill-scarce economy.

Next consider the case where endowments differ sufficiently to give rise to complete specialization without equalizing factor prices in the two economies. Suppose that the home economy has a comparative advantage in producing a given good; then it has an advantage in producing all goods that are more skill-intensive. Thus, when the two economies trade, the home economy specializes in producing skill-intensive goods, while the foreign economy specializes in producing goods that are less skill-intensive. The threshold commodity, $\overline{z}$, is determined in equilibrium, such that its cost of production is equal in both economies:

$$P\left(\overline{z}\right) = P^{*}\left(\overline{z}\right) \Rightarrow w_{L} a\left(\overline{z}\right) + w_{H} c\left(\overline{z}\right) = w_{L}^{*} a^{*}\left(\overline{z}\right) + w_{H}^{*} c^{*}\left(\overline{z}\right), \tag{13}$$

where producing one unit of commodity z requires $a(z)$ units of low-skilled labor and $c(z)$

units of high-skilled labor. Since the skill premium in the home economy is lower ($\omega < \omega^*$), the threshold commodity is produced with a higher skill intensity in the home economy. The home economy produces the goods in the range $[0, \overline{z}]$ and imports goods in the range $[\overline{z}, 1]$. The total income in the home economy is $w_L L + w_H H$ and a fraction $\theta \equiv \int_0^{\overline{z}} b(z)dz$ of this income is spent on imported goods. The condition of balanced trade is therefore

$$(w_L L + w_H H)\,\theta = (w_L^* L^* + w_H^* H^*)\,(1-\theta)\,. \tag{14}$$

The equilibrium conditions in the markets for low- and high-skilled labor in the home economy are

$$\begin{aligned} L &= \int_0^{\overline{z}} \frac{a(z)b(z)\left[(w_L L + w_H H) + (w_L^* L^* + w_H^* H^*)\right]}{P(z)} dz \text{ and} \\ H &= \int_0^{\overline{z}} \frac{c(z)b(z)\left[(w_L L + w_H H) + (w_L^* L^* + w_H^* H^*)\right]}{P(z)} dz. \end{aligned} \tag{15}$$

Combining these results with the balanced trade equation we get

$$\int_0^{\overline{z}} \frac{b(z)(1+\omega h)}{1+\omega h(\omega; z)} \left[h(\omega; z) - h\right] dz = 0. \tag{16}$$

The unique equilibrium is characterized by the wage ratio in the home economy, $\omega = \omega^T$.[9] Skill intensity declines in z, so $h(\omega; z) < h$ for all $z > \overline{z}$. We therefore conclude that

$$\phi(\omega^T; h) = \int_0^1 \frac{b(z)(\omega^T + \omega^T h)}{\omega^T + \omega^T h(\omega^T; z)} \left[h(\omega; z) - h\right] dz < 0. \tag{17}$$

Comparing this result with the closed economy equilibrium and using the fact that ϕ is decreasing in the skill premium, we conclude that $\omega^T > \omega^A$. Hence opening to trade

[9]There is a unique equilibrium skill premium because the competitive equilibrium maximizes total output with respect to (L, H, L^*, H^*), subject to the two resource constraints for each economy. The four endogenously determined variables are the the skill premia in the two economies (ω^T,ω^{T*}), the ratio of skilled wages in the two economies (w_H/w_H^*), and the threshold commodity ($\overline{z}$). The problem is convex, and hence there is a unique solution.

increases the skill premium in the home economy. A similar calculation for the foreign economy indicates that it, too, has a unique skill premium, ω^{T*}, and that $\omega^{T*} < \omega^{A*}$. In other words, when the foreign economy opens to trade the skill premium in this economy declines.

Since factor endowments are assumed constant in this case, trade increases the relative wage bill of the locally abundant factor. In addition, since trade leads to specialization, each economy clearly shifts production towards goods intensive in the locally abundant factor in response to opening to trade.

1.2.3 The Effect of Trade on the Demand for Skill

The qualitative effects of opening to trade on labor markets are similar regardless of whether factor prices are equalized. First, trade increases the demand for the abundant factor. When trade becomes possible the wage-bill share of skilled workers in the skill-abundant economy increase relative to that of less-skilled workers. The opposite is true for an economy with a low endowment of skill. Second, trade shifts the composition of employment towards industries intensive in the relatively abundant factor. These predictions persist in the borderline case with complete specialization and factor price equalization.

The two predictions outlined above are robust to allowing free migration of workers between the two economies (see Appendix B). Suppose each economy has a fixed supply of housing and people spend a constant fraction of their income on housing. Assume that, all else equal, skilled workers prefer to live in the home economy.[10] If migration is possible but trade is too costly, the home economy is more skill-abundant, and therefore has a lower skill premium and higher price of housing. Opening to trade increases the demand for skilled labor in the home economy, raising the relative wage-bill share of skilled workers through inflows of high-skilled workers and outflows of low-skilled workers.

[10] If workers do not prefer either economy and migration is costless then the skill endowments of both economies are identical (assuming both are populated). The assumption of different preferences yields differences in endowment and motivates trade.

The model presented here may be extended to include capital, and its predictions hold as long as free flow of capital equalizes the interest rate. Including land as a factor of production can change the model's predictions, so in the following sections I test whether controlling for differences in land abundance affects the outcomes of interest.[11]

To examine the implications of the H-O framework for the relationship between an exogenous reduction in trade barriers and relative wages, consider two potential trading blocs. Each bloc consists of two economies, one of which is more skill-abundant. To make the link with the empirical work in the next section, I now propose to think of the economies in each trading bloc as counties. Initially, the counties in each bloc are autarkic, and relative wages are determined by local supply and demand. Highways are then constructed between the economies of one bloc, allowing them to trade. First, highways increase trade flows between those counties; trade flows equalize commodity prices, though not necessarily factor prices. Second, the trade flows increase the wage-bill of high-skilled workers relative to low-skilled workers where skill is relatively abundant, and decrease it elsewhere. Finally, trade shifts the composition of employment in skill-abundant counties towards more skill-intensive industries, and vice versa in counties where skill is scarce.

1.3 History of the Interstate Highway System

The Interstate Highway System provides a natural experiment that I envision as inducing an exogenous reduction in trade barriers. During the first half of the 20th century much of the economic activity in the United States was highly localized, as distances were long and transcontinental travel was slow. Lewis (1997) describes President Franklin Delano Roosevelt's early interest in constructing a national network of highways to reduce travel

[11] Trade may also affect the demand for skill in other ways. For example, technology embodied in goods that are traded may affect firms production possibilities. Matsuyama (2004) argues that exporting firms are more skill-intensive, so reducing trade barriers could favor skilled workers everywhere. If a similar argument applies in the domestic U.S. setting, trade may increase the demand for skilled workers even where skill is scarce. A similar result may occur if the assumption of identical and homothetic preferences is violated. For example, Leonardi (2003) argues that wealthier and more educated workers tend to consume more skill-intensive goods, so if trade increases income it could favor skilled workers.

time:

> Given his interest in road building, it is little wonder that early in 1937 [President] Roosevelt called Thomas MacDonald, chief of the Bureau of Public Roads, to the White House. On a map of the United States, the president had drawn three lines north and south and three lines east and west. These would be the routes for a new transcontinental system of interstate toll highways, he explained.

This grid pattern persisted in all the subsequent modifications of the highway plan; the next section describes how I use it to construct an instrument for the highway system.

The Federal-Aid Highway Act of 1944 laid out a plan for a system of highways designed "To connect by routes as direct as practicable the principal metropolitan areas, cities and industrial centers, to serve the national defense and to connect suitable border points with routes of continental importance in the Dominion of Canada and the Republic of Mexico."[12] Although rural areas were not considered by the planners, highways were designed to cross many rural counties as an unintended consequence of meeting these policy goals.

The construction of the Interstate Highway System began following the approval of the Federal-Aid Highway Act of 1956, which also changed planned routes of the highways.[13] The legislation stipulated that access to the highways be free, except for a few existing toll highways incorporated into the Interstate Highway System. The federal government bore 90 percent of the cost of construction, while the states financed the remaining 10 percent. Figure 3 shows that in 1966 the highways were still mostly disconnected–thick lines show constructed sections, while thin lines show planned sections. By 1975, however, almost all the sections had been completed (see Figure 4).

[12]Public Roads Administration press release (1947). Figure 2 shows the layout of the plan.

[13]In subsequent years further changes were made to the design of the system, but they were relatively minor.

1.4 Data and Samples

I use a number of data sources to construct the sample of Interstate Highways. First, the National Transportation Atlas Databases (2002) identifies the exact routes of the highways. Second, I use historical data to restrict the sample to highways that were mostly constructed from 1959-1975. I exclude state-interstate highway cells for which the 1975 mileage was less than 80 percent of the 2002 mileage.[14] Using maps issued by the Bureau of Public Roads and the Federal Highway Administration, I exclude state-interstate highway cells where the 1959 mileage exceeded 20 percent of the 1975 mileage. This selection criterion excludes toll highways, which were constructed before 1959 and later incorporated into the Interstate Highway System. Third, I restrict the sample to longer highways, which more likely connect distant locations (as envisioned by the early planners), and are therefore less affected by local economic conditions. I therefore exclude all 3-digit highways, which serve metropolitan areas, and restrict the sample to highways whose total remaining length exceeds 500 miles. This leaves most segments of 18 highways, half of which run primarily north and south and half of which run primarily east and west. Together these segments extend over more than 24,000 miles, more than half of the total length of the Interstate Highway System.

The Interstate Highways were constructed in all 48 contiguous states, but they only crossed some counties, affording substantial within-state variation. Counties are a meaningful geographic unit for the analysis of labor markets, since from 1970-1990 only about 20-30 percent of workers in rural counties commuted to work outside their county of residence. Publicly available micro data do not identify individuals' county of residence, so I use aggregate county-level data from County and City Data Books, County Business Patterns, Bureau of Economic Analysis Regional Economic Accounts, and the National Transportation Atlas Database. I limit the sample to counties whose population in 1950 was more than 50 percent rural and whose land area changed by no more than 5 percent from 1950-1980. I also exclude

[14]To determine the length of each highway in December 1975 in every state, I use the Interstate Gap Study (1976); this study is a report to Congress by the Department of Transportation.

counties that had one or more highway segments running through them, but no segment was constructed between 1959-1975.

Table 1 shows descriptive statistics for the sample of counties. Sample counties were predominantly rural in 1950, so they were more sparsely populated and somewhat poorer than non-sample counties. About three-quarters of the mileage was planned for construction on new right-of-way, most likely due to the high cost of land adjacent to existing highways. This suggests that highway counties may have been negatively selected compared to non-highway counties. But Table 1 shows that highway counties were somewhat richer and experienced faster population growth even before the construction of the highways. These differential rates of population growth motivate an analysis that compares counties in per capita terms and examines the possibility of pre-existing trends in key variables.

Although the Interstate Highways were not intended to serve rural counties, their routes may have been changed by political considerations correlated with the economic conditions that prevailed after World War II. I therefore use an indicator for having a highway planned in 1944 (z_{1c}) as an instrument for the location of the interstate highway system.[15]

I use the geographic variation in the allocation of highways to counties to generate a second instrument. Figure 5 shows a key feature of the Interstate Highway System, dating back to President Roosevelt: routes are mostly along lines of latitude and longitude. Since highways were also planned to connect cities, I calculate the orientation of the nearest large city with respect to each county's geographic centroid:[16]

$$A_c = \frac{90}{(\pi/2)} \arcsin\left(\left(\widetilde{y}_c - y_c\right) / \sqrt{\left(\widetilde{x}_c - x_c\right)^2 + \left(\widetilde{y}_c - y_c\right)^2} \right), \tag{18}$$

where (x_c, y_c) and $(\widetilde{x}_c, \widetilde{y}_c)$ are the coordinates of the county centroid and the nearest city.

[15] In concurrent and independent research, Baum-Snow (2004) looks at the effect of highways on population growth in suburban areas. He uses a 1947 map of the Interstate Highway System to construct an instrument for the routes of highways in metropolitan areas. Lahr et al. (2005) also examine the effect of highways on the size of metropolitan areas.

[16] The sample of cities is constructed using 1950 population data. It includes the most populous city in each state and any city that had at least 100,000 persons. The resulting sample includes 119 cities. I calculated the geographic centroid of each county using the Geographic Information System.

I use this measure to construct an instrument for the probability that a county received a highway: $z_{2c} = \frac{|45-|A_c||}{45}$.

Figure 6 plots a kernel regression of the probability that a highway crosses a county as a function of the orientation. If you live in a rural county and the nearest major city is to your north, east, or west, the odds of having an interstate run through your county are much better than if the city's orientation if off one of the major axes.

To test if the two instruments affect the probability that a highway crosses a county, I estimate the following cross-section regressions of the form:

$$h_c = \alpha z_c + \beta x_c + \varepsilon_c, \tag{19}$$

where h_c is an indicator for a segment of the Interstate Highway System crossing county c, z_c includes either (or both) instruments z_{1c}, z_{2c}, and ε_c is a residual. The county level controls, x_c, vary across specifications, and include region fixed effects and the distance from the county centroid to the nearest city. Table 2 shows that the instrument based on the 1944 plan is a very strong predictor of the routes along which highways were eventually constructed. The instrument based on the direction to the nearest city also has substantial predictive power for the location of highways.[17]

1.5 The Effect of Highways on Trade

In this section I estimate the effect of the Interstate Highway System on domestic trade. By allowing traffic to flow more rapidly, the highways reduced barriers to domestic trade, facilitating cross-county commerce. Since I have no data on county level imports and exports to the rest of the nation, I measure correlates of domestic commerce - trucking and retail sales. I find that trucks used rural highways very intensively, and aggregate data suggests that the

[17]I later use the instruments to test if outcomes change differentially over time for highway and non-highway counties. For that purpose I interact each instrument with a dummy for post-1975, and use this interaction term to instrument for the interaction of the highway dummy with post-1975. The first stage for the interacted regression is essentially identical to the results in Table 2.

Interstate Highway System contributed to the growth of the trucking industry. Next, I show that highway counties experienced a large increase in trucking and retail sales relative to other counties after the Interstate Highway System was completed. While highways appear to have affected trade, I show that cross-county commuting did not change differentially for highway counties relative to other counties. Finally, I discuss the implications of highways for the equalization of prices and wages.

The Interstate Highway System consists almost entirely of four-lane, divided highways with controlled and limited access. As such, it allows vehicles to travel more safely and at higher speeds in rural areas. Data from 1982-1991 suggests that the average speed of vehicles on rural Interstate Highways was at least 6-9 percent higher than the average speed on other rural principal and minor arterials and 10-15 percent higher than the average speed on rural major collectors.[18] In addition, rural Interstate Highways allow traffic to bypass small urban areas, allowing even larger time gains. Since the late 1970s, rural Interstate Highways have carried about 8 percent of the total passenger car traffic and 11 percent of single-unit truck traffic in the U.S. In contrast, rural Interstate Highways have borne over 30 percent of the total traffic of combination trucks, which are typically designed to transport large volumes over long distances (Table 3).[19] In fact, in the past couple of decades trucks account for almost one-fifth of the traffic on rural Interstate Highways. It thus appears that the Interstate Highway System has proved very important for the trucking industry.

During the 1970s, as the Interstate Highway System opened for traffic, the use of combination trucks expanded much more rapidly than in previous or subsequent decades (see Figure 7).[20] In 1969, the ratio of earnings in the trucking and warehousing industry over earnings in the railroad industry was about 1.7; by 1997 this ratio increased to almost 4.8 (Regional Economic Accounts 2004). By then trucks transported more than 71 percent of

[18] The data are from the National Transportation Statistics 1993, Table 13.

[19] A combination truck consists of a truck tractor and at least one trailer unit.

[20] Federal regulations that govern the weight and dimensions of trucks and other motor vehicles were first enacted in the Federal-Aid Highway Act of 1956. They were subsequently revised in 1975 and 1983 (U.S. House of Representatives, 2002). It is therefore highly unlikely that the increased use of combination trucks in the first half of the 1970's was caused by such regulation.

the value of domestic trade in the United States.[21] Thus, the aggregate evidence suggests that the Interstate Highway System facilitated domestic trade by allowing a more extensive use of trucks.

My interpretation of the effect of the Interstate Highway System on economic outcomes assumes that they reduced barriers to trade across counties. In 1997 most of the domestic trade in the U.S. – about 58 percent – was conducted across state borders; this is clearly a very low bound on cross-county trade. In fact, almost two-thirds of the value of commodities transported by truck were shipped for at least 50 miles and therefore, most likely, across county borders.[22] These figures are consistent with the view that the highways are important for cross-county trade.

The evidence presented thus far pertains to aggregate trends, but we can also examine the effect of highways on rural counties they crossed, relative to other rural counties. I have no county-level measure of real trucking activity, such as miles traveled or value of goods transported. The Bureau of Economic Analysis does, however, provide county-level data on earnings in the trucking and warehousing industry. I use this data to estimate specifications of the form:

$$T_{ct} = \psi_c + \rho_t + \alpha_t h_c + \varepsilon_{ct}, \tag{20}$$

where T_{ct} is log earnings in the trucking and warehousing industries per capita in county c at time t; ψ_c and ρ_t are county fixed effects and year effects; α_t is a time-varying coefficient on the indicator for highway counties, h_c; and ε_{ct} is a residual. Recall from Table 1 that highway counties are on average closer to large cities. To check that the estimates are not driven by differential trends for counties with different locations or land-abundance, some specifications control for time-varying effects of region, distance from the county centroid to the nearest city, and 1950 population density.

[21]Commodity Flow Survey (1997). If we include commodities transported by multiple modes of transportation these figures are even higher.

[22]If we the median county in the sample were a square, it would measure about 25 miles on a side. The maximum linear distance to traverse within such a square is about 35 miles.

The results (Table 4 and Figure 8) indicate that earnings in the trucking and warehousing industry, per capita, increased in highway counties relative to non-highway counties. Most of the increase took place during the 1970s, consistent with the timing of the construction of the Interstate Highway System. It is possible that non-highway counties also benefitted from the highways, although to a lesser extent. Conversely, some trucking activity may have shifted from non-highway counties to highway counties. Thus we can only identify the differential effect of highways on highway counties relative to non-highway counties.[23]

The results in Table 4 suggest that highways indeed facilitated the flow of commodities. However, these findings do not rule out the possibility that truckers reside in highway counties and transfer goods used in other counties. To further substantiate the hypothesis that highways increased the flow of commerce in counties they crossed, I estimate their effect on retail sales. Specifically, I regress log retail sales per capita on the same regressors as in (20); similar specifications control for other covariates. The results (Table 5 and Figure 8) show that highway counties experienced a rapid increase in retail sales relative to non-highway counties since the 1970s. The results also indicate that highway and non-highway counties displayed similar trends before and during the highway construction. Given that highways also increased trucking, it appears very likely that the increase in retail sales is due to goods "imported" from outside the county.[24]

In order to address the concern that the results may be affected by selection, I also estimate the effect of highways on trucking and retail sales using the instrumental variables described in the previous section. Table 6 presents estimates using specification of the form:

$$Y_{ct} = \psi_c + \rho_t + \beta d_{1975} h_c + \varepsilon_{ct}, \tag{21}$$

[23]The available data also does not account for trucks used by firms outside the trucking industry.

[24]Highways could increase retail sales per capita for a number of reasons. First, H-O theory predicts that trade will increase income, thereby raising average consumption (see Dornbusch, Fischer, and Samuelson 1980). Second, if highways facilitate market integration of rural areas, sales may shift to formal establishments, further raising sales of retailers. Finally, it is possible that retailers will make capital investments complementary to the highways, further increasing their sales.

where Y_{ct} is the outcome and d_{1975} is an indicator for post-1975, when the highway segments in the sample were mostly complete. Another specification substitutes a state-level index of highway completion for the post-1975 indicator.[25] Finally, I also estimate this equation using IV, where $z_{c,1}d_{1975}$ or $z_{c,2}d_{1975}$ serve as instruments for $h_c d_{1975}$.

OLS estimates suggest that highways increased retail sales per capita by 8-10 percentage points and IV estimates using the 1944 plan give a similar result. Using the direction instrument, rather than the 1944 plan, gives estimates that are larger and less precise. When I limit the sample to the Midwest and South, where the first stage is larger, the IV estimates are twice as large as the OLS estimates and marginally significant. The OLS estimates for earnings in trucking and warehousing per capita also increased by about 8-10 percentage points. The IV estimates are less precise, but they are similar to the OLS estimates. These results are consistent with the view that highways affected economic outcomes by changing the patterns of trade.

In the following analysis, I interpret the effect of highways on the labor market as a consequence of the removal of trade barriers. One potential concern about this interpretation is that highways may also affect patterns of commuting, thereby changing the geographic skill distribution of employment. Figure 7, however, suggests that rural Interstate Highways had little aggregate effect on passenger traffic. The final outcome in Table 6 is the fraction of workers, who commute to work outside their county of residence. The results show that commuting did not significantly increase in rural highway counties, compared to other rural counties.

A related concern is that migration patterns may be correlated with highway location for reasons other than a change in labor demand. My estimates suggest that highway counties

[25] The benchmark specification assumes that highways affected outcomes only after 1975, when they were essentially complete. In order to relax this assumption, I calculate a state-level index of highway completion using the length of rural interstate highways with four lanes and restricted access control that were open to traffic. Since most of the interstate highways that were open to traffic in 1960 were toll roads incorporated into the system, I exclude them from my analysis. Thus, the state-level index measures if the mileage of highways in a given year, net of the 1960 mileage, accounts for more than 90 percent of the difference between the 1975 mileage and the 1960 mileage.

experienced a faster rate of population growth both before and after the highways were constructed, with no evidence of a change in trend (results not shown). In the next section I discuss the possibility that highways changed the relative supply of skill, rather than the relative demand.

The theoretical framework also predicts that where costless trade is possible, commodity prices (though not necessarily factor prices) will be equal. The Interstate Highway System appears to have significantly reduced the cost of trade in commodities. While I have no direct evidence on price changes in rural areas, Parsley and Wei (1996) use data from 1975-1992 and find rapid convergence of commodity prices across U.S. cities. This finding is consistent with the theory, since all major U.S. cities are connected to the Interstate Highway System. Bernard, Redding, and Schott (2005) find that wage differences persist across geographically disparate labor markets in the U.S.; this suggests that factor endowment differences may have prevented factor price equalization.

1.6 The Effect of Highways on the Relative Demand for Skilled Labor

This section tests the H-O prediction that by facilitating trade, highways increase the relative wage-bill of non-production workers in counties with a highly skilled workforce and decrease it in counties with a less educated workforce.[26] To test this prediction I interact the exogenous reduction in the cost of trade caused by the Interstate Highway System with pre-existing differences in human capital endowment. As explained in Section IV, there are no micro data that identify individuals' county of residence during the relevant time period. I therefore use the fraction of high school educated workers among persons 25 years and older in 1950 (before the Interstate Highway System was constructed) as a measure of a county's skill endowment. I use non-production and production workers in manufacturing as proxies for

[26] As Table 1 shows, the sample counties are on average less skill abundant than the rest of the US. However, there is considerable variation in the sample counties' skill endowment, and about a quarter of the sample of counties had a higher fraction of high-school graduates than the US average in 1950.

high- and low-skilled labor, respectively.[27]

In order to examine this prediction I estimate a regression of the form:

$$\ln\left(S_{ct}^{H}\right) = \psi_c + \rho_t + \beta d_{1975} h_c + \gamma d_{1975} h_c s_{c,1950} + \delta d_{1975} s_{c,1950} + \varepsilon_{ct}, \tag{22}$$

where $\ln\left(S_{ct}^{H}\right) = \ln\left(\omega_{ct}^{H} h_{ct}\right) = \ln\left(w_{ct}^{H}/w_{ct}^{L}\right) - \ln\left(H_{ct}/L_{ct}\right)$ denotes the wage-bill of non-production workers in manufacturing, relative to production workers. The fraction of high school graduates among persons 25 years and older in 1950 is $s_{c,1950}$, and d_{1975} is a dummy for post-1975.[28] Other specifications include county-level covariates, and IV estimates using $z_{1,c}$ and $z_{2,c}$, interacted with appropriate terms, to instrument for terms that include the highway dummy, h_c.

The first column of Panel A in Table 7 presents estimates of this equation under the constraint that $\gamma = \delta = 0$. The results show that on average, highways did not increase the relative demand for skill. Subsequent columns relax this constraint, and test the prediction that $\beta < 0$ and $\gamma > 0$. These results do suggest that highways significantly increased the relative demand for non-production workers in counties that had a highly skilled labor force and reduced it elsewhere. The results are robust to controlling for the contemporaneous fraction of high school graduates in the labor force and for time-varying coefficients on region dummies, distance from the county to the nearest city, and 1950 population density. The IV estimates using the 1944 plan are somewhat larger estimates than the OLS estimates, while the direction instrument is not precise enough to identify the effect of highways on labor demand.[29]

[27] Census data for 1960 and 1980 indicate that non-production workers in manufacturing industries had about 2-3 more years of education than production workers. For further discussion of the differences between production and non-production workers see Berman, Bound and Griliches (1994).

[28] These regressions are weighted by 1950 population, since data for low population counties are less precise.

[29] To test the hypothesis that the effect of highways on the relative wage-bill of skilled manufacturing workers varies by land endowment I add the interaction (1950 population density)*(post 1975)* highway to the specification in Table 7, panel A, column 3. This changes beta from -0.168 (0.068) to -0.177 (0.070) and gamma from 0.609 (0.241) to 0.605 (0.244), suggesting that land endowment has little effect on the outcome of interest. Taking the same baseline specification and controlling for a county-level index of skill intensity in manufacturing in 1967 (described in the Appendix) interacted with year dummies yields estimates of -0.153 (0.074) and 0.462 (0.253) for beta and gamma, indicating that these results are also not driven by differential

The results in Panels B and C of Table 7 suggest that highways increased both the relative wages and the employment share of non-production workers in high-skill counties, although most estimates are not precise. Similarly, highways appear to have reduced the wages and employment of non-production workers where skill was relatively scarce. These results are consistent with the H-O view, that trade shifts the relative demand curve for skilled labor. The change in wages and employment shares may reflect a movement along the relative supply curve for skilled workers. The positive and finite elasticity of the relative supply of skill may reflect endogenous cross-county migration as well as changes that took place within counties, such as occupational transition and entry or exit of workers from the market.

I use the results in Table 7 to evaluate the possibility that the effect of highways on the wage-bill are due to changes in relative supply correlated with highway location, rather than a change in relative demand. Assume that the aggregate elasticity of substitution between high- and low-skilled workers is locally a constant, η, then the elasticity of the relative wage-bill of skilled workers with respect to their wages is:

$$\frac{\partial \ln \left(w_{ct}^{H}/w_{ct}^{L}\right)}{\partial \left[\ln \left(H_{ct}/L_{ct}\right)\right]} = -\frac{1}{\eta}. \tag{23}$$

There is a consensus in the literature that η is higher than 1 (e.g. Freeman 1986; Katz and Murphy 1992; and the elasticity implied in Angrist 1995). Using these estimates, the effect of highways (and highways interacted with 1950 schooling) on relative employment should have been bigger in magnitude than their effect on relative wages, and opposite in sign. Thus, the estimates in Table 7 suggest that highways changed the relative demand for skill, rather than the relative supply of skill.

As a further check, I test if highway counties experienced changes in the wage-bill of non-production workers, relative to production workers, before the highways were completed or after they were already in place. By time-differencing (22) (post-1975 minus pre-1975) we

time effects on counties with different industry compositions.

get an expression for the change in the relative wage-bill of non-production workers:

$$\Delta \ln \left(S_c^H\right) = \rho + \beta h_c + \gamma h_c s_{c,1950} + \delta s_{c,1950} + \varepsilon_c. \tag{24}$$

I estimate this equation using OLS (with and without county-level controls) and IV, instrumenting h_c and $h_c s_{c,1950}$ using $z_{i,c}$ and $z_{i,c} s_{c,1950}$, where $i = 1, 2$. Since data is not available for all counties and all years, I restrict myself to a constant sample of counties, for which I have data in 1947, 1967, 1982, and 1992. The results (Table 8) show that before the construction of highways was complete (1947-1967) the changes in the relative demand for skill did not vary significantly between highway and non-highway counties. This is true for both the OLS estimates (with and without county-level controls) and the IV estimates using the 1944 plan. As before, the direction instrument is not powerful enough to give precise results. The estimates for 1967-1982 are similar in magnitude to those found in Table 7. The estimate of the main effect of highways, γ, is statistically significant only in the IV estimate, while the estimate of δ is significant in all specifications. Finally, the changes that took place from 1982-1992 do not vary significantly across highway and non-highway counties. These results lend support to the hypothesis the Interstate Highway System was indeed the cause of the changes in the relative demand for skilled labor.

In order to assess the magnitude of the effect of trade I compare my estimates to those of Borjas, Freeman and Katz (1997). Borjas et al. use a factor-content approach to measuring the effect of trade on wage inequality. They find that imports to the U.S. from less-developed countries as a fraction of GDP increased by about 1.6 percentage points from 1980-1995. Using factor-content analysis they calculate that this increase could have raised the skill premium by about $0.9 - 1$ percentage points.[30] This suggests that the elasticity of the skill premium with respect to (Imports/GDP) was about 0.6.

Using my estimates for trucking (Table 4) I assume that highways increased trade by

[30] This figure reflects the change in wages of college graduates relative to high-school graduates and of high-school graduates compared to high-school dropouts.

about 7 percentage points in counties they crossed, relative to other counties. Data from the commodity flow survey of 1997 suggests that value of goods traded domestically in the U.S. was roughly equal to the GDP. I assume that during the 1970s the ratio of domestic trade to GDP was about 0.9. This suggests that the change in (domestic trade/local GDP) induced in highway counties, relative to other counties, was about 0.063. In 1950, the fraction of high school graduates in the mean county in the sample was 0.251 with a standard deviation of 0.103. The OLS estimate of the effect of highways on the relative wage-bill of high-skilled workers (Table 7 Panel A, column 2) in a county that is one standard deviation above the mean level of education is 0.072 (p-value 0.03). The IV estimate (Table 7 Panel A, column 6) is 0.061 (p-value 0.12). This suggests that the elasticity of the relative wage-bill with respect to (domestic trade/local GDP) for a county that is one standard deviation above the mean level of education is close to 1.[31] The estimates of the skill premium are somewhat smaller in magnitude and less precise.

The assumptions used in this paper are quite different from those used in the factor-content analysis of Borjas et al. (1997).[32] However, despite the differences in assumptions and sources of variation, the magnitude of our estimates appears quite comparable. My results therefore support the view that while trade may contribute to changes in labor market inequality, its effects are limited in magnitude.

1.7 The Effect of Highways on the Industrial Composition

The theoretical framework also predicts that trade will change the industry composition of employment. Specifically, it predicts that trade causes a skill-abundant economy to shift its production towards more skill-intensive goods and vice versa for an economy where skill is scarce. In order to test this prediction I construct a measure of the skill intensity of

[31] Similar calculations yield estimates that are close to 0 for a county with the mean level of education and about -1.3 for a county that is one s.d. below the mean level of education.

[32] Factor-content analysis typically assumes the absence of non-competing imports, no endogenous response of factor supplies to trade, and identical elasticities of substitution across all production functions and the utility function (Panagariya 2000).

each industry. I match the two-digit SIC codes to the 1950 classification of manufacturing industries in the household census and compute the fraction of non-production workers each industry's labor force. Using the County Business Patterns data I compute an index of the skill intensity of the manufacturing workforce: $I_{ct} = \sum_i n_{i,1960} I_{ict}$, where $n_{i,1960}$ is the fraction of non-production workers employed in industry i in 1960 and I_{ict} is the fraction (or estimated fraction) of the manufacturing employees in county c employed in industry i at time t. See Appendix C for details on the construction of the data.

To test whether trade changed the industrial composition of employment in manufacturing, I estimate regressions of the following form:

$$I_{ct} = \psi_c + \rho_t + \beta d_{1975} h_c + \gamma d_{1975} h_c s_{c,1950} + \delta d_{1975} s_{c,1950} + \varepsilon_{ct}. \tag{25}$$

I estimate the equation using OLS and IV, instrumenting the various interactions of the highway dummy with corresponding interactions of the two instruments. The estimated coefficients of interest, β and γ, are of the expected sign, but they are not statistically significant in any of the specifications (Table 9). Thus we cannot reject the hypothesis that trade has no effect on the industrial composition of employment.[33]

There are two possible ways to interpret the absence of significant effects of removing trade barriers on industrial composition. One approach is to interpret these results as evidence that the theoretical framework outlined in section II may be incomplete. For example, there may be frictions that restrict the mobility of labor across industries. My findings may also suggest that endogenous migration may have played only a limited role, since migration is likely to have reinforced the effects of trade on industrial composition (see Appendix B).[34] In related work, Goldberg and Pavcnik (2004) survey a number of recent studies that find very little effect of tariff reductions on industry composition in developing countries. These

[33]Note that even if factors are not perfectly mobile across industries, trade could still affect the relative demand for skill by changing the relative prices of commodities.

[34]However, Card and Lewis (2005) find that even inflows of low-skilled workers had limited effects on the industrial composition in U.S. cities. Lewis (2005) argues that firms may vary the skill intensity of the production technique in response to migration.

studies attribute their findings to imperfections of product markets or labor markets.

Alternatively, it is possible that my estimation strategy is not precise enough to estimate such effects.[35] For example, it may be that changes in labor demand have taken place at lower levels of industry aggregation or even within product classes (Schott 2004). Moreover, the absence of accurate employment data in many county-industry cells requires a process of imputation that may have resulted in non-negligible measurement error (see Appendix C). Further research may be needed to determine the extent to which the removal of trade barriers affects the industrial composition.

1.8 Concluding Remarks

The literature on international trade suggests that trade may affect the demand for skill, but it has proved difficult to identify this effect, since identification requires exogenous variation in the barriers to trade. In this paper I use the advent of the U.S. Interstate Highway System as a source of exogenous variation in trade barriers. The Interstate Highway System was built to better connect large cities, to serve national defense, and to connect with major routes in Canada and Mexico. As an unintended consequence of meeting these objectives, the highways crossed many rural counties. I find that the rural Interstate Highways were particularly important for the flow of large trucks. These highways increased trucking activity and retail sales by about 7-10 percentage points per capita in rural counties they crossed, relative to other rural counties.

Using the Interstate Highway System as a source of variation for trade, I test whether trade affected the demand for skill in rural areas. I find that on average, highways had no effect on the demand for high-skilled workers relative to low-skilled workers in manufacturing. However, highways increased the wage-bill of high-skilled workers relative to low-skilled workers in counties where skill was abundant, and reduced it where skill was scarce. This finding is consistent with the Heckscher-Ohlin view that trade increases the relative demand

[35]Revenga (1992) finds that U.S. industries faced with increasing import competition due to changes in exchange rates did reduce their employment and relative wages.

for the abundant factor. However, the magnitude of the effect is quite small: in a county that exceeds the mean level of education by one standard deviation, the elasticity of the wage-bill of non-production workers relative to production workers in manufacturing with respect to the ratio of domestic trade to local GDP is roughly equal to 1. In addition, I find no evidence for the prediction of the Heckscher-Ohlin model that trade significantly shifts the industrial composition of employment towards industries intensive in the abundant factor. This result suggests that changes in skill composition in response to reduced trade barriers may have taken place within industries or product classes, or that trade may have increased the demand for skill in skill-abundant counties through other channels.

My findings suggest that the ongoing expansion of trade between economies that differ in their skill endowment, such as trade between the developed world and the less-developed world, may continue to contribute to changes in labor market inequality. However, my results also indicate that opening to trade is not likely to explain a great deal of the variation in the demand for skill experienced by many countries in recent years.

1.9 Appendix A. Open Economy with Factor-Price Equalization

In this appendix I derive the open-economy equilibrium with equalized factor prices. Denote the home economy's share of low-skilled labor by $\xi = L/(L+L^*)$. The ratio of the stock of high-skilled workers to low-skilled workers in the two economies together

$$\widehat{h} \equiv \frac{H+H^*}{L+L^*} = \xi h + (1-\xi)h^*. \tag{26}$$

Since factor prices are equal in both economies and the production technology is assumed to be identical, the factor requirements in producing both goods are the same in the two countries.

The equilibrium conditions for high-skilled labor in the two countries are:

$$\begin{aligned} H &= \int_0^1 \frac{\alpha(z)b(z)c(z)\left[(w_L L + w_H H) + (w_L L^* + w_H H^*)\right]}{P(z)} dz \text{ and} \\ H^* &= \int_0^1 \frac{(1-\alpha(z))\, b(z)c(z)\left[(w_L L + w_H H) + (w_L L^* + w_H H^*)\right]}{P(z)} dz, \end{aligned} \tag{27}$$

where $\alpha(z)$ is the fraction of total output of good z produced in the home economy. Producing one unit of commodity z requires $a(z)$ units of low-skilled labor and $c(z)$ units of high-skilled labor.

Combining the equations above we get an expression for the stock of high-skilled workers relative to low-skilled workers:

$$\int_0^1 \frac{b(z)h(\omega;z)(1+\omega\widehat{h})}{1+\omega h(\omega;z)} dz = \widehat{h}. \tag{28}$$

Similarly, by adding the equilibrium conditions for low-skilled labor we find that

$$\int_0^1 \frac{b(z)(1+\omega\widehat{h})}{1+\omega h(\omega;z)} dz = 1. \tag{29}$$

Putting together the results for both factors we have

$$\phi(\omega_{FPE}^T;\widehat{h}) = \int_0^1 \frac{b(z)(1+\omega_{FPE}^T\widehat{h})}{1+\omega h(\omega_{FPE}^T;z)}\left[h(\omega_{FPE}^T;z)-\widehat{h}\right]dz = 0, \tag{30}$$

where $\omega = \omega_{FPE}^T$ is the equilibrium skill premium with trade and factor price equalization.

1.10 Appendix B. Open Economy with Endogenous Migration

This Appendix extends the analysis to account for the possibility that workers migrate in response to the change in relative wages induced by the opening to trade. I assume that people differ in their preferences for living in either of the two economies, and that their preferences are correlated with their skill:

$$U_{ijk} = \theta_1 \int_0^1 b(z)\ln d_i(z)dz + (1-\theta_1)\ln\widetilde{q}_i + \theta_2 I_{ij} I_{ik}, \tag{31}$$

where U_{ijk} is the utility of a person i with skill level $j \in \{H, L\}$, and the subscript k denotes home or foreign economy. The consumption of good z is denoted by $d_i(z)$ and housing is denoted by $\widetilde{q}_i$. I_{ij} is an indicator for whether person i is skilled, I_{ik} is an indicator for living in the home economy and θ_2 is the preference of skilled workers for living in the home economy. I assume that the supply of housing in each of the economies is constrained by a fixed supply of land, $\widetilde{q}_k$:

$$\int_i I_{ik}\widetilde{q}_i di = \widetilde{q}_k. \tag{32}$$

A group of high-skilled workers of measure 1 residing in economy k consumes $d^H(z) = \theta_1 b(z)\frac{w_k^H}{P_k(z)}$ units of good z and $\frac{\omega_k\widetilde{q}_k}{(1+\omega_k h)L_k}$ units of housing. Similarly, a group of low-skilled workers of measure 1 residing in economy k consumes $d^L(z) = \theta_1 b(z)\frac{w_k^L}{P_k(z)}$ units of good z and $\frac{\widetilde{q}_k}{(1+\omega_k h)L_k}$ units of housing. If all goods are produced at home (or if trade equalizes factor prices) then $d^H(z) = \theta_1 b(z)\frac{\omega_k f_z(h(\omega_k;z))}{1+\omega_k h(\omega_k;z)}$ and $d^L(z) = \theta_1 b(z)\frac{f_z(h(\omega_k;z))}{1+\omega_k h(\omega_k;z)}$.[36] Hence their

[36]To derive the last two equations notice that $P_k(z) = a(z)w_k^L + c(z)w_k^H$, $h(z) = c(z)/a(z)$ and $a(z) = [f_z(h(\omega_k;z))]^{-1}$.

indirect utility of type j residing in country k as a function of the skill premium and the population of low-skilled workers is:

$$U_{ijk}(\omega_k, L_k) = \theta_1 \int_0^1 b(z) \ln\left(\theta_1 b(z) \frac{(\omega_k I_{ij} + (1 - I_{ij})) f_z(h(\omega_k; z))}{1 + \omega_k h(\omega_k; z)}\right) dz + \\ + (1 - \theta_1) \ln\left(\frac{(\omega_k I_{ij} + (1 - I_{ij}))}{(1 + \omega_k h) L_k} \widetilde{q}_k\right) + \theta_2 I_j I_k, \tag{33}$$

where $f_z(h(\omega_k; z)) = F_z(H(z), L(z))/L(z)$ is output per low-skilled worker in producing good z.

In order to analyze the effect of trade on the utility of both types, consider the utility of a representative agent in economy k, who owns 1 unit of low-skilled labor and h units of high-skilled labor.[37]

$$U_{ik}(\omega_k, L_k) = \theta_1 \int_0^1 b(z) \ln\left(\theta_1 b(z) \frac{(\omega_k h + 1) f_z(h(\omega_k; z))}{1 + \omega_k h(\omega_k; z)}\right) dz + \\ + (1 - \theta_1) \ln\left(\frac{\widetilde{q}_k}{L_k}\right). \tag{34}$$

Assume that production functions are Cobb-Douglas and trade equalizes factor prices (before migration takes place), so the elasticity of substitution between high- and low skilled labor is one. In this case the relative wage bill of skilled workers in each industry, $\omega_k h(\omega_k; z)$, is constant. The first welfare theorem implies that the choice of inputs maximizes the welfare of the representative agent (34). Since $\omega_k h(\omega_k; z)$ is constant, the choice of inputs maximizes: $\int_0^1 b(z) \ln f_z(h(\omega_k; z))\, dz$ given ω_k. Therefore, using the envelope theorem:

$$\frac{\partial}{\partial \omega_k}\left[\int_0^1 b(z) \ln f_z(h(\omega_k; z))\, dz\right] = 0. \tag{35}$$

Using (35), we conclude that when production functions are Cobb-Douglas, $\frac{\partial U_{iHk}(\omega_k)}{\partial \omega_k} > 0$ and $\frac{\partial U_{iLk}(\omega_k)}{\partial \omega_k} < 0$, so high-skilled workers prefer a higher skill premium, while low-skilled

[37]To simplify the analysis I assume that the representative agent derives no utility from living in either economy.

workers prefer a lower skill premium. Since $\frac{\partial U_{iHk}(\omega_k)}{\partial L_k} < 0$ and $\frac{\partial U_{iLk}(\omega_k)}{\partial L_k} < 0$, both types prefer a lower population of low-skilled workers in their economy (holding relative wages fixed), because a higher population implies higher housing prices. By continuity, these results also hold if the elasticity of substitution in production of all goods is sufficiently close to one.

In order to analyze the equilibrium, it is convenient to begin by considering the case where free migration is possible but there is no trade between the two economies. I focus on the case where workers choose to live in both economies.[38] If $\theta_2 = 0$ there is a symmetric equilibrium in which both economies have an equal number of workers. To prove the existence and uniqueness of this equilibrium, consider the conditions for indifference of the two types of workers between the two economies. As figure A1 shows, the indifference curve of high-skilled workers between the two economies (U_H) is upward sloping in the space of the home economy skill premium, ω, and the relative supply of low-skilled labor L/L^*, while the indifference curve for low-skilled workers (U_L) is downward sloping. Holding all else constant, as ω approaches zero high-skilled workers require a lower level of L/L^* than low-skilled workers to be indifferent between the two economies. Similarly, if ω is high enough, low skilled workers require a lower level of L/L^* than high-skilled workers for indifference. Since the utility functions are continuous, there is a unique equilibrium combination of ω and L/L^*. Given a fixed aggregate supply of high- and low-skilled labor in both economies together, ω and L/L^* determine a unique level of h in the home economy. This, in turn, implies a unique level of employment of skilled workers in the foreign economy and hence a unique skill premium ω^*.

Now consider the case where high-skill workers prefer to live in one of the two economies ($\theta_2 \neq 0$). Without loss of generality I assume that $\theta_2 > 0$, so ceteris paribus high-skilled workers prefer to reside in the home economy. The indifference curve for the high-skilled

[38] There are also two other equilibria in which all the workers reside in either of the two economies. In this case no worker has an incentive to migrate, because production (and positive wages) are only possible with both types of workers.

workers is below the indifference curve corresponding to $\theta_2 = 0$ (see indifference curve U'_H Figure 1). In equilibrium the skill premium ω and the relative employment of low-skilled labor (L/L^*) are lower. High-skill workers are indifferent between the two economies because the price and wage differentials offset their preference for the home economy. Low-skilled workers are also indifferent because in the home economy they have higher wages and a higher price of housing.

Suppose that $\theta_2 > 0$ and the economies open to trade with each other. Assume that there is no initial response of migration and that factor endowments do not differ too much, so initially trade equalizes factor prices. The equilibrium analysis is identical to the cases outlined without migration, except that now the fraction of income spent on each good z is a fraction θ_1 of its share when there was no expenditure on housing. As we saw, trade raises the skill premium in the home economy and decreases it in the foreign economy. Since factor prices are equalized, high-skilled workers migrate to the home economy, raising the price of housing. Therefore low-skilled workers migrate to the foreign economy. In equilibrium, the wage premium in the home economy must be lower than in the foreign economy (otherwise all high-skilled workers choose to reside in the home economy, as do all low-skilled workers). However, if we assume that on aggregate high- and low-skilled labor are not gross complements, trade increases the wage-bill share of the abundant factor in each economy.

In summary, when migration is costless trade increases the wage-bill share of high-skilled workers in the skill intensive economy by increasing their relative employment. To the extent that migration is costly, the increase is mainly due to a change in relative wages. Migration also induces a larger change in the composition of production due to Rybczynski effects: an increase in the share of high-skilled workers increases the skill content of production and exports of the home economy.

1.11 Appendix C. Industrial Composition Data

The data on each county's employment in 2-digit SIC industries are from the County Business Patterns data set. For 1977-1997 I use publicly available data from the University of Virginia, and for 1967 and 1972 I use data from the University of California Berkeley Data Center. There are two important differences between the two sources of data. First, industries were re-classified in the 1970s, so I exclude from the data SIC 19 (Ordinance), which only exists for the earlier years. Second, the earlier data are reported only for county-industry cells with 100 employees or more or at least 10 establishments. Moreover, the data reports employment by establishment size categories, which have changed slightly over time, and exact overall employment is not reported for all counties. In order to solve these problems I assume that the employment in an establishment with a given size category is the geometric average of the category's two limits. I then use a regression to predict the employment in establishments in the largest size category for each year. When total employment is not available I predict it using a regression. Finally, to ensure comparability over time, I exclude county-industry cells with less than 10 establishments or with fewer than 100 workers (or fewer than 100 predicted workers) for 1977-1997. Using this data I calculate the fraction of manufacturing employees in each county, industry, and year.

References

Acemoglu, Daron. "Patterns of Skill Premia." *Review of Economic Studies,* 2003, 70(2), pp. 199-230.

Acemoglu, Daron; Autor, David H. and Lyle, David. "Women, War, and Wages: The Effect of Female Labor Supply on the Wage Structure at Midcentury." *Journal of Political Economy,* June 2004, 112(3), pp. 497-551.

Angrist, Joshua D. "The Economic Returns to Schooling in the West Bank and Gaza Strip." *American Economic Review,* December 1995, 85(5), pp.1065-87.

Antras, Pol; Garicano, Luis; Rossi-Hansberg, Esteban. "Offshoring in a Knowledge Economy." National Bureau of Economic Research (Cambridge, MA) Working Paper No. 11094, January 2005.

Attanasio, Orazio; Goldberg, Pinelopi and Pavcnik, Nina, "Trade Reforms and Wage Inequality in Colombia." *Journal of Development Economics,* August 2004, 74(2), pp. 331-66.

Baum-Snow, Nathaniel. "The Effects of Changes in the Transportation Infrastructure on Suburbanization: Evidence from the Construction of the Interstate Highway System." Unpublished, 2004.

Berman, Eli; Bound, John and Griliches, Zvi. "Changes in the Demand for Skilled Labor within U.S. Manufacturing: Evidence from the Annual Survey of Manufactures." *Quarterly Journal of Economics,* May 1994, 109(2), pp. 367-97.

Berman, Eli; Bound, John and Machin, Stephen. "Implications of Skill-Biased Technological Change: International Evidence." *Quarterly Journal of Economics*, November 1998, 113(4), pp. 1245-79.

Bernard, Andrew; Redding, Stephen; Schott, Peter K. "Factor Price Equality and the Economies of the United States." Centre for Economic Performance, London School of Economics, CEP Discussion Papers, July 2005.

Borjas, George J.; Freeman, Richard B. and Katz, Lawrence F. "How Much Do Immigration and Trade Affect Labor Market Outcomes?" *Brookings Papers on Economic Activity,* 1997, 0(1), pp. 1-67.

Card, David and DiNardo, John. "Do Immigrant Inflows Lead to Native Outflows?" *American Economic Review,* May 2000, 90(2), pp. 360-67.

Card, David and Lewis, Ethan G. "The Diffusion of Mexican Immigrants During the 1990s: Explanations and Impacts." Unpublished, January 2005.

Chandra, Amitabh and Thompson, Eric. "Does Public Infrastructure Affect Economic Activity? Evidence from the Rural Interstate Highway System." *Regional Science and Urban Economics,* July 2000, 30(4), pp. 457-90.

Davis, Donald R.; Weinstein, David E.; Bradford, Scott C. and Shimpo, Kazushige. "Using International and Japanese Regional Data to Determine When the Factor Abundance Theory of Trade Works." *American Economic Review,* June 1997, 87(3), pp. 421-46.

Dornbusch, Rudiger; Fischer, Stanley and Samuelson, Paul A. "Heckscher-Ohlin Trade Theory with a Continuum of Goods." *Quarterly Journal of Economics,* September 1980, 95(2), pp. 203-24.

Fogel, Robert W. *Railroads and American Economic Growth: Essays in Econometric History.* Baltimore: Johns Hopkins Press, 1964.

Freeman, Richard. "Demand For Education," in Orley Ashenfelter and Richard Layard, eds., Handbook of Labor Economics, Vol 1, Amsterdam: Elsevier North Holland, 1986, pp. 357-386.

Freeman, Richard B. "Trade Wars: The Exaggerated Impact of Trade in Economic Debate." *World Economy,* January 2004, 27(1), pp. 1-23.

Gey, Fredric C. *County Business Patterns Data (1967 and 1972).* Berkeley: UC DATA Survey Research Center, 2004.

Horiba, Yutaka and Kirkpatrick, Rickey C. "Factor Endowments, Factor Proportions, and the Allocative Efficiency of U.S. Interregional Trade." *Review of Economics and Statistics,* May 1981, 63(2), pp. 178-87.

Katz, Lawrence F. and Murphy, Kevin M. "Changes in Relative Wages, 1963-1987: Supply and Demand Factors." *Quarterly Journal of Economics*, February 1992, 107(1), pp. 35-78.

Kremer, Michael and Maskin, Eric. "Globalization and Inequality." Unpublished, 2003.

Krugman, Paul R. "Growing World Trade: Causes and Consequences." *Brookings Papers on Economic Activity*, 1995, 0(1), pp. 327-62.

Lahr, Michael L.; Duran, Rodrigo; and Varughese, Anupa. "Estimating the Impact of Highways on Average Travel Velocities and Market Size." Unpublished, 2005

Leonardi, Marco. "Product Demand Shifts and Wage Inequality." IZA (Bonn) Discussion Paper No. 908, October 2003.

Lewis, Ethan. "Immigration, Skill Mix, and the Choice of Technique." Federal Reserve Bank of Philadelphia, Working Paper No. 5(8), 2005.

Lewis, Tom. *Divided Highways: Building the Interstate Highways, Transforming American Life.* New York, NY: Viking Penguin, 1997.

Matsuyama, Kiminori. "Beyond Icebergs: Modeling Globalization as Biased Technical Change." Unpublished, 2004.

Parsley, David C. and Wei, Shang-Jin. "Convergence to the Law of One Price without Trade Barriers or Currency Fluctuations." *Quarterly Journal of Economics,* November 1996, 111(4), pp. 1211-36.

Revenga, Ana L. "Exporting Jobs? The Impact of Import Competition on Employment and Wages in U.S. Manufacturing." *Quarterly Journal of Economics,* February 1992, 107(1), pp. 255-84.

Ruggles, Steven, et al. *Integrated Public Use Microdata Series: Version 3.0.* Minneapolis, MN: Minnesota Population Center, 2004.

Schott, Peter K. "Across-Product versus Within-Product Specialization in International Trade." *Quarterly Journal of Economics*, May 2004, 119(2), pp. 647-78.

U.S. Department of Commerce, Bureau of Economic Analysis. *Regional Economic Ac-*

counts. 2004, http://www.bea.gov/bea/regional/reis.

U.S. Department of Commerce, Bureau of the Census. *County and City data books,* 1947-2000 [Computer files].

U.S. Department of Transportation, Bureau of Transportation Statistics. *Commodity Flow Survey.* 1993 and 1997.

U.S. Department of Transportation, Bureau of Transportation Statistics. *National Transportation Atlas Databases,* Washington, D.C.: The Bureau, 2002.

U.S. Department of Transportation, Bureau of Transportation Statistics. *National Transportation Statistics.* Annual Reports, 1993.

U.S. Department of Transportation, Federal Highway Administration. *Interstate Gap Study,* 1976.

U.S. Department of Transportation, Federal Highway Administration, Office of Highway Policy Information. *Highway Statistics.* 1995, 1996, 1998, 2000 http://www.fhwa.dot.gov/policy/ohpi/hss/hsspubs.htm.

U.S. House of Representatives. *A Report of the National Interregional Highway Committee, Outlining and Recommending a National System of Interregional Highways.* U.S. Government Printing Office, Washington: 1944.

U.S. House of Representatives, The Subcommittee on Highways and Transit. *Hearing on Trucking Safety,* 2002 http://www.house.gov/transportation/highway/07-09-02/07-09-02memo.html.

U.S. Federal Works Agency, Public Roads Administration. *Press release,* 1947.

University of Virginia library, Geospatial and Statistical Data Center. *County Business Patterns* (1977-1997), 2005 http://fisher.lib.virginia.edu/collections/stats/cbp/county.html.

Verhoogen, Eric. "Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector: Theory and Evidence from an Exchange-Rate Shock." Unpublished, 2004.

Table 1. County descriptive statistics

Mean	Earliest data (post-WWII)	All counties	Sample counties		
			Full Sample	With highway	Without highway
Land Area	1950	959	988	1,238	906
Population	1950	48,699	19,378	24,858	17,590
Population density	1950	213	32	38	30
Population growth	1930-1950	0.004	0.001	0.004	-0.001
Per capita income	1959	1,352	1,237	1,319	1,210
Earnings in trucking and warehousing per capita	1969	44	42	41	42
Retail sales per capita	1948	693	630	678	614
High-school graduates (of 25+ years old)	1950	0.27	0.25	0.27	0.25
Fraction commuting to work outside county	1970	0.20	0.20	0.22	0.19
Distance to nearest large city (miles)		84	92	78	96
Northeast		0.07	0.04	0.04	0.04
Midwest		0.34	0.36	0.30	0.38
South		0.46	0.46	0.44	0.47
West		0.13	0.14	0.22	0.11
Observations		3,101	2,000	492	1,508

Notes: The summary statistics are from the County and City Data books and from authors calculations using the National Transportation Atlas Databases, and they are calculated for all counties for which land area in known for 1950. Income, earnings, and sales data are in nominal US dollars.

Table 2. Determinants of highway assignment to counties

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Highway planned in 1944 legislation	0.795	0.798	0.842	0.789	0.785						0.780	
	(0.024)	(0.023)	(0.016)	(0.024)	(0.025)						(0.025)	
Direction to nearest city instrument						0.218	0.221	0.186	0.231	0.232	0.055	0.270
						(0.051)	(0.052)	(0.034)	(0.051)	(0.048)	(0.040)	(0.048)
1950 population weights	Yes	Yes	No	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes
Distance to nearest city	No	No	No	Yes	Yes	No	No	No	Yes	Yes	Yes	Yes
Geographic indicators	None	Region	Region	Region	State	None	Region	Region	Region	State	Region	Region

Notes: Cross-section regressions for sample counties. Columns 1-11 use full sample of counties (2000 observations) and column 12 uses only counties in the Midwest and the South (1647 observations). Robust standard errors are in parenthesis.

Table 3. Vehicle miles traveled on rural interstate highway, by vehicle type

	1970			1980			1990		
Vehicle miles traveled (billions)	Rural Interstate Highways	Other highways	Fraction traveled on rural highways	Rural Interstate Highways	Other highways	Fraction traveled on rural highways	Rural Interstate Highways	Other highways	Fraction traveled on rural highways
Combination trucks	8.1	27.0	0.23	21.1	47.6	0.31	30.1	64.2	0.32
Single-unit trucks	2.0	25.1	0.07	4.0	35.8	0.10	5.7	46.2	0.11
Passenger cars and motorcycles	62.3	857.3	0.07	89.5	1,032.3	0.08	117.5	1,300.3	0.08
Other vehicles	7.1	120.7	0.06	20.5	276.5	0.07	46.9	533.4	0.08
Total	79.5	1,030.2	0.07	135.1	1,392.2	0.09	200.2	1,944.2	0.09

Notes: The data are from Highway Statistics, 1995-2000. The figures include vehicle miles traveled on toll highways incorporated into the Interstate Highway System, since data for those segments are not reported separately.

Table 4. Effect of highways on ln(earnings in trucking and warehousing per capita)

	(1)	(2)	(3)	(4)	(5)
1972	0.022	0.026	0.009	0.028	0.027
	(0.021)	(0.021)	(0.023)	(0.021)	(0.021)
1977	0.074	0.084	0.069	0.083	0.077
	(0.031)	(0.032)	(0.031)	(0.032)	(0.032)
1982	0.061	0.079	0.083	0.093	0.082
	(0.037)	(0.038)	(0.040)	(0.039)	(0.037)
1987	0.078	0.090	0.090	0.100	0.088
	(0.043)	(0.043)	(0.045)	(0.044)	(0.043)
1992	0.115	0.149	0.158	0.154	0.144
	(0.049)	(0.048)	(0.049)	(0.049)	(0.048)
1997	0.041	0.086	0.117	0.093	0.089
	(0.052)	(0.050)	(0.051)	(0.051)	(0.051)
Observations	12,220	12,220	12,220	12,220	12,220
Weights	Yes	Yes	No	Yes	Yes
Region*year	No	Yes	Yes	Yes	Yes
Distance*year	No	No	No	Yes	Yes
1950 Pop. density*year	No	No	No	No	Yes

Notes: All estimates are from a panel of the sample counties that includes county and year dummies. Each column reports highway*year interactions from a seperate regression, and the omitted interaction is highway*1969. The weights are 1950 population. Robust standards errors in parenthesis are clustered at the county level. Distance denotes the distance in miles from the county centroid to the nearest large city.

Table 5. Effect of highways on ln(retail sales per capita)

	(1)	(2)	(3)	(4)	(5)
1954	-0.013	-0.002	-0.010	-0.003	0.000
	(0.010)	(0.009)	(0.009)	(0.009)	(0.009)
1958	-0.031	-0.018	-0.032	-0.016	-0.011
	(0.012)	(0.010)	(0.010)	(0.010)	(0.010)
1963	-0.029	-0.011	-0.033	-0.007	-0.003
	(0.014)	(0.011)	(0.011)	(0.011)	(0.011)
1967	-0.027	-0.001	-0.019	-0.005	-0.002
	(0.017)	(0.014)	(0.014)	(0.014)	(0.014)
1972	0.023	0.042	0.032	0.033	0.032
	(0.020)	(0.015)	(0.016)	(0.015)	(0.016)
1977	0.034	0.053	0.051	0.045	0.043
	(0.020)	(0.016)	(0.017)	(0.016)	(0.017)
1982	0.057	0.078	0.078	0.072	0.070
	(0.023)	(0.018)	(0.020)	(0.019)	(0.019)
1987	0.076	0.102	0.107	0.086	0.080
	(0.025)	(0.021)	(0.021)	(0.020)	(0.020)
1992	0.087	0.110	0.118	0.095	0.089
	(0.025)	(0.021)	(0.022)	(0.020)	(0.021)
1997	0.101	0.135	0.149	0.123	0.117
	(0.027)	(0.022)	(0.024)	(0.022)	(0.022)
Observations	21,839	21,839	21,839	21,839	21,839
Weights	Yes	Yes	No	Yes	Yes
Region*year	No	Yes	Yes	Yes	Yes
Distance*year	No	No	No	Yes	Yes
1950 Pop. density*year	No	No	No	No	Yes

Notes: All estimates are from a panel of the sample counties that includes county and year dummies. Each column reports highway*year interactions from a seperate regression, and the omitted interaction is highway*1948. The weights are 1950 population. Robust standards errors in parenthesis are clustered at the county level. Distance denotes the distance in miles from the county centroid to the nearest large city.

Table 6. Effect of highways on trade and commuting

	OLS						IV		
							Instrument		
							1944 plan	Direction to city	
	Full Sample					Midwest and South	Full Sample	Full Sample	Midwest and South
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	A. Dependent variable: ln(retail sales per capita): 1963-1997								
Highway*(post-1975)	0.082	0.086	0.107	0.081	0.071	0.063	0.082	0.121	0.144
	(0.013)	(0.012)	(0.014)	(0.012)	(0.012)	(0.013)	(0.015)	(0.086)	(0.082)
Observations	15,854	15,854	15,854	15,854	15,854	13,053	15,854	15,854	13,053
	B. Dependent variable: ln(earnings in trucking and warehousing per capita): 1969-1997								
Highway*(post-1975)	0.063	0.084	0.098	0.074	0.082	0.088	0.054	0.070	0.180
	(0.033)	(0.033)	(0.033)	(0.032)	(0.032)	(0.039)	(0.043)	(0.228)	(0.217)
Observations	12,220	12,220	12,220	12,220	12,220	10,099	12,220	12,220	10,099
	C. Dependent variable: fraction commuting to work outside their county of residence: 1970-1990								
Highway*(post-1975)	0.005	0.008	0.007	0.011	0.008	0.008	0.007	0.031	0.024
	(0.005)	(0.005)	(0.004)	(0.006)	(0.005)	(0.005)	(0.006)	(0.030)	(0.028)
Observations	5,874	5,874	5,874	5,874	5,874	4,815	5,874	5,874	4,815

Notes: All estimates are from a panel of the sample counties that includes county and year dummies. Robust standards errors in parenthesis are clustered by county. Each cell reports the coefficient on highway*(post-1975) interaction from a separate regression. Columns 2-9 control for region-specific year effects. All regressions are weighted using the 1950 population, except column 3 which is unweighted. Column 4 uses an index of highway completion, by state, instead of a post-1975 indicator. Columns 5-9 control for (distance to nearest city)*year and (1950 population density)*year interactions. Columns 6 and 9 limit the sample to the regions where the first stage is significant for the direction instrument (the Midwest and the South).

Table 7. Effect of highways on the demand for skill in manufacturing

	OLS					IV	
						Instrument	
						1944 plan	Direction to city
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
A. Dependent variable: ln(relative wage-bill of non-production workers)							
(Post-1975)*highway	0.006	-0.149	-0.168	-0.151	-0.101	-0.223	-1.370
	(0.024)	(0.069)	(0.068)	(0.077)	(0.069)	(0.094)	(2.049)
(Post-1975)*highway*(1950 hs)		0.623	0.609	0.564	0.456	0.802	4.177
		(0.249)	(0.241)	(0.262)	(0.247)	(0.312)	(5.960)
(Post-1975)*(1950 high school)		-0.443	-0.323	-0.208	-0.077	-0.278	-1.200
		(0.148)	(0.213)	(0.255)	(0.093)	(0.271)	(1.683)
Observations	5,795	5,795	5,793	4,455	4,455	4,455	4,455
B. Dependent variable: ln(relative wage of non-production workers)							
(Post-1975)*highway	-0.051	-0.129	-0.113	-0.113	-0.069	-0.136	-1.259
	(0.020)	(0.059)	(0.059)	(0.059)	(0.058)	(0.077)	(1.573)
(Post-1975)*highway*(1950 hs)		0.313	0.274	0.274	0.130	0.312	3.098
		(0.202)	(0.206)	(0.206)	(0.207)	(0.262)	(4.563)
(Post-1975)*(1950 high school)		-0.251	-0.479	-0.479	0.022	-0.484	-1.145
		(0.134)	(0.220)	(0.220)	(0.085)	(0.230)	(1.285)
Observations	4,456	4,456	4,455	4,455	4,455	4,455	4,455
C. Dependent variable: ln(relative employment of non-production workers)							
(Post-1975)*highway	0.063	-0.004	-0.037	-0.038	-0.032	-0.087	-0.111
	(0.027)	(0.081)	(0.079)	(0.079)	(0.075)	(0.096)	(1.646)
(Post-1975)*highway*(1950 hs)		0.264	0.289	0.290	0.326	0.490	1.079
		(0.278)	(0.267)	(0.267)	(0.268)	(0.320)	(4.808)
(Post-1975)*(1950 high school)		-0.134	0.276	0.272	-0.099	0.206	-0.055
		(0.179)	(0.257)	(0.258)	(0.103)	(0.276)	(1.373)
Observations	4,461	4,461	4,460	4,455	4,455	4,455	4,455

Notes: All estimates are from a panel of the sample counties that includes county and year dummies. All estimates use data for 1967-1982, and include 1950 population weights. Robust standards errors in parenthesis are clustered by county. Columns 1-3 use the full sample, and columns 4-7 use a fixed sample size across panels. Columns 3-7 control for region*year, (distance to nearest city)*year, and (1950 population density)* year interactions, and the fraction of high-school graduates among 25+ year-olds. Column 5 uses a state-level index of the fraction of highways completed.

Table 8. Effect of highways on the demand for skill in manufacturing

	OLS		IV	
			Instrument	
Dependent variable: change in ln(relative wage-bill of non-production workers)			1944 plan	direction to city
	(1)	(2)	(3)	(4)
A. Before highway construction was complete: 1947-1967				
Highway	-0.069	-0.052	0.049	0.209
	(0.120)	(0.119)	(0.159)	(1.385)
Highway*(1950 high school)	0.476	0.322	-0.062	-0.814
	(0.415)	(0.414)	(0.524)	(4.200)
(1950 high school)	-0.547	-0.754	-0.639	-0.391
	(0.248)	(0.381)	(0.394)	(1.178)
B. When highway construction was being completed: 1967-1982				
Highway	-0.140	-0.134	-0.228	0.719
	(0.089)	(0.089)	(0.106)	(1.307)
Highway*(1950 high school)	0.736	0.655	0.929	-2.202
	(0.309)	(0.309)	(0.360)	(3.920)
(1950 high school)	-0.335	-0.098	-0.167	0.702
	(0.199)	(0.280)	(0.298)	(1.091)
C. After the construction of highways was complete: 1982-1992				
Highway	0.047	0.082	0.077	-0.114
	(0.076)	(0.075)	(0.102)	(1.156)
Highway*(1950 high school)	-0.076	-0.233	-0.125	1.171
	(0.275)	(0.261)	(0.344)	(3.464)
(1950 high school)	0.243	-0.070	-0.116	-0.592
	(0.169)	(0.241)	(0.245)	(0.932)

Notes: Cross section regression using a fixed subsample of 1,072 counties for which data exists in 1947,1967,1982 and 1992. Columns 2-4 control for region dummies, distance to nearest city, and 1950 population density. Robust standards errors are in parenthesis.

Table 9. Effect of highways on industrial composition of manufacturing employment

	OLS				IV	
					Instrument	
					1944 plan	Direction to city
Dependent variable: index of non-production worker intensity	(1)	(2)	(3)	(4)	(5)	(6)
(Post-1975)*highway	-0.002	-0.007	-0.011	-0.012	-0.014	0.032
	(0.003)	(0.009)	(0.009)	(0.008)	(0.012)	(0.126)
(Post-1975)*highway*(1950 hs)		0.020	0.030	0.032	0.035	-0.075
		(0.029)	(0.031)	(0.029)	(0.040)	(0.386)
(Post-1975)*(1950 high school)		0.001	0.031	0.001	0.031	0.052
		(0.017)	(0.026)	(0.009)	(0.027)	(0.096)
Observations	5,818	5,818	5,813	5,813	5,813	5,813

Notes: All estimates are from a panel of the sample counties that includes county and year dummies. All estimates use data for 1967-1982, and include 1950 population weights. Robust standards errors in parenthesis are clustered by county. Columns 3-6 control for region*year, (distance to nearest city)*year, and (1950 population density)*year interactions, and the fraction of high-school graduates among 25+ year-olds. Column 4 uses a state-level index of the fraction of highways completed.

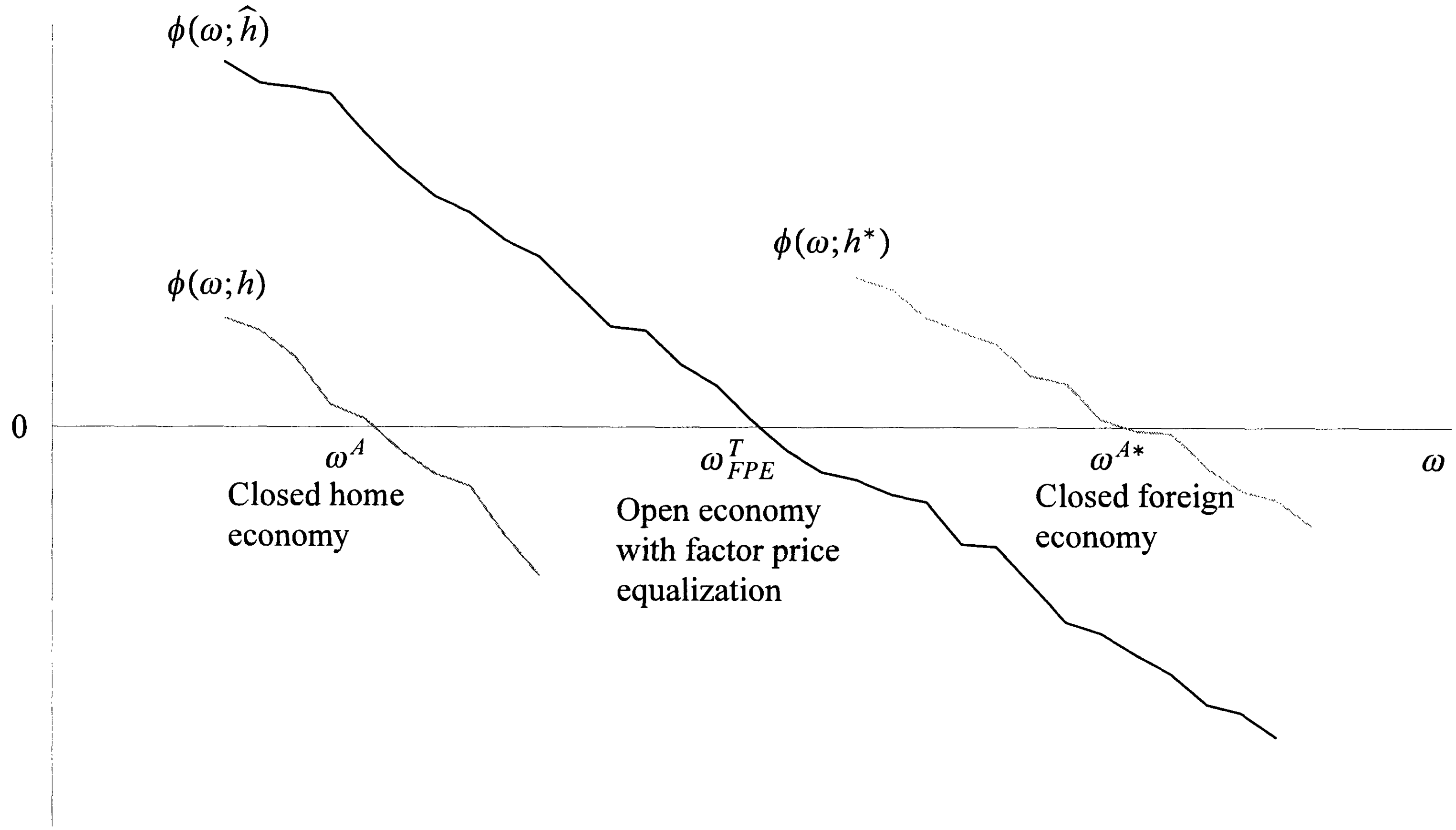

Figure 1. Existence and uniqueness of the equilibrium skill premium without migration. Holding the endowment fixed (along each of the curves), ϕ is positive for low skill premia and negative for high skill premia. Since it is continuous and decreasing in the skill premium there is a unique equilibrium skill premium such that $\phi = 0$.

Figure 2. Routes of the recommended interregional highway system, 1944 plan. Source: U.S. House of Representatives, *Interregional Highways*, Washington D.C.: Government Printing Office, House document no. 379, 78th congress, 2nd session, January 1944

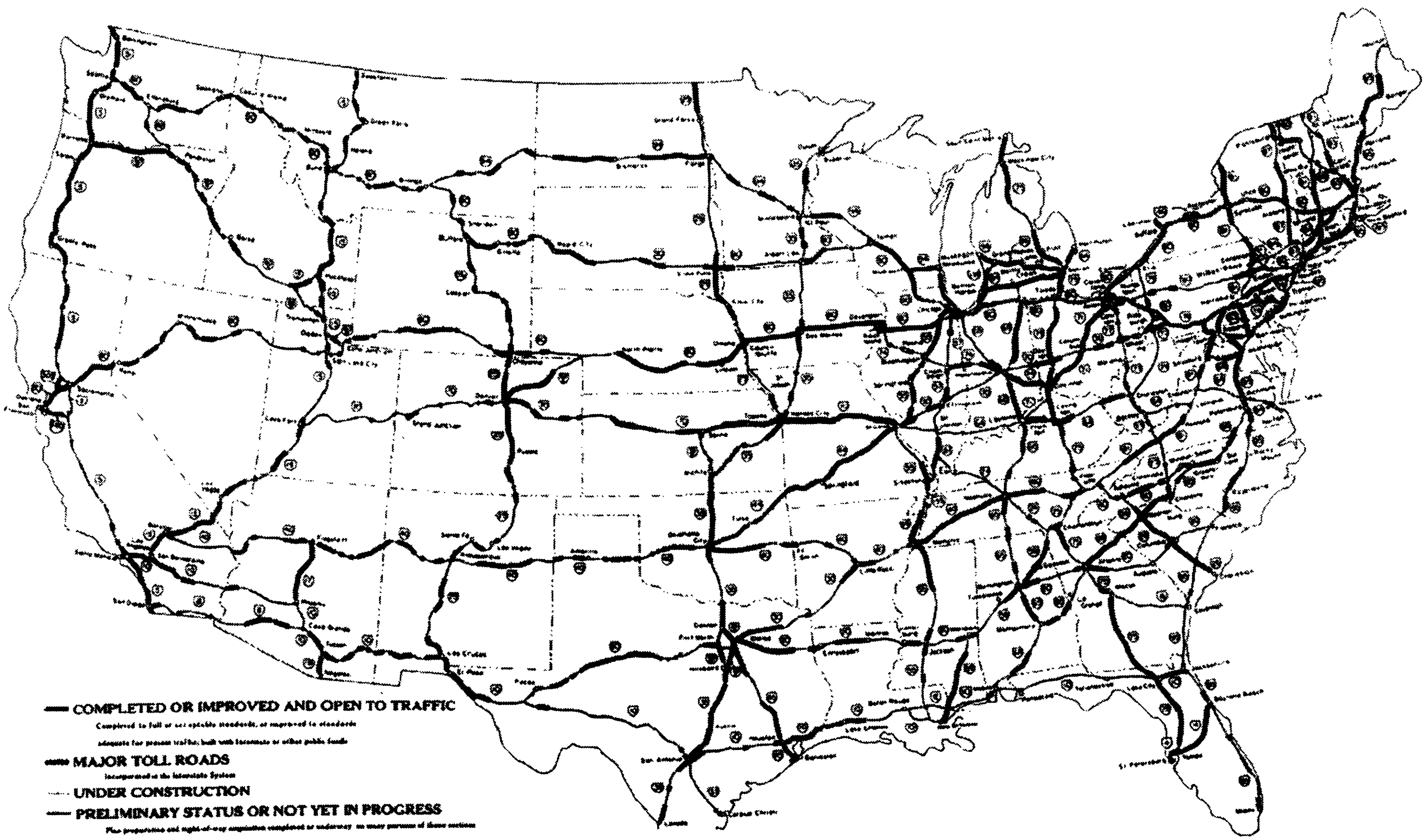

Figure 3. The Interstate Highway System in September 1966. Source: Bureau of Public Roads.

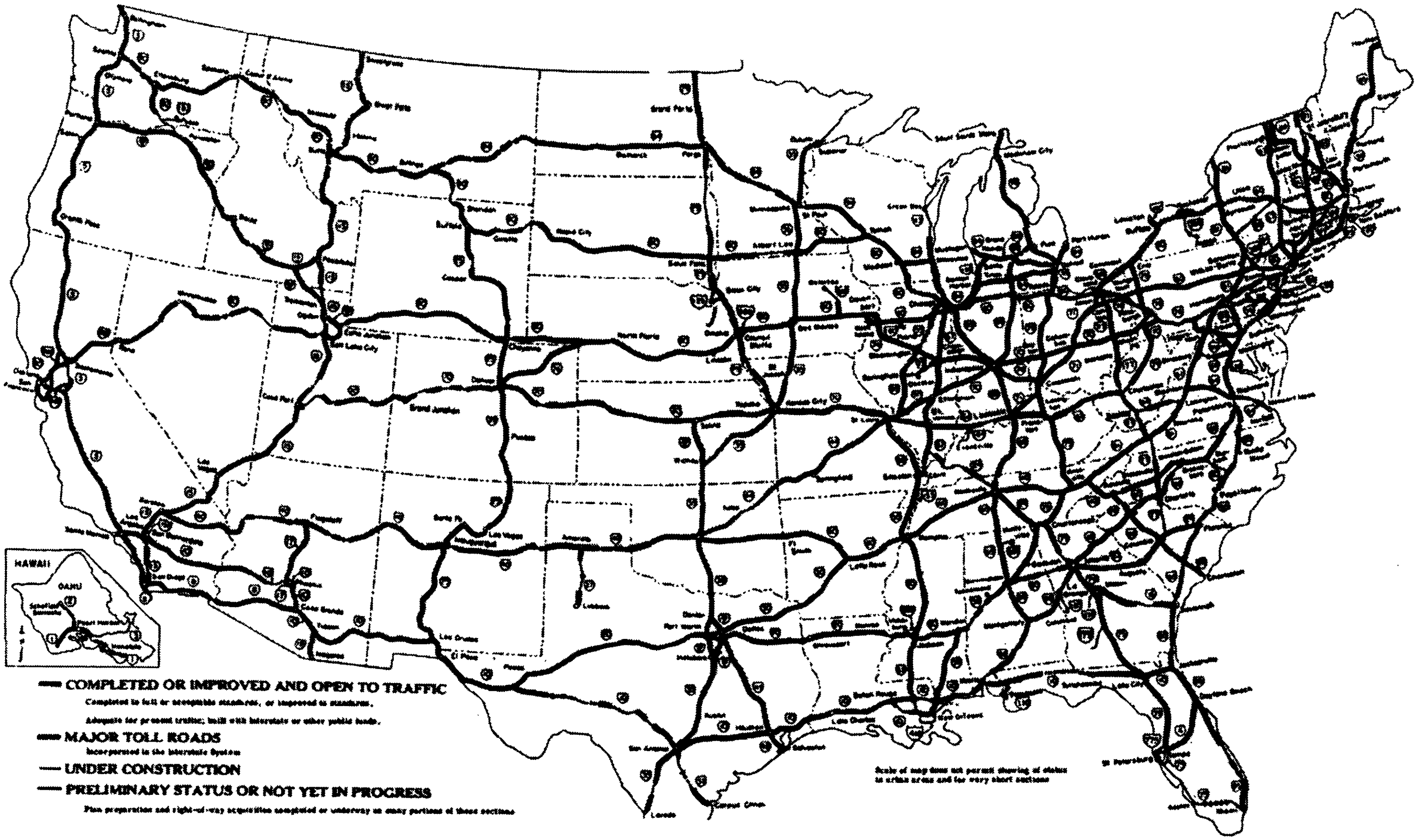

Figure 4. The Interstate Highway System in December 1975. Source: Federal Highway Administration.

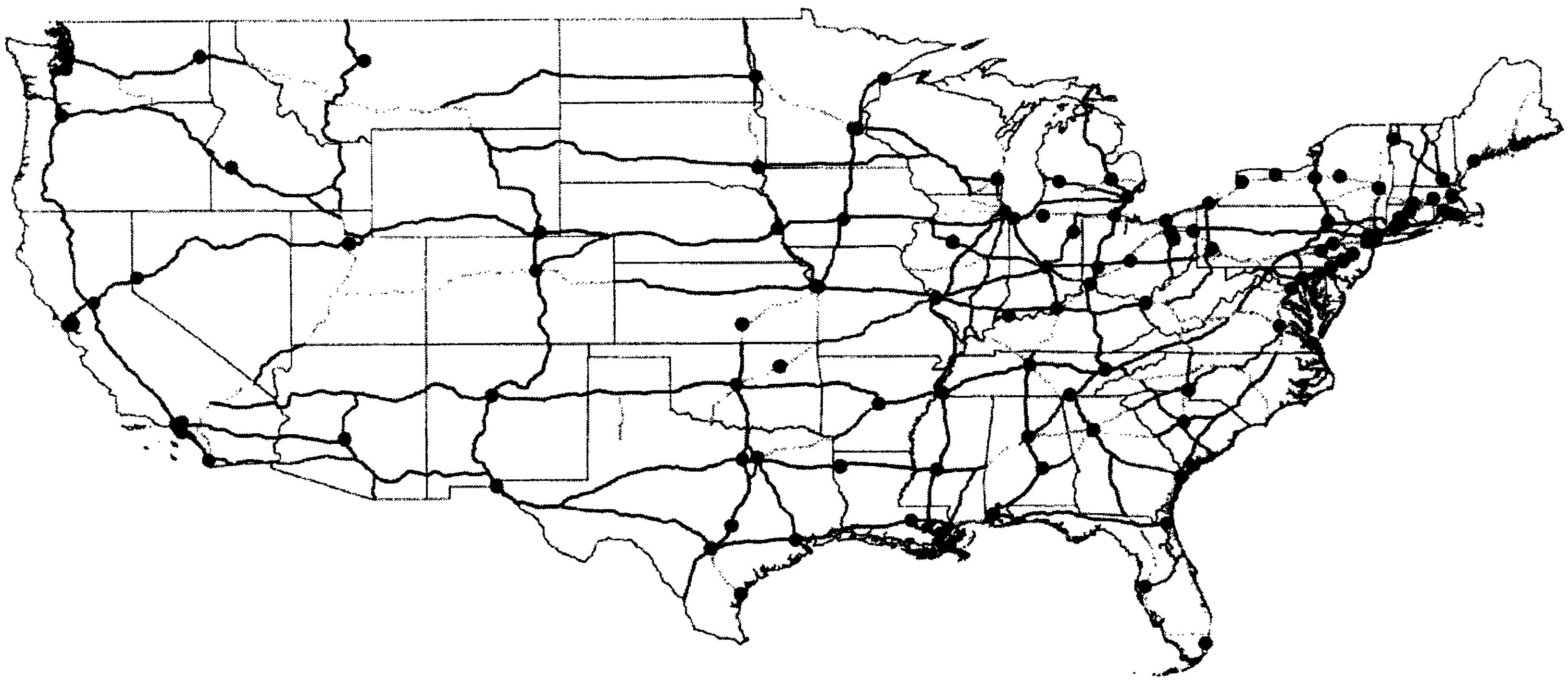

Figure 5. The Interstate Highway System in 2002. Black lines denote highways segments included in the sample and grey lines denote highway segments excluded from the sample. Black dots denote cities that had a population of 100,000 or more or were the largest in their state in 1950.

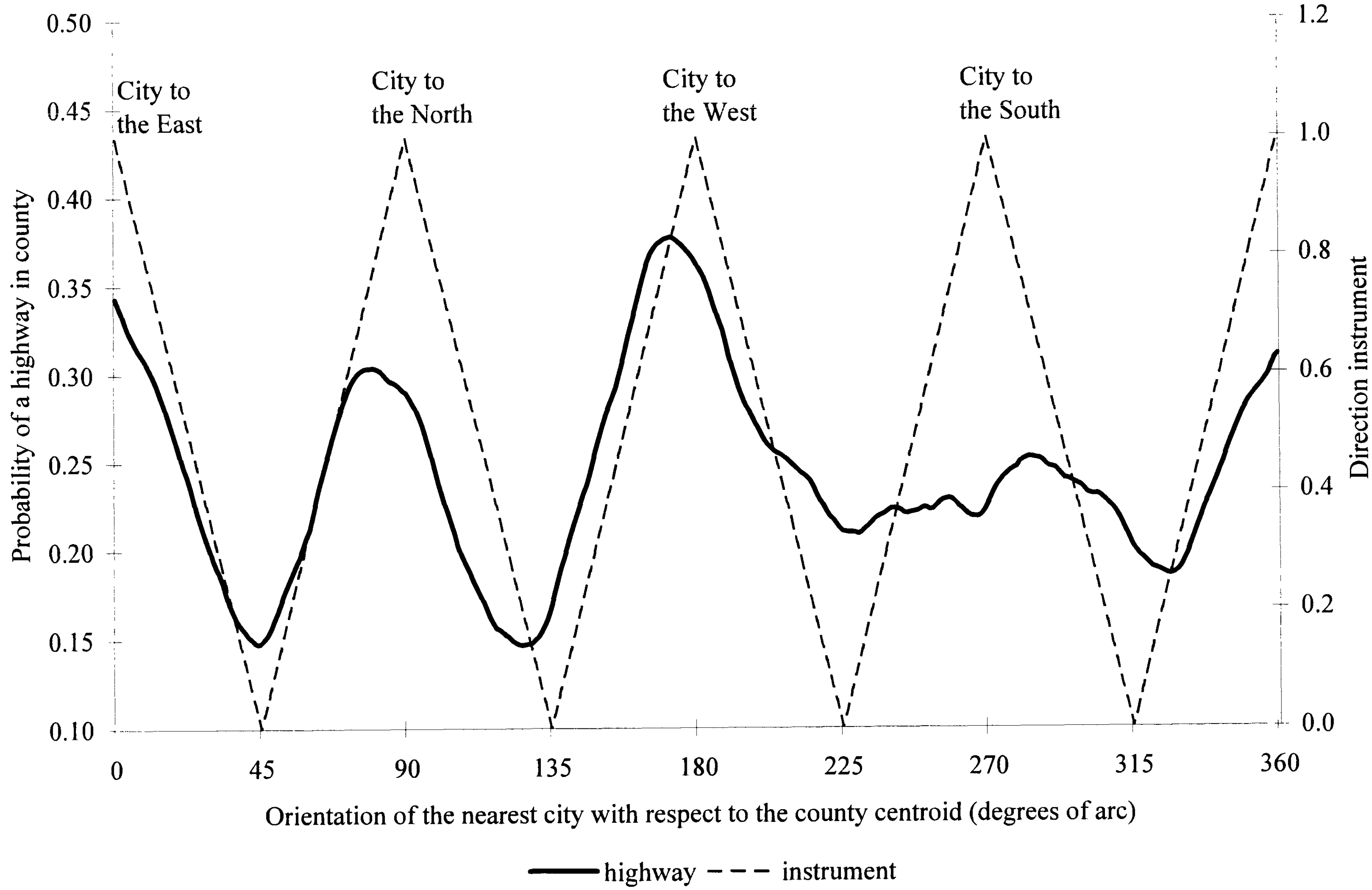

Figure 6. The direction to the nearest city and the probability an interstate highway crosses a rural county. The probability is estimated using a kernel regression with an Epanechnikov kernel and a bandwidth of 20.

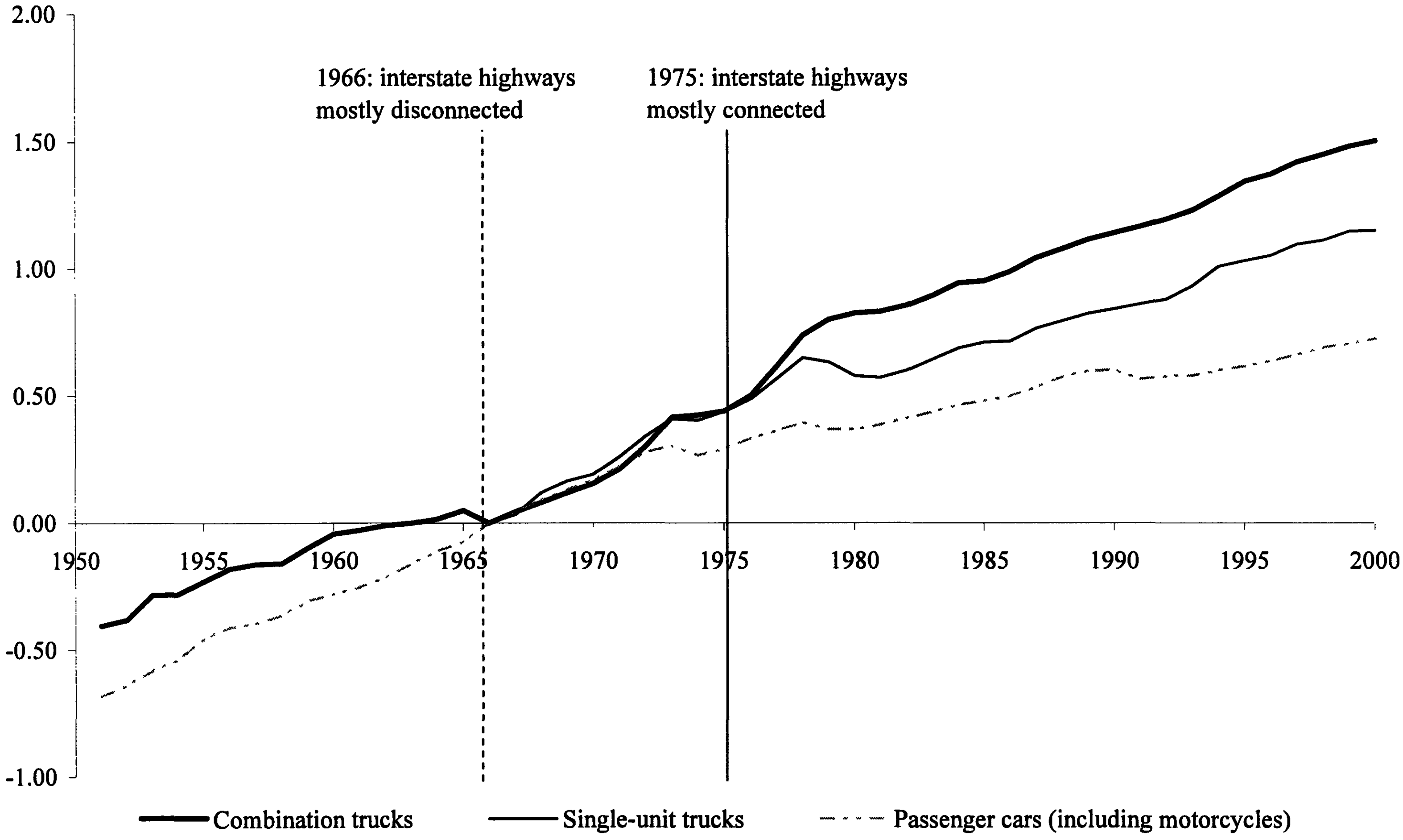

Figure 7. Ln(vehicle miles traveled), by vehicle type (base year is 1966). Source: U.S. Department of Transportation, Federal Highway Administration, Office of Highway Policy Information, Highway Statistics

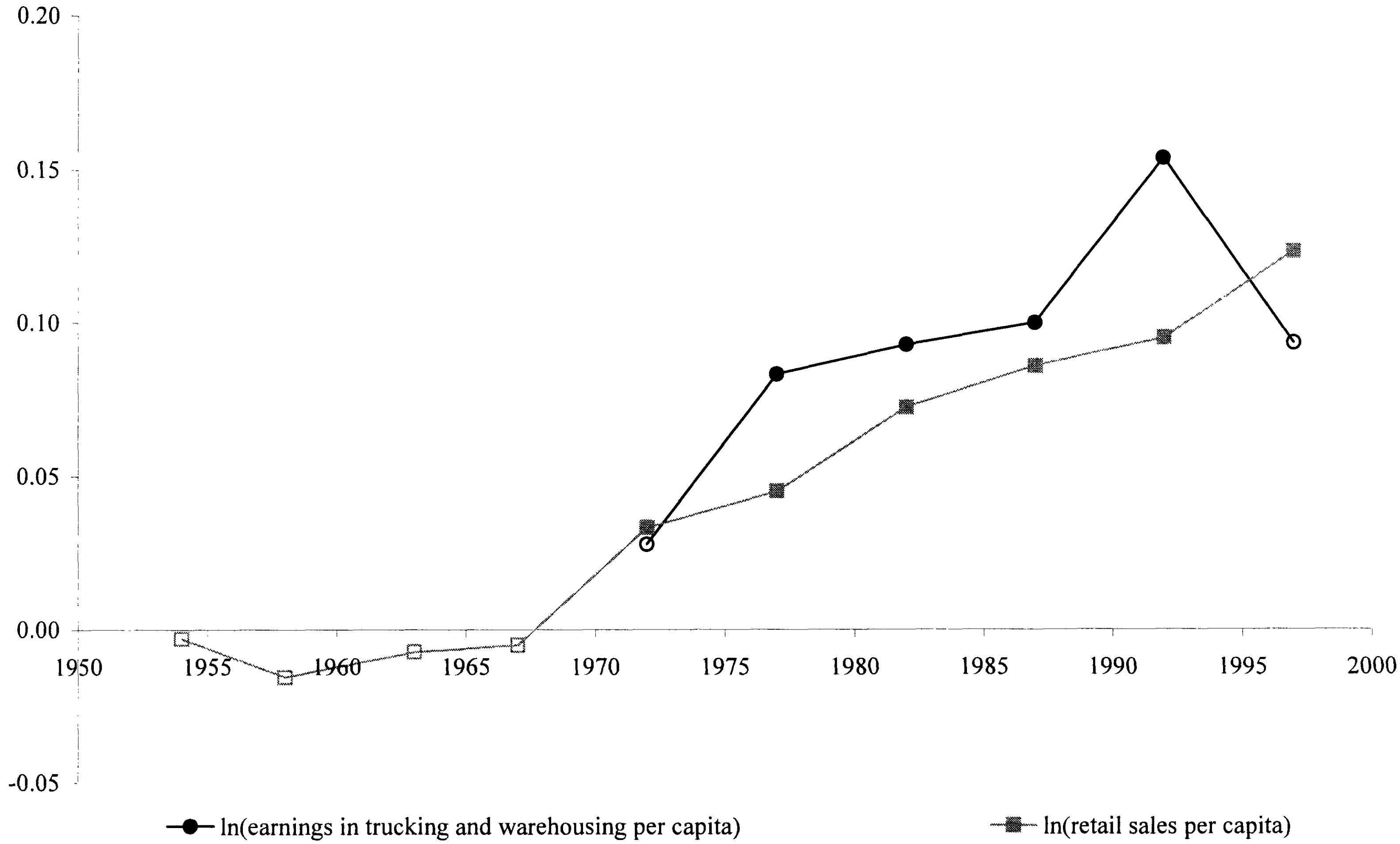

Figure 8. The effect of highways on trade in rural counties. The figure reports the coefficients on highway*year interactions from Column 4 in Tables 4 and 5. Open points represent coefficients not significant at the 5 percent level.

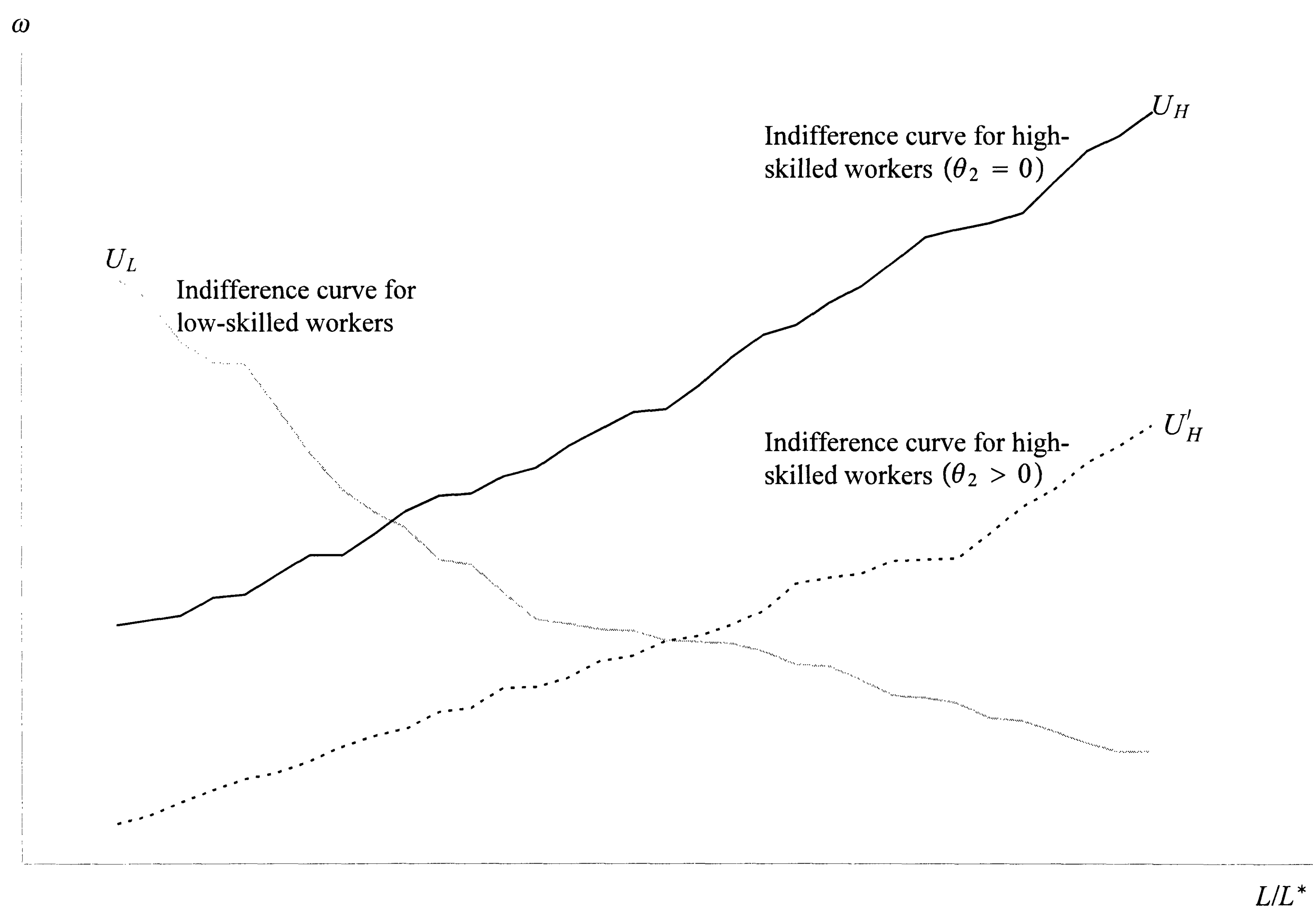

Figure A1. Existence and uniqueness of the equilibrium with migration.

Chapter 2

Technology, Complexity and Information: The Evolution of Demand for Office Workers

2.1 Introduction

Information processing is an important part of work in all firms and organizations, yet we know relatively little about its determinants. In this paper, I analyze manufacturing firms' demand for workers who process information - clerical office workers (clerks). Clerks are skilled workers, who accounted for over 11 percent of the US manufacturing workforce in 2000. Clerks use information technology equipment to generate, store, and communicate information. The information that clerks process is an intermediate input that facilitates the coordination of production.

Production technologies vary in the extent of the division of labor they require. Studies of the division of labor date back to Adam Smith (1776), and recent theoretical work argues that the coordination of specialized workers who perform different tasks is costly (Becker and Murphy 1992; Garicano 2000). This paper examines the hypothesis that firms' costs of coordination, proxied by the fraction of clerks in their workforce, increase in the extent of their division of labor.

In order to test this hypothesis empirically, I characterize a production process as complex when it requires a diverse set of occupations. Specifically, I define the complexity of a manufacturing industry as one minus the Herfindahl index of the occupations of its employees, excluding managers, clerks, accountants, and auditors. The ranking of industries by complexity is stable over time, likely reflecting differences in their technology of production. For example, textile industries employ an occupationally homogenous workforce and therefore have low complexity. By contrast, manufacturers of transportation equipment employ an occupationally heterogeneous workforce, so their complexity is high. Figure 1 shows that manufacturing industries' complexity has risen over the 20th century.

Figure 1 also shows that more complex industries have consistently employed relatively more clerks. In 2000, an increase of 1 standard deviation in an industry's measure of complexity was associated with an increase of about 1 percentage point (or about 8 percent) in the fraction of clerks in its workforce. In earlier decades this relationship was even stronger - from 1910-1980 an increase of 1 s.d. in complexity was associated with an increase of 1.3-2.9 percentage points (16-25 percent) in clerks' fraction of employment.

In order to further test the hypothesis that complexity increases firms' demand for information, I study an early information technology (IT) revolution, which took place around the turn of the 20th century. Before 1880, information technology was very limited, and so was the use of clerks. However, from 1880-1910 telephones, typewriters, and vertical filing techniques came into widespread use (Yates 1989). Figure 2 shows that the diffusion of telephones and typewriters during those years was rapid. These innovations dramatically increased the productivity of clerks and made it cost-effective for firms to use information more intensively. Interestingly, Figure 3 shows that the fraction of clerks in the manufacturing workforce increased from little more than 1 percentage point in 1880 to over 6 percentage points in 1910.[1] At the same time, clerks earned about twice as much as production workers, and the clerical earnings premium appears to have been stable from 1890-1910 (Figure 4).[2] This evidence suggests that the early information technology (IT) revolution increased the overall demand for clerks relative to other manufacturing workers. Importantly, I find that industries that were more complex in 1880 increased their demand for clerks significantly

[1] Figure 1 also shows that the employment share of managers and other white-collar employees increased much more slowly over the same period.

[2] Wage data before 1939 are not nationally representative and should therefore be treated with some caution.

more than other industries over the next 30 years.

The increased demand for clerks during the early information technology revolution had two important implications. First, the changes in office technology increased the demand for clerks, who were highly skilled workers, at a time when high-school graduation rates were low (Figure 5). The rise of the High-School Movement after 1910 (Goldin 1998) may have been, at least in part, a response to this increased demand for skill. Second, technological change in the office can help explain the integration of women into the labor force during the 20th century. Although men outnumbered women among clerks employed in manufacturing in 1910, this relationship reversed over the subsequent decades. Thus, women's increased labor force participation during the 20th century was facilitated by the technological change that increased the demand for clerks.

In contrast to the early IT revolution, the computer revolution that took place in recent decades had very different implications for the employment of clerks. In fact, the availability of improved IT equipment from 1960-2000 caused firms to substitute computers for clerks. Figure 3 shows the decline in the share of clerks in total manufacturing employment, and Figure 4 shows a decline in their relative wages. Three mechanisms facilitated the substitution of computers for clerks: direct replacement of clerical jobs by machines, indirect replacement due to increased productivity in the performance of clerical tasks, and a re-organization of tasks that allowed all workers to process information. This finding provides a concrete case in which technology and a specific set of skills are substitutes rather than complements. Interestingly, I find that firms in more complex industries increased their demand for computer personnel and decreased their demand for clerks more rapidly than other firms.

The remainder of the paper is organized as follows. Section 2 presents a simple framework that illustrates the main arguments. Section 3 considers the relationship between firm organization and clerical employment. Section 4 analyzes the early IT revolution and the increased demand for clerical workers. Section 5 discusses the decline in clerical employment during the recent computer revolution, and section 6 concludes.

2.2 A Theory of Demand for Information and Clerks

To frame the key questions of this paper, I construct a framework for analyzing the effect of technology on the demand for information processing. I conceptualize the process of manufacturing as consisting of production and coordination. Production includes all tasks required to physically make the product. The workers who take part in production are operatives, craftsmen, laborers, and professional workers. Coordination involves generating, storing, and communicating information that is used to organize production. The information used to coordinate production is processed by clerks.[3]

All manufacturing firms need to perform the tasks required for production, but they can differ in their approach to coordination. One approach, which I call low information intensity, relies on informal information to coordinate production. For example, this may involve foremen supervising different sections of a plant. This approach typically involves costly misallocation of resources in production. The second approach relies on high information intensity in coordination. This requires the firm to employ clerks, who use IT equipment to provide information that ensures a precise allocation of inputs in production.

Assume firms use labor inputs to produce output using a constant elasticity of substitution (CES) production function:

$$Y_{i,j} = \left[\int_0^{t_{i,j}} L_{i,j}\left(x\right)^{(\sigma-1)/\sigma} dx \right]^{\sigma/(\sigma-1)}, \qquad (1)$$

where the information intensity i is either high (h) or low (l) and the industry is indexed by j. The complexity of the production function, $t_{i,j}$, is the measure of the set of different tasks employed in production. For simplicity, this model considers only labor inputs, rather than capital inputs, so complexity is closely related to the division of labor.

I assume that firms using high information intensity execute all tasks correctly, so $t_{h,j} = t_j$. Firms using low information intensity do not execute all tasks correctly. This may be due

[3]Managers' role can be seen as using the information processed by clerks to coordinate production. For simplicity, their tasks are not modeled in this framework.

to problems in selecting appropriate workers, assigning them adequate tasks, providing them adequate complementary inputs, monitoring them or giving them incentives. I assume that $t_{l,j} = t_j^{\gamma}$, where γ measures the efficiency of informal coordination ($\gamma \in [0,1)$). In industry j, a set of workers of measure $L_{i,j}(x)$ is assigned to task x. The elasticity of substitution between the different tasks is σ, and I assume that tasks are gross substitutes ($\sigma > 1$).

A firm using high information intensity needs to gather information I_j in proportion to the size of its workforce and the complexity of the technology it uses:

$$I_j = \left(\int_0^{t_j} L_{h,j}(x)\, dx \right) t_j \tag{2}$$

I assume that information is an intermediate input produced using CES technology:

$$I_j = \left[\alpha \left(L_{h,j}^c \right)^{\eta} + \beta \left(K_{h,j}^c \right)^{\eta} \right]^{1/\eta}, \tag{3}$$

where $L_{h,j}^c$ and $K_{h,j}^c$ are clerks and IT equipment. The productivity of clerks and IT equipment in producing information is captured by α and β, and the elasticity of substitution between clerks and information technology is $\sigma_I = 1/(1-\eta)$. Given a technology t_j, a firm producing with high information intensity solves the following optimization problem:

$$Min \left\{ wL_{h,j} + w_c L_{h,j}^c + pK_{h,j}^c \right\} \quad s.t.: (1) - (3), \tag{4}$$

where the wage of labor is w, the wage of clerks is w_c and the price of information technology equipment is p. Using the assumption that all tasks are equally important, $L_{h,j}(x) = L_{h,j}/t_j$, and total labor demand is:

$$L_{h,j} = t_j^{-1/(\sigma-1)} Y_j \tag{5}$$

The demand for clerks and for information technology equipment is:

$$L_{h,j}^c = (\alpha + \beta\phi)^{-1/\eta}\, t_j^{(\sigma-2)/(\sigma-1)} Y_j \tag{6}$$

$$K^c_{h,j} = (\beta + \alpha/\phi)^{-1/\eta} t_j^{(\sigma-2)/(\sigma-1)} Y_j, \tag{7}$$

where $\phi \equiv (\beta w_c/\alpha p)^{\eta/(1-\eta)}$. The cost of production with high information intensity is:

$$\begin{aligned} C_{h,j} = & t_j^{-1/(\sigma-1)} w Y_j + (\alpha + \beta\phi)^{-1/\eta} t_j^{(\sigma-2)/(\sigma-1)} w_c Y_j + \\ & + (\beta + \alpha/\phi)^{-1/\eta} t_j^{(\sigma-2)/(\sigma-1)} p Y_j. \end{aligned} \tag{8}$$

As long as the elasticity of substitution is sufficiently low ($\sigma \leq 2$), the cost function is decreasing in the level of complexity t_j, because the increased efficiency of production with a more complex technology outweighs the additional cost of information associated with it.[4]

A firm may choose to rely on low information intensity for supervising work. In this case it fails to execute some tasks correctly, but it does not need to employ clerks or IT equipment. The labor demand of a firm producing with low information intensity is:

$$L_{l,j} = Y t_j^{(\sigma-\gamma\sigma-1)/(\sigma-1)}, \tag{9}$$

and its cost function is:

$$C_{l,j} = t_j^{(\sigma-\gamma\sigma-1)/(\sigma-1)} w Y_j. \tag{10}$$

Cost decreases in t_j as long as the elasticity of substitution between tasks is sufficiently low ($\sigma < \frac{1}{1-\gamma}$). If $\sigma > \frac{1}{1-\gamma}$ then firms using low information intensity prefer the simplest technology available.

Using the constant returns to scale property of the production function we can define the net benefit of switching from low to high information intensity per unit

$$\begin{aligned} I_{\Delta,j} \equiv & MC_{l,j} - MC_{h,j} = (t_j^{(\sigma-\gamma\sigma-1)/(\sigma-1)} - t_j^{-1/(\sigma-1)}) w - \\ & - (\alpha + \beta\phi)^{-1/\eta} t_j^{(\sigma-2)/(\sigma-1)} w_c - (\beta + \alpha/\phi)^{-1/\eta} t_j^{(\sigma-2)/(\sigma-1)} p \end{aligned} \tag{11}$$

[4] I asume that $\sigma < 2$, so firms always prefer a higher level of technology.

Firms employ clerks if and only if $I_{\Delta,j} > 0$. Only the firms using high information intensity employ clerks and information technology equipment, so a decline in the price of these inputs makes high information intensity more advantageous:

$$\frac{\partial I_{\Delta,j}}{\partial w_c} < 0, \ \frac{\partial I_{\Delta,j}}{\partial p} < 0. \tag{12}$$

Hence, firms may switch to using high information intensity if w_c or p decline.

Now consider the case where all firms use high information intensity. We can compute the ratio of clerks to other workers:

$$c_j \equiv L_{h,j}^c / L_{h,j} = \left[\alpha + \beta^{1/(1-\eta)} \left(w_c/\alpha p\right)^{\eta/(1-\eta)}\right]^{-1/\eta} t_j. \tag{13}$$

Therefore, when all firms use high information intensity:

$$\frac{\partial c_j}{\partial t_j} > 0, \frac{\partial c_j}{\partial w_c} < 0, \frac{\partial c_j}{\partial p} > 0, \frac{\partial^2 c_j}{\partial p \partial t_j} > 0. \tag{14}$$

In other words, the fraction of clerks in the workforce rises as the production function becomes more complex. An increase in clerks' wages or a decrease in the price of information technology causes firms to substitute IT equipment for clerks. And finally, the fraction of clerks in the workforce is more responsive to changes in the price of IT equipment in more complex industries.

The remainder of this paper uses the framework I outlined to explain the variation in clerical employment across manufacturing industries and over time. Section 3 focuses on the cross-sectional prediction, that industries with a more complex production technology employ more clerks relative to other workers. Cross-industry regressions for all decades from 1910-2000 support the view that a more complex production technology is associated with more clerical employment.

Section 4 considers the effect of information technology, rather than production technology, on the demand for clerks.I show that before 1880 few clerks were employed in manufacturing industries and firms typically relied on informal information. The IT revolution of 1880-1910 induced firms to use information more intensively and employ clerks. Assuming that the ranking of industries by complexity are stable over time, the model also predicts that the increase in clerical employment be led by firms that were more complex in 1880. I compare these predictions with two alternative hypotheses, that an increase in either firm size or in the supply of educated workers caused the increase in clerical employment.

Finally, the model predicts that when all firms use high information intensity, a decline in the price of IT equipment reduces the fraction of clerks in the workforce.[5] Moreover, we expect that more complex industries substitute IT equipment for clerks at a faster rate. Section 5 shows evidence for these predictions using the recent computer revolution.

2.3 Technological Complexity and Demand for Clerks

Technological complexity is a measure of the variety of tasks required to physically make products. In terms of labor inputs, complex production processes involve an occupationally diverse workforce. Accordingly, I define complexity as 1 minus the Herfindahl index of occupations of employees in each industry. This measure is designed to reflect occupational heterogeneity in production, but not in coordination; therefore, I calculate this measure excluding managers, clerks, accountants and auditors. My main source of data is the 1 percent Integrated Public Use Microdata Series (IPUMS) of the decennial censuses (Ruggles et al. 1997). Since occupational definitions change over time, I use the IPUMS 1950 classification of occupations and industries throughout the paper.[6]

Appendix Table A1 shows the industries with particularly high or low levels of complexity from 1880-2000. During the early 20th century complexity was positively correlated with using continuous-process or batch-production technologies (Chandler 1977; Goldin and Katz

[5] Assuming that effect of the fall in the price of IT on reducing the demand for clerks dominates the effect of an increase in complexity on the demand for clerks.

[6] Computer-related occupations are the only exception to this rule, since they are only defined from 1970 onwards. The appendix discusses various measurement issues pertaining to the data.

1998), as industries using these technologies ranked above the median level of complexity.[7] In subsequent decades transportation industries ranked among the most complex industries, while textile and clothing industries were among the least complex.

Comparisons of industries and occupations over time should be taken with caution, even though they rely on a fixed classification of occupations. With this caveat in mind, note that complexity appears to have risen over time (Table 1). Complexity rankings are also very consistent over time: from 1880-1940 the decade-by-decade correlation of complexity ranged from 0.5-0.9, and since 1940 it has exceeded 0.9.

In order to examine the relationship between complexity and information processing, I investigate the employment of clerks. An examination of clerical occupations shows that they use IT equipment to generate, store, and communicate information, which is used for coordination. Throughout the 20th century clerical occupations in manufacturing included stenographers, typists, secretaries, bookkeepers, shipping and receiving clerks, and office machine operators. In recent decades, clerks have often also been employed in record processing, material recording, scheduling and distribution, providing information, and message distribution (Hunt and Hunt 1986).

Clerks differ from other white collar workers, since their main task is to process information that is used to coordinate production. By contrast, engineers and other professional workers typically process information that is used to design products or a processes. Sales workers may also process information as part of their work, but their main objective is to sell products. Clerks also differ from managers, since managers perform actual coordination and supervision. However, during the 19th century the distinction between clerks and managers was not as clear as it is today, as managers often processed information themselves. Therefore, in the empirical analysis I consider the differential effect of complexity on managers and clerks.

The task of processing information are typically skill-intensive, so clerks are more skilled

[7] Industries using continuous-process or batch-production technologies include beverages, dairy products, grain-mill product, paints and varnishes and petroleum refining.

than most other manufacturing workers. The earliest nationally representative micro data that identifies the education and occupation of individuals is the 1940 Census. Table 2 shows the education of clerks and other workers, over age 50, who were employed in manufacturing in 1940. Clerks had much higher levels of education than other workers, suggesting that even if we account for changes in occupation and industry and attrition from the labor force, clerks were highly educated workers in the early decades of the 20th century.[8] Table 2 also shows that the skill gap between clerks and other workers persisted at least through 1980, although it has narrowed over time.

The model predicts that when firms use high information intensity, more complex industries employ more clerks relative to other workers. In order to test this hypothesis I use individual-level census data to estimate a linear probability model:

$$c_{ij} = \alpha t_j + \beta k_j + z_i' \gamma + \varepsilon_{ij}, \tag{15}$$

where c_{ij} is an indicator for person i, working in industry j, being employed as a clerk (or a manager). The regressor of interest is industry j's complexity, t_j. I also include individual level geographic controls (z_i) and some specifications also control for the industry's capital/labor ratio, k_j.

The top panel of Table 3 shows results of cross-section regressions without industry-level controls for each decade from 1900-2000. These results suggest that since 1910, more complex industries have employed more clerks. An increase of one standard deviation in complexity was associated with an increase of about 1.6 percentage points in the employment of clerks in 1910. This partial correlation increased to about 2.9 percentage points in 1960 and declined to about 1 percentage point in 2000. In other words, from 1910-1980 a change in one standard deviation of complexity corresponded to 15-25 percent of the average clerical employment in manufacturing, declining to 8-13 percent in 1990 and 2000. The results are similar in

[8]Goldin and Katz (1999) show that clerks were much better educated than blue collar workers in Iowa in 1915.

magnitude, though less precise, after controlling for the capital/labor ratio. Table 3 also shows that managerial employment, unlike clerical employment, was not correlated with higher complexity in 1910 and 1920, and it was substantially weaker throughout the first half of the 20th century.[9] This lends evidence that clerks and managers performed different tasks even in the early 20th century, so the demand for their services was driven by different factors.

The evidence presented thus far is consistent with the hypothesis that the coordination of a more complex production process requires relatively more clerks. In order to further examine this idea, I examine how more complex industries respond to large changes in the cost of IT equipment.

2.4 The Early IT Revolution and the Rise of Clerks

This section discusses the early IT revolution, which took place from 1880-1910. I show that this revolution increased the demand for clerks, and that industries that were more complex to begin with increased their demand more rapidly. Next, I consider two alternative explanations for the growth in the demand for clerks. The first hypothesis attributes the increased demand for clerks to the growth in plant size. The second hypothesis is that increased supply of educated workers caused firms to develop strategies for employing them as clerks. I find little support for either hypothesis. Finally, I consider the implications of the increased demand for clerks on the accumulation of human capital and the employment of women.

Yates (1989) and Levenstein (1991) describe the IT equipment that came into widespread use in offices from 1880-1910. The new IT equipment had substantial impact on four different aspects of information processing: production, copying, storage, and communication.

First, the production of information was enhanced by the advent of the typewriter, which

[9] The relationship between complexity and managerial employment became significant from 1940 onwards, with a change of one standard deviation in complexity associated with 0.4-1.5 percentage points (10-20 percent change) in managerial employment. Interestingly, the correlation between complexity and managerial employment has strengthened in recent decades.

allowed rapid and inexpensive generation of neat documents. The patent that led to the first commercially viable typewriter was registered in 1868. During the 1870s the first typewriters were sold to court reporters and telegraphers. Once the value of typewriters was realized by firms, their production expanded rapidly, and in 1900 alone over 140,000 machines were sold. The average annual growth rate of the real value of typewriters produced in the U.S. from 1890-1909 was 9 percent, and it fell to 3 percent over the subsequent five years (see Figure 2). Complementary innovations, such as the visible typewriter and the dictaphone, allowed typists to see their output as they typed and to record dictations. Cash registers and other office machines also became popular, and from 1890-1909 the real value of such equipment grew at an average rate of 16 percent; the growth rate fell to a mere 3 percent over the subsequent five years.[10]

Second, the technology for copying information also improved from 1880-1910. Earlier copying technologies, such as the book press and the rolling press, suffered from substantial drawbacks, and carbon paper was incompatible with contemporary pens. With the advent of typewriters, carbon paper allowed clerks to produce several copies at the point of origin. This innovation reduced copying cost and made documents neater and more durable. Stencils, first patented in 1876, became useful for mass duplication of internal communication. The first photocopying machine was invented in 1900, and in 1911 the Taft Commission evaluated it and found it useful for repeated copying of complicated documents and diagrams. The widespread use of technologies for producing and copying information increased the use of paper. From 1890-1910, the tonnage of fine writing paper manufactured in the U.S. grew at a rate of 6.4 percent per capita, compared to 4.8 percent from 1910-1930 (Feenberg and Miron 1997).

Third, the technology for storage and retrieval of information was quite limited before the last decades of the 19th century. Documents were typically stored in pigeonhole desks that could only accomodate few documents, or in cabinets and flat files that made retrieval

[10]The data on output comes from Roy (1990). Following Goldin and Katz (1995), Table 5, I use the cost-of-living index due to Rees as the price deflator (Historical Statistics 1997, series E186).

difficult. Vertical filing was presented to the business world at Chicago's World Fair of 1893. It allowed systematic grouping of documents, quick rearrangement and retrieval, and substantial saving in storage space.

Finally, the technology for communicating information also improved from 1880-1910. The telegraph was introduced to America as early as the 1840s, but it proved less useful in manufacturing than in the railroad industry (Chandler 1977). More important for manufacturing was the telephone, which was invented in 1876 and came into commercial use shortly thereafter. By the 1890s private branch exchanges were used in cities across the U.S., and around 1900 long distance communication become possible. Figure 2 shows that the annual growth rate in phones per capita from 1880-1910 was about 16 percent; in the next decade it fell to about 4 percent.[11] According to data from the 1920s to the 1970s, about one-third of the phones in the U.S. were used by businesses, and this fraction declined only slowly over time (Historical Statistics 1997). The use of printed communication also grew rapidly: the annual growth rate of the number of items sent by the U.S. Postal Service was about 5.4 percent from 1874-1910 and only about 1.6 percent from 1910-1970.[12]

Taken together, this evidence shows that an early information technology revolution took place from 1880-1910. The key period for this revolution was the first decade of the 20th century, when technologies for producing, copying, storing and communicating information matured. This early IT revolution created the potential to use information much more intensively than before.

Throughout most of the 19th century manufacturing firms made little systematic use of information for coordination of production. Rather, firms were often managed by owners, typically with the help of skilled artisans or foremen (Chandler 1977; Hounshell 1984; Yates 1989; Nelson 1995). To be sure, there were exceptional firms that employed innovative

[11]Similarly, Sorkin (1980) finds that the number of telephone connections made by the Bell Company increased from 1.87 billion to 8.14 billion from 1900-1910. In per capita terms, this reflects an average annual growth rate of about 14 percent during the first decade of the 20th century, compared to 2.3% in the subsequent decade and an average growth rate of about 3.5% from 1920-1970.

[12]Author's calculations are based on data from Bangs (1875) and Sorkin (1980).

techniques of cost accounting; but these techniques were typically firm-specific and suffered from substantial shortcomings (Fleischmen and Tyson 1993).

During the late 19th century, a group of managers and engineers led a movement for "Systematic Management" (Nelson 1980, 1995).[13] They promoted systematic cost accounting, use of job cards and time clocks, inventory control, centralized purchasing, and incentive wages. Management literature had barely existed before 1870, but it expanded rapidly during the next thirty years. It was also around the turn of the century that several American universities first established programs in business administration. The next step in expanding the use of information was taken by the movement for "Scientific Management", led by Frederick W. Taylor. Taylor and his associates served as consultants and initiated changes in plant management practices. These changes included mechanical innovations, improved cost accounting and purchasing systems, the introduction of storerooms and planning departments, use of time studies, and more sophisticated incentive wage systems. While neither universally adopted nor entirely successful, these changes were a sign of the attempts to use information for managerial purposes.

Most of the literature on this organizational transformation has focused on the managers who initiated them (e.g. Chandler 1977 and Nelson 1995), but the role of clerks in facilitating these changes was no less important. As Table 1 shows, production workers accounted for about 95 percent of the manufacturing workforce in 1880, while most of the remaining 5 percent were managers and owners. But by 1910 clerks emerged as a new class of white collar employees in the manufacturing industries, outnumbering managers by about two-to-three.[14]

[13]The availability of new power sources (steam and later electric power) also allowed more flexibility in organizing the factory floor, and complemented these changes.

[14]Estimates using the IPUMS data and the 1950 occupational definitions suggest that there were about 18,000-20,000 clerks in manufacturing in 1870 and 1880. This figure corresponds to about 1 percent of the manufacturing labor force. The 1950 classification of clerical occupations appears to match earlier classification systems well only for data from 1880 onwards, so estimates prior to 1880 are potentially subject to substantial measurement error. The Census of Manufactures uses a narrower definition of clerical occupations than the population census. According to these definitions there were fewer than 6,000 and 12,000 clerks in manufacturing in 1870 and 1880.

Existing data on wages suggest that on average clerks earned almost twice as much as other manufacturing workers in 1890, reflecting their higher level of education at a time when high school graduates were scarce.[15] As Figure 4 shows, the wage of clerks relative to other manufacturing workers appears to have remained constant or even increased slightly from 1890-1910, while clerical employment increased rapidly. This consistent with the view that the changes in clerical employment reflect a large increase in the relative demand for clerks.

In order to examine how the changes in demand varied by industry, I use the measure of complexity introduced in the previous section. The model predicts that more complex industries increase their employment of clerks more rapidly than other industries, assuming that complexity ranking of industries is stable over time. To test this hypothesis I estimate a pooled cross-section regression for 1880, 1900, and 1910:

$$c_{ijt} = \alpha(t_{j,1880} * I_t) + x_j'\beta + z_i'\gamma + d_j'\delta + T_t'\lambda + \varepsilon_{ijt}, \tag{16}$$

where $t_{j,1880}$ is the complexity of industry j in 1880, I_t is a dummy for 1910 (after the information technology revolution had matured), x_j includes industry-level controls, d_j is a vector of industry fixed effects and T_t is a vector of year dummies.

As the results in Table 4 show, industries that were more complex in 1880 increased their employment of clerks more rapidly than other industries over the next three decades. These results are robust to different weighting methods (columns 1 and 2). Column 3 limits the sample to industries with large plants. The coefficient of interest, α, is larger and more precise, suggesting that my results are not driven by industries with small and heterogenous plants. In a similar vein, I exclude industries that are loosely defined (e.g. miscellaneous wood products), and the results are similar to the benchmark. The results are also robust to including only industries with more than 100 observations in 1880 (column 5). Finally I use

[15] These data should be taken with some caution, since they are not nationally representative. See Appendix Table A2 calculations, based on Goldin and Katz (1998).

an alternative measure of complexity, computed using only production workers; once again, the results are similar (column 6). These results suggest that an increase of one standard deviation of complexity in 1880 corresponds to an increase of about 1-2 percentage points (or about about 15-30 percent) in the employment of clerks in 1910, similar to the cross-sectional estimates. By comparison, the estimated effect of complexity on the relative employment of managers is positive but substantially smaller, and only marginally significant.

The evidence presented thus far is consistent with the predictions of the model; I now consider two alternative explanations for the increase in clerical employment from 1880-1910. The first alternative explanation is that plant size, rather than complexity, accounts for the increase in the employment of clerks.[16] Chandler (1977) discusses the growth in the scale and scope of corporations around the turn of the 20th century. Indeed, my calculations show that the average number of workers per plant in the median industry increased from about 16 to 40 from 1880-1910. However, the results in Table 4 and 5 indicate that plant size is uncorrelated with the employment of clerks.

The second alternative explanation is that information technology equipment and clerical employment increased in response to a rise in the level of education of the labor force.[17] But Goldin and Katz (1995) estimate that in 1910 only about 5.4 percent of the workforce had graduated from high school, compared to about 4 percent in 1890. In contrast, after 1910 the high school graduation rates in the U.S. accelerated dramatically (see Figure 5). The changes in the wage premium for clerks relative to production workers also appear to reflect this pattern: after remaining stable from 1890-1914, they declined rapidly in subsequent decades (Table A2). Thus, the aggregate evidence suggests that the increase in high school graduation rates appears to be an outcome, rather than a cause of the increase in clerical employment.

To further test the hypothesis that education increased clerical employment in manufac-

[16]Industry level regressions of complexity on ln(number of workers per plant), weighted by the number of observations in each industry show no statistically significant correlation between these two measures.

[17]This argument echoes Acemoglu (2002), who argues that the skill-bias of technological change in recent decades may be a result of an endogenous response to the growing number of college graduates.

turing, I use state-level variation in the level of education. I estimate the following regression for 1880, 1900 and 1910:

$$c_{is} = \alpha h_s + x_i'\beta + \varepsilon_{is}, \tag{17}$$

where i denotes a person, s denotes her state and c_{is} is an indicator for clerks. The measure of human capital in each state, h_s, measures the fraction of children aged 16-18 in school when the current census and the previous census were held.[18] Finally, x_i is a set of industry level controls. The results in columns 4-6 of Table 5 suggest no significant correlation of state-level education and the share of clerks in states' manufacturing labor force. I also test whether the share of clerks in manufacturing employment increased more rapidly from 1880-1910 in states that had more educated workers in 1880 by estimating the following specification:

$$c_{ist} = \alpha(h_{s,1880} * I_t) + z_s'\beta + d_i'\gamma + T_t'\delta + \varepsilon_{ijs}, \tag{18}$$

where z_s is a vector of state dummies and d_i is a vector of city size dummies and industry level controls (complexity and plant size). The results in the last column of Table 5 show some evidence that human capital stocks may have contributed to the subsequent rise in the employment of clerks, although the estimate is not statistically significant.

The rise in clerical demand appears to have contributed not only to the increase in the demand for skill, but also to the employment of women. Rotella (1979), Davis (1982), Fine (1990) and Hartman Strom (1989, 1992) show that the increased availability of clerical positions allowed many women to enter the labor force. Indeed, from 1880-1910 the share of women among clerks employed in manufacturing industries rose from a mere 5.3 percent to over 36 percent. However, even as late as 1900-1910, men took up more net new clerical jobs than women. Hence, the feminization of office work appears to have been a response to a growing demand for qualified and inexpensive workers.

The results in this section show that clerks and information technology equipment were

[18]This measure is designed to reduce the measurement error as much as possible.

complements around the turn of the 20th century. The next section examines whether they were sill complements in recent decades.

2.5 The Recent IT Revolution and the Decline of Clerks

The computer revolution that has taken place in recent decades shares much in common with the IT revolution that occurred around the turn of the 20th century. First, computer software (e.g. word processors and electronic spreadsheets) and hardware (e.g. printers) substantially reduce the cost of producing information, just as the typewriter had done almost one century earlier. Second, computers allow us to copy information at almost no cost. Third, computers allow effective storage and retrieval of vast amounts of information. And finally, computer networks make communication cheaper and more effective, as the telephone had done in the past.

In 1970 there were fewer than 75,000 computers in the U.S., but over the next decade about 875,000 computers were purchased (Phister 1979; Hunt and Hunt 1986). Substantial reduction in the cost of processing power (Jorgenson 2001; Nordhaus 2001) induced a transition from mainframe computers to microcomputers. The range of computer applications also increased over time. By 1984 nearly half of the clerks in the U.S. used computers, compared to only a quarter of other workers; more recent data show that the fraction of clerks (and other workers) using computers has since continued to rise (Friedberg 2003). I therefore conclude that during the 1970s (and possibly the early 1980s) the price of office computers declined sharply, and we can examine the effect of this decline on the demand for clerks.

The new IT equipment reduced the demand for clerks in three different ways (Hunt and Hunt 1986; Osterman 1986). First, new equipment has directly replaced clerical workers . For example, automatic switching replaced telephone operators; office dictation equipment took the place of stenographers; computer software for pricing insurance supplanted specialized clerks (raters); automatic mail sorting devices replaced mail clerks; and computerized inventory management software took the place of shipping and receiving clerks. Second, indirect replacement occurred as the new equipment raised the productivity of clerical work, reducing the number of clerks required to perform a given task. Finally, reorganization of information processing has also reduced the demand for clerks. For example, some stenographers and typists lost their jobs when word processing allowed workers to produce information cheaply and efficiently.

Technological change may also have an opposite effect of increasing the demand for clerks. Theory predicts that if the complexity of production increases, firms may demand more information, thereby raising their relative demand for clerks. In addition, Osterman (1986) argues that a tendency for "information deepening" may also act as a countervailing force. In other words, the lower cost of processing information may cause firms to produce more information relative to other intermediate inputs.[19]

As Table 1 shows, the fraction of clerks in the manufacturing workforce fell from 14.4 percent in 1960 to 11.4 percent in 2000. This decline does not reflect an increase in the relative wages of clerks. As Appendix Table A2 shows, the wage of clerks relative to other manufacturing workers has gradually fallen over this period. Taken together, this evidence suggests manufacturers' demand for clerk, relative to other workers, has been falling.

Of particular interest is the effect is the differential response of complex industries to the fall in the price of IT equipment. The first four columns in Table 6 show the cross-sectional relation between complexity and the employment of computer personnel from 1970-2000.[20] The coefficient is positive in all years, although it is statistically significant at the 5 percent level only for 1980 and 2000 and marginally significant for 1970 and 1990. These results are consistent with the model's prediction that more complex industries have higher demand for information technology equipment.

Given the persistence of complexity, the model predicts that more complex industries substitute IT equipment for clerks more rapidly than other industries. Assuming that the

[19]The framework I use assumes that when all firms use high information intensity, information and other inputs are used in fixed ratios as the price of IT falls.
[20]See Data Appendix for a definition of the category of "computer personnel."

number of computer operators required for each computer does not vary across industries in each year, we expect that higher complexity leads to greater employment of computer personnel. Table 6 shows the results of the following regression using data from 1960-1990:

$$c_{ijt} = \alpha(t_{j,1960} * I_t) + x_j'\beta + z_i'\gamma + d_j'\delta + T_t'\lambda + \varepsilon_{ijt}, \tag{19}$$

where $t_{j,1960}$ is the complexity of industry j in 1960, I_t is a dummy for 1980 and 1990 (after the computer revolution had matured), x_j includes industry-level controls, d_j is a vector of industry fixed effects and T_t is a vector of year effects. The results suggest that more complex industries significantly reduced their employment of clerks. An increase of one s.d. in complexity in 1960 is associated with a decrease of about 1.3 percentage points in clerical employment in 1980 and 1990.

Table 6 also shows that complex industries increased the demand for computer personnel more than other industries. However, this differntial increase was only about 1/20 in magnitude compared to the decline in their employment of clerks. In part, this finding may reflect the fact that a single computer operator may be in charge of multiple computers or multiple user interfaces. It may also reflect the capacity of a single computer operator to substitute multiple clerks that are not using computers. Finally, it is also possible that some jobs may have been outsourced to service providers outside manufacturing.[21]

In summary, these findings show that manufacturing industries substituted computer personnel for clerks, and that this substitution was significantly stronger among more complex industries.

2.6 Conclusions

This paper analyzes the evolution of the demand for clerks in manufacturing over more than a century. Clerks are skilled workers who process information that is used to coordinate

[21] Interestingly, complex industries also appear to have increased their demand for managers more than other industries. This finding suggests that they may face more challenges in coordinating complex production processes, even when plenty of information is available.

production. In order to explore the determinants of the costs of coordination I define the complexity of an industry's production process as a measure of the occupational heterogeneity of its workforce. I show that throughout the 20th century more complex industries consistently employed relatively more clerks.

I also find that the price of information technology affected the demand for clerks. An information technology revolution that took place from 1880-1910 introduced to the office innovations such as the typewriter, the telephone, and improved filing techniques. During that period clerks and information technology equipment were complements, since in conjunction they allowed firms to use information much more intensively than they had previously done. Importantly, industries that were more complex in 1880 increased their demand for clerks more rapidly than other industries.

The technology-driven increase in the demand for clerks from 1880-1910 was rapid, and it had important implications for the US labor force. First, the increased availability of clerical positions raised the demand for skill at a time when high school graduation rates were low. This finding suggests that technological change may have contributed to the onset of the High School Movement from 1910 onwards. Second, the increased demand for clerks contributed to the rise in labor force participation of women. Like the increased demand for skill, the rise of women's labor force participation was an important trend that has persisted throughout the 20th century.

Whereas in the past clerks and IT equipment were complements, in recent decades they have become substitutes. I find that manufacturing firms have substituted computers for clerks, reducing the employment share of clerks despite the decline in their relative wages. This result provides a concrete case in which technology and a specific set of skills are substitutes rather than complements. Moreover, my findings suggest that the replacement of clerks by computers was more rapid in more complex industries.

2.7 Data Appendix

The data for this paper comes primarily from the Integrated Public Use Microdata Series (IPUMS) samples of the U.S. decennial household census. This data includes samples of approximately 1 percent of the decennial US census from 1860-2000, excluding 1890 and 1930, for which no data is currently available. This data allows a systematic cross-sectional and longitudinal analysis of the manufacturing workforce, while controlling for geographic variation. Using the IPUMS 1950 classification of occupations and industries mitigates some of the problems that could arise from changes in the classification of jobs throughout the period.[22] Nevertheless, before 1880 people were not directly asked for both their occupation and industry, so the distinction between the two had to be imputed and this created a potential for substantial measurement error. This data also lacks the wage and educational attainment of workers before 1940.[23] In order to obtain wage data for that period, I rely on the estimates of Goldin and Katz (1995), which are based on partial data.

I include workers in all manufacturing industries, except for 'miscellaneous manufacturing industries' and 'not specified manufacturing industries'. Dropping these observations reduced the sample by about 5 percent. Specific occupations that appeared in the data but seem incompatible with employment in manufacturing were excluded from the sample. These occupations include: chiropractors; clergymen; dancers and dancing teachers; entertainers; farm and home management advisors; foresters and conservationists; funeral directors and embalmers; recreation and group workers; religious workers; therapists and healers; farmers (owners and tenants); farm managers; auctioneers; hucksters and peddlers; real estate

[22] In the classification of industries, I merged together several industries, in order to facilitate the merging of the household census data with the census of manufactures data. Specifically, I merged industries 306 (logging) and 307 (sawmills, planing mills, and mill work); industries 336 (blast furnaces, steel works, and rolling mills) and 337 (other primary iron and steel industries); industries 346 (fabricated steel products) and 347 (fabricated nonferrous metal products); industries 439 (yarn, thread, and fabric mills), 436 (miscellaneous textile mill products) and 449 (miscellaneous fabricated textile products); and industries 399 (miscellaneous manufacturing industries) and 499 (not specified manufacturing industries). I also dropped all the observations that are classified as: employed, unclassifiable; non-occupational response; occupation missing/unknown; and N/A (blank).

[23] The schooling variable does make it possible to estimate high school participation across geographical regions.

agents and brokers; stock and bond salesmen; all service workers; all farm laborers; people whose occupation was unclassified; people with a non-occupational response; and all people with occupation missing/unknown or non-available. Dropping all these occupations further reduced the sample by about 3 percent.

The only industry-level variable constructed using the IPUMS data is the complexity measure. It is defined as one minus the Herfindahl index of occupations, excluding managers, clerks and accountants and auditors. In some specifications I also use an alternative measure of complexity, calculated using only production workers (Craftsmen, Operatives, and Laborers not elsewhere classified).

I construct other industry-level variables using the Census of Manufactures data for the years 1880-1909 and 1960-1990. The data for the early period includes the number of workers and establishments, as well as the value of output and cost of various inputs for over 250 industries, as classified by Roy (1990). The more recent data is from the NBER-CES Manufacturing Industry Database (Bartelsman et al. 2000).

The household census and the Census of Manufactures data use different classifications of industries and the latter classification is typically much more detailed. In order to match the data I pooled together several of the IPUMS industries: logging was merged with sawmills, planing mills, and mill work; blast furnaces, steel works, and rolling mills was merged with other primary iron and steel industries; yarn, thread, and fabric mills was merged with miscellaneous textile mill products and miscellaneous fabricated textile products. Merging these industries yielded a total of 52 industries, although a few had no observations in some census years.

Under the IPUMS classification of occupations, all white collar workers are classified into one of four one-digit categories: clerks, managers, professional workers and sales workers. I constructed the category of computer personnel based on the IPUMS occupations for 1970-2000 since no such category existed prior to 1970. In 1970 this category includes computer and peripheral equipment operators. In 1980 and 1990 it includes supervisors, computer equipment operators; computer operators; and peripheral equipment operators. In 2000 it includes computer and information systems managers; database administrators; network and computer systems administrators; and computer operators.

References

Acemoglu, Daron, “Directed Technical Change”, Review of Economic Studies, 2002, p. 782-809.

Autor, David H., Frank Levy and Richard J. Murnane, "The Skill Content of Recent Technological Change: An Empirical Exploration", Forthcoming Quarterly Journal of Economics, 118(4), November 2003.

Bangs, George S., "A Letter to the Postmaster General showing the Cost, Revenues and Comparative Profit and Loss of Carrying the Different Classes of Mail Matter", Washington Government Printing Office, 1875.

Bartelsman, Eric J., Randy A. Becker, and Wayne B. Gray, "NBER-CES Manufacturing Industry Database", National Bureau of Economic Research, June 2000
http://www.nber.org/nberces/nbprod96.htm

Becker, Gary S. And Kevin M. Murphy, "The Division of Labor, Coordination Costs, and Knowledge", *Quarterly Journal of Economics* v107, n4 (November 1992): 1137-60

Carter, Susan B. Et al., ed. *Historical statistics of the United States on CD-ROM: colonial times to 1970*, U.S. Bureau of the Census.

Chandler, Alfred D., Jr., *The Visible Hand: The Managerial Revolution in the American Business*, The Belknap Press of the Harvard University Press, 1977.

Davis, Margery W., *Women's Place Is at the Typewriter – Office Work and Office Workers 1870-1930*, Temple University Press, 1982.

Feenberg Daniel and Jeff Miron "Improving the Accessibility of the NBER's Historical Data", Journal of Business and Economic Statistics, Volume 15 Number 3 (July 1997) pages 293-299. Data series: m01104a U.S. Fine Writing Paper Production 01/1918-12/1933

Fine, Lisa M., *The Souls of the Skyscraper – Female Clerical Workers in Chicago, 1870-1930*, Temple University Press, Philadelphia, 1990.

Fleischman Richard K. and Thomas N. Tyson, "Cost Accounting during the Industrial Revolution: The Present State of Historical Knowledge", The Economic History Review, New Series, Volume 46, Issue 3 (Aug. 1993), p. 503-517.

Friedberg, Leora, "The Impact of Technological Change on Older Workers: Evidence from Data on Computer Use", *Industrial and Labor Relations Review*, Vol. 56, No. 3 (April 2003).

Garicano, Luis, "Hierarchies and the Organization of Knowledge in Product'ion", *Journal of Political Economy* v108, n5 (October 2000): 874-904.

Goldin, Claudia, "America's Graduation from High School: The Evolution and Spread of Secondary Schooling in the Twentieth Century", *Journal of Economic History*, Vol. 58 (June 1998): 345-374

Goldin, Claudia and Katz, Lawrence F. "Education and Income in the Early 20th Century: Evidence from the Prairies." National Bureau of Economic Research, Inc, NBER Working Papers: 7217, 1999

Goldin, Claudia and Lawrence F. Katz "The Decline of Non-Competing Groups: Changes in the Premium to Education, 1890 to 1940.", National Bureau of Economic Research Working Paper: 5202, 1995.

Goldin, Claudia and Lawrence F. Katz "The Origins of Technology-Skill Complementarity", Quarterly Journal of Economics, 1998.

Hartman Strom, Sharon, "'Light Manufacturing': The Feminization of American Oddice Work, 1900-1930", *Journal of Industrial and Labor Relations Review*, Volume 49, Issue 1, 1989, p. 53-71.

Hartman Strom, Sharon, *Beyond the Typewriter: Gender, Class and the Origins of Modern American Office Work, 1900-1930*, University of Illinois Press, 1992.

Hounshell, David A., *From the American System to Mass Production, 1800-1932*, Johns Hopkins University Press, 1984.

Hunt, H. Allan and Timothy L. Hunt, *Clerical employment and technological change*, W.E. Upjohn Institute for Employment Research, 1986.

Jorgenson, Dale W., "Information Technology and the U.S. Economy", *American Eco-*

nomic Review, March 2001, p. 1-32.

Levenstein, Margaret, "The Use of Cost Measures: The Dow Chemical Company, 1890-1914" in *Inside the Business Enterprise: The Use and Transformation of Information*, ed. Peter Temin (University of Chicago Press), 1991.

Nelson, Daniel, *Fredrick W. Taylor and the Rise of Scientific Management*, The University of Wisconsin Press, 1980.

Nelson, Daniel, *Managers and Workers – origins of the Twentieth-Century Factory System in the United States, 1880-1920*, The University of Wisconsin Press, second edition (1995).

Nordhaus, William, "The Progress of Computing", Cowles Foundation Discussion Paper No. 1324, 2001.

Osterman, Paul, "The Impact of Computers on the Employment of Clerks and Managers", *Industrial and Labor Relations Review*, Volume 39, Issue 2, Jan. 1986, p. 175-186.

Phister, Montgomery, Jr., *Data Processing Technology and Economics*, Second Edition, Digital Press, 1979. [Table II.1.28]

Rotella, Elyce J., Women's Labor Force Participation and the Growth of Clerical Employment in the United States, 1870-1930 (in Summaries of Doctoral Dissertations), *The Journal of Economic History*, Vol. 39, No. 1, The Tasks of Economic History. (Mar., 1979), pp. 331-333.

Roy, William G., "Rise of American Industrial Corporations, 1880-1914" [Computer file]. Los Angeles, CA: University of California, Dept. Of Sociology [producer], 1990. Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor], 1990.

Ruggles, Steven et. Al., *Integrated Public Use Microdata Series: Version 2.0*, Minneapolis: Historical Census Projects, University of Minnesota, 1997 (www.ipums.org).

Smith, Adam, *The Wealth of Nations*, 1776 (Reprinted in New York: Modern Library, 1965).

Sorkin, Alan L., *The Economics of the Postal System*, Lexington Books, 1980.

U.S. Department of Commerce, Bureau of the Census, *Historical Statistics of the United States, Bicentennial edition*, Cambridge University Press, 1997 (CD).

U.S. Census Bureau, Population Division, Education and Social Stratification Branch, "Computer Use in the Unitd States", October 1989 (P23-171) http://www.census.gov/population/www/socdemo/computer/p23-171.html.

Yates, JoAnne, "Control through Communication – The Rise of System in American Management", Johns Hopkins University Press, 1989.

Table 1. Descriptive Statistics

Year	Total manufacturing workforce (millions)	Fraction of white collar occupations in manufacturing: Clerks	Managers	Professional workers	Sales workers	Computer administrators and operators (see note)	All white collar workers	Fraction of women among clerks	Average industry complexity
1860	0.9	0.001	0.022	0.004	0.000	0.000	0.028	0.000	0.400
1870	1.6	0.012	0.031	0.004	0.000	0.000	0.047	0.089	0.438
1880	1.8	0.011	0.033	0.007	0.001	0.000	0.052	0.046	0.361
1900	3.8	0.026	0.035	0.010	0.001	0.000	0.071	0.173	0.583
1910	7.6	0.064	0.043	0.011	0.010	0.000	0.128	0.357	0.684
1920	9.1	0.083	0.039	0.015	0.002	0.000	0.139	0.450	0.688
1940	11.1	0.103	0.041	0.030	0.031	0.000	0.205	0.428	0.679
1950	14.6	0.113	0.047	0.042	0.023	0.000	0.223	0.542	0.714
1960	21.4	0.144	0.044	0.062	0.026	0.000	0.275	0.690	0.696
1970	24.4	0.140	0.045	0.081	0.023	0.002	0.290	0.697	0.716
1980	25.5	0.131	0.066	0.090	0.022	0.005	0.309	0.691	0.755
1990	23.4	0.124	0.096	0.121	0.030	0.007	0.370	0.696	0.791
2000	23.1	0.114	0.099	0.138	0.030	0.007	0.381	0.667	0.775

NOTES: The sample includes workers in manufacturing industries, and it is described in detail in the Data Appendix. Using the IPUMS 1950 classification of occupations, all white collar workers fall into one of the following categories: clerks, managers, professional workers and sales workers. The category of computer administrators and operators includes some clerks, managers and professional workers (see details in Data Appendix). The complexity measure is computed as one minus the Herfindahl index of occupations, excluding managers, clerks and accountants and auditors. The means are calculated using census person weights.

Table 2. Education of Clerks and Other Workers in Manufacturing

	1940				1980	
	Age 50 and over		No age restriction		No age restriction	
	Clerks	Non-clerks	Clerks	Non-clerks	Clerks	Non-clerks
Education level:						
No high school	0.46	0.77	0.20	0.57	0.04	0.14
Some high school	0.18	0.10	0.20	0.21	0.12	0.22
High school graduates	0.24	0.08	0.44	0.16	0.49	0.37
Some college	0.07	0.03	0.11	0.04	0.27	0.17
College graduates	0.05	0.03	0.05	0.03	0.07	0.11

NOTES: The sample includes workers in manufacturing industries, and it is described in detail in the Data Appendix. The means are calculated using census person weights.

Table 3. Cross-Section Regressions for Clerks and Managers: 1900-2000

	1900	1910	1920	1940	1950	1960	1970	1980	1990	2000
					A. Excluding industry-level controls					
Clerks										
Complexity	-0.003	0.102	0.086	0.112	0.107	0.136	0.151	0.157	0.139	0.106
	(0.020)	(0.042)	(0.029)	(0.029)	(0.029)	(0.031)	(0.033)	(0.046)	(0.049)	(0.044)
Managers										
Complexity	-0.017	-0.01	-0.034	0.039	0.023	0.039	0.052	0.083	0.128	0.156
	(0.028)	(0.025)	(0.018)	(0.010)	(0.012)	(0.009)	(0.010)	(0.018)	(0.022)	(0.023)
Observations	19,080	30,018	90,504	112,424	170,925	214,951	243,983	255,021	237,953	233,649
					B. Including industry-level controls					
Clerks										
Complexity		0.088				0.118	0.131	0.149	0.129	
		(0.052)				(0.033)	(0.034)	(0.046)	(0.049)	
Capital/Labor		0.013				0.01	0.011	0.005	0.004	
		(0.011)				(0.007)	(0.006)	(0.005)	(0.005)	
Managers										
Complexity		-0.025				0.04	0.049	0.078	0.12	
		(0.027)				(0.011)	(0.011)	(0.018)	(0.026)	
Capital/Labor		0.011				0	0.002	0.003	0.004	
		(0.009)				(0.003)	(0.003)	(0.003)	(0.004)	
Observations		29,489				214,245	242,389	254,621	237,407	

NOTES: The sample includes workers in all manufacturing industries, according to the IPUMS 1950 classification of industries, except miscellaneous manufacturing industries and not specified manufacturing industries. Specific occupations that were likely the result of classification error are excluded from the sample (see Data Appendix). The complexity measure is computed as one minus the herfindahl index of occupations, excluding managers, clerks and accountants and auditors. All the regressions include state dummies. Dummies for the size of place in which the person lived were included in the years for which this information was available (1900, 1910, 1940, 1950, 1980, 1990). All regressions are weighted (using person-weights) and the standard errors are clustered at the industry level.

Table 4. The Effect of Complexity in 1880 on the Change in Employment of Clerks and Managers in 1910

	(1)	(2)	(3)	(4)	(5)	(6)
			A. Clerks			
(Complexity in 1880)* (Year=1910)	0.048	0.051	0.082	0.037	0.049	0.066
	(0.019)	(0.019)	(0.015)	(0.019)	(0.020)	(0.017)
Ln(capital / labor)	-0.012	-0.014	-0.047	-0.01	-0.012	-0.019
	(0.012)	(0.015)	(0.018)	(0.012)	(0.012)	(0.012)
Ln(average stablishment size)	-0.008	-0.005	-0.024	-0.01	-0.009	-0.009
	(0.006)	(0.007)	(0.005)	(0.006)	(0.006)	(0.006)
Year=1900	0.019	0.018	0.039	0.021	0.019	0.023
	(0.008)	(0.011)	(0.009)	(0.008)	(0.008)	(0.008)
Year=1910	0.034	0.032	0.062	0.039	0.033	0.036
	(0.012)	(0.015)	(0.013)	(0.013)	(0.012)	(0.014)
			B. Managers			
(Complexity in 1880)* (Year=1910)	0.017	0.018	0.015	0.009	0.017	0.017
	(0.008)	(0.009)	(0.009)	(0.009)	(0.009)	(0.010)
Ln(capital / labor)	0.007	0.007	-0.009	0.011	0.008	0.006
	(0.007)	(0.008)	(0.016)	(0.006)	(0.007)	(0.007)
Ln(average stablishment size)	0.005	0.002	0.002	0.005	0.005	0.005
	(0.002)	(0.003)	(0.004)	(0.002)	(0.002)	(0.002)
Year=1900	-0.006	-0.004	-0.002	-0.004	-0.006	-0.005
	(0.004)	(0.005)	(0.007)	(0.003)	(0.004)	(0.004)
Year=1910	-0.011	-0.008	-0.002	-0.01	-0.012	-0.009
	(0.006)	(0.007)	(0.012)	(0.006)	(0.007)	(0.006)
Observations	66,473	66,473	39,950	50,309	60,452	66,473

NOTES: The sample includes all manufacturing industries, except miscellaneous and not-specified manufacturing industries and occupations likely misreported (see Data Appendix). The complexity measure is computed as one minus the herfindahl index of occupations, excluding managers, clerks and accountants and auditors. All regressions include industry fixed effects and city size dummies. The column specifications are: (1) Benchmark specification (sum of person weights is one for each year); (2) Standard person weights; (3) Only industries above median in terms of 1880 estaqblishment size; (4) Excludes "miscellaneous" industries; (5) Includes only industries with more than 100 observations in 1880; (6) Complexity as calculated based only on production workers (Craftsmen, Operatives, Laborers n.e.c.). Robust standard errors in parentheses adjusted for clustering by industry*(Year=1910).

Table 5. Testing Alternative Hypotheses

Effects of Establishment Size on the Employment of Clerks

	(1)	(2)	(3)
	1880	1900	1910
size)	0.000	-0.006	-0.011
	(0.002)	(0.003)	(0.006)
Complexity	0.003	-0.007	0.072
	(0.012)	(0.025)	(0.053)
Ln(Capital/Labor)	0.012	0.017	0.022
	(0.006)	(0.008)	(0.012)
State fixed effects	Yes	Yes	Yes
Observations	18,132	18,852	29,489

Effects of State-Level Education on the Employment of Clerks

	(4)	(5)	(6)	(7)
	1880	1900	1910	1880-1910
State schooling	-0.019	-0.005	0.014	
	(0.011)	(0.029)	(0.050)	
(State schooling in 1880)				0.039
				(0.026)
Ln(Capital/Labor)	0.012	0.017	0.021	0.016
	(0.002)	(0.006)	(0.003)	(0.002)
Ln(average stablishment size)	0	-0.006	-0.011	-0.006
	(0.001)	(0.002)	(0.001)	(0.001)
Complexity	0.004	-0.005	0.072	0.017
	(0.005)	(0.013)	(0.017)	(0.009)
year=1900				0.007
				(0.003)
year=1910				0.023
				(0.009)
State fixed effects	No	No	No	Yes
Observations	18,100	18,771	29,334	67,009

NOTES: The sample includes workers in all manufacturing industries, according to the IPUMS 1950 classification of industries, except miscellaneous manufacturing industries and not specified manufacturing industries. Specific occupations that were likely the result of classification error are excluded from the sample (see Data Appendix). The complexity measure is computed as one minus the herfindahl index of occupations, excluding managers, clerks and accountants and auditors. All regressions are weighted using person-weights, and in the last column these weights are normalized so that they add up to one in each year. In columns (1)-(3) the standard errors are clustered at the industry level, in columns (4)-(6) they are clustered at the state level, and in column (7) they are clustered at the state*(year=1910) level.

Table 6. Regressions for Clerks, Managers and Computer Personnel: 1960-2000

	Dependent variable: computer				Pooled cross-section regressions (1960-1990)					
					Dependent variable: clerks		Dependent variable: managers		Dependent variable: computer personnel	
	1970	1980	1990	2000						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Complexity	0.002	0.008	0.011	0.017						
	(0.001)	(0.003)	(0.007)	(0.006)						
(Complexity in 1960)*					-0.063	-0.065	0.023	0.024	0.003	0.003
(year=1980 or 1990)					(0.008)	(0.008)	(0.006)	(0.006)	(0.001)	(0.001)
Ln(capital/labor)	0.000	0.000	-0.001		0	0	0.028	0.028	0.003	0.003
	(0.000)	(0.001)	(0.001)		(0.008)	(0.009)	(0.008)	(0.008)	(0.001)	(0.001)
Year=1970					-0.009	-0.01	-0.006	-0.006	0.001	0.001
					(0.004)	(0.004)	(0.003)	(0.003)	(0.000)	(0.000)
Year=1980					0.023	0.025	-0.012	-0.012	0.001	0.001
					(0.005)	(0.005)	(0.005)	(0.005)	(0.001)	(0.001)
Year=1990					0.01	0.011	0.007	0.006	0.002	0.002
					(0.007)	(0.007)	(0.007)	(0.007)	(0.001)	(0.001)
City size dummies	No	Yes	Yes	No	No	No	No	No	No	No
Industry dummies	No	No	No	No	Yes	Yes	Yes	Yes	Yes	Yes
Observations	242,389	254,621	237,407	233,649	948,662	948,662	948,662	948,662	948,662	948,662

NOTES: The sample includes workers in all manufacturing industries, except miscellaneous manufacturing industries and not specified manufacturing industries. Specific occupations that were likely the result of classification error are excluded from the sample (see Data Appendix). The complexity measure is computed as one minus the herfindahl index of occupations, excluding managers, clerks and accountants and auditors. All regressions are weighted (using person-weights) and they all include state dummies. Dummies for the size of place in which the person lived were included in the years for which this information was available (1980 and 1990). All regressions are weighted - columns (5),(7) and (9) use standard person weights and specification while columns (6), (8) and (10) use normalized weights, such that the sum of weights in each year is equal to one. Standard errors for columns the columns (1)-(4) are clustered at the industry level and for columns (5)-(10) they are clustered at the industry*(year=1980 or 1990) level.

Appendix Table A1. Industries with High and Low Levels of Complexity: 1880-2000

	1880	1910	1940	1970	2000
High levels of complexity					
1	Blast furnaces, steel works and rolling mills	Printing, publishing, and allied industries	Ship and boat building and repairing	Ship and boat building and repairing	Ship and boat building and repairing
2	Drugs and medicines	Petroleum refining	Petroleum refining	Petroleum refining	Aircraft and parts
3	Primary nonferrous industries	Fabricated steel products	Railroad and misc transportation equipment	Blast furnaces, steel works and rolling mills	Petroleum refining
4	Printing, publishing, and allied industries	Railroad and misc transportation equipment	Aircraft and parts	Aircraft and parts	Drugs and medicines
5	Professional equipment	Agricultural machinery and tractors	Agricultural machinery and tractors	Cement, concrete, gypsum and plaster products	Railroad and misc transportation equipment
Low levels of complexity					
1	Cement, concrete, gypsum and plaster products	Tobacco manufactures	Footwear, except rubber	Apparel and accessories	Apparel and accessories
2	Knitting mills	Leather products, except footwear	Knitting mills	Knitting mills	Carpets, rugs, and other floor coverings
3	Petroleum refining	Footwear, except rubber	Apparel and accessories	Footwear, except rubber	Leather products, except footwear
4	Canning and preserving fruits, vegetables, and seafoods	Knitting mills	Leather products, except footwear	Leather products, except footwear	Paperboard containers and boxes
5	Confectionary and related products	Paperboard containers and boxes	Tobacco manufactures	Canning and preserving fruits, vegetables, and seafoods	Rubber products

NOTES: The sample includes the manufacturing industries, accoring to the IPUMS 1950 classification of industries. Industries with vague classification and industries that had fewer than 10 observations in a given year were excluded from this table.

Appendix Table A2. Wages of Clerks in Manufacturing

	Wage premium of clerks relative to production workers		Annual earnings of clerks relative to production workers	
	Men	Women	Men	Women
		A. 1890-1939		
1890				1.85
1895			1.69	1.94
1909			1.65	1.96
1914			1.70	2.07
1919			1.20	1.53
1924			1.10	1.40
1929			1.13	1.53
1939			1.15	1.56
		B. 1939-1999		
1939	0.89	1.12	1.12	1.70
1949	0.97	0.99	1.02	1.30
1959	0.90	0.91	1.10	1.23
1969	0.89	0.86	1.10	1.21
1979	0.87	0.87	1.06	1.19
1989	0.82	0.80	0.97	1.01
1999	0.77	0.74	0.91	0.89

NOTES: The estimates in Panel A are taken from Table 5 in Goldin and Katz (1995), and they are based in part on sources that are not nationally representative. The estimates in Panel B are based on the author's calculations from IPUMS data. The earnings premium for clerks is defined as the ratio of the median annual earnings of clerks and non-clerks in manufacturing industries.

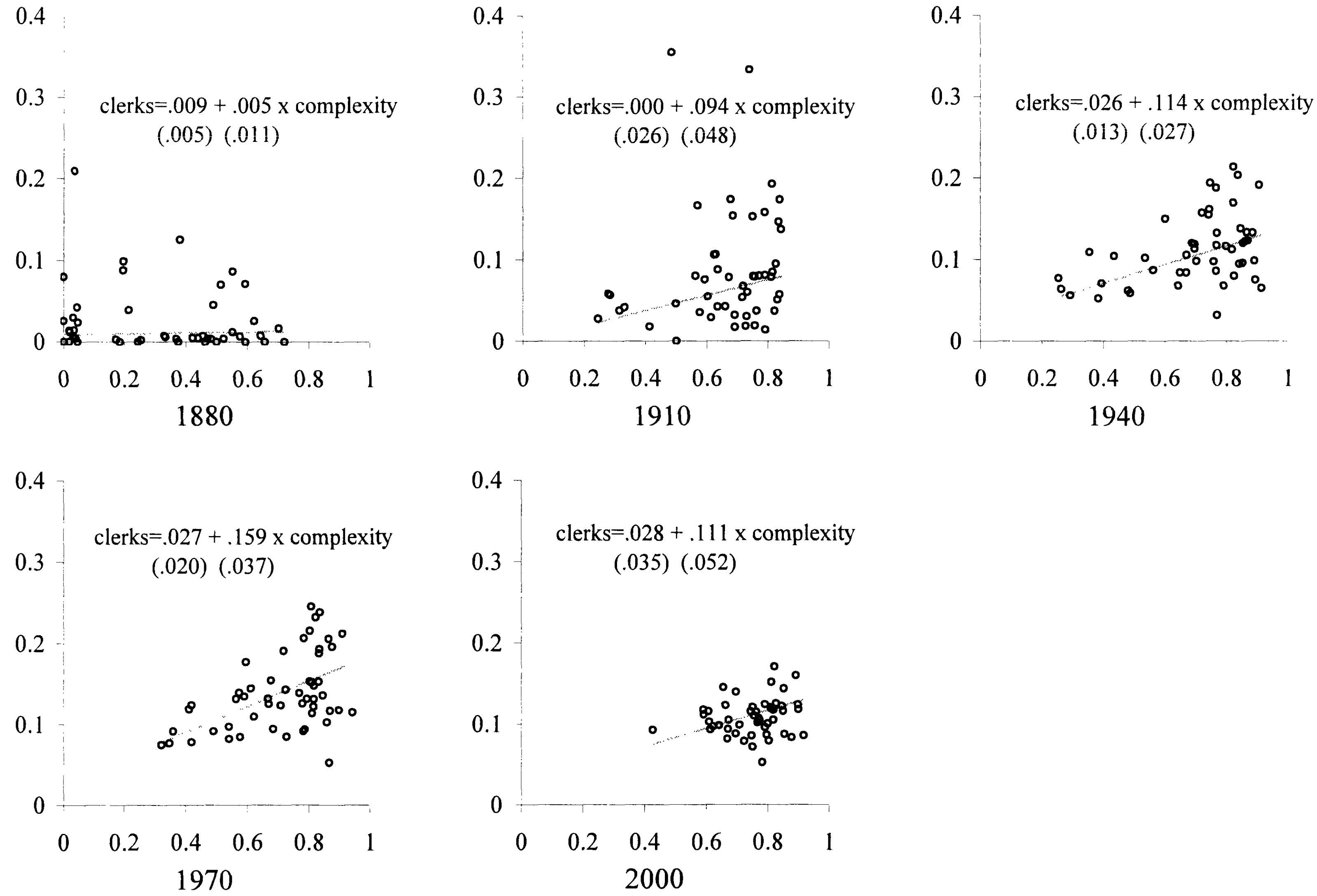

Figure 1. Percentage of clerks in workforce as a function of complexity: 1880-2000. Source: author's calculations.

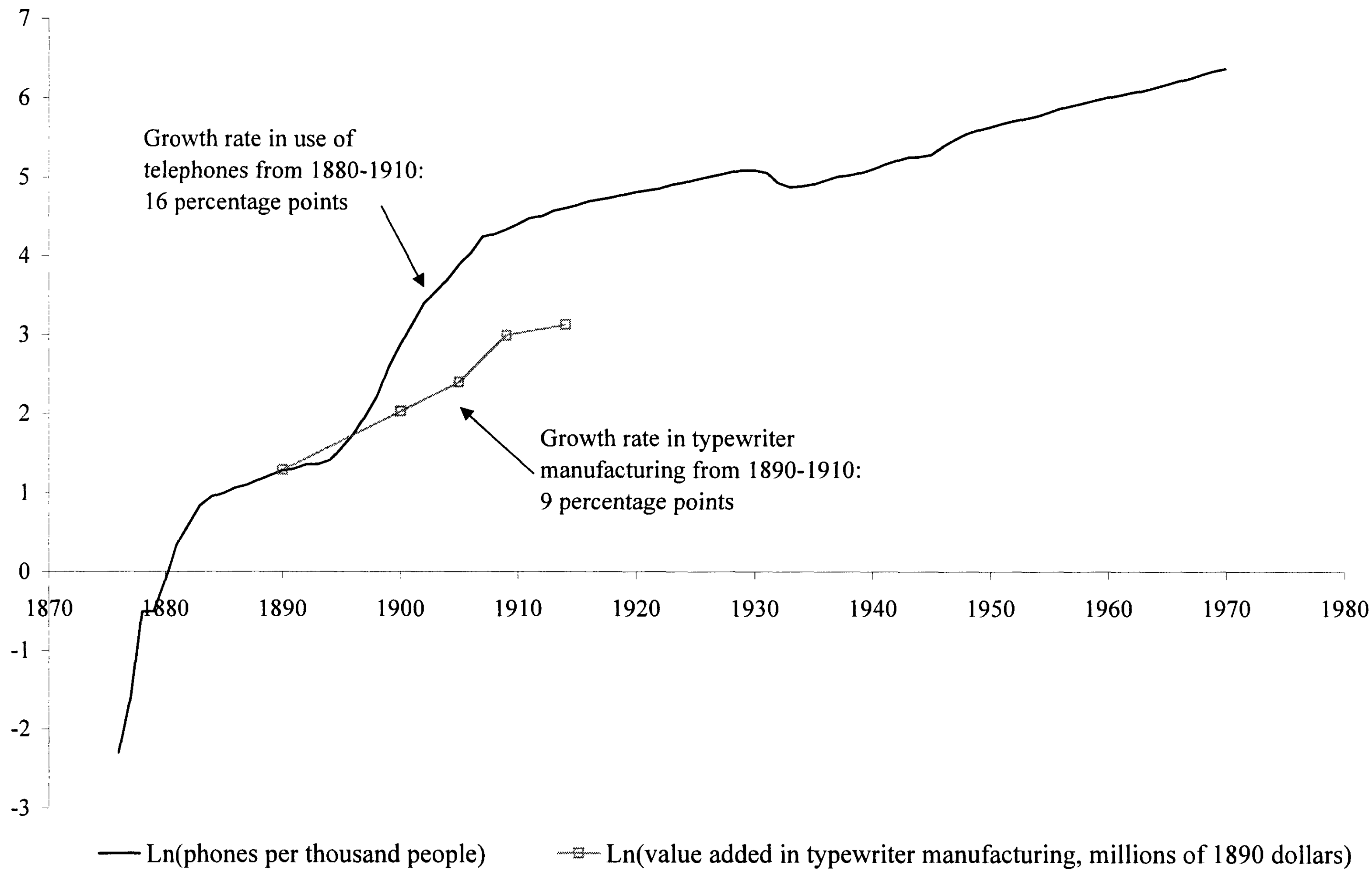

Figure 2. Diffusion of information technology equipment during the early IT revolution. Source: author's calculations based on the Census of Manufactures.

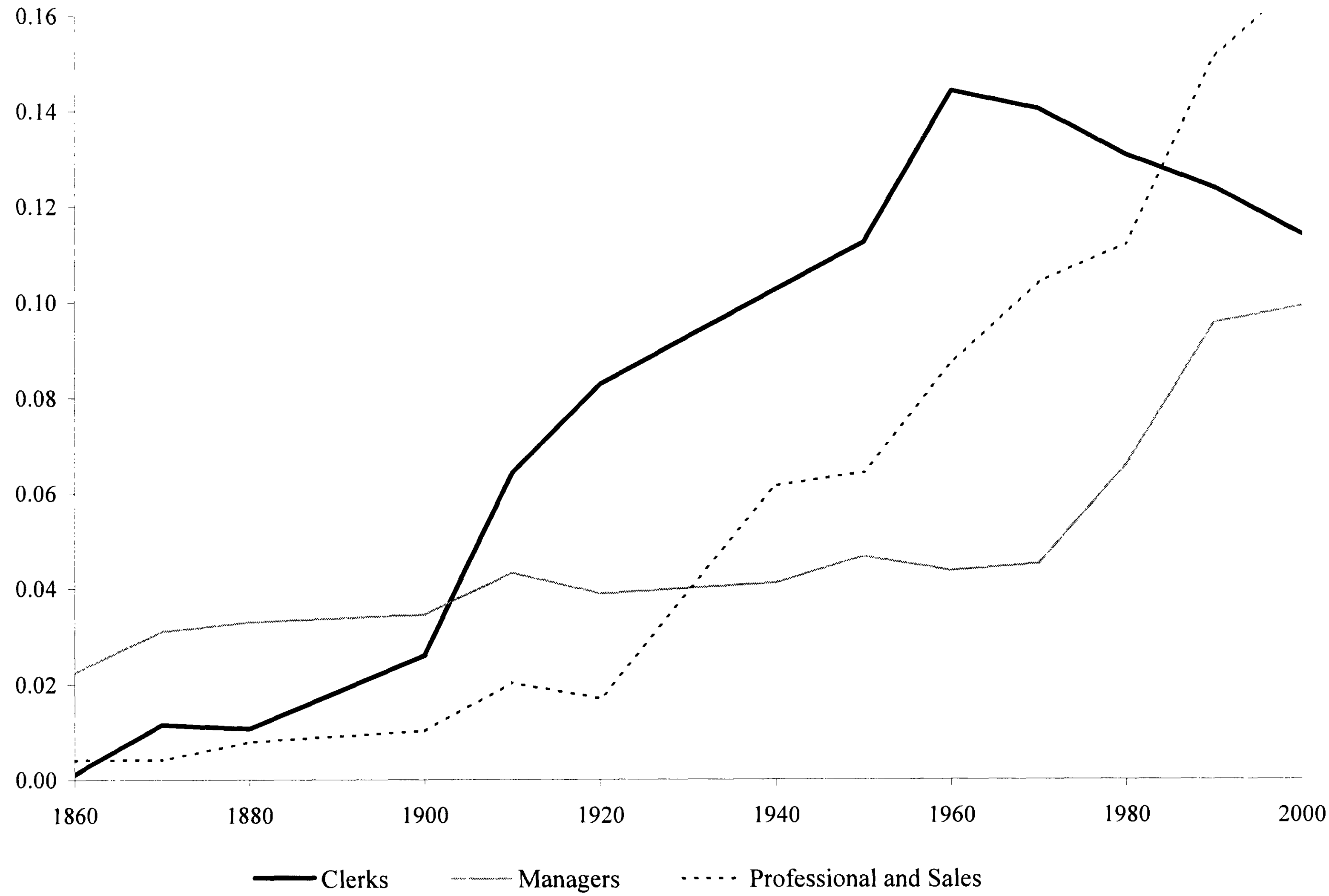

Figure 3. Fraction of white-collar workers in manufacturing, by occupational class. Source: author's calculations.

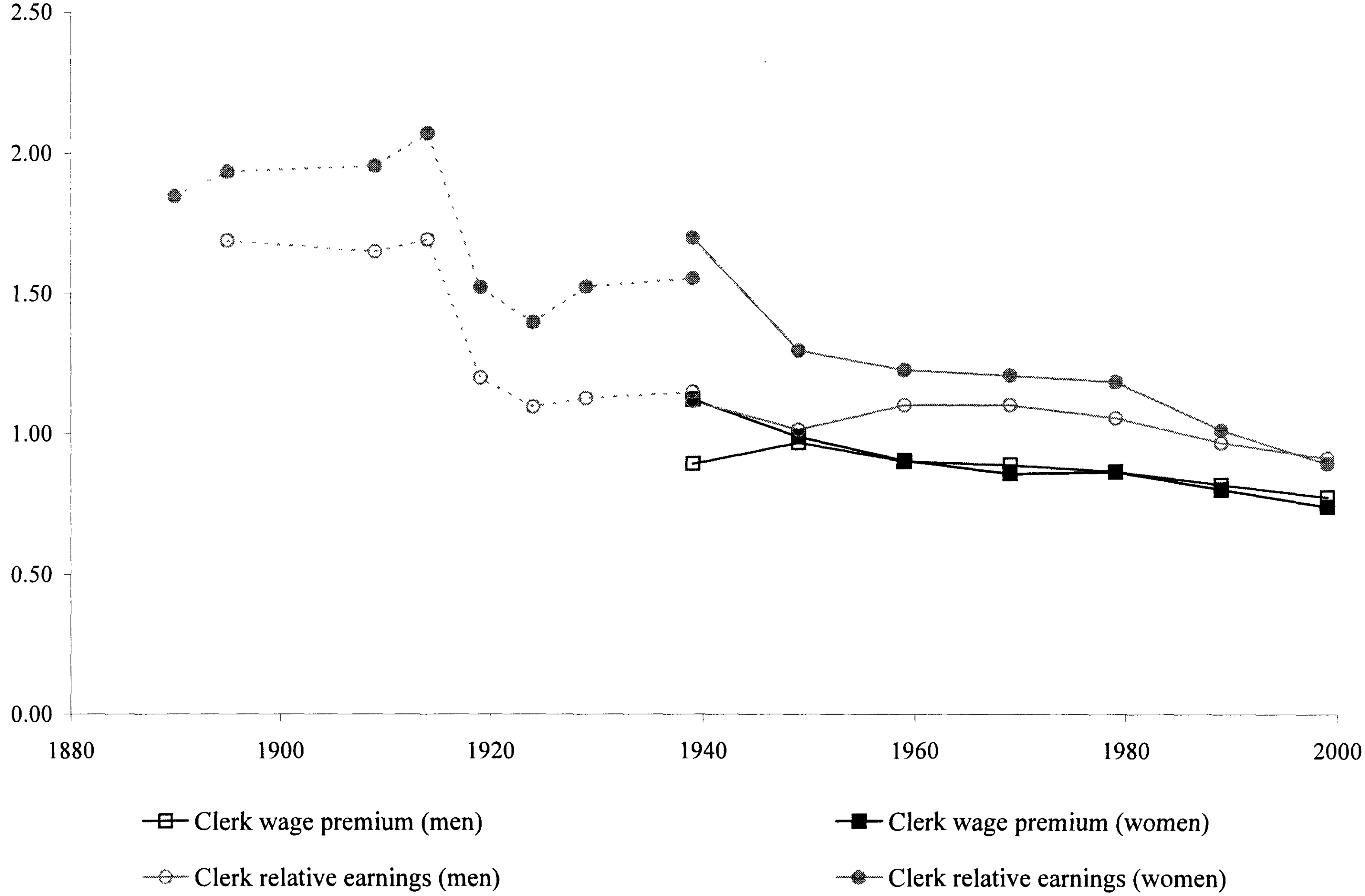

Figure 4. Wages of clerks relative to production workers in manufacturing: 1880-2000. Broken lines denote data from before 1940, which are not nationally representative. Sources: see Appendix Table A2.

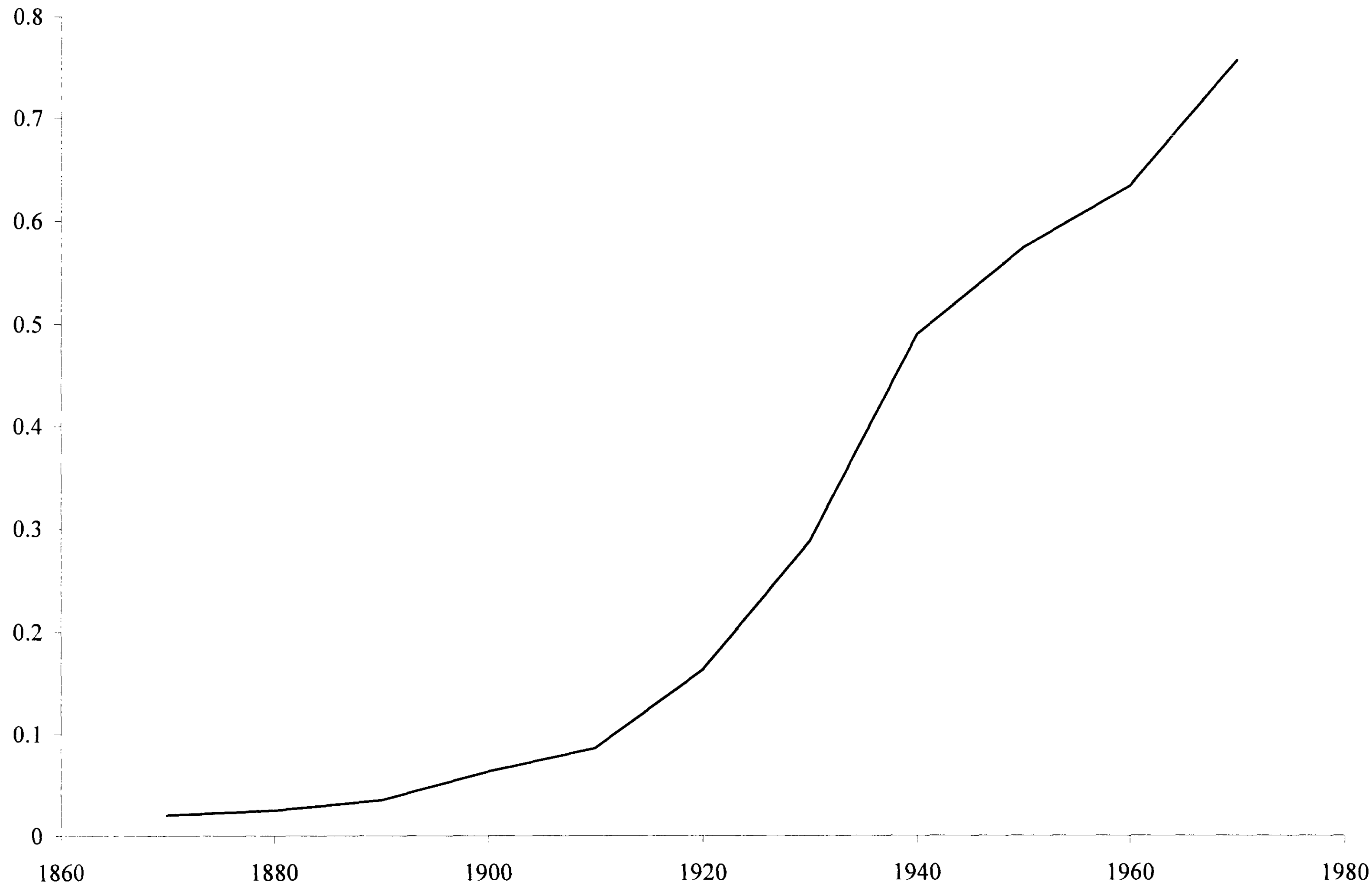

Figure 5. High-school graduates as fraction of 17 year-olds: 1870-1970. Source: Historical Statistics of the United States (1997).

Chapter 3

The Effect of Marital Breakup on the Income and Poverty of Women with Children

(Joint with Elizabeth O. Ananat, MIT)

3.1 Introduction and Motivation

The poverty rate for single mothers fell substantially between 1974 and 2002. Over the same period, the poverty rate for married mothers remained virtually unchanged. Given these facts, one might assume that women with children are less likely to be poor than they were 30 years ago. In fact, however, the overall poverty rate for women with children rose slightly over this period (Figure 1).

A clue to resolving this puzzle may be found in the fact that divorce and single parenthood have been increasing dramatically over the past several decades throughout the developed world. In the United States, the proportion of mothers who are single rose from about 16 percent in 1974 to roughly 26 percent in 2002 (Figure 2). It appears that in the absence of this trend, overall poverty rates would have decreased, rather than increased.

In order to establish this argument, however, one would need to demonstrate a causal relationship between divorce and poverty. Current political discussions commonly assume that marriage has causal beneficial effects on women and children. In particular, recent welfare legislation encourages marriage as a method of increasing income and reducing the need or eligibility for welfare. President Bush's 2005 budget proposal earmarks $1.2 billion of its Temporary Assistance to Needy Families (TANF, or welfare) budget for a five-year initiative "supporting healthy marriages"[1]; five states allocate some of their general TANF budget to marriage promotion activities, and ten states provide cash marriage incentives in the structure of their TANF benefits[2].

Despite the assumptions made in popular debate, a causal relationship between economic well-being and marriage preservation has not been established. Indeed, there are reasons to suspect that the correlation between marital breakup and poverty overstates the causal effect of divorce. For example, negative household income shocks such as job loss or falling real wages may increase marital tension and thereby increase the likelihood of divorce, so that poverty causes divorce rather than divorce driving poverty. In addition, women with worse earnings opportunities may be less attractive marriage partners, which would mean that women who get divorced have lower income, but not simply because they got divorced. Of course, the possibility also exists that the causal relation is understated.

This paper uses an instrumental variables (IV) approach to separate the causal effects of divorce from its well-known correlations. Using the 1980 U.S. Census, we document that having a female first-born child slightly, but robustly, increases the probability that a woman's first marriage breaks up. We find that the likelihood that a woman's first marriage is broken is 0.63 percentage points higher if her first child is a girl, representing a 3.7 percent increase from a base likelihood of 17.2 percentage points. This result, which we will argue implies that sons have a stabilizing effect on the family unit, is consistent with the finding (established in the psychology and sociology literature) that fathers bond more with sons.

When we use child sex as an instrumental variable to measure the effect of divorce on economic outcomes, we find that the marital breakup driven by having a girl does not

[1] U.S. Department of Health and Human Services, "FY2005 Budget in Brief."
http://www.hhs.gov/budget/05budget/acf.html

[2] Gardiner, Karen, Michael Fishman, Plamen Nikolov, Asaph Glosser, and Stephanie Laud. "State Policies to Promote Marriage: Final Report." U.S. Department of Health and Human Services Assistant Secretary for Planning and Evaluation, 2002.
http://aspe.hhs.gov/hsp/marriage02f/report.htm

significantly affect mean household income. Marital breakup does, however, significantly affect the household income distribution; in particular, it dramatically increases the probability that a mother will end up with very low income. While child support, welfare, and an increase in her own earnings (due to a substantially increased labor supply) go some way towards alleviating the loss of her husband's earnings, they often do not prevent poverty status. We also find some evidence that women who divorce may be slightly more likely to end up at the top of the income distribution, possibly due to re-marriage.

Several concerns might reasonably be raised about causal interpretation of the relationships between child sex, marital breakup, and economic outcomes. First, one might be concerned that the observed relationship between child sex and divorce is in fact an artifact of differential custody rates; women are more likely to lose custody of their boys in divorce, leading to divorced women being observed with fewer boys. However, we find highly comparable estimates in Current Population Survey (CPS) supplements that provide detailed fertility histories, which lends evidence that our results are not due to bias in observed household composition. Second, biologists (Trivers and Willard 1973) argue that child sex is endogenous to resources—that is, that natural selection has favored the ability of parents to have more girls in bad economic times. We find, however, that in our sample there is no difference between the economic resources of mothers of boys and mothers of girls when their first child is born; rather, the difference in income builds up over time, as does the difference in divorce rates. Third, one might worry that sex of the first-born child affects economic outcomes through channels other than divorce. For example, parents may work harder when they have a son. But we find a negligible effect of child sex on economic outcomes in a sample of women with a low exogenous probability of divorce, which provides evidence that child sex is not affecting economic outcomes through a channel other than marital breakup.

After discussing our findings, we consider the potential macro implications of the relationship we identify between divorce and poverty. In recent decades, mothers' poverty rates have failed to decline as much as the overall rate. Can the increase in the proportion of divorced women in the U.S. explain this stagnation? We calculate that, in fact, the poverty rate of women with children today is substantially higher than it would be if divorce had not risen over the post-1980 period.

The paper proceeds as follows. In section 2, we set our findings in the context of previous research. In section 3, we describe the data and the sample we use. In section 4, we describe the estimation strategy for analyzing mean outcomes and describe the results. In section 5, we develop a framework for estimating distributional outcomes and discuss the results. In section 6, we conclude.

3.2 Divorce and Children's Gender

Previous research has established the need for instrumental variables when examining the effect of marital status on women's outcomes, both in theory (e.g. Becker, Landes, and Michael 1977; Becker 1985) and empirically (e.g. Angrist and Evans 1998). In particular, Gruber (2000) emphasizes the necessity of alternatives to cross-sectional analysis when measuring the effect of divorce on child outcomes.

We instrument for the breakup of a woman's first marriage using the sex of the first-born child. Having a first-born girl could be positively correlated with breakup of the first marriage for one or more reasons involving both the husband and the wife. First, a man with a daughter may be less interested in sustaining marriage than a man with a son. He may intrinsically value a son more, perceiving him as an "heir." Or he may have more in common with a boy than with a girl, and therefore bond better over time. Second, a woman with a daughter may be less interested in sustaining marriage than a woman with a son. For example, she may believe a son needs his father as a role model, and therefore work harder to retain her husband. Or she may get better companionship from a daughter, and therefore place less weight on her relationship with her husband.

The previous literature supports the idea that fathers tend to bond more with their sons. Aldous, Mulligan and Bjarnason (1998), using data from the U.S. National Surveys of Families and Households, find that fathers spend more time with boys than with girls, and that the gender gap in paternal attention increases as the child ages. This supports the idea of increased bonding over time. Harris, Furstenberg and Marmer (1998) construct a "parental involvement index" and find that fathers' involvement with their adolescent children is higher with sons than with daughters, while mothers' involvement does not differ by child sex. This supports the claim that fathers bond more with sons, but does not lend evidence that mothers get better companionship from daughters. Thus the literature is more consistent with the theory that fathers bond more with sons, and that this has a stabilizing effect on the family unit.

Previous work (e.g. Angrist and Evans 1998) considers child sex as an exogenous variable when analyzing U.S. data and after conditioning on race. This seems particularly plausible when considering only births that took place prior to the 1980's, when ultrasound was limited to problematic pregnancies (Campbell 2000). Nevertheless, concern has been raised by some that child sex is endogenous to resource levels. Trivers and Willard (1973) argue that a male is likely to out-reproduce a female if both are in good condition, and vice versa if both are in bad condition. They argue that natural selection may therefore have favored parental ability to adjust sex ratios so that more girls are born in bad times, and they argue that data from mammals support their hypothesis. Our findings, however, are not consistent with the Trivers and Willard hypothesis. We find no relationship between economic circumstances at birth and child sex. Our findings are consistent with the hypothesis that the sex of the eldest child is exogenous.

Dahl and Moretti (2004), in simultaneous research, examine the relationship between child sex and divorce—that is, our first stage. They too find a significant relationship between having female children and divorce, although our estimates of the effect size are smaller because we address serious selection problems that otherwise overstate the relationship.

In related research, Bedard and Deschenes (2003) use the same instrumental variable for divorce: gender of the first child. When we use specifications similar to theirs, our results are generally consistent with their findings: they, like we, find that divorce has no significant negative causal effect on women's mean economic outcomes and that the cross-sectional relationship between divorce and household income at the mean is due to selection.

However, when we look beyond the mean to consider the way in which marital breakup affects the entire income distribution of women with children, we arrive at a different conclusion. Namely, we find that marital breakup has significant negative consequences for many women; in particular, it increases the probability that a woman lives in poverty.

We also find some evidence that women who remarry may actually become "better off" in terms of income after the end of their first marriages. This is consistent with an earlier literature that does not attempt to establish causality or deal with selection into remarriage: Mueller and Pope (1980) and Jacobs and Furstenberg (1986) find that women who remarry do better in terms of their husbands' education level and SES score. Duncan and Hoffman (1985) find that five years after their divorce, women who remarried gained in terms of family income.[3]

Previous literature has not emphasized the relationship between divorce and inequality, although both have increased substantially over the past three decades. Much of the recent

[3] Both Jacobs and Furstenberg and Duncan and Hoffman argue that this is largely due to life cycle effects, so that *if* the man these women divorced is similar to the average person in his age group, he would have made similar gains from the time of his divorce until the time we observe the woman's second husband. We suspect, however, that in our case the men who divorce are negatively selected, so it is plausible that women who re-married did somewhat improve their economic situation.

literature on the causes of inequality has focused on wage inequality and the forces that may be affecting it, such as: technology (Acemoglu 2002); the decline of labor market institutions (DiNardo, Fortin, and Lemieux 1996); and the rise of international trade. Our findings suggest that the decline of the traditional family unit as an institution may have contributed to the rise in income inequality.

3.3 Data

We use data on women living with minor children from the 5 percent 1980 Census file, which allow us sufficient power to identify the effect of sex of the first-born child on marital breakup.[4] We limit our sample to white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. These limitations are necessary in order to create a sample for which measurement error in the sex of the observed first-born child has a classical structure. In particular it is important to consider only women who live with all their children, since by doing so we avoid any spurious correlation between child sex and household structure due to differential maternal custody rates of boys and girls. For further discussion of the sample construction, see the Appendix.

Table 1 shows summary statistics for our full sample and subsamples relative to the overall population of women with minor children. Our sample is quite similar to the overall population except in terms of age and marital status. Our sample is younger than average, consistent with the requirement that a woman's eldest child is under 17. The women in our sample are slightly less likely to be divorced, both because they are younger and because we require that they have custody of all children. And of course, unlike the overall population,

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[4] While earlier censuses have these measures, in previous decades the divorce rate was very low. Subsequent censuses, on the other hand, don't have all of the necessary measures. We did, however, try similar specifications using 1990 data and found similar patterns to those reported here; this finding lends evidence that cohort-specific effects are not driving the relationship between firstborn sex and divorce.

women in our sample cannot be never-married. On other characteristics, however, the two groups differ little: women in our sample have slightly more education and household income than the overall population and work and earn slightly less.

In addition to estimating our model on the full sample, we look specifically at two subsamples that we use in specification checks: those who are at high risk of having ever divorced and those at low risk of having ever divorced. We create an index of exogenous risk for divorce, which is orthogonal to the sex of the eldest child (see Appendix). The high-risk subsample includes women whose predicted risk is in the top quartile; the low-risk subsample includes women whose predicted risk is in the bottom quartile.

The high-predicted divorce subsample is much older on average than the low-predicted divorce subsample, consistent with more years of exposure to risk of divorce. They are also less educated, but they have higher own and household income—probably due to lifecycle effects.

The two subsamples allow us to test the validity of our IV strategy. As Table 1 shows, women in the high-predicted divorce subsample are much more likely to have undergone divorce by the time we observe them, so if the responsiveness to the child-sex instrument is proportional to the overall level of divorce then we would expect the first stage to be bigger for this group. Moreover, the high predicted-divorce subsample has lower years of schooling and thus may be, on average, of lower SES; below, we conjecture that our instrument may play a larger role in marital breakup decisions for this group. Thus we expect the first stage to be stronger for the high-divorce subsample than for the low-divorce subsample. As a consequence, the two-stage effect of child sex on outcomes should also be more precise for the high-divorce subsample, if indeed our IV strategy is valid and child sex is affecting outcomes through divorce rather than through other channels.

3.4 The Effect of Marital Breakup on Average Economic Outcomes

3.4.1 Estimation Framework

We begin our investigation of the causal effects of divorce by considering a simple econometric model. In this model, income is affected by the breakup of the first marriage, which is treated as a classic endogenous regressor:

(1) $$D = \alpha_1 Z + X\alpha_2 + U\alpha_3 + u$$

(2) $$Y = \beta_1 D + X\beta_2 + U\beta_3 + \varepsilon$$

The right-hand-side variable of interest in equation (2), D, is a dummy for the breakup of the first marriage. We define a woman as having her first marriage intact if she reported both that she was "currently married with spouse present" and that she had been married exactly once. We define as having her first marriage broken any woman who: has been married multiple times, is married but currently not living with her husband, is currently separated from her husband, is currently divorced, or is currently widowed.[5]

Our outcomes, Y, are total others' income, defined as total household income less total own income; household income; a measure of household poverty; and in some specifications hours worked last year. The change in others' income measures the direct effect on a woman of losing her husband as a source of income (to the extent that the husband is not replaced by other wage earners). The change in total household income captures this direct effect but also includes the indirect effects of divorce on income: transfers from the ex-husband in the

[5] We have run the analysis with ever-divorced, rather than first marriage broken, as the explanatory variable: in this case, widows and those separated from or not living with their first husband are coded as 0 rather than 1. Our results are not sensitive to this difference in categorization.

It is not possible to systematically remove widows from the sample, since the data do not allow us to identify those whose first marriage ended in death among those who have had multiple marriages. In any event, since widowhood is endogenous to both socioeconomic status and marital duration, it is probably not desirable to exclude widows.

form of alimony and child support[6]; transfers from the state in the form of cash assistance; and income generated by the woman's own labor supply response.[7]

Our controls, denoted by X, are a vector of pre-determined demographic variables including age, age squared, age at first birth and a dummy for high-school dropouts.[8] While our Ordinary Least Squares (OLS) estimates of the relationship between marital breakup and income may be sensitive to their inclusion, our IV specifications are robust to controls (results without controls are available upon request). We think of U as representing unobserved factors such as human capital, views on gender roles, and taste for non-market work relative to market work and leisure. Finally, our instrumental variable, Z, is an indicator for having a girl as one's firstborn child.

As discussed above, there are two major problems with estimating equation (2) with OLS. The first, correlation of U and D, can be seen as omitted variables bias or a selection problem, and it induces a bias of indeterminate direction. Women with worse earnings opportunities may be less attractive marriage partners: this would imply negative selection into marital breakup and upward bias in the estimated cost of marital breakup for women. On the other hand, women with worse earnings opportunities may be more interested in sustaining marriage. Or a stronger preference for household work relative to market work may cause a woman to work harder at staying married. In this case, the women who stay married would have supplied fewer hours of labor had their marriage broken up, and their earnings would have been lower.

[6] The 1980 Census question reads: "Unemployment compensation, veterans' payments, pensions, alimony or child support, or any other sources of income received regularly… Exclude lump-sum payments such as money from an inheritance or the sale of a home."

[7] We have also looked at various intermediate outcomes, such as the level of alimony and child support, the level of welfare, the woman's own earnings, and the woman's hours. All of the analyses gave results highly consistent with the results presented here, and are available from the authors upon request.

[8] Since we look only at women who gave birth to their first child after age 19, we can reasonably assume that the decision on whether to graduate from high school is made prior to the realization of the sex of the first-born child.

The second major problem with OLS estimates of equation (2) is reverse causality. Conditional on a woman's observed and unobserved characteristics, a negative income shock to the household may increase marital tension and thereby raise the likelihood of divorce.

To address concerns both about omitted variable bias and about reverse causality, we estimate equation (2) using two-stage least squares, using the sex of the eldest child (Z) as an instrument for whether the first marriage is broken (D). Angrist and Imbens (1994) show that in the absence of covariates and given the standard two-stage least squares assumptions,[9] the IV approach identifies the local average treatment effect:

$$(3) \qquad \beta_{1,IV} = \frac{Cov(D,Y)}{Var(D)} = \frac{E[Y \mid Z=1] - E[Y \mid Z=0]}{E[D \mid Z=1] - E[D \mid Z=0]} = E[Y_1 - Y_0 \mid D_1 > D_0]$$

Here Y_1 and Y_0 denote the income (or other dependent variable) for women whose first marriage is broken and intact, respectively. D_1 is an indicator for whether a woman would divorce if her first child were a girl; D_0 is an indicator for whether she would divorce if her first child were a boy. D_1 and D_0 are, of course, just hypothetical constructs; in practice we can only observe the indicator for the child sex that is realized.

β_{1IV} therefore measures the change in income due to divorce for women whose first marriage breaks up if they have a girl and remains intact if they have a boy. That is, two-stage least squares estimates the average effect of divorce only for that part of the population for whom the treatment is equal to the instrument ($D_1 > D_0$), the group known as "compliers." As one might imagine, this group is typically not a random subsample of the

[9] The assumptions are: conditional independence of Z, exclusion of the instrument, existence of a first stage and monotonicity – see Abadie (2002). The result can be generalized for the case with covariates.

population, and in fact we will show below that those who comply with sex of the first-born child tend to be of low socio-economic status (SES).[10]

3.4.2 First Stage Results: Child Sex and Marital Breakup

The first stage results for the entire population as well as for the full sample and various subsamples are presented in Table 2. When equation (1) is estimated for all women living with minor children, the coefficient on child sex is 1.13 percent. We believe that this figure overstates the actual effect of child sex on marital breakup, because endogeneity in the sex of the eldest child residing with a woman will create a spurious correlation between living with an eldest girl and being divorced.[11] That is, since women are more likely to end up with custody of girls than boys in the event of divorce, living with a girl is cross-sectionally correlated with being divorced above and beyond the causal effect of having a girl on marital breakup.

The second column gives the estimated relationship for our sample, which is constructed in order to close the custody channel, so that the relationship between child sex and marital breakup is causal. We find that the average effect of the first child being a girl on the breakup of the first marriage is about 0.63 percent, which is substantially smaller than the full-population estimate, but still highly significant. Furthermore, as discussed below, we replicate our result using data from the Current Population Survey, which identifies the sex

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[10] Note that β_{1IV} can only be interpreted as the causal effect of divorce for the average complier to the extent that the sample is not contaminated with "defiers". Defiers are women for whom $D_0 > D_1$, that is, women whose first marriage would dissolve only if their firstborn is a boy. We have run first-stage regressions for a variety of sub-populations and found no significant sign reversals in the coefficient of child sex on divorce, suggesting that defiers, if any, are a very small group.

[11] In fact, a regression similar to those in Table 2 for a sample of women with all of our restrictions except that she must reside with all of her children gives an estimate of .010. This result is much closer to the estimate in column 1 than in column 2, suggesting that the factor driving the difference between the estimates in the full population and in our sample is the endogeneity of child residence.

of a woman's actual firstborn child (and hence requires no sample restrictions). The CPS results are highly similar to those from our Census sample.

The effect increases with the age of the eldest child, suggesting, sensibly, that the sex of the first child affects the hazard rate of divorce and separation.[12] This can also be seen in Figures 3 and 4, each presenting the results from 17 separate regressions by the age (0 to 16) of the eldest child. We also find that the effect size varies by the woman's education—high-school dropouts exhibit the strongest effect, and the effect size diminishes as education increases. This suggests that our instrument primarily provides us insight on the effects of marital breakup particularly for low-SES women.

Finally, we find a larger and more significant effect on those at higher exogenous risk of divorce: the quarter of our sample with the highest predicted risk of divorce has a first stage of 1.13 percent (highly significant), and the group that is least likely to divorce has a first stage of 0.36 percent (marginally significant). The differential effect by underlying probability of divorce lends evidence that the instrument is acting in the way we would anticipate.

One further implication of the figures should also be noted: a comparison between Figures 3 and 4 reveals that the effects of sex of the eldest child on the breakup of the first marriage are bigger than those on being currently divorced. Much of this difference comes from the fact that some of the women who divorced due to the instrument have since remarried. In addition, some of those who have not re-married have instead moved in with other adults (such as their parents). We find that having an eldest girl increases the probability that a woman is the head of household by only about half as much as the increase in probability that her first marriage is broken (result not shown).[13] It is therefore important

[12] The base hazard rate of divorce may also change with the length of the relationship, which is correlated with the age of the eldest child.

[13] Note that this will not capture all avenues of combining households, since a woman may share a household and be identified as its head.

to keep in mind that we are not estimating the effect of being currently divorced or of residing in a mother-only household on economic outcomes. Rather, we are estimating the effect of having ever been divorced on current outcomes.

3.4.3 Specification Checks

While the lack of over-identification implies that we cannot formally test the validity of our instrument, we do bring to bear some evidence that the instrument is not correlated with other omitted variables. First, we test the Trivers and Willard hypothesis that adverse conditions cause women to give birth to a higher proportion of girls. We have examined the income of mothers who were never married, and found that the income distribution in this group was very similar whether the eldest was a girl or a boy. We have also examined the income of women who were married when their first child was born and whose first marriage was still intact, and similarly found that the income distribution in this group was virtually the same whether the eldest was a girl or a boy. (Results not shown.)

This is still not a conclusive refutation of the endogenous sex hypothesis for our sample: the first group, those who were never married, are not included in our sample (we limit to those whose first birth occurred after their first marriage); the second group includes women who remain married after having a girl (i.e. those who do not respond to our instrument, and thus for whom we cannot measure the causal effect of divorce). But we have also found that the reduced-form effect of the instrument on household income and on others' income is small and statistically insignificant shortly after the child is born. Rather, the effect of the instrument increases with time.[14] This suggests that child sex is not the result of resources, but rather that later resources are the result of child sex, through its effect on marital breakup—note that the effect of sex on breakup similarly builds over time (see Figures 3 and 4).

[14] We also regressed the sex of the eldest child on household income and controls for mothers whose children were aged 0-1 and found no significant effect of household income on child sex.

If child sex is not the result of resources, we can interpret as causal the reduced-form relationship between having an eldest girl and economic outcomes. But child sex may affect outcomes through paths other than divorce, so that our IV estimate of the effect of divorce on outcomes is misstated. There are various possible reasons why either or both parents might work more or less in response to the sex of the eldest child.[15] We again cannot refute these possibilities directly. However, when we looked at the sample of women whose marriages remain intact, the income distribution and labor supply of mothers did not substantially differ by the sex of the eldest child. Similarly, if child sex affects economic outcomes through channels other than marital breakup, we would expect differences in outcomes by firstborn sex for the group that is predicted to have a low divorce rate. But we do not find any significant effects for this subsample.

We might also overestimate the cost of divorce if fathers pay differential child support by the sex of the eldest child. It is possible that fathers of boys transfer more money to the mother of their child, even if they are no longer married to her. In this case, again, having a girl would be negatively correlated with women's outcomes, but not solely through marital breakup. To check this possibility, we examined the Census "other income" variable (which includes child support and alimony) for women who are currently divorced and for women who are ever divorced. We found that other income did not differ by the sex of the eldest child (results not shown).[16]

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[15] For example, parents may work harder in the labor market when they have a boy. Lundberg and Rose (2002) find that men's labor supply and wage rates increase more in response to the births of sons than to the births of daughters. In addition, having a girl may, even in the absence of divorce, cause increased marital conflict, which could in turn negatively impact economic outcomes. In either case, having a girl would be negatively correlated with women's outcomes, but not solely through marital breakup, and using it as an instrument would cause an overstatement of the cost of divorce. Alternatively, women may increase their labor supply as a "defensive investment" in anticipation that their marriage may be less stable (Weiss 1997) if their first child is a girl.

[16] Dahl and Moretti (2003) find that divorced mothers of multiple children who are all girls are marginally significantly less likely to receive child support than those whose children are all boys. They, however, use a much less restricted sample, and use the gender of all children. Even within their sample and methodology, when they examine all family sizes they do not find a significant effect.

A final caveat remains for using the sex of the eldest child as an instrument for marital breakup: having a son as one's eldest child may affect the likelihood of re-marriage after the first marriage was broken. Note, again, that this would not invalidate the causal interpretation of the reduced-form estimates, but could cause us to overstate the direct effects of divorce. Again, we cannot rule out the existence of such a causal effect. However, when we regress a dummy for being currently married on having an eldest girl for ever-divorced women in our sample, the coefficient is indeed negative, but is not statistically significant (t-statistic=0.67).

As an additional check on the validity of the relationship between child sex and marital breakup, we replicate our Census estimates using CPS data. We do so to address potential concern about the reliability of our Census sample—that despite our best effort to restrict our sample to women for whom error in observed eldest child sex is uncorrelated with marital status, some error remains and is contaminating our estimate. We address this concern by using CPS fertility supplements. Because they record a woman's fertility history regardless of whether her children are still in the household, we do not have to worry about bias in the sex of the children still in the household.

Four June CPS supplements (1980, 1985, 1990, and 1995) record the sex of a woman's first-born child. In this sample, we need only restrict to women who are white, had their first birth after their first marriage, and had a single first birth—we need not make any of the Census restrictions that were geared towards assuring that we accurately observe the sex of the first child.[17] When we pool these supplements (N=107,519), we again find that having an eldest girl significantly increases the probability of marital breakup. In fact, we find that

[17] Compared to the census data, the CPS has drawbacks as well as advantages. On the one hand, the CPS allows us to look at older women, whose eldest children left their house after their full impact on marital stability had been realized, thus increasing the potential causal effect. On the other hand, older women are less likely to have experienced divorce (due to cohort effects) and widowhood may be higher, attenuating the estimated effect of child sex on marital breakup. Finally, in order to get a large enough sample, we needed to look at more recent data, since the CPS recorded women's fertility history only in the four samples we used.

having an eldest girl causes a 0.68 percent increase in the probability that the first marriage breaks up (t-statistic=2.31). We consider the similarity of the point estimates from both samples (the Census estimate is 0.63 percent) an encouraging sign that our results reflect a real and significant effect of the sex of the first child on the probability of marital breakup.

3.4.4 The Effect of Marital Breakup on Mean Outcomes

Cross-sectional (OLS) regressions of income and labor supply on marital breakup (Table 3), which admit no causal interpretation, show that breakup of the first marriage is correlated with large losses in income and large increases in labor supply. Both of these relationships confirm the conventional view that women whose first marriages end are significantly worse off than women whose first marriages remain intact.

Two-stage least squares estimates, on the other hand, suggest that the mean effect of marital breakup on material well-being is quite different from the cross-sectional results. The two-stage estimate of the effect of divorce on others' income is negative but insignificant, though still within two standard deviations of the OLS estimates, as is the two-stage estimate for log household income.

The two-stage estimate of the effect of divorce on household income level, however, is significantly more positive than the OLS estimate—while the coefficient is not statistically different from zero, it is more than two standard deviations from the negative OLS estimate. The same is true for the high predicted-divorce subsample. For the low predicted-divorce subsample our first stage is barely significant, so the two-stage least squares estimate cannot rule out zero effect or an effect equal to the cross-sectional effect. (The absence of significant estimates for the low predicted divorce subsample in all of our IV specifications is consistent with the fact that the first stage is only marginally significant for this group, supporting our hypothesis that child sex affects outcomes only through divorce).

Taken together, these results imply that on average there is negative selection into divorce—women who would have had low income anyway are more likely to divorce, creating a negative cross-sectional correlation between income and divorce—and that there is no significant causal effect of divorce on mean income. The lack of significant effect on others' income is probably due to the fact that most women compensate for the loss of their husbands' income by re-marriage or by moving in with other adults. Moreover, we find a very large increase in mean hours worked (about 1,000 hours a year), which is more than double the cross-sectional effect and clearly helps offset the loss of husband income in the measure of household income.[18]

As the labor supply results indicate, the stability of mean income should not be taken to imply that there is no mean welfare loss from divorce. The increase in work without a significant increase in income could imply a welfare loss for women, since they experience a mean decrease in leisure without an expansion of consumption possibilities. In addition, it may imply adverse consequences for children; for example, if parents are not perfectly altruistic they may divorce even if doing so disrupts joint household production, including childrearing. Our results indicate that on average children receive no gain in consumption but likely experience a decrease in time with their mother (and probably with their father as well).

3.5 The Effect of Marital Breakup on the Income Distribution

3.5.1 Estimation Framework

The difference in sign between the IV estimates of the effect of divorce on mean log income and on mean level of income leads us to investigate the possibility that there are important effects of marital breakup on the income distribution, particularly at the bottom,

[18] Our results thus far are broadly consistent with those of Bedard and Deschenes (2003), despite our use of different sample restrictions and specifications, lending further evidence of the robustness of the relationships we analyze.

that aren't evident at the mean. In fact, it makes intuitive sense that the effect of divorce on the income distribution would be to fatten the lower tail, rather than to shift the entire distribution uniformly downward. After all, divorce—and the implied withdrawal of the husband's income—represents a discrete fall in income. To the extent that women remarry, move in with other relatives, or have high earning potential, many may end up as well off financially as before (or even better off)—resulting in little effect of divorce on the mean of the distribution. And yet a subset of women who cannot recover from the loss would experience a much greater than average effect of divorce, falling near the bottom of the distribution.

To address this possibility, we consider more flexible specifications in which we estimate the effect of marital breakup on the probability that a woman's income is below various thresholds. This enables us to identify the marginal effect of marital breakup on the cumulative distribution function (CDF). In addition, we obtain estimators of the CDF itself for those whose first marriage is broken and for those whose first marriage is intact. Doing so allows us to evaluate the magnitude of the effect of divorce on the income distribution.

Recall that the local average treatment effect formula in equation (3) gave the causal effect of treatment on the compliers. Similarly, Abadie (2002, 2003) demonstrates that in absence of covariates and with the same assumptions as the standard two-stage least squares model:[19]

$$(4) \qquad E[Y_0 \mid D_1 > D_0] = \frac{E[(1-D)Y \mid Z=1] - E[(1-D)Y \mid Z=0]}{E[(1-D) \mid Z=1] - E[(1-D) \mid Z=0]}$$

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[19] Proof of equation (4): $Z = i \Rightarrow D = D_Z \Rightarrow Y = Y_Z$ for $i = 0,1$, so the numerator is:

$$E[(1-D)Y \mid Z=1] - E[(1-D)Y \mid Z=0] = E[(1-D_1)Y_1 - (1-D_0)Y_0]$$

By the monotonicity assumption this equals:

$$E[(1-D_1)Y_1 - (1-D_0)Y_0 \mid D_1 > D_0]P[D_1 > D_0] = E[Y_0 \mid D_1 > D_0]P[D_1 > D_0]$$

Dividing by the denominator yields: $E[Y_0 \mid D_1 > D_0]$.

Since X is discrete, this proof can be extended to the case where we add controls.

We can similarly estimate the following equation using two-stage least squares with controls, using the sex of the eldest child as an instrument:

(5) $(1-D)Y = \gamma(1-D) + X\delta + \mu$

This strategy gives a consistent estimator:

(6) $\hat{\gamma}_{IV} \to \gamma = E[Y_0 \mid X; D_1 > D_0]$

That is, γ gives the expected value of Y for compliers who have a boy (and whose marriage therefore remains intact). When we apply this method to the case where Y is the CDF of the income distribution we can effectively trace out the CDF for compliers who have boys. Similarly, we could estimate the CDF for compliers who have girls.[20] But a standard two-stage least squares estimate of the effect of marital breakup on an indicator for a given income level gives the difference between the two CDFs. Thus, by summing the standard coefficient and the estimate of equation (6), we can likewise trace out the CDF for compliers who have girls. Finally, we also estimate equation (6) using OLS, and compare it to the instrumental variables estimate.

3.5.2 The Effect of Marital Breakup on the Distribution of Outcomes

In each of Tables 4, 5, and 6, the columns of results labeled "First marriage intact" report the CDF for those who remain married, estimated using the method described above. (Recall that in the two-stage least squares regressions, the CDFs are estimated for the specific group of women—"compliers"—who divorce or stay married in response to the sex of the first-born child.) The columns labeled "Difference in CDFs: broken – intact" report the estimated difference in CDFs between the compliers who have divorced and those who stay married, which is the standard coefficient on breakup in the OLS or two-stage least squares estimation.

[20] By estimating the equation: $DY = \theta D + X\rho + \eta$ we can get $\hat{\theta}_{IV} \to \theta = E[Y_1 \mid X; D_1 > D_0]$. The proof is similar.

OLS estimates of the difference in distribution of income by marital breakup (first two columns in Tables 4 and 5) tell a familiar story. In cross-section, those with broken first marriages have income distributions —both household and others'— that are first-order stochastically dominated by those of women with intact first marriages. That is, divorce is correlated with a uniform shift downward in income at every point in the income distribution.[21]

The two-stage least squares estimates in Tables 4 and 5 give a more nuanced picture.[22] We do in fact find a large effect of marital breakup on the probability of being at the bottom of the income distribution: women whose first marriage is broken are 42 percentage points more likely (about twelve times as likely) to have less than $5000 in others' income, and 23 percentage points more likely (about 80 percent more likely) to have less than $10,000. While legal transfers (which include child support and mean-tested transfers) reduce the number of compliers with no income, over a quarter of the divorced compliers have no unearned income even after accounting for transfers—compared to virtually none of the compliers who stay married (results not shown).

The analysis of household income similarly shows a large effect of marital breakup on the density at the bottom of the income distribution. Roughly one in six of those who experience marital breakup have less than $5000 in household income, compared to virtually none of those who remain married.[23] Results follow a similar pattern for the high predicted-

[21] Note also that the estimated income distribution for compliers whose first marriage remains intact is first-order stochastically dominated by the income distribution for the full sample. This result, which holds throughout Tables 4-6, is consistent with the view that the people who respond to the instrument are typically of relatively low SES.

[22] Note that the estimation of linear probability models results in some estimates that are slightly outside the [0,1] interval, though always well within two standard errors of this interval. Despite this drawback, we prefer to follow this methodology because of its transparency.

[23] Although they cannot earn enough to make up for the loss in others' income, divorced compliers do have a very large labor supply response to the loss. Since the distributional effect on hours does not differ markedly from the mean responsiveness, a separate analysis is not included here.

divorce subsample, as expected. For the low predicted-divorce subsample, the effects are not significant and do not demonstrate any consistent pattern, also as expected.

Interestingly, the two-stage estimates also show that those who divorce due to the instrument (in both the full sample and the high predicted-divorce subsample) are somewhat more likely to have income near the top of the distribution, although the differences are not statistically significant. The reversal in the sign of the difference occurs at (in the case of others' income) or below (in the case of household income) the mean of the distribution, explaining why models for the mean find little effect of marital breakup on income. Previous literature (e.g. Mueller and Pope 1980, Jacobs and Furstenberg 1986) finds that when divorced women remarry, their second husband is typically more educated and has a higher occupational SES score (although at least some of this is due to life-cycle effects). In addition, Bedard and Deschenes (2003) argue that many divorced mothers co-reside with their parents, who likely have higher joint incomes than their husbands. The reversal in sign is more substantial for total household income than for others' income, probably because that also includes top-end variation in women's earnings as well as sources of other income (such as alimony and child support).

The results discussed thus far ignore one important aspect of divorce—namely, it reduces family size. Thus even though income decreases at the bottom with divorce, women may not necessarily end up worse off; on the other hand, if income gains at the top come primarily through remarriage, the effect may be neutralized by increased family size. To estimate the effect of divorce on the ratio of income to needs, we divide each woman's total household income by the poverty line for a household of that size.

As shown in Table 6, we find that changes in family size do not mitigate our results. The OLS estimates still indicate that marital breakup decreases normalized household income at all levels. The two-stage least squares results indicate that virtually none of those still in their first marriage have household income below the poverty line, while nearly a quarter of those whose first marriage ended are below poverty. Compliers whose first marriage ended are, however, significantly more likely to be above 400 percent of poverty than are compliers whose first marriage remains intact. The results for the high-predicted divorce subsample are similar – divorce significantly increases the probability of poverty and the effect reverses at the top of the distribution, although the effect is not significant for the subsample. As always, the low-poverty subsample shows no significant effect and no discernable pattern, consistent with our hypothesis.

3.6 Discussion

Our results suggest that negative selection into divorce accounts for the observed relationship between marital breakup and lower mean income. Yet marital breakup does have a significant causal effect on the distribution of income: divorce increases the percent of women at the bottom—and perhaps at the top—tail of the income distribution. In net, divorce causally increases poverty, and perhaps inequality more generally, for women with children.

What would have been the poverty rate of women with children in 1995 had the fraction of ever-divorced mothers not changed since 1980? In 1995 (the most recent year for which we have data on marital history, from the CPS), the poverty rate for women with our sample characteristics who are still in their first marriage was 8.7 percent, while the rate for women who have ever divorced was 21.7 percent. Thus if divorce had stayed at its 1980 prevalence (17.2 percent ever divorced), the overall poverty rate in this sample would be:

Counterfactual poverty rate = (% ever divorced 1980)*(poverty rate | ever divorced 1995)

+ (% never divorced 1980)*(poverty rate | never divorced 1995)

= 0.172*0.217 + (1-0.172)*0.087 = 0.109, or 10.9 percent. Because the prevalence of divorce rose to 28.6 percent, the poverty rate became:

Actual poverty rate = (% ever divorced 1995)*(poverty rate | ever divorced 1995)

+ (% never divorced 1995)*(poverty rate | never divorced 1995)

= 0.286*0.217 + (1-0.286)*0.087 = 0.124,

or 12.4 percent. Thus the increase in divorce may potentially have caused an increase of 1.5 points, or 13.8 percent, in the poverty rate for women with our sample characteristics. If we assume that the effect holds outside of those with our sample characteristics, we can conclude that nearly 1.4 million more women and children were in poverty in 1995 than would have been if the divorce rate had remained at its 1980 level.

We conclude by noting that our findings do not have clear-cut policy implications. While some might interpret them as evidence for the importance of traditional family structures, others may conclude that they emphasize the need for more generous welfare policies.

3.7 Data Appendix

The 5 percent 1980 Census data contain several measures that allow us to analyze a woman's fertility history. These include the number of children ever born to a woman, the number of marriages, the quarter as well as year of first birth, and the quarter and year of first marriage. This information permits us to identify the sex of the first-born child for most women, although not for women whose eldest child has left the household.

A substantial drawback of using cross-sectional data is the fact that we can only observe the sex of the eldest child residing with a woman, whereas our true characteristic of interest is the sex of the firstborn child. It is imperative that we create a sample of women for whom measurement error in the sex of the observed first-born child has a classical structure.

To that end, we attempt to restrict the sample to those women observed with all their biological children. We do so in order to limit the risk that our results will be affected by differential attrition of boys and girls. In particular, we are concerned that boys may be differentially likely to end up in the custody of their fathers in the event of marital breakup. In fact, as Table A, panel 1, shows, girls are significantly less likely than boys to be living with their father and not their mother. This pattern could lead to endogeneity of our instrument if the sample were left uncorrected. If, in the event of divorce, fathers keep the sons and mothers keep the daughters, there will be a spurious positive correlation in the overall sample between marital breakup and the eldest observed child being a girl.

To address this issue, we exclude any woman for whom the number of children ever born does not equal the number of children living with her. If a mother lives with stepchildren or adopted children in a number that exactly offsets the number of her own children that are not living with her, this rule will fail to exclude her. We therefore further minimize the possibility of including women who have non-biological children "standing in" for biological children by including only women whose age at first birth is measured as between 19 and 44.

Limiting our sample to women who are living with all of their children reduces but does not eliminate the threat that differential custody rates could bias our result—because we could still be more likely to include divorced women with two girls than those with one boy and one girl or those with two boys. We find, however, that mothers living with all their children and mothers in the overall population are equally likely to be observed with a girl as the eldest child, which suggests that sex of the eldest child is not a major determinant of living with all of one's children (See Table A, panel 2).

Second, we limit our sample to women whose first child was born after their first marriage, since breakup of the first marriage is our focus. If instead we included out-of-wedlock births, we would be concerned that the sex of the first child affected selection into the first marriage. To the extent that people could learn the sex of the child before it was born and thereby select into "shotgun marriages," we may still have selection into first marriage. But ultrasound technology was not yet widely used in 1980 (Campbell 2000), so this threat is not of particular concern. In addition, because we can only identify the beginning of the first marriage and not the end, we may include some women whose first child was born after the breakup of the first marriage. But this should create only classical measurement error in our first stage estimation.

Third, we look only at mothers whose eldest child is a minor, since those who still live with their adult children may be a select group. Further, since girls are differentially likely to leave home early (at ages 17 and 18), we restrict to mothers whose eldest child is under age 17. Fourth, we limit our sample to white women because black women's childbearing and marital decisions may be quite different, and fully modeling the differences would greatly complicate the analysis. Finally, we leave out women whose first child was a twin, both because different-sex twins would complicate our instrument and because twins increase the number of children a woman has.

Selection into our restricted sample might be cause for concern, just as selection into the labor market is cause for concern when measuring labor outcomes. To test for such a problem, we generated the predicted probability that a woman was included in our sample based on age and birthplace dummies. Then we re-ran our two-stage estimates, treating a woman's predicted probability of inclusion as an endogenous regressor. Our first- and second-stage estimates were quite stable with and without this variable.

To create our high- and low-predicted divorce subsamples, we create an exogenous risk index (Table A, panel 3) by predicting "ever divorced" using age, age squared, age at first birth, and dummies for place of birth and for high school dropout (recall that our sample includes only women whose first birth occurs after age 19, so the decision on whether to graduate from high school can be treated as pre-determined).

In summary, the limitations we place on the sample are designed to create a group of women for whom we can measure the sex of the firstborn child with only classical measurement error. These restrictions weaken the power of our first stage, but we believe this compromise is necessary in order to minimize concerns about endogeneity of our instrument.

References

Acemoglu, Daron. "Technical Change, Inequality, and the Labor Market." *Journal of Economic Literature* 40 (March 2002): 7-72.

Abadie, Alberto. "Semiparametric Instrumental Variable Estimation of Treatment Response Models." *Journal of Econometrics* 113 (April 2003): 231-63.

Abadie, Alberto. "Bootstrap Tests for Distributional Treatment Effects in Instrumental Variable Models." *Journal of the American Statistical Association* 97 (March 2002): 284-292.

Aldous, Joan; Mulligan, Gail M.; and Bjarnason, Thoroddur. "Fathering over Time: What Makes the Difference?" *Journal of Marriage and the Family* 60 (November 1998): 809-820.

Angrist, Joshua D. "Estimation of Limited Dependent Variable Models With Dummy Endogenous Regressors: Simple Strategies for Empirical Practice." *Journal of Business and Economic Statistics* 19 (January 2001): 2-16.

Angrist, Joshua D. and Evans, William S. "Children and their Parents' Labor Supply: Evidence from Exogenous Variation in Family Size." *American Economic Review* 88 (June 1998): 450-477.

Angrist, Joshua D. and Imbens, Guido. "Identification and Estimation of Local Average Treatment Effects." *Econometrica* 62 (March 1994): 467-476.

Becker, Gary S. "Human Capital, Effort, and the Sexual Division of Labor." *Journal of Labor Economics* 3 (January 1985): S33-58.

Becker, Gary S. *A Treatise on the Family.* Cambridge, MA and London: Harvard University Press, 1981

Becker, Gary S.; Landes, Elizabeth M.; and Michael, Robert T. "An Economic Analysis of Marital Instability." *Journal of Political Economy* 85 (December 1977): 1141-87

Bedard, Kelly, and Deschenes, Olivier. "Sex Preferences, Marital Dissolution, and the Economic Status of Women.", mimeo, University of California at Santa Barbara, 2003. http://www.econ.ucsb.edu/~olivier/papers.html

Campbell, Stuart. "History of Ultrasound in Obstetrics and Gynecology." presented at 2000 International Federation of Obstetrics and Gynecology, Washington, D.C. http://www.obgyn.net/avtranscripts/FIGO_historycampbell.htm

Dahl, Gordon, and Moretti, Enrico. "The Demand for Sons: Evidence from Divorce, Fertility, and Shotgun Marriage.", Working Paper No. 10281, Cambridge, MA: National Bureau of Economic Research, 2004.

DiNardo, John; Fortin, Nicole M.; and Lemieux, Thomas. "Labor Market Institutions and the Distribution of Wages, 1973-1992: A Semiparametric Approach." *Econometrica* 64 (September 1996): 1001-44.

Duncan, Greg J., and Hoffman, Saul D. "A Reconsideration of the Economic Consequences of Marital Dissolution." *Demography* 22 (November 1985): 485-497.

Gruber, Jonathan. "Is Making Divorce Easier Bad for Children? The Long Run Implications of Unilateral Divorce." Working Paper no. 7968. Cambridge, MA: National Bureau of Economic Research, 2000.

Harris, Kathleen M.; Furstenberg, Frank F.; and Marmer, Jeremy M. "Paternal Involvement with Adolescents in Intact Families: The Influence of Fathers over the Life Course." *Demography* 35 (May 1998): 201-216

Jacobs, Jerry A., and Furstenberg, Frank F. "Changing Places: Conjugal Careers and Women's Marital Mobility." *Social Forces* 64 (March 1986): 714-732.

Lundberg, Shelly, and Rose, Elaina "The Effects of Sons and Daughters on Men's Labor Supply and Wages." *Review of Economics and Statistics* v84, n2 (May 2002): 251-68

Mueller, Charles W., and Pope, Hallowell. "Divorce and Female Remarriage Mobility: Data on Marriage Matches after Divorce for White Women." *Social Forces* 3 (March 1980): 726-738.

Ruggles, Steven; Sobek, Matthew; et al. *Integrated Public Use Microdata Series: Version 3.0.* Minneapolis: Historical Census Projects, University of Minnesota, 2003 .
http://www.ipums.org

Trivers, Robert L., and Willard, Dan E. "Natural Selection of Parental Ability to Vary the Sex of Offspring." *Science* 179 (January 1973): 90-92.

US Bureau of the Census.
http://www.census.gov/population/socdemo/hh-fam/tabMS-1.xls

Weiss, Yoram "The Formation and Dissolution of Families: Why Marry? Who Marries Whom? and What Happens upon Divorce." Amsterdam; New York and Oxford: Elsevier Science, North-Holland. *Handbook of population and family economics*. Volume 1A (1997): 81-123.

Table 1. Descriptive Statistics—Mothers with Minor Children Living at Home

		Samples		
	All women with children		Predicted-divorce index	
Variable		Full sample	High	Low
Demographics				
Age	35.1	31.6	36.0	25.2
Years of schooling	12.0	12.8	11.5	13.0
Household income	22,747	23,114	23,598	19,549
Total own income	4,905	4,458	4,625	3,805
Weeks worked last year	23.9	22.5	23.1	21.4
Usual hours worked	21.1	19.9	19.8	21.2
Marital status				
Currently married, spouse present	0.802	0.891	0.869	0.913
Currently separated	0.040	0.024	0.028	0.025
Currently divorced	0.085	0.072	0.084	0.052
Ever divorced	0.215	0.172	0.202	0.118
Never married	0.043	0.000	0.000	0.000
Number of observations	1,610,516	619,499	154,041	155,355

NOTE: The full sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. The high predicted-divorce subsample includes only women in the top quartile of the predicted divorce distribution; the low predicted-divorce subsample includes only those in the bottom quartile. Income is in 1980 dollars.

Table 2. The Effect of Eldest Child Sex on Marital Breakup

	All women with children	Samples					
			Age of oldest child			Predicted-divorce index	
		Full sample	0-5	6-11	12-16	High	Low
All education levels	0.0113	0.0063	0.0038	0.0057	0.0106	0.0113	0.0036
	(0.0007)	(0.0010)	(0.0015)	(0.0018)	(0.0021)	(0.0022)	(0.0018)
Observations	1,605,339	619,499	236,522	221,937	161,040	154,041	155,355
High school dropouts		0.0149	0.0106	0.0090	0.0264		
		(0.0033)	(0.0054)	(0.0057)	(0.0059)		
Observations		73,426	25,111	24,849	23,466		
High school graduates		0.0057	0.0041	0.0053	0.0084		
		(0.0015)	(0.0024)	(0.0027)	(0.0030)		
Observations		277,155	100,601	99,816	76,738		
At least some college		0.0045	0.0023	0.0052	0.0075		
		(0.0015)	(0.0020)	(0.0026)	(0.0034)		
Observations		268,918	110,810	97,272	60,836		

NOTE: The sample includes white women who are living with all of their children, whose oldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. The table reports estimates of the coefficient on (eldest child is a girl) in equation (1) in the text. Other covariates are: Age, Age squared, Age at first birth and High school dropouts. Robust standard errors in parentheses.

Table 3. The Effect of Marital Breakup on Mean Economic Outcomes

			Predicted divorce index (Specification check)			
	Full Sample		High		Low	
Dependent Variable	OLS	2SLS	OLS	2SLS	OLS	2SLS (Falsification check)
Ln(Household income)	-0.501	-0.230	-0.547	-0.867	-0.439	1.514
	(0.005)	(0.437)	(0.009)	(0.542)	(0.011)	(1.693)
Others' income	-9241	-1041	-10561	-8450	-5014	8605
	(42.57)	(5178)	(78.68)	(5939)	(87.82)	(15369)
Household income	-5577	6548	-6721	-981	-2649	9147
	(43.41)	(5570)	(80.57)	(6224)	(89.76)	(15817)
Hours worked last year	420	1053	347	937	369	951
	(2.85)	(360)	(5.39)	(408)	(6.31)	(1180)
Number of observations	619,499		154,041		155,355	

NOTE: The full sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. High predicted-divorce subsample includes only women in the top quartile of the predicted divorce distribution; the low predicted-divorce subsample includes only those in the bottom quartile. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. There were 4,766 observations, or about 0.77% of our observations, for which household income was zero or negative. For those observations we set Ln(Household income) to zero. Income is in 1980 dollars. Robust standard errors are in parentheses.

Table 4. Cumulative Distribution of Income of Women's Other Household Members, by Marital Status

		OLS		2SLS					
						Predicted divorce index (Specification check)			
		Full Sample		Full Sample		High		Low (Falsification check)	
Dependent Variable		CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact
Others' income	≤$5000	0.0507	0.3928	0.0378	0.4177	0.0800	0.5249	0.1587	0.2041
		(0.0003)	(0.0014)	(0.0785)	(0.1202)	(0.0929)	(0.1457)	(0.3061)	(0.4067)
	≤$10000	0.1458	0.3850	0.2770	0.2284	0.2932	0.2727	0.8545	-0.4356
		(0.0005)	(0.0015)	(0.1270)	(0.1567)	(0.1413)	(0.1785)	(0.6081)	(0.6870)
	≤$20000	0.5459	0.2226	0.8751	-0.0416	0.5735	0.1968	1.0290	-0.0898
		(0.0008)	(0.0014)	(0.1814)	(0.1937)	(0.1866)	(0.2054)	(0.5957)	(0.6209)
	≤$30000	0.8412	0.0754	1.0369	-0.0821	0.9996	-0.0195	0.8886	0.1992
		(0.0006)	(0.0010)	(0.1323)	(0.1399)	(0.1554)	(0.1629)	(0.3166)	(0.3620)
Number of observations		619,499		619,499		154,041		155,355	

NOTE: The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Income is in 1980 dollars. Robust standard errors in parentheses.

Table 5. Cumulative Distribution of Household Income, by Marital Status

	OLS		2SLS					
					Predicted divorce index (Specification check)			
	Full Sample		Full Sample		High		Low (Falsification check)	
Dependent Variable	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact
Household income ≤$5000	0.0294	0.0982	-0.0155	0.1900	0.0082	0.3086	0.2163	-0.0367
	(0.0005)	(0.0007)	(0.0603)	(0.0865)	(0.0739)	(0.1137)	(0.2437)	(0.3082)
≤$10000	0.0921	0.1999	0.1502	0.2875	0.1941	0.3994	0.6388	-0.2840
	(0.0008)	(0.0010)	(0.1024)	(0.1312)	(0.1193)	(0.1591)	(0.4912)	(0.5475)
≤$20000	0.3994	0.2153	0.7760	-0.2547	0.5994	-0.0094	1.7693	-0.8853
	(0.0013)	(0.0015)	(0.1806)	(0.2048)	(0.1821)	(0.2086)	(0.8697)	(0.8331)
≤$30000	0.7496	0.0788	1.1079	-0.2360	0.8813	-0.0610	0.8815	0.2333
	(0.0012)	(0.0013)	(0.1621)	(0.1729)	(0.1740)	(0.1897)	(0.4202)	(0.4758)
Number of observations	619,499		619,499		154,041		155,355	

NOTE: The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Income is in 1980 dollars. Robust standard errors in parentheses.

Table 6. Cumulative Distribution of Poverty, by Marital Status

	OLS		2SLS					
					Predicted divorce index (Specification check)			
	Full Sample		Full Sample		High		Low (Falsification check)	
Dependent Variable	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact	CDF when first marriage is intact	Difference in CDFs: broken-intact
Percentage of the poverty threshold ≤100	0.0565	0.1322	0.0041	0.2407	0.0714	0.2982	0.2015	-0.0188
	(0.0004)	(0.0011)	(0.0806)	(0.1081)	(0.1038)	(0.1407)	(0.2983)	(0.3736)
≤200	0.2224	0.2069	0.2958	0.2487	0.3695	0.2160	1.0154	-0.2754
	(0.0006)	(0.0015)	(0.1449)	(0.1684)	(0.1659)	(0.1941)	(0.6610)	(0.6492)
≤300	0.5051	0.1550	0.8077	-0.1800	0.5079	0.1726	2.2120	-1.2861
	(0.0008)	(0.0015)	(0.1824)	(0.2011)	(0.1869)	(0.2078)	(1.0217)	(0.9594)
≤400	0.7349	0.0780	1.2086	-0.4442	0.8366	-0.0519	1.1202	-0.1784
	(0.0007)	(0.0013)	(0.1751)	(0.1909)	(0.1673)	(0.1842)	(0.5366)	(0.5563)
Number of observations	619,499		619,499		154,041		155,355	

NOTE: The sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. All the regressions include the following controls: age, age squared, age at first birth and a dummy for high school dropouts. Poverty is calculated using 1980 measures. Robust standard errors in parentheses.

Table A. Specification Checks and Sample Construction

Panel 1	Dependent variable: Child is a girl
Child lives with father and not with mother	-0.0501
	(0.0022)
Constant	0.4888
	(0.0003)
Observations	2,810,256
Panel 2	Dependent variable: Eldest child is a girl
Mother lives with all her children	0.0000
	(0.0013)
Constant	0.4871
	(0.0011)
Observations	790,746
Panel 3	Dependent variable: Ever divorced
Age	0.0264
	(0.0003)
Age squared ÷ 100	-0.0305
	(0.0005)
Age at first birth	-0.0025
	(0.0001)
Dummy for high school dropout	0.0947
	(0.0012)
Observations	1,092,433

NOTES: The sample in panel 1 includes white children under age 19 (the unit of observation is a child)
The sample in panel 2 includes women meeting the set of restrictions imposed throughout the paper, except that we do not restrict to mothers living with all their children. That is, it includes white mothers who had their first birth after marriage after age 18 and before age 45, had a single first birth, and whose eldest child is under 17.
The sample in panel 3 includes white women who had their first birth after marriage, after age 18, and before age 45, and had a single first birth. Controls also include birthplace dummies (147 categories)
Robust standard errors in brackets

Figure 1. Mothers' Poverty Rates, by Marital Status

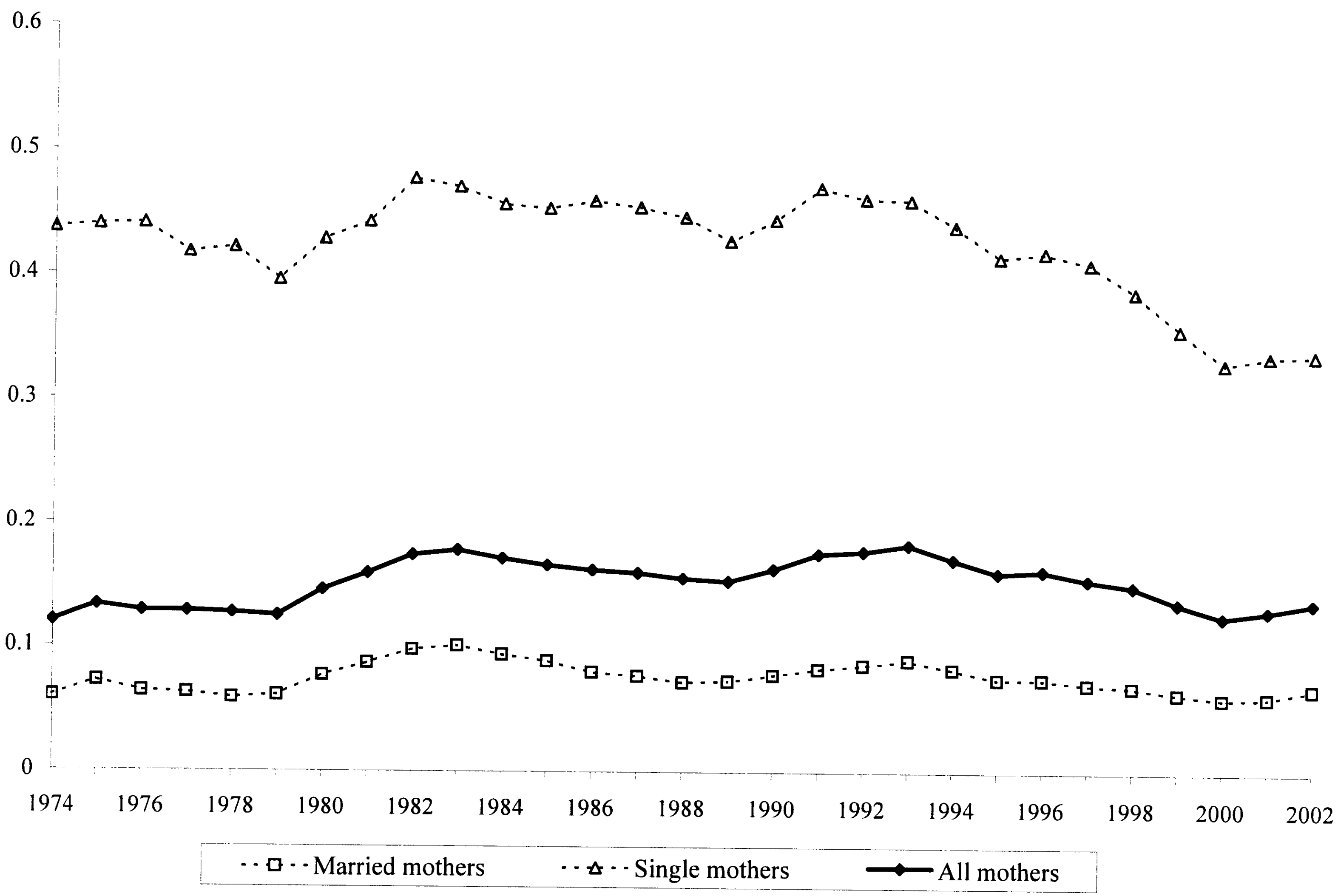


Figure 2. Percent of Mothers Who are Single

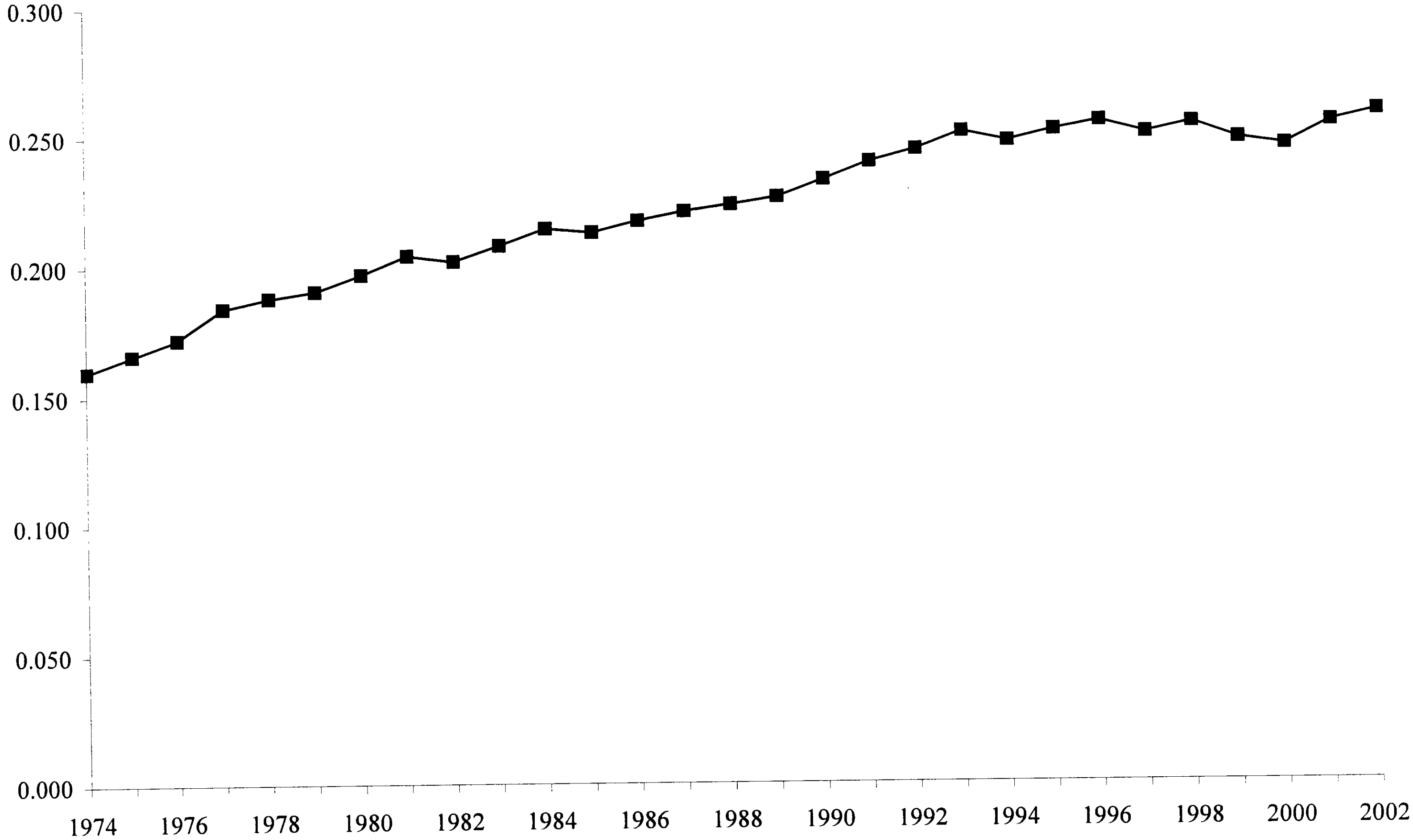
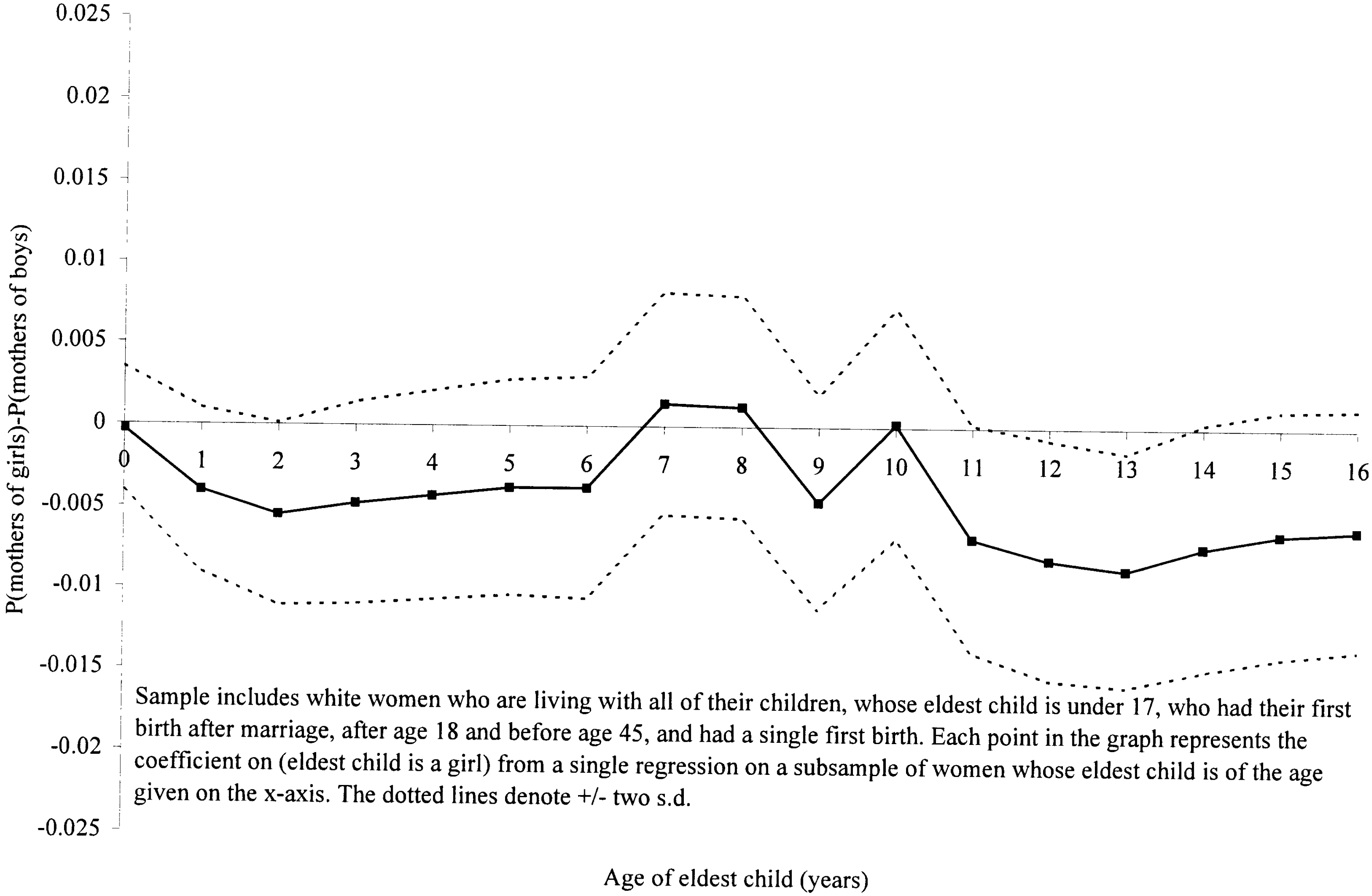
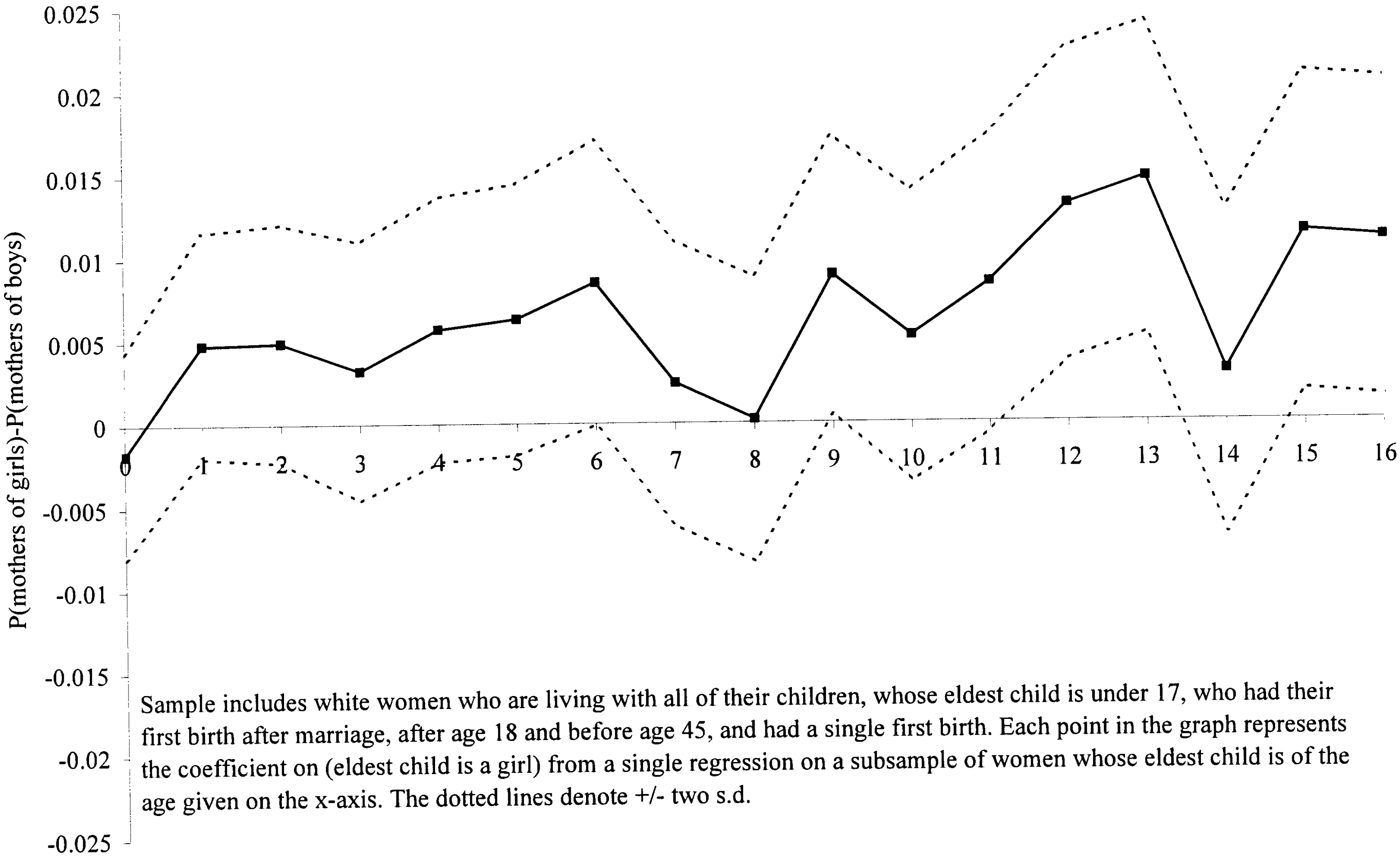


Figure 3. The Effect of Eldest Child Sex on the Probability of Being Currently Married



Sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. Each point in the graph represents the coefficient on (eldest child is a girl) from a single regression on a subsample of women whose eldest child is of the age given on the x-axis. The dotted lines denote +/- two s.d.

Figure 4. The Effect of Eldest Child Sex on the Probability that First Marriage is Broken



Sample includes white women who are living with all of their children, whose eldest child is under 17, who had their first birth after marriage, after age 18 and before age 45, and had a single first birth. Each point in the graph represents the coefficient on (eldest child is a girl) from a single regression on a subsample of women whose eldest child is of the age given on the x-axis. The dotted lines denote +/- two s.d.

Age of eldest child (years)