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# What do computers really do? Computerization, fading pay-setting institutions and rising wage inequality

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ABSTRACT

In this paper we advance the argument that the widespread assumption that computerization and institutional changes are independent explanations for the resurgence of wage inequality is inaccurate. Instead we posit for complex dynamics between computerization and fading pay-setting institutions, arguing that the latter is a mechanism by which the former operates. To test our argument that computerization increases wage inequality not only via the mechanisms specified by skill-biased Technological Change, but also indirectly through structural processes, we utilize longitudinal U.S. industrial-level data on computerization, pay-setting institutions, and wage inequality. Estimating Error Correction Models, we find a stronger longitudinal association between computerization and wage inequality in industries where labor processes were subject to both computerization and the breakup of pay-setting institutions (such as labor unions) than in industries where these institutions never had much of a presence. These findings provide some evidence that computerization operates also through the mechanism of weakening labor market institutions.



## 1. Introduction

The resurgence of inequality since the late 1970s in rich countries, which was most prompt, substantial, and prominent in the US, is one of today's most widely discussed and controversial issues. Most economists argue that market forces have been primarily responsible for the rise in wage inequality. Pointing to technological changes, they maintain that computer technology is complementary to human capital, meaning that at the same level of human capital, productivity is much higher when computer technology is used. That being the case, the diffusion of computers has led to an increase in the relative demand for high-skilled workers that tend to use computers, thereby raising their wages relative to less-skilled workers that do not use computers (Acemoglu & Autor, 2011). At the same time that demand for skilled workers rose, there was a slowdown in the growth in numbers of college graduates, thereby raising the wages of highly educated workers even more (Goldin & Katz, 2008). The demand for less-skilled workers, on the other hand, has stagnated or even declined as computers enhanced processes of automation (Autor, Levy, & Murnane, 2003). This explanation, known as Skill-Biased Technological Change (SBTC), implies that computerization is the main explanation for rising wage inequality, at least among the bottom "99 percent" of wage and salary workers (Autor, 2014).

On the other side of the argument are sociologists and political scientists who stress the role of political forces – especially the weakening of labor unions – as driving rising wage inequality in the United States (Card, Lemieux, & Riddell, 2004; Lin & Tomaskovic-Devey, 2013; Kristal & Cohen, 2015; Moller, Alderson, & Nielsen, 2009; Western & Rosenfeld, 2011), the United Kingdom (Machin, 1997), Germany (Card, Heining, & Kline, 2012), and Israel (Kristal & Cohen, 2007). Supporting the political argument, comparative studies show that the weakness of pay-setting institutions in the US, labor unions in particular, explain why income inequality is the highest in the US (Alderson & Nielsen, 2002; Beckfield, 2006; Bradley, Huber, Moller, Nielsen, & Stephens, 2003; Brady, 2003; Garnero, Kampelmann, & Rycx, 2014; Iversen & Soskice, 2006; Rueda & Pontusson, 2000).

There is a consensus among researchers that both fading pay setting institutions and computerization are the main factors responsible for rising inequality in the US and Europe (Lemieux, 2008; Piketty, 2014; Western & Rosenfeld, 2011). The disagreement is about the relative importance of the two factors. Results of empirical studies comparing the size of these effects appear to depend on the specific research design. A cross-countries study found institutions to be more important (OECD, 2011), while a time-series US study found the opposite (Wolff, 2006). A recent

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US study (Kristal & Cohen, 2015) that measured these effects at the industry level found institutions (de-unionization and the stagnation in minimum wage) to be twice as important as computerization for explaining rising inequality.

While the distinction between technology and political forces is assumed to be clear-cut, we put forward a new perspective that underlines the complex dynamics between technology and politics in the wage determination process, arguing that the latter is a mechanism by which the former operates. Specifically, we posit that computer-based technologies have changed the social relations at workplace in fundamental ways that has enhanced union decline and the weakening of other pay-setting institutions such as internal labor markets, which in turn has boosted wage inequality. We test our argument in the US, in which these processes should be more evident than in European countries with strong deliberative institutions (Hall & Soskice, 2001) where less adversarial labor unions and employers are better equipped to deal with labor-saving technological changes than their U.S. counterparts. To get as close as possible to the dynamics of the workplace with aggregate data, we utilize longitudinal industrial-level data on computerization, pay-setting institutions and wage inequality[1]. We empirically test our thesis by estimating whether there is an interaction between computerization and fading pay-setting institutions in the wage determination process. In support of our complementary thesis to the canonical SBTC thesis, we present evidence for variation in the longitudinal relations between computerization and wage inequality across industries experiencing more and less institutional change; variation that indicates that computerization operates also through the mechanism of weakening institutions.

This paper's contribution, then, is to clarify the question regarding the mechanisms through which computerization affects inequality. The core notion of this new "structural perspective" (Kalleberg, Wallace, & Althauser, 1981; Tomaskovic-Devey, 2014) we advance is that computerization increases wage inequality not only via the mechanisms specified by SBTC, but also through structural processes related to institutional factors. In her study of the decline in labor's share of national income, Kristal (2013b) demonstrates that computerization reduced the labor's share (and increased corporate profits) also indirectly by exacerbating union decline. Here we develop and expand this thesis that computerization operates through the mechanism of weakened institutions to explain the surge in wage inequality, a different dimension of economic inequality[2]. Although we do not test directly the mechanisms through which computerization enhanced fading pay-setting institutions, our findings for an interaction effect provide an essential step toward validating the feasibility of such mechanisms.

The remainder of this paper is structured as follows. In Section 1 we elaborate the new structural perspective that computerization inhibits unionization and harms other pay-setting institutions, thereby indirectly affecting inequality via institutional mechanisms rather than merely due to "market forces". In Section 2 we describe the longitudinal industry data, measures, and method of analysis. In Section 3 we estimate the associations between computerization and wage inequality across industrial sectors and detailed industries that faced more and less institutional change. In the Conclusions, we summarize and discuss the implications of the findings for the canonical accounts for rising wage inequality.

## 2. Computerization and institutional change

While there is a consensus that computerization explains part of rising inequality, there is less agreement on the precise mechanisms through which it affects wage inequality (Card & DiNardo, 2002; Handel, 2007). Recently, a few studies have questioned the assumption that the invisible hand of the market is the main mechanism through which computerization increases inequality (DiMaggio & Bonikowski, 2008; Hanley, 2014; Kristal, 2013b; Guy & Skott, 2008). Indeed, the mechanisms through which fading pay-setting institutions have driven wage inequality are self-explicating and supported by empirical evidence. By contrast, the mechanisms of supply, demand, and returns to productivity, through which according to the SBTC computers have led to the growth in wage inequality, are relatively vague and difficult to measure, and it is therefore nearly impossible to accumulate direct empirical evidence for such mechanisms.

While SBTC surely has a role in explaining rising inequality, it is rather restrictive to assume that computers have impacted the labor market and wage inequality solely via skills and productivity, as it had profound impact on various structural aspects of the production process and on the relations among workers. We have some evidence that it has affected union decline (Kristal, 2013b). Plausibly enough, computerization may have accentuated many additional post-1970 developments in rich countries's labor markets, especially in the US, from the rise of nonstandard employment relations, to outsourcing abroad and the substitution of foreign merchandise for domestic products. While computer technology is not the main cause of these labor market transformations, we posit that computer-based technologies enhanced these processes, thereby increasing wage inequality via various mechanisms.

What are the possible mechanisms through which computerization contributes to fading pay-setting institutions? Begin with union decline. Two main mechanisms are discussed by past research. First is the well-known effect of downsizing manufacturing jobs: computers enhanced automation of the production process and prompted firms to utilize computer equipment in tasks previously performed manually by blue-collar, mostly unionized workers, downsizing many unionized manufacturing jobs and leading to union decline (Fligstein & Shin, 2007; Kristal, 2013b; Milkman, 1995). A second plausible mechanism links union decline to skill polarization of the workforce. Previous studies suggest that new computer technologies had highly polarizing effects on the workforce: skilled workers experienced up-skilling, while many production workers underwent de-skilling (Burris, 1998; Vallas & Beck, 1996). This skill polarization has not only increased wages at the top via SBTC, but also fueled wage gaps via structural factors: it deepens divisions among workers and may have undermined workers' solidarity, thereby reducing the likelihood of working-class cohesion and collective action (Kristal, 2013b). One manifestation of this polarization among organized workers has been the ongoing process of decentralization in collective bargaining agreements that have contributed to rising wage inequality since the early 1980s (Western & Rosenfeld, 2011).

[1] To test the effect of computerization on inequality we need data on computer technology at the establishment, industry, or country level for a long period of time. While there is a huge variance between industries in the use of computer technology, there is less variance between OECD countries in their use of computer technology. Evidently, the ideal design would be a cross national time-series study within industries. However, to the best of our knowledge, no country other than the US provides data on computerization, unionization, and inequality for a large enough number of industries nor for a long enough period of time.

[2] While both wage inequality and the share of capitalists' profits in national income (compared to labor compensation) have increased since the late 1960s, these dimensions of inequality differ substantially in the sources of income and classes of people to which they apply (Kristal, 2013a; Piketty, 2014). Wages at the top of the wage distribution, for example, have fueled wage inequality among workers but grew at a much slower pace than corporate profits, allowing capitalists to grab the lion's share of the fruits of (relatively slow) economic growth. Empirically, too, these two dimensions of inequality are not highly correlated.

A third possible mechanism is that the strengthening of management control due to the computer revolution (Burris, 1993; Crowley, Tope, Chamberlain, & Hodson, 2010; Hanley, 2014; Kristal, 2013b; Vallas, 1993; Zuboff, 1988) may have empowered employers and management, allowing them to use more antiunion tactics, especially those related to the surveillance of union leaders and activists as well as dissemination of antiunion propaganda., While there are no studies that explore the relations between computerization and the prevalence of employers' antiunion actions, there is some evidence that Information communication Technology (ICT) has facilitated better monitoring of workers (Guy & Skott, 2008). According to the 2007 Electronic Monitoring and Surveillance Survey of the American Management Association and the ePolicy institute[3], at least three-quarters of companies monitor their employees' email, Internet use, phone calls, computer files, and use surveillance cameras to monitor work. It may be the case, then, that employers are taking advantage of such monitoring technologies to fight unions. Consider, for example Wal-Mart, a firm that uses computer technology to create work environment that is hostile to unionization. The retail giant provides managers computer-based learning modules on what to do and how to handle union organizing, and shows workers videos and PowerPoint presentations on the (terrible) consequences of union formation during organizing drives. In addition, when necessary, managers can reposition surveillance cameras to monitor union supporters[4]. Surely, there are also plausible positive effects of computerization on union organizing, as the same technologies became available to trade unionists as well, facilitate the digital connections among workers and making the information on unions much more accessible, for example via mobile applications. Yet, in the more adversarial relations between employers and unions in the US, there is evidence that in some cases access to computerized email lists of employees was denied from unions[5], and the National Labor Relations Board (NLRB) road to union representation still relies on secret-ballot process instead of electronic voting methods.

Computerization has contributed to the transformation of the world of work in a way that not only exacerbated union decline, but may have also enhanced a decline in the proportion of workers in large firms (where internal labor markets and other standard employment relations are prevalent), greater flexibility in practices of employment, as well as a reliance on imports. One notable mechanism for the link between computerization and mounting nonstandard employment relations is the declining use of efficiency wages. The underlying assumption of the efficiency wage model is that due to managers' limited ability to monitor workers' efforts and productivity, workers are being paid more than their market wage to maximize their efforts. Workers, aware that they are being paid above market wages (and enjoy relative job security), are therefore more reluctant to shirk and lose such high wages (Ackerlof & Yellen, 1986). As Guy and Skott (2008) argue, the new ICT technologies that have diffused since the 1970s have allowed firms to monitor workers more closely, thereby reducing employers' need and willingness to pay efficiency wages.

We may speculate about other mechanisms that link computerization to the acceleration of structural changes in the labor market. ICT eliminated the need for some layers of middle management and provided an infrastructure that enabled outsourcing (Baumol, Blinder, & Wolff, 2003). As a result, the size of firms declined, especially in manufacturing (Baumol et al., 2003; Brynjolfsson, Malone, Gurbaxani, & Kambil, 1994), and the use of part-time workers, many of whom would prefer to work fulltime, increased (Tilly, 1996). Rapid changes in technology have also driven companies to prefer employing contingent workers in an effort to stay competitive. This could be achieved by increasingly using staffing arrangements such as working at home, which ICT made possible, and by contracting companies, independent contractors, or temporary workers to supply business services that cannot be produced economically in-house or specialized services that organizations need for only a short period of time, such as computer programming, computer-system analysis, or software design. Likewise, improvements in computer hardware, software, and networks made production information less personalized and localized by storing most of it in databases and making them accessible to newly recruited workers, temporary workers, as well as workers stationed at home or overseas. This, in turn, may have boosted the externalization of employment relationships, physically and socially isolating workers, as well as enabling offshoring and the importation of goods from developing countries.

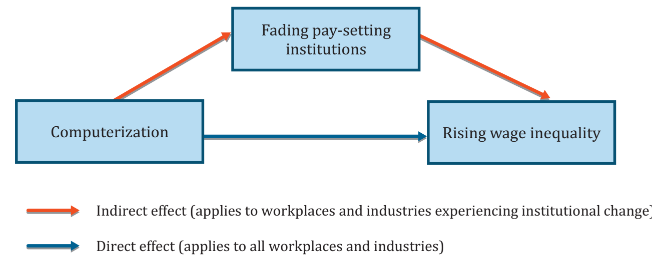


**Fig. 1.** Study's hypotheses on direct and indirect effects of computerization on wage inequality.

These mechanisms are but a few plausible examples of the complex processes whereby computerization may have affected wage inequality indirectly, via its effect on pay-setting institutions. While some of these mechanisms are admittedly speculative, for our purpose the more important task is to empirically test if indeed there is a robust indirect effect of computerization on rising wage inequality that is channeled, at least partly, through the decline of unions and other institutional factors. If there is indeed such an indirect effect, then we should find variation in the longitudinal effect of computerization on wage inequality across workplaces and industries: a stronger effect in workplaces and industries experiencing more institutional change (where both direct and indirect effect operated) and a weaker effect in workplaces and industries experiencing less institutional change (where only a direct effect operated). The rational of our argument is illustrated in Fig. 1.

## 3. Data, variables, and method

### 3.1. Data

We test the association between computerization, fading pay-setting institutions, and rising wage inequality using longitudinal data on U.S. industries. We use a pooled cross-sectional time-series design (i.e., yearly observations for each industry) to test the study's arguments. The combined industry-year datasets include 43 comparable (two-digit) industries that cover the entire nonagricultural private sector. Due to the major change in the industry classification structure in 1997 from Standard Industrial Classification (SIC) system to the North American Industry Classification System (NAICS), we have one dataset on 43 industries for the years 1969 to 1997 and another dataset on 43 industries for the years 1988 to 2012.

[3] Available at http://www.epolicyinstitute.com/survey.asp.

[4] Based on Human Rights Watch report (May 2007).

[5] See Register Guard, 351 NLRB 1110 (2007). This policy has recently been changed, in favor of unions. The two Republican members (out of 5) on the Board strongly dissented from the majority opinion that overruled the Board's 2007 decision. See Purple Communications, Inc., 361 NLRB No. 126 (Dec. 11, 2014).

The analyses are based on data drawn from several governmental and census publications on U.S. industries. In short, we combine data on annual and hourly wage inequality from the Current Population Survey (CPS) samples with data on the magnitude of each industry's investments in computers from the Bureau of Economic Analysis (BEA) Industry Economic Accounts data, with data on unionization, education, and part-time employment from the Bureau of Labor Statistics (BLS) and CPS samples, with data on employment in large firms from the Census Bureau County Business Patterns, and with data on import penetration available only in manufacturing from 1973 to 1997 from Schott (2010). Annual data on the federal minimum hourly wage is from the U.S. Department of Labor and is applied to all industries.

To measure the changes in the U.S. wage structure over the last four decades, we draw on data from the CPS's large and representative household data source. The most widely used source of data for studies of wage inequality is the March CPS that is available from the mid-1960s onward. The March CPS has comparable data on total annual earnings and weeks of work in the previous year. Since data on hourly wage is available only from 1976, many researchers have focused on "fulltime, full-year" (FTFY) workers. We use the March files from 1969 to 2013 (covering earnings from 1968 to 2012) to compile a sample of annual earnings for wage and salary workers aged 18 to 65 who participate in the labor force on an FTFY basis, defined as working 35-plus hours per week and fifty-plus weeks per year. Starting in 1976 (earnings year 1975), the March survey began collecting information on hours worked in the previous year. This allows us to create a second sample of hourly wage data for all wage and salary workers for the earnings years 1975 to 2012.

For both the annual wage sample and the hourly wage sample, we follow standard practice and replace top-coded wages with 1.5 times the top-coded value (Card & DiNardo, 2002), which means that the study excludes the top 1 percentile of wage and salary workers. The samples of annual wage exclude all observations whose estimated annual earnings are below $2000 in 1979 dollars, and the samples of hourly wage exclude all observations whose estimated hourly wage is less than one dollar or greater than $100 per hour in 1979 dollars. In constructing statistics for FTFY workers, we use the CPS sample weights. In constructing statistics for all workers, we use the CPS sample weights, multiplied by the number of hours worked in the previous year (divided by 2000). Weighting by hours worked allows the inclusion of part-time workers.

Employment in large firms is computed using the County Business Patterns published by the Census Bureau. County Business Patterns provides annual statistics for businesses with paid employees within the U.S. at a detailed industry level and employment size class since 1964. Information is available on the number of establishments, employment during the week of March 12, first quarter payroll, and annual payroll. County Business Patterns cover all private nonagricultural industries excluding rail transportation. We imputed the average numbers for the transformation industry.

### 3.2. Variables

We follow previous studies and measure overall *wage inequality* by the standard deviation of log wages and the log of the ratio of the ninetieth percentile of wages to the 10th percentile (i.e., the 90/10 log wage differential). The two measures are usually very close, with differences mainly reflecting top-coding and the treatment of very low-wage observations. Following studies that make use of BEA data on the magnitude and composition of each industry's capital investments to explain rising wage inequality (Autor, Katz, & Krueger, 1998; Fligstein & Shin, 2007; Kristal, 2013b; Wolff, 2006), we employ a simple measure of *computer technology* by measuring real investments in computers, computer-peripheral equipment, and software as a share of total nonresidential fixed assets investments. While the BEA data do not directly measure the kind of technology implemented in the production process and it is affected by both the quantity and prices of computer technology, we assume that when firms invest in computing equipment they are most likely to use this new equipment at different stages of the production process. A broader measure for computer technology that was used by Autor et al. (1998) and Wolff (2006) also includes investments in office and accounting equipment (i.e., electric and nonelectric office machines). The correlation between this measure and the one we use is 0.982 and the estimations results are appreciably the same. Another common measure for computerization relies on net stocks instead of investments in computer technology. Since net stocks are subject to tax manipulations more than investments, we prefer to use the latter. At any rate, research that used both investments and stocks as measures for computerization (Autor et al., 1998) reached similar results, probably because computer stocks are highly correlated with computer investments (0.902). The correlations between these two measures of computerization and inequality (and also with union density) are similar. Additionally, since we utilize fixed-effect models that estimate changes over-time within industries, any differences between stocks and investments are considerably reduced. Following Autor (2014), we use college share of hours work as a proxy for *college supply* in each year and added its annual growth rate to the models.

We employ several measures for pay-setting institutions, for our argument is that the effect of computerization on inequality varies according to industries' institutional setting. *Union density* is measured by dividing the number of union members in each industry by the number of wage and salary workers (see Kristal, 2013b). Union membership figures have been compiled for all employed civilian wage and salary workers, aged 16 and over. Not included are employed 14–15 year old, self-employed workers, and a small number of unpaid family workers. We have two measures for organizational change that serves as proxies for nonstandard employment relations: part time employment, and employment in large firms, where the likelihood for the persistence of internal labor market is greater than in smaller firms. *Part-time* employment is measured by dividing the number of part-time workers employed in each industry by the number of employed workers. We measure *large firms* by the percentage of the workforce that is employed in firms employing more than 500 workers, which is computed using the County Business Patterns provided by the Census Bureau. We use import data by industry and country from Schott (2010) to construct a measure of the imports in manufacturing industries originating in low-wage countries' industrial product. *Import penetration* is thereby measured by imports from low-wage countries as a share of industry's value added[6].

### 3.3. Method

To estimate whether the effect of computerization on wage inequality vary by institutional change, we analyze OLS models in time-series cross-sectional dynamic specification (a lagged dependent variable is included among the predictors), where the dependent variable is wage inequality and the independent variables include indicators for computer technology and pay-setting

[6] We classify a country as low-wage in year t if its per capita GDP is less than 20 percent of U.S. per capita GDP (data on countries' per capita GDP are from Penn World Table). This cutoff captures an average of 80 countries per year. The list of countries that are classified as low wage includes China and India as well as relatively small exporters such as Angola. We choose a 20 percent cutoff to classify countries as low wage since it represents the world's most labor-abundant cohort of countries and therefore the set of countries most likely to have an effect on U.S. manufacturing plants.

institutions. Results from OLS models in which we assume that the independent variables are exogenous may not represent a causal relationship and may result in inconsistent estimates of the effect of computerization on wage inequality. We utilized two workable strategies to minimize the challenge of endogeneity given the industrial data. Yet, to be on the safe side, we describe the findings in terms of the relations between computerization and wage inequality as opposed to the casual effect of computerization on wage inequality.

First, we apply fixed-effects models. Fixed effects estimators, which exploit within-industry variation as a means of purging unit heterogeneity, make it possible to obtain unbiased and consistent estimates of parameters when industry effects are arbitrarily correlated with measured explanatory variables. By applying fixed effects estimators, the models focus on the within-industry variation over time, and the coefficients represent a cross-industry average of the longitudinal effect. Second, to estimate the long- and short-run relations between computerization, pay-setting institutions, and wage inequality, we analyze single-equation error correction models (ECMs) that can accommodate stationary and non-stationary variables, given that the errors are stationary (Beck & Katz, 2011; De Boef & Keele, 2008)[7], and fit cointegrated data[8]. One potential difficulty in analyzing cross-section time-series data over a relatively short period of time is that the empirical data is likely to trend over time, i.e., to be non-stationary. Consequently, OLS regression can produce "spurious relations" as a result of the variables trending together over time. While first-differencing is a convenient technical solution to the non-stationary problem, it throws out any long-run information about the variables and restricts the type of relationship that can be uncovered to those in which the effect of an explanatory variable is constrained to a single point in time (Beck & Katz, 2011).

We therefore specify the cross-section time-series variant of the single-equation error correction model for the dynamic relationships:

$$\Delta wage\_inequality_{i,t} = \alpha_0 + \beta_1 \Delta X_{i,t} - \beta_2 \left( wage\_inequality_{i,t} - \beta_3 X_{i,t-1} \right) + \varepsilon_{i,t}$$

In this model, current changes in wage inequality (measured in first difference, i.e., $Y_t - Y_{t-1}$) are a function of both the short-term changes (i.e., first differences) in the independent variables and their long-term levels. Specifically, $\beta_1$ captures any short-term effects on wage inequality, while the long-term effects are captured by $\beta_3$. The long-term effect occurs at a rate dictated by the value of $\beta_2$ that captures the rate of return to equilibrium. In all models the estimates are weighted by industry size to make sure that the results are not biased by small industries representing only a small fraction of the workforce.

To compare the relative size of the effects, we calculate the long-run multipliers for the variables of interest. The long-run multiplier is the total long- and short-run effect on wage inequality for a one-point increase in the independent variable, and it is measured by dividing the coefficient of the lagged independent variable by the coefficient of the lagged wage inequality. To measure how much a given independent variable may have affected wage inequality within the history of a single industry, we calculated a coefficient of maximum longitudinal impact (Alderson & Nielsen, 2002). This is the long-run multiplier coefficient multiplied by the average within-industry range of the independent variable. The coefficient reflects how much change in inequality could have taken place over time in a single industry given the typical range of variation in the independent variable within an industry.

## 4. Findings: Why computerization increases wage inequality

The first step in estimating the likelihood that computerization has caused rising inequality partly through waning pay-setting institutions is to contend with the question of timing – did computerization partly precede the decline in labor unions and standard employment relations, and growing import penetration? To answer this question, Fig. 2 presents the timing of the growth in computer investments and decline of pay-setting institutions by decade and industrial sector. The figure supports the common knowledge about the timing of the major changes in the labor market. Computer-related technological change has been going on since at least the 1970s, continued throughout the 1980s (an increase of 6 percentage points), accelerated in the 1990s (an increase of 9 percentage points), and declined in the 2000s. In all sectors the sharpest decline in union density occurred in the 1980s and continued well into the 1990s and 2000s. Employment by large firms declined in the 1970s and 1980s, as well as the 2000s (but increased in the 1990s), and import penetration in manufacturing industries accelerated primarily in the 1990s. Computerization, then, was not only a precursor of growing import penetration, but also partly preceded the decline in labor unions and standard employment relations. These trends imply that in terms of timing, the diffusion of computer technologies across industries and workplaces, which began as early as the late 1960s (Bresnahan, 1999), may have contributed to union decline, a decrease in the share of workers in large firms, and growing import penetration in the following years.

To be sure, two types of interpretations for the interaction effect between computerization and pay-setting institutions on inequality are plausible. The first is that firms and industries first de-unionized, and were only then penetrated by computers, as weak unions and workers employed in precarious work were no longer able to resist labor-saving technological change. On the face of it, Fig. 2 provides some support for this claim. It illustrates that computer investments considerably increased in the nonunion industries of trade and services in the 1980s, earlier than in the more unionized industries that more commonly employ internal labor market practices (i.e., mining, construction, manufacturing, and transportation), where the significant growth in computer investments was delayed until the 1990s. If indeed the impact of fading pay-setting institutions on inequality preceded the effect of computerization, then we should find a weaker longitudinal effect of computerization on inequality in industries where labor processes were subject to the breakup of unions, employment externalization, and computerization than in industries where organized labor never had much of a presence. In other words, in the latter industries the effect of computerization on rising wage inequality should be greater because it has continued for a longer period. As illustrated in Fig. 3, the industrial sectors of mining, construction, manufacturing and transportation are part of the first category – industrial sectors experiencing more institutional decline – and the industrial sectors of trade, FIRE (Finance, Insurance, and Real Estate) and services are part of the second category – industrial sectors experiencing less institutional change.

Alternatively, the second interpretation of the interaction – that computerization, together with other processes, has contributed

[7] Based on the results from stationary tests for the error terms in all models.

[8] To test whether the data series are cointegrated, we performed the standard two-step cointegration test by regressing $Y$ on $X$ (in levels) and then testing whether the residual is stationary. We find that we can reject the null hypothesis of no cointegration at the 5 percent confidence level or better for all independent variables, concluding that there is equilibrium between wage inequality and the independent variables.

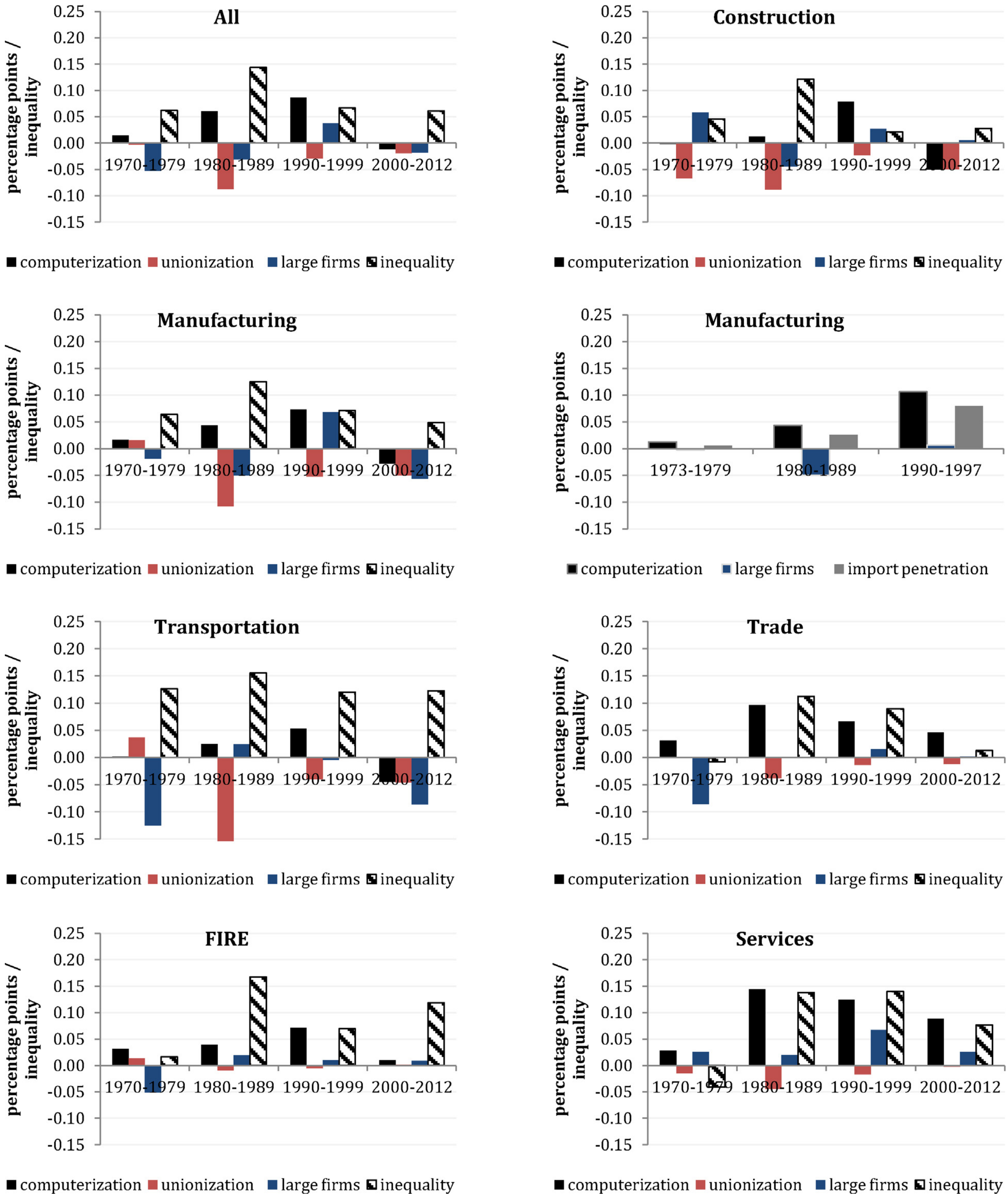


**Fig. 2.** Changes in computerization, unionization, large firms' employment, and wage inequality by industrial sector and decade. Wage inequality is measured by the 90/10 (ln)annual wage differential. Computerization, unionization and large firms' employment are measured in percentages, and therefore the change is indicated in percentages points. *Note*: In mining industries computerization occurred mainly in the 1990s (an increase of four percentage points), union density declined over the entire period but mainly during the 1980s (a decline of 15 percentage points) and the 1990s (a decline of 7 percentage points), and large firms' employment declined in the 1980s (by 9 percentage points).

to the decline in pay-setting institutions – is consistent with the opposite result, namely, a stronger longitudinal effect of computers on inequality in industries where labor processes were subject to both computerization and the breakup of unions than in sectors where organized labor never had much of a presence. Another words, in the latter industries there was only a direct effect of computerization on rising wage inequality while in former we should expect that computerization exert both a direct effect and an indirect effect on inequality which is channeled through the decline of unions and other institutional factors.

To test which of the two interpretations for the interaction effect between computerization and pay-setting institutions on

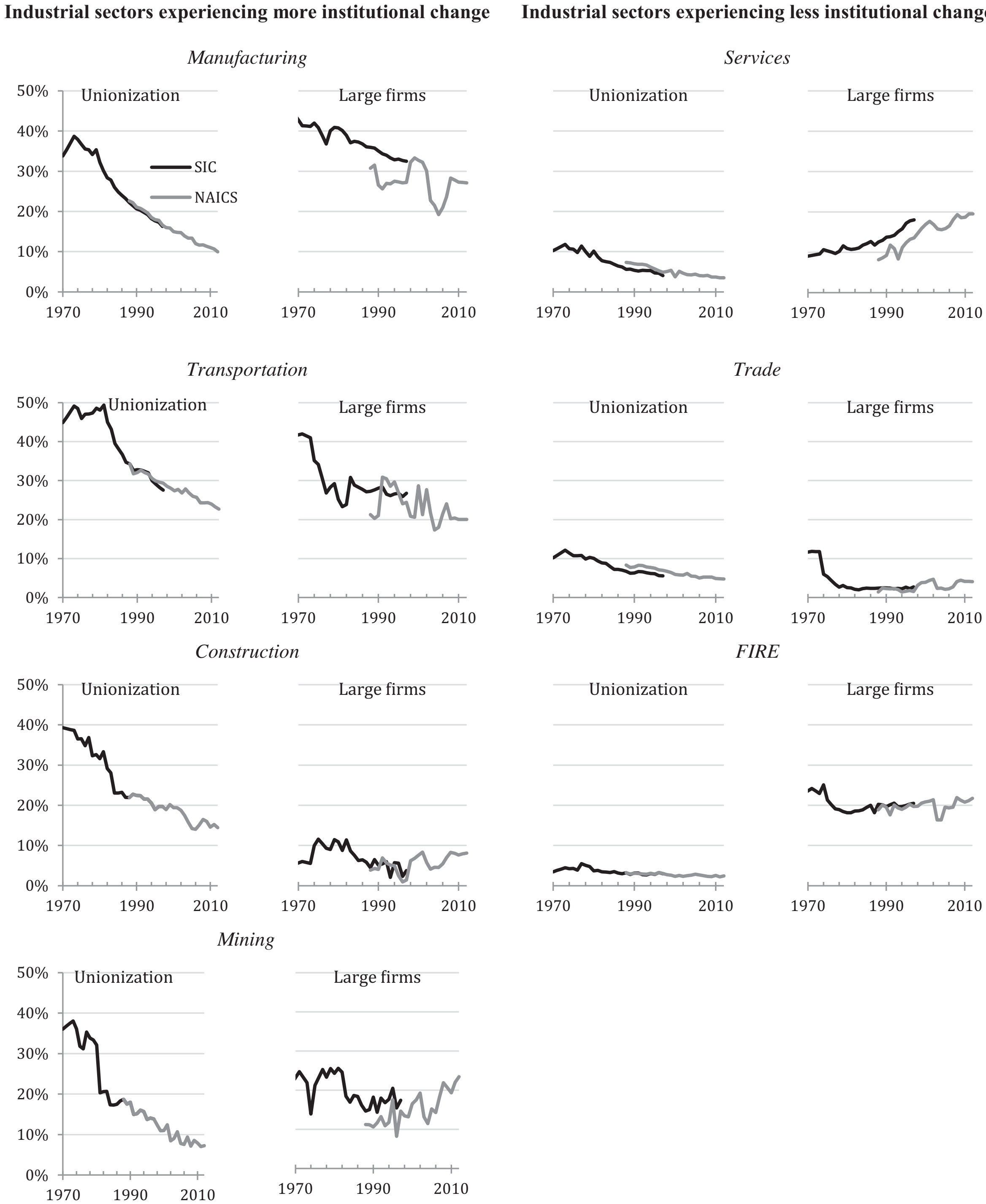


**Fig. 3.** Unionization and large firms' employment by industrial sector, 1970–2012. *Note*: It is not possible to provide comparable time-series trends by industrial sectors for the entire period. We therefore present one line for 1970–1997 (SIC in the black line) and another line for 1987–2012 (NAICS in the grey line).

inequality is supported by the data, we first estimated separate models for the relations between computerization and change in wage inequality in highly unionized core industries and scantly unionized industries as well as by industrial sector. For the sake of clarity, Fig. 4 graphically presents the results for the long-run multiplier of computers on rising inequality by sector.

In all industrial sectors (except construction) we find a positive association between computerization and rising inequality. Yet in less unionized industries, such as trade, FIRE and services, the association is significantly smaller in the years 1969 to 1997 than in the more unionized industries such as mining, manufacturing and transportation. These findings are consistent with the interpretation that part of the positive effect of computerization on wage inequality over the 1980s and early 1990s was channeled through the erosion of pay-setting institutions. They are inconsistent, however, with the first interpretation, namely, that highly unionized workplaces first de-unionized, and only then introduced computers. In other words, that the relations between computerization and rising wage inequality are much larger in mining, manufacturing and transportation than in other industries, suggests that the

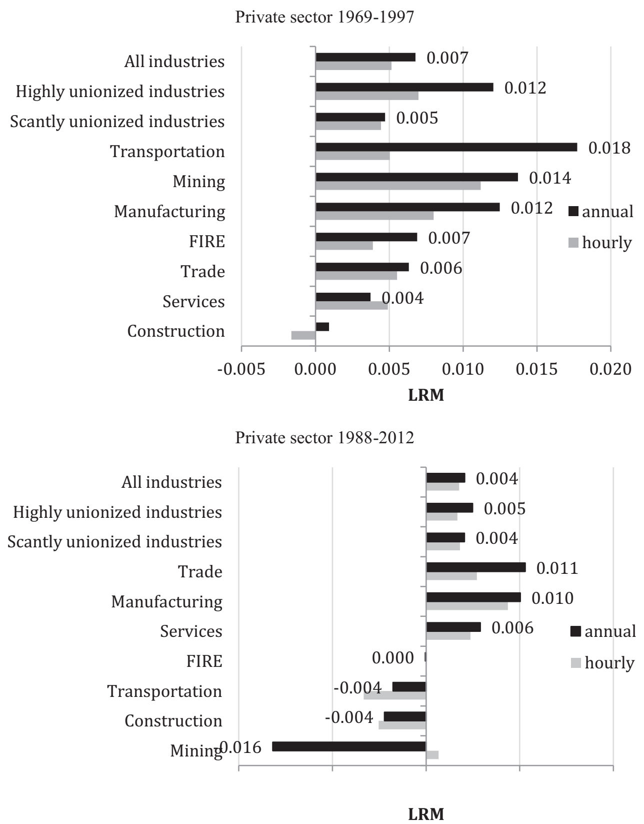


**Fig. 4.** Long-run multipliers for the effect of computers on rising wage inequality by industrial sector. The long-run multiplier (LRM) is the total long- and short-run effect on wage inequality for a one-point increase in the independent variable, and it is measured by dividing the coefficient of the lagged independent variable by the coefficient of the lagged wage inequality. *Source*: Appendices A–C.

second interpretation for the interaction effect between computerization and pay-setting institutions on inequality is supported by the data.

To further test these two possibilities, we estimate the relations between computerization and wage inequality within detailed 43 private industries with an interaction dummy for unionization, employment in large firms, and part-time employment. In the dummy variables for unionization and employment in large firms "1" denotes a decline of more than five percentage points between the first year and the last year. In the dummy variable for part-time employment "1" denotes an increase of more than 1.5 percentage points between the first year and the last year. All models control for indicators to other pay-setting institutions. We also estimate models for only 18 manufacturing industries, in which the value one in the dummy variable for union decline denotes a decline of more than 18 percentage points between the first year and the last year. The full models are shown in Appendices D, E and F and the comparable results for the interaction terms displayed by the long-run multipliers are illustrated in Fig. 5.

First and foremost, we find that computer investments had positive relations with changes in wage inequality between 1969 and 2012 mainly in industries where unionization declined. For example, in mining, construction, manufacturing and 6 industries within transportation, where unionization significantly declined, the rise in computer investments from 1969 to 1997 led to a rise of about 0.005 (the long-run multiplier) points in annual wage inequality. By calculating the maximum longitudinal impact (i.e., the long-run multiplier multiplied by the average within-industry range of the independent variable) of computerization in industries where unionization declined and how much of overall rising inequality between 1969 and 1997 this maximum longitudinal impact accounts for, we obtained that in industries where unionization declined computerization accounts for 21 percent of overall rising inequality. In industries where unionization stayed constant (and low), computerization led to a rise of only 0.002 points. Applying the same calculations as above, we obtained that in these later industries computerization explains 10 percent of overall rising inequality.

Based on the results for the interactions of computers with the share of employment in large firms and in part-time employment, we also find that the association between computerization and wage inequality was much higher in industries where standard employment relations dwindled[9]. For example, in industries where employment by large firms declined (parts of manufacturing, communication, and banking), the rise in computer investments from 1969 to 1997 led to an increase of about 0.004 points in annual wage inequality (23 percent of overall rising inequality), while in industries where the share of workers in large firms stayed relatively constant (construction, trade, FIRE, and services), computerization led to an increase of only 0.001 points (7 percent of overall rising inequality). The differences in the effect of computerization on inequality between industries where part-time employment increased and industries where it remain constant are relatively minor. In industries where part-time employment increased, the rise in computer investments from 1969 to 1997 led to an increase of about 0.003 points in annual wage inequality (14 percent of overall rising inequality), while in industries where the share of workers in part-time employment stayed relatively constant, computerization led to an increase of 0.002 points (12 percent of overall rising inequality).

The results for manufacturing industries between 1973 and 1997 are even stronger and more pronounced. Computer investments had a positive relation with wage inequality (almost) only in manufacturing industries (such as motor vehicles, machinery, apparel) where import penetration significantly increased, unionization declined, and employment in large firms declined too. By contrast, in industries such as chemicals and paper products, where import penetration was relatively minor, computerization had a *negative* effect on rising hourly wage inequality; likewise, in industries where employment in large firms has never been high (such as wood products and furniture), computerization also had a *negative* effect on annual and hourly wage inequality. Calculating how much of overall rising inequality computerization accounts for, we find that in manufacturing industries, where import penetration significantly increased, computerization accounts for 25 percent of overall rising inequality, while in industries where import penetration was relatively minor, computerization explains 11 percent of overall rising inequality. Similar differences are between manufacturing industries where employment in large firms declined compared to industries where employment in large firms has never been high. In the former, computerization accounts for 28 percent of overall rising inequality, compared to none in the latter.

The finding that a large portion of the overall relations between computerization and rising wage inequality was in specific industries in which pay-setting institutions had eroded holds mostly for the 1970s, 1980s and early 1990s and less so for the late 1990s and early 2000s. The bottom graph in Fig. 5 shows the results of models with interactions between computerization and pay-setting institutions for the years 1988 to 2012. Here the results are less robust and do not clearly show in which industries the relations between

[9] Although there are industries that experienced both union decline and a rise in nonstandard employment relations, the correlations between the dummy variables for unionization, employment by large firms and part-time employment are only 0.3 and 0.2, respectively.

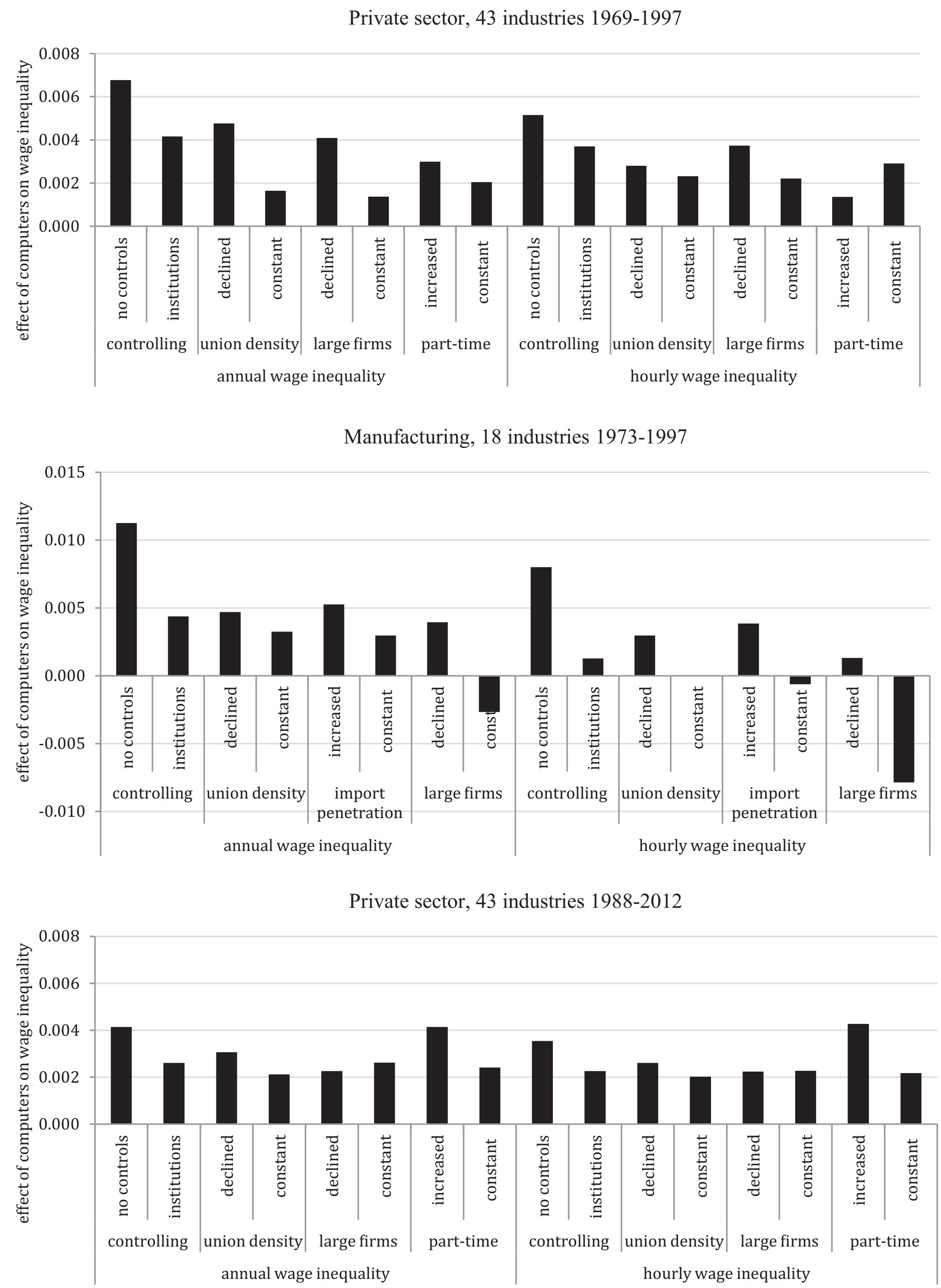


**Fig. 5.** Long-run multipliers for the direct and indirect effects of computers, channeled through pay-setting Institutions, on rising wage inequality. *Source*: Appendices D–F.

computerization and inequality were stronger. That the effect of computerization on rising inequality during the late 1990s and 2000s was less likely to be channeled through the erosion of pay-setting institutions than in the 1970s, 1980s and early 1990s, may be due in part to the fact that by the 1990s unions were already weak.

## 5. Discussion and conclusion

This study reveals that the distinction between technology and institutional factors in explaining rising inequality is not as clear-cut as has been assumed. One important question in the inequality literature about the causes of the resurgence of wage inequality relates to the mechanisms behind the well-known positive correlation between computerization and earnings. Our findings suggest that the introduction of computers affects wage inequality not only via the well-known mechanism of SBTC, but also via a second mechanism – weakening pay setting institutions. In this paper we do not test directly this second mechanism but rather engage in the essential step of testing whether politics have a mediating role in the nexus between computerization and wage inequality.

Our study focuses on the US, where both the rise in inequality and the decline in unions were sharper than in Europe. What may explain these differences between the US and continental

Europe? One plausible explanation that has not been tested yet is that computerization contributed to the decline of unions in the U.S. more than in Europe, where less adversarial labor unions and employers (in particular those in Scandinavian countries) are better equipped to deal with labor-saving technological changes than their U.S. counterparts. This can explain the more modest rise in inequality in continental Europe than in the US. Institutional features – leftist parties that create union-friendly political and regulative environments and centralized collective bargaining that reduces the incentives for employers to oppose unionization – were found to attenuate the negative effects of business cycles on unions (Western, 1997). It is likely that these institutions also mitigated the negative effect of computerization on unions through a variety of practices, which are absent in the adversarial system of labor relations in the U.S, such as setting of rules and norms that make it more difficult for employers to exploit technology that facilitates the decline of pay setting institutions. These include the protection of organized blue-collar workers from layoffs in industries where workers are being replaced by computers, and more centralized wage-setting that lessens the fragmentation among organized workers due to skill polarization.

That a substantial part of the association between computerization and rising wage inequality can be linked to the dwindling powers of wage-setting institutions implies that computerization role in the wage determination process works not only through the interplay of demand and supply factors, workers' skills and productivity. By hampering union efforts to organize workers, downsizing companies, stimulating short-term and flexible employment arrangements, and encouraging outsourcing and offshoring, computers have contributed to transformations in the social relations of the workplace that most likely have played a significant role in increasing earnings inequality.

## Acknowledgments

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## Appendix A.

Effects of computers on wage inequality (1969–2012) in nonagricultural private-sector industries by broad industrial sector.

| Dep. variable | ΔAnnual wage inequality 90/10 | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| Sector | Union declined | Union constant | Core[a] | Services[b] | Union declined | Union constant | Core[a] | Services[b] |
| No. of industries | 31 | 12 | 30 | 13 | 26 | 17 | 26 | 17 |
| Years | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 |
| Model | I | II | III | IV | V | VI | VII | VIII |
| ΔComputer investments | 0.001 (0.002) | 0.000 (0.002) | 0.002 (0.001) | −0.000 (0.002) | −0.003 (0.004) | 0.003 (0.002) | **−0.009**$^{**}$ **(0.003)** | 0.004 (0.002) |
| Computer investments$_{(t-1)}$ | **0.006**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.001)** | **0.006**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.002)** | **0.003**$^{**}$ **(0.001)** |
| Dependent variable$_{(t-1)}$ | **−0.507**$^{**}$ **(0.062)** | **−0.574**$^{**}$ **(0.037)** | **−0.479**$^{**}$ **(0.061)** | **−0.621**$^{**}$ **(0.045)** | **−0.522**$^{**}$ **(0.052)** | **−0.784**$^{**}$ **(0.123)** | **−0.536**$^{**}$ **(0.049)** | **−0.774**$^{**}$ **(0.123)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.266 | 0.293 | 0.246 | 0.324 | 0.275 | 0.385 | 0.293 | 0.380 |
| Modified DW | 2.00 | 1.84 | 2.08 | 1.81 | 1.99 | 2.02 | 2.07 | 1.98 |
| $N$ | 899 | 348 | 870 | 377 | 650 | 425 | 650 | 425 |
| Dep. variable | ΔHourly wage inequality 90/10 | | | | | | | |
| Sector | Union declined | Union constant | Core[a] | Services[b] | Union declined | Union constant | Core[a] | Services[b] |
| No. of industries | 31 | 12 | 30 | 13 | 26 | 17 | 26 | 17 |
| Years | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 |
| Model | I | II | III | IV | V | VI | VII | VIII |
| ΔComputer investments | 0.001 (0.002) | −0.000 (0.003) | 0.001 (0.002) | −0.000 (0.003) | −0.001 (0.003) | 0.001 (0.002) | **−0.005 (0.002)** | 0.002 (0.002) |
| Computer investments$_{(t-1)}$ | **0.004**$^{**}$ **(0.001)** | **0.004**$^{**}$ **(0.001)** | **0.004**$^{**}$ **(0.001)** | **0.004**$^{**}$ **(0.001)** | **0.002**$^{**}$ **(0.001)** | **0.002**$^{**}$ **(0.001)** | 0.002 (0.002) | **0.003**$^{**}$ **(0.001)** |
| Dependent variable$_{(t-1)}$ | **−0.556**$^{**}$ **(0.054)** | **−0.794**$^{**}$ **(0.118)** | **−0.525**$^{**}$ **(0.048)** | **−0.833**$^{**}$ **(0.112)** | **−0.557**$^{**}$ **(0.063)** | **−0.782**$^{**}$ **(0.147)** | **−0.596**$^{**}$ **(0.062)** | **−0.757**$^{**}$ **(0.173)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.289 | 0.389 | 0.273 | 0.410 | 0.289 | 0.389 | 0.313 | 0.375 |
| Modified DW[c] | 2.02 | 1.82 | 2.09 | 1.86 | 2.04 | 1.98 | 2.07 | 1.98 |
| $N$ | 682 | 264 | 660 | 286 | 650 | 425 | 650 | 425 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ indicates the annual change in the variable.

[*] $P < 0.10$.

[**] $P < 0.05$, two-tailed test.

[a] Core—Manufacturing, transportation, construction, and mining industries.

[b] Services—FIRE, services, and trade industries.

[c] The modified version of the Bhargava, Franzini, and Narendranathan (1982) Durbin Watson test for fixed effects panel models.

## Appendix B.

Effects of computers on wage inequality (1969–1997) in nonagricultural private-sector industries by industrial sector.

| Dep. variable | ΔAnnual wage inequality 90/10 | | | | | |
|---|---|---|---|---|---|---|
| Sector | Construction | Manufacturing | Transportation | Trade | FIRE | Services |
| No. of industries | 1 | 18 | 9 | 2 | 4 | 7 |
| Years | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 |
| Model | I | II | III | IV | V | VI |
| ΔComputer | 0.024 (0.020) | 0.001 (0.001) | −0.009 (0.012) | −0.009 (0.011) | 0.001 (0.003) | 0.001 (0.002) |
| $Computer_{(t-1)}$ | 0.001 (0.004) | **0.005**$^{**}$ **(0.001)** | **0.009**$^{**}$ **(0.002)** | 0.004 (0.002) | 0.004 (0.002) | **0.003**$^{**}$ **(0.000)** |
| Dependent $variable_{(t-1)}$ | −**0.754**$^{**}$ **(0.197)** | −**0.425**$^{**}$ **(0.077)** | −**0.528**$^{**}$ **(0.102)** | −**0.563**$^{**}$ **(0.009)** | −**0.548**$^{**}$ **(0.089)** | −**0.747**$^{**}$ **(0.057)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | No | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.299 | 0.219 | 0.280 | 0.417 | 0.219 | 0.437 |
| Modified DW | 1.95 | 2.25 | 2.01 | 1.87 | 1.77 | 1.95 |
| *N* | 29 | 522 | 261 | 58 | 116 | 203 |
| Dep. variable | ΔHourly wage inequality 90/10 | | | | | |
| Sector | Construction | Manufacturing | Transportation | Trade | FIRE | Services |
| No. of industries | 1 | 18 | 9 | 2 | 4 | 7 |
| Years | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 |
| Model | VII | VIII | IX | X | XI | XII |
| ΔComputer | 0.006 (0.012) | −0.001 (0.002) | 0.006 (0.014) | −0.008 (0.002) | 0.003 (0.006) | **0.004**$^{**}$ **(0.001)** |
| $Computer_{(t-1)}$ | −0.002 (0.003) | **0.004**$^{**}$ **(0.001)** | 0.003 (0.004) | 0.003 (0.001) | 0.003 (0.002) | **0.005**$^{**}$ **(0.000)** |
| Dependent $variable_{(t-1)}$ | −**0.976**$^{**}$ **(0.251)** | −**0.465**$^{**}$ **(0.059)** | −**0.605**$^{**}$ **(0.082)** | −**0.620**$^{**}$ **(0.023)** | −**0.873**$^{**}$ **(0.227)** | −**1.033**$^{**}$ **(0.081)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | No | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.354 | 0.247 | 0.312 | 0.362 | 0.395 | 0.525 |
| Modified DW | 1.87 | 2.19 | 1.99 | 1.99 | 1.75 | 2.01 |
| *N* | 22 | 396 | 198 | 44 | 88 | 154 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ indicates the annual change in the variable.

* $P < 0.10$.

** $P < 0.05$, two-tailed test.

## Appendix C.

Effects of computers on wage inequality (1988–2012) in nonagricultural private-sector industries by industrial sector.

| Dep. variable | ΔAnnual wage inequality 90/10 | | | | | |
|---|---|---|---|---|---|---|
| Sector | Construction | Manufacturing | Transportation | Trade | FIRE | Services |
| No. of industries | 1 | 17 | 6 | 2 | 5 | 10 |
| Years | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 |
| Model | I | II | III | IV | V | VI |
| ΔComputer | −**0.018**$^{*}$ **(0.006)** | −0.003 (0.003) | −**0.018**$^{**}$ **(0.007)** | −0.001 (0.003) | 0.000 (0.004) | **0.010**$^{**}$ **(0.001)** |
| $Computer_{(t-1)}$ | −0.003 (0.002) | **0.006**$^{**}$ **(0.002)** | −0.002 (0.002) | 0.006 (0.004) | −0.000 (0.002) | **0.004**$^{**}$ **(0.001)** |
| Dependent $variable_{(t-1)}$ | −**0.655**$^{**}$ **(0.159)** | −**0.553**$^{**}$ **(0.065)** | −**0.529**$^{**}$ **(0.127)** | −**0.561**$^{**}$ **(0.160)** | −**1.005**$^{**}$ **(0.156)** | −**0.704**$^{**}$ **(0.090)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | No | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.443 | 0.301 | 0.297 | 0.455 | 0.499 | 0.358 |
| Modified DW | 2.05 | 2.05 | 2.09 | 2.06 | 1.91 | 1.84 |
| *N* | 25 | 425 | 150 | 50 | 125 | 250 |
| Dep. variable | ΔHourly wage inequality 90/10 | | | | | |
| Sector | Construction | Manufacturing | Transportation | Trade | FIRE | Services |
| No. of industries | 1 | 17 | 6 | 2 | 5 | 10 |
| Years | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 |
| Model | VII | VIII | IX | X | XI | XII |
| ΔComputer | −**0.012 (0.004)** | −0.002 (0.002) | −0.011 (0.007) | −0.000 (0.004) | −0.000 (0.003) | **0.004**$^{*}$ **(0.001)** |
| $Computer_{(t-1)}$ | −**0.004 (0.002)** | **0.005**$^{**}$ **(0.002)** | −**0.004**$^{**}$ **(0.001)** | 0.005 (0.001) | 0.000 (0.001) | **0.003**$^{**}$ **(0.001)** |
| Dependent $variable_{(t-1)}$ | −**0.762**$^{**}$ **(0.174)** | −**0.612**$^{**}$ **(0.074)** | −**0.659**$^{**}$ **(0.182)** | −**0.638 (0.045)** | −**1.004**$^{**}$ **(0.263)** | −**0.656**$^{**}$ **(0.080)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | No | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.434 | 0.328 | 0.342 | 0.427 | 0.505 | 0.333 |
| Modified DW | 2.05 | 2.13 | 1.97 | 1.90 | 1.88 | 1.95 |
| *N* | 25 | 425 | 150 | 50 | 125 | 250 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ indicates the annual change in the variable.

* $P < 0.10$.

** $P < 0.05$, two-tailed test.

## Appendix D.

Direct and channeled effects of computers on earnings inequality in nonagricultural private-sector industries (1969–1997).

| Dep. variable | ΔAnnual wage inequality 90/10 | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Sector | Private sector | | | | | Manufacturing | | | |
| No. of industries | 43 | 43 | 43 | 43 | 43 | 18 | 18 | 18 | 18 |
| Years | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1969–1997 | 1973–1997 | 1973–1997 | 1973–1997 | 1973–1997 |
| Model | I | II | III | IV | V | VI | VII | VIII | IX |
| ΔComputer investments | 0.001 (0.002) | −0.001 (0.002) | | | | 0.002 (0.002) | | | |
| Computer investments$_{(t-1)}$ | **0.003**$^{**}$ **(0.001)** | **0.002**$^{**}$ **(0.001)** | | | | **0.006**$^{**}$ **(0.001)** | | | |
| ΔCollege supply growth | | | 0.002 (0.001) | 0.001 (0.001) | 0.001 (0.001) | | −0.001 (0.002) | −0.001 (0.002) | −0.001 (0.002) |
| College supply growth$_{(t-1)}$ | | | 0.000 (0.000) | 0.000 (0.001) | 0.001 (0.001) | | −0.000 (0.002) | −0.000 (0.002) | −0.000 (0.002) |
| ΔUnion density | | **−0.004**$^{**}$ **(0.001)** | **−0.003**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** | | **−0.004**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** |
| Union density$_{(t-1)}$ | | **−0.004**$^{**}$ **(0.001)** | **−0.002**$^{**}$ **(0.001)** | **−0.003**$^{**}$ **(0.001)** | **−0.003**$^{**}$ **(0.001)** | | **−0.004**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** |
| ΔLarge firms | | −0.001 (0.001) | −0.000 (0.001) | 0.000 (0.001) | −0.001 (0.001) | | 0.000 (0.003) | 0.001 (0.003) | 0.000 (0.003) |
| Large firms$_{(t-1)}$ | | −0.001 (0.001) | −0.000 (0.001) | −0.000 (0.001) | −0.000 (0.001) | | −0.001 (0.001) | 0.000 (0.001) | −0.000 (0.001) |
| ΔPart-time | | −0.003 (0.002) | −0.003 (0.002) | −0.003 (0.002) | −0.003 (0.002) | | 0.001 (0.003) | 0.002 (0.003) | 0.002 (0.003) |
| Part-time$_{(t-1)}$ | | −0.001 (0.001) | −0.002 (0.001) | **−0.003**$^{**}$ **(0.001)** | **−0.003**$^{**}$ **(0.001)** | | 0.004 (0.005) | 0.003 (0.005) | 0.004 (0.005) |
| ΔMinimum wage | | | −0.008 (0.006) | −0.007 (0.006) | −0.006 (0.006) | | −0.009 (0.012) | −0.009 (0.012) | −0.009 (0.012) |
| Minimum wage$_{(t-1)}$ | | | **−0.020**$^{**}$ **(0.004)** | **−0.019**$^{**}$ **(0.004)** | **−0.018**$^{**}$ **(0.004)** | | **−0.019**$^{**}$ **(0.008)** | **−0.019**$^{**}$ **(0.009)** | **−0.019**$^{**}$ **(0.008)** |
| ΔImport penetration | | | | | | | 0.001 (0.006) | 0.002 (0.006) | 0.001 (0.006) |
| Import penetration$_{(t-1)}$ | | | | | | | −0.000 (0.001) | 0.001 (0.001) | −0.000 (0.001) |
| Interactions | | | | | | | | | |
| ΔComputer × UD declined | | | −0.002 (0.002) | | | | −0.003 (0.002) | | |
| ΔComputer × UD constant | | | −0.000 (0.002) | | | | 0.003 (0.002) | | |
| Computer$_{(t-1)}$ × UD declined | | | **0.003**$^{**}$ **(0.001)** | | | | **0.003**$^{**}$ **(0.001)** | | |
| Computer$_{(t-1)}$ × UD constant | | | **0.001**$^{**}$ **(0.000)** | | | | **0.002**$^{**}$ **(0.001)** | | |
| Union density declined[a] | | | **−0.544**$^{**}$ **(0.054)** | | | | −0.148 (0.079) | | |
| ΔComputer × LF declined | | | | −0.001 (0.002) | | | | −0.001 (0.002) | |
| ΔComputer × LF constant | | | | −0.001 (0.003) | | | | −0.001 (0.003) | |
| Computer$_{(t-1)}$ × LF declined | | | | **0.003**$^{**}$ **(0.001)** | | | | **0.003**$^{*}$ **(0.001)** | |
| Computer$_{(t-1)}$ × LF constant | | | | **0.001**$^{**}$ **(0.000)** | | | | −0.002 (0.004) | |
| Large firms declined[a] | | | | **−0.441**$^{**}$ **(0.054)** | | | | **−0.161**$^{**}$ **(0.063)** | |
| ΔComputer × PT increased | | | | | −0.001 (0.003) | | | | |
| ΔComputer × PT constant | | | | | −0.001 (0.002) | | | | |
| Computer$_{(t-1)}$ × PT increased | | | | | **0.002**$^{*}$ **(0.001)** | | | | |
| Computer$_{(t-1)}$ × PT constant | | | | | **0.001**$^{**}$ **(0.001)** | | | | |
| Part-time increased[b] | | | | | **−0.119**$^{**}$ **(0.029)** | | | | |
| ΔComputer × IP increased | | | | | | | | | −0.001 (0.002) |
| ΔComputer × IP constant | | | | | | | | | 0.001 (0.003) |
| Computer$_{(t-1)}$ × IP increased | | | | | | | | | **0.004**$^{**}$ **(0.001)** |
| Computer$_{(t-1)}$ × IP constant | | | | | | | | | **0.002**$^{**}$ **(0.001)** |
| Import penetration increased[a] | | | | | | | | | **−0.125**$^{**}$ **(0.062)** |
| Dependent variable$_{(t-1)}$ | **−0.492**$^{**}$ **(0.044)** | **−0.615**$^{**}$ **(0.036)** | **−0.658**$^{**}$ **(0.038)** | **−0.661**$^{**}$ **(0.037)** | **−0.654**$^{**}$ **(0.036)** | **−0.496**$^{**}$ **(0.085)** | **−0.730**$^{**}$ **(0.079)** | **−0.733**$^{**}$ **(0.086)** | **−0.730**$^{**}$ **(0.090)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.259 | 0.314 | 0.336 | 0.337 | 0.334 | 0.256 | 0.371 | 0.371 | 0.371 |
| Modified DW | 1.95 | 1.90 | 1.89 | 1.88 | 1.88 | 2.17 | 1.99 | 1.99 | 2.00 |
| Stationary test, *P*-value[c] | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *N* | 1.247 | 1.247 | 1.247 | 1.247 | 1.247 | 450 | 450 | 450 | 450 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ Indicates the annual change in the variable. UD denotes union density, LF denotes large firms, PT part-time, and IP import penetration.

* $P < 0.10$.

** $P < 0.05$, two-tailed test.

[a] Dummy variable 1 denotes a decline of more than five percentage points between 1969 and 1997.

[b] Dummy variable 1 denotes an increase of more than 1.5 percentage points between 1969 and 1997.

[c] The null hypothesis in the panel stationary test is that the error terms are nonstationary.

## Appendix E.

Direct and channeled effects of computers on earnings inequality in nonagricultural private-sector industries (1976–1997).

| Dep. variable | ΔHourly wage inequality 90/10 | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| Sector | Private sector | | | | | Manufacturing | | | |
| No. of industries | 43 | 43 | 43 | 43 | 43 | 18 | 18 | 18 | 18 |
| Years | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 | 1976–1997 |
| Model | I | II | III | IV | V | VI | VII | VIII | IX |
| ΔComputer investments | 0.000 (0.002) | −0.001 (0.002) | | | | −0.001 (0.002) | | | |
| Computer investments$_{(t-1)}$ | **0.003**$^{**}$ **(0.001)** | **0.003**$^{**}$ **(0.000)** | | | | **0.004**$^{**}$ **(0.001)** | | | |
| ΔCollege supply growth | | | 0.001 (0.002) | 0.001 (0.002) | 0.001 (0.002) | | 0.002 (0.002) | 0.002 (0.002) | 0.001 (0.002) |
| College supply growth$_{(t-1)}$ | | | 0.002 (0.002) | 0.002 (0.002) | 0.002 (0.002) | | 0.003 (0.003) | 0.003 (0.003) | 0.003 (0.003) |
| ΔUnion density | | **−0.006**$^{**}$ **(0.001)** | **−0.005**$^{**}$ **(0.001)** | **−0.005**$^{**}$ **(0.001)** | **−0.005**$^{**}$ **(0.001)** | | **−0.005**$^{**}$ **(0.001)** | **−0.006**$^{**}$ **(0.001)** | **−0.006**$^{**}$ **(0.001)** |
| Union density$_{(t-1)}$ | | **−0.004**$^{**}$ **(0.001)** | −0.001 (0.001) | −0.001 (0.001) | −0.001 (0.001) | | **−0.003**$^{**}$ **(0.001)** | **−0.004**$^{*}$ **(0.001)** | **−0.004**$^{**}$ **(0.001)** |
| ΔLarge firms | | −0.001 (0.001) | −0.001 (0.001) | −0.001 (0.001) | −0.001 (0.001) | | −0.003 (0.003) | −0.003 (0.003) | −0.003 (0.003) |
| Large firms$_{(t-1)}$ | | −0.001 (0.001) | −0.001 (0.001) | −0.001 (0.001) | −0.001 (0.001) | | −0.002 (0.002) | −0.002 (0.002) | −0.002 (0.001) |
| ΔPart-time | | 0.002 (0.002) | 0.002 (0.002) | 0.002 (0.002) | 0.002 (0.002) | | **0.006**$^{*}$ **(0.003)** | **0.006**$^{**}$ **(0.003)** | **0.006**$^{*}$ **(0.003)** |
| Part-time$_{(t-1)}$ | | 0.002 (0.003) | 0.001 (0.002) | 0.001 (0.002) | 0.002 (0.002) | | 0.006 (0.004) | **0.006**$^{*}$ **(0.003)** | **0.006**$^{*}$ **(0.003)** |
| ΔMinimum wage | | | **−0.019**$^{**}$ **(0.007)** | **−0.019**$^{**}$ **(0.007)** | **−0.018**$^{**}$ **(0.007)** | | −0.016 (0.010) | −0.015 (0.010) | −0.016 (0.010) |
| Minimum wage$_{(t-1)}$ | | | **−0.020**$^{**}$ **(−0.005)** | **−0.020**$^{**}$ **(−0.005)** | **−0.020**$^{**}$ **(−0.005)** | | −0.017 (0.010) | −0.015 (0.010) | −0.016 (0.009) |
| ΔImport penetration | | | | | | | −0.002 (0.004) | 0.000 (0.004) | −0.002 (0.004) |
| Import penetration$_{(t-1)}$ | | | | | | | **0.002**$^{**}$ **(0.001)** | **0.002**$^{**}$ **(0.001)** | **0.001**$^{*}$ **(0.000)** |
| Interactions | | | | | | | | | |
| ΔComputer × UD declined | | | −0.002 (0.002) | | | | −0.004 (0.003) | | |
| ΔComputer × UD constant | | | −0.001 (0.002) | | | | −0.001 (0.002) | | |
| Computer$_{(t-1)}$ × UD declined | | | **0.002**$^{*}$ **(0.001)** | | | | **0.002**$^{*}$ **(0.001)** | | |
| Computer$_{(t-1)}$ × UD constant | | | **0.002**$^{**}$ **(0.000)** | | | | 0.000 (0.000) | | |
| Union density declined[a] | | | **−0.564**$^{**}$ **(0.061)** | | | | 0.223 (0.139) | | |
| ΔComputer × LF declined | | | | −0.003 (0.002) | | | | **−0.004**$^{**}$ **(0.002)** | |
| ΔComputer × LF constant | | | | −0.001 (0.002) | | | | 0.009 (0.008) | |
| Computer$_{(t-1)}$ × LF declined | | | | **0.003**$^{**}$ **(0.001)** | | | | 0.001 (0.001) | |
| Computer$_{(t-1)}$ × LF constant | | | | **0.002**$^{**}$ **(0.000)** | | | | −0.006 (0.004) | |
| Large firms declined[a] | | | | **0.123**$^{**}$ **(0.046)** | | | | 0.027 (0.061) | |
| ΔComputer × PT increased | | | | | −0.001 (0.002) | | | | |
| ΔComputer × PT constant | | | | | −0.001 (0.003) | | | | |
| Computer$_{(t-1)}$ × PT increased | | | | | **0.001**$^{*}$ **(0.001)** | | | | |
| Computer$_{(t-1)}$ × PT constant | | | | | **0.002**$^{**}$ **(0.001)** | | | | |
| Part-time increased[b] | | | | | **−0.357**$^{**}$ **(0.035)** | | | | |
| ΔComputer × IP increased | | | | | | | | | −0.003 (0.003) |
| ΔComputer × IP constant | | | | | | | | | 0.002 (0.002) |
| Computer$_{(t-1)}$ × IP increased | | | | | | | | | **0.003**$^{**}$ **(0.001)** |
| Computer$_{(t-1)}$ × IP constant | | | | | | | | | −0.001 (0.001) |
| Import penetration increased[a] | | | | | | | | | −0.003 (0.039) |
| Dependent variable$_{(t-1)}$ | **−0.632**$^{**}$ **(0.064)** | **−0.736**$^{**}$ **(0.058)** | **−0.764**$^{**}$ **(0.059)** | **−0.764**$^{**}$ **(0.060)** | **−0.768**$^{**}$ **(0.059)** | **−0.465**$^{**}$ **(0.059)** | **−0.749**$^{**}$ **(0.071)** | **−0.753**$^{**}$ **(0.066)** | **−0.762**$^{**}$ **(0.076)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.322 | 0.370 | 0.390 | 0.391 | 0.392 | 0.247 | 0.393 | 0.397 | 0.400 |
| Modified DW | 1.98 | 1.93 | 1.92 | 1.92 | 1.92 | 2.19 | 2.01 | 2.00 | 2.01 |
| Stationary test, $P$-value[c] | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| $N$ | 946 | 946 | 946 | 946 | 946 | 396 | 396 | 396 | 396 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ indicates the annual change in the variable. UD denotes union density, LF denotes large firms, PT part-time, and IP import penetration.

* $P < 0.10$.

** $P < 0.05$, two-tailed test.

[a] Dummy variable 1 denotes a decline of more than five percentage points between 1969 and 1997.

[b] Dummy variable 1 denotes an increase of more than 1.5 percentage points between 1969 and 1997.

[c] The null hypothesis in the panel stationary test is that the error terms are nonstationary.

## Appendix F.

Direct and channeled effects of computers on earnings inequality in nonagricultural private-sector industries (1988–2012).

| Dep. Variable | ΔAnnual wage inequality 90/10 | | | | | ΔHourly wage inequality 90/10 | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Sector | Private sector | | | | | | | | | |
| No. of industries | 43 | 43 | 43 | 43 | 43 | 43 | 43 | 43 | 43 | 43 |
| Years | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 | 1988–2012 |
| Model | I | II | III | IV | V | VI | VII | VIII | IX | X |
| ΔComputer investments | 0.001 (0.002) | 0.002 (0.002) | | | | 0.000 (0.001) | 0.001 (0.001) | | | |
| Computer investments$_{(t-1)}$ | **0.003$^{**}$ (0.001)** | **0.002$^{**}$ (0.001)** | | | | **0.002$^{**}$ (0.001)** | **0.002$^{**}$ (0.001)** | | | |
| ΔCollege supply growth | | | **−0.004$^{**}$ (0.002)** | **−0.004$^{**}$ (0.002)** | **−0.004$^{**}$ (0.002)** | | | −0.003 (0.002) | −0.003 (0.002) | −0.003 (0.002) |
| College supply growth$_{(t-1)}$ | | | **−0.008$^{**}$ (0.003)** | **−0.008$^{**}$ (0.003)** | **−0.009$^{**}$ (0.003)** | | | −0.003 (0.002) | −0.002 (0.002) | −0.002 (0.002) |
| ΔUnion density | | **−0.007$^{**}$ (0.002)** | **−0.006$^{**}$ (0.002)** | **−0.006$^{**}$ (0.002)** | **−0.006$^{**}$ (0.002)** | | **−0.005$^{**}$ (0.002)** | **−0.005$^{**}$ (0.002)** | **−0.005$^{**}$ (0.002)** | **−0.005$^{**}$ (0.002)** |
| Union density$_{(t-1)}$ | | **−0.010$^{**}$ (0.002)** | **−0.009$^{**}$ (0.002)** | **−0.010$^{**}$ (0.002)** | **−0.009$^{**}$ (0.002)** | | **−0.008$^{**}$ (0.002)** | **−0.008$^{**}$ (0.002)** | **−0.008$^{**}$ (0.002)** | **−0.009$^{**}$ (0.001)** |
| ΔLarge firms | | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) | | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) |
| Large firms$_{(t-1)}$ | | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) | 0.001 (0.001) | | 0.000 (0.001) | 0.000 (0.001) | 0.000 (0.001) | 0.000 (0.001) |
| ΔPart-time | | **−0.005$^{**}$ (0.002)** | **−0.004$^{*}$ (0.002)** | **−0.004$^{*}$ (0.002)** | **−0.004$^{*}$ (0.002)** | | **−0.003$^{**}$ (0.002)** | **−0.003$^{*}$ (0.002)** | **−0.003$^{*}$ (0.002)** | **−0.003$^{*}$ (0.002)** |
| Part-time$_{(t-1)}$ | | **−0.004$^{*}$ (0.002)** | −0.003 (0.002) | −0.003 (0.002) | −0.003 (0.002) | | **−0.004$^{**}$ (0.002)** | **−0.004$^{*}$ (0.002)** | **−0.004$^{*}$ (0.002)** | **−0.004$^{*}$ (0.002)** |
| ΔMinimum wage | | | 0.002 (0.009) | 0.001 (0.009) | 0.002 (0.009) | | | −0.004 (0.007) | −0.004 (0.007) | −0.003 (0.007) |
| Minimum wage$_{(t-1)}$ | | | −0.004 (0.007) | 0.004 (0.007) | 0.004 (0.007) | | | −0.001 (0.006) | −0.001 (0.006) | −0.001 (0.006) |
| Interactions | | | | | | | | | | |
| ΔComputer × UD declined | | | 0.001 (0.003) | | | | | 0.002 (0.002) | | |
| ΔComputer × UD constant | | | 0.002 (0.002) | | | | | 0.001 (0.001) | | |
| Computer$_{(t-1)}$ × UD declined | | | **0.002$^{**}$ (0.001)** | | | | | **0.002$^{**}$ (0.001)** | | |
| Computer$_{(t-1)}$ × UD constant | | | **0.002$^{*}$ (0.001)** | | | | | **0.002$^{**}$ (0.001)** | | |
| Union density declined[a] | | | **0.271$^{**}$ (0.115)** | | | | | **0.453$^{**}$ (0.075)** | | |
| ΔComputer × LF declined | | | | 0.003 (0.002) | | | | | 0.002 (0.002) | |
| ΔComputer × LF constant | | | | 0.002 (0.002) | | | | | 0.001 (0.001) | |
| Computer$_{(t-1)}$ × LF declined | | | | **0.002$^{*}$ (0.001)** | | | | | **0.002$^{*}$ (0.001)** | |
| Computer$_{(t-1)}$ × LF constant | | | | **0.002$^{**}$ (0.001)** | | | | | **0.002$^{**}$ (0.001)** | |
| Large firms declined[a] | | | | **0.568$^{**}$ (0.108)** | | | | | **−0.081$^{*}$ (0.023)** | |
| ΔComputer × PT increased | | | | | −0.005 (0.006) | | | | | −0.003 (0.006) |
| ΔComputer × PT constant | | | | | 0.002 (0.002) | | | | | 0.001 (0.001) |
| Computer$_{(t-1)}$ × PT increased | | | | | 0.003 (0.002) | | | | | 0.003 (0.002) |
| Computer$_{(t-1)}$ × PT constant | | | | | **0.002$^{**}$ (0.001)** | | | | | **0.002$^{*}$ (0.001)** |
| Part-time increased[b] | | | | | **0.290$^{**}$ (0.108)** | | | | | **−0.121$^{**}$ (0.033)** |
| Dependent variable$_{(t-1)}$ | **−0.677$^{**}$ (0.096)** | **−0.786$^{**}$ (0.089)** | **−0.797$^{**}$ (0.090)** | **−0.797$^{**}$ (0.091)** | **−0.798$^{**}$ (0.089)** | **−0.686$^{**}$ (0.100)** | **−0.792$^{**}$ (0.098)** | **−0.797$^{**}$ (0.099)** | **−0.797$^{**}$ (0.100)** | **−0.798$^{**}$ (0.099)** |
| Constant | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Industry dummies | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| $R^2$ | 0.338 | 0.392 | 0.397 | 0.397 | 0.398 | 0.345 | 0.399 | 0.401 | 0.401 | 0.402 |
| Modified DW | 2.05 | 1.99 | 2.00 | 2.00 | 2.00 | 2.05 | 2.00 | 2.01 | 2.01 | 2.02 |
| Stationary test, *P*-value[c] | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *N* | 1075 | 1075 | 1075 | 1075 | 1075 | 1075 | 1075 | 1075 | 1,075 | 1,075 |

Each column represents a pooled regression of changes in wage inequality. Table entries are OLS estimates. Robust standard errors in parentheses are heteroskedasticity and autocorrelation consistent. Estimates are weighted by mean industry share of total employed workers over the years. All models control for recession years. Δ indicates the annual change in the variable. UD denotes union density, LF denotes large firms, and PT part-time.

* $P < 0.10$.

** $P < 0.05$, two-tailed test.

[a] Dummy variable 1 denotes a decline of more than five percentage points between 1988 and 2012.

[b] Dummy variable 1 denotes an increase of more than 1.5 percentage points between 1988 and 2012.

[c] The null hypothesis in the panel stationary test is that the error terms are nonstationary.

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