Changes in the relative structure of wages and employment: a comparison of the United States, Canada, and France

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Abstract. Standard economic models suggest that adverse demand shocks will lead to bigger employment losses if institutional factors prevent real wages from declining. Some analysts have argued that this insight explains the dichotomy between the United States, where real wages of less-skilled workers fell over the 1980s and aggregate employment expanded vigorously, and Europe, where real wages of less-skilled workers were constant and employment was stagnant. We find little support for this hypothesis when we compare recent changes in wages and employment rates for different age and education groups in the United States, Canada, and France. JEL Classification: J21, J31

Changements dans la structure relative des salaires et de l'emploi: une comparaison des Etats-Unis, du Canada et de la France. Dans le modèle concurrentiel standard, une baisse de la demande de travail entraîne une chute plus marquée de l'emploi lorsque que des facteurs institutionnels empêchent les salaires de s'ajuster à la baisse, que lorsque ces derniers sont parfaitement flexibles. Certains analystes ont suggéré que cette observation permet d'expliquer l'apparente dichotomie entre les États-Unis, où les salaires réels des travailleurs moins qualifiés chutèrent dans les années 1980 alors que la croissance de l'emploi y était robuste, et l'Europe, où ces salaires demeurèrent constants pendant que l'emploi ne montrait aucun signe de croissance. Nous trouvons peu de support pour cette hypothèse en comparant l'évolution récente des salaires et des taux d'emploi pour différents groupes de travailleurs classifiés selon leur âge et leur scolarisation aux États-Unis, au Canada et en France.

1. Introduction

The U.S. labour market displayed two prominent trends during the 1980s: robust

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employment growth and widening wage inequality. The corresponding trends in many Western European countries (e.g., Germany and France) were very different. While indexes of wage inequality showed little or no change, employment was stagnant, leading to massive increases in unemployment.[1] Some economists have argued that the contrast between the United States and Europe illustrates a fundamental tradeoff between wage inequality and employment growth in the face of declining demand for less-skilled labour.[2] According to this hypothesis, job growth in the United States has been aided by flexible institutions that allowed real wages of younger and less-educated workers to fall. In Western Europe, on the other hand, minimum wages, union wage setting, and generous unemployment benefits have propped up the wages of less-skilled workers, preventing a rise in wage inequality but also severely limiting job growth.

In this paper we attempt to evaluate the evidence in favour of this 'tradeoff hypothesis' using data on detailed age and education groups for the United States, Canada, and France. We argue that similar negative shocks have affected the relative demand for less-skilled workers in all three countries. In the United States, where labour markets are flexible, the tradeoff hypothesis predicts that adverse shocks will primarily affect the relative *wages* of less-skilled workers. In France, where labour markets are relatively inflexible, the same shocks will primarily affect the relative *employment* of less-skilled workers. Finally, in Canada, where labour market institutions lie somewhere between those of the United States and France, adverse demand shocks will lead to smaller relative wage adjustments than in the United States and smaller relative employment changes than in France.

To test the tradeoff hypothesis we require comparable indexes of the relative demand shocks that have affected different skill groups in the three countries. We consider two alternatives. The first builds on the fact that wages of different skill groups tended to 'fan out' over the 1980s (see Juhn, Murphy, and Pierce 1993, or Card and Lemieux 1996). This pattern suggests that the level of wages for a particular skill group in the early 1980s can serve as a proxy for the relative demand shocks faced by the group over the decade. Our second index is linked to a more specific explanation for widening wage inequality: skill-biased technical change driven by innovations in computer technology (see, e.g., Krueger 1993; Berman, Bound, and Griliches 1994). We use the fraction of each skill group that used a computer on the job in the late 1980s as an alternative index of relative demand shifts over the preceding decade.

We compare the effects of these two alternative demand indexes on the structure of relative wages and employment across different age and education groups over the 1980s. For the United States, we find that groups with lower wages at the beginning of the 1980s – or lower rates of computer use at the end of the 1980s – had

1 An important exception to the Western European pattern is the United Kingdom, where inequality and employment trends were closer to those in the United States. See Freeman and Katz (1995).

2 See Krugman (1994) for a clear statement of this view and Freeman (1994, 14) and Katz, Loveman, and Blanchflower (1995, 57) for more guarded statements.

significantly slower wage growth over the decade and somewhat slower employment growth. A simple interpretation of these patterns is that negative demand shocks put downward pressure on the relative wages of less-skilled workers, ultimately leading to reductions in relative labour supply (this interpretation is suggested by Juhn 1992, among others). In contrast to the patterns in the United States, relative wages of French workers with lower wages at the beginning of the 1980s – or lower rates of computer use at the end of the decade – were constant or even rose slightly over the decade. Nevertheless, the relative employment of less-skilled French workers declined at roughly the same rate as in the United States. Our results for Canada are less clear cut, although on balance we find smaller relative wage adjustments across skill groups than in the United States. As in the French case, this relative wage inflexibility does not seem to have led to a systematic widening of employment outcomes across skill groups relative to the situation in the United States.[3]

2. Labour markets in the United States, Canada, and France

2.1. Sources of Relative Labour Demand Shocks

There is now an extensive U.S. literature documenting the fact that real wages of younger and less-educated workers fell over the 1980s and attempting to explain these trends (see e.g., Murphy and Welch 1991; Bound and Johnson 1992; Katz and Murphy 1992; Borjas and Ramey 1994, 1995; Lawrence and Slaughter 1993; Kosters 1994). A primary conclusion of this literature is that the declines are attributable to relative demand shifts.[4] The two leading explanations for these shifts are skill-biased technical change and trade.

With respect to technological change, many analysts have suggested that computers differentially raise the productivity of more highly skilled workers, leading to a decline in relative demand for less-skilled labour (see Bound and Johnson 1992 and Krueger 1993). With respect to trade, standard theory suggests that the opening up of trade will lower the demand for workers in industries that must compete with the newly available import goods. If younger or less-educated workers are disproportionately employed in trade-impacted industries, rising trade may raise skill premiums throughout the labour force.[5]

To what extent have the same kinds of trade and technology shocks affected labour markets in France and Canada? To get a rough impression we assembled

3 Our results are consistent with Nickell and Bell (1996), who find that changes in the relative structure of unemployment were similar in OECD countries. Kuhn and Robb (1997) also find that changes in the relative structure of employment by skill level were relatively similar in Canada and the United States. One contribution of our paper relative to these two other papers is that we look explicitly at the impact of several indexes of the relative demand shocks like computer usage rates.

4 Some authors (e.g., DiNardo, Fortin, Lemieux 1996) have argued that institutional changes, such as declining unionization and falling minimum wages, have contributed to widening wage inequality in the United States.

5 The empirical plausibility of the case that trade has affected relative wages is a matter of some dispute. See, for example, Bhagwati and Dehejia (1994), Lawrence (1994), and Leamer (1994).

TABLE 1
Utilization of computers at work in the United States, Canada, and France

	United States 1989	Canada 1989	France 1991
1. Overall	37.4	33.8	34.0
2. By gender			
Women	43.4	37.6	37.6
Men	32.3	30.8	31.2
Ratio: men/women	0.74	0.82	0.83
3. By education			
Less than high school[a]	7.8	12.4	14.6
High school or diploma[b]	29.3	37.2	40.2
College degree[c]	58.2	54.5	54.9
Ratio: Less than high school / college	0.13	0.23	0.27
Ratio: high school / college	0.50	0.68	0.73

NOTES
U.S. data are taken from Krueger (1993). Canadian data are taken from Lowe (1991). French data are taken from unpublished tabulations of the 1991 Labor Force Survey. In all three countries, use of a computer includes word-processing, micro- and mainframe computing.
a In France, this category includes those with no degree, and those whose highest qualification is a CEP (elementary school) certificate.
b In France, this category includes those whose highest qualification is either BEPC (roughly, junior high school) or CAP (vocational or technical school), or baccalauréat (academic high school).
c In the United States this category includes those with sixteen years of education. In Canada, it includes those with a university degree. In France, it includes those with an études supérieures qualification.

the comparative data in tables 1 and 2. Table 1 shows the fraction of the workforce who report that they use a computer at work as of the late 1980s in the United States, Canada, and France. Computer use is slightly higher in the United States than in Canada or France, but the diffusion of computer technology seems to have proceeded at a fairly similar pace in the three economies. Women in each country are more likely than men to use a computer at work, perhaps reflecting occupational differences in the adoption of computers.

As shown in row 3 of table 1, better-educated workers are also more likely to use a computer, and, in fact, usage rates for college graduates are fairly similar in the three countries. For those without a college degree computer usage rates are 30–45 per cent higher in Canada and France than in the United States. Conditional on education, computer use is therefore *lower* in the United States than in Canada or France, but higher levels of education in the United States lead to a higher overall level of computer use.

When we turn to the issue of trade, table 2 presents a series of import penetration ratios for the three countries in 1973, 1983, and 1993. The first column of the table shows the ratio of total imports to GDP in each country. Compared with Canada or

TABLE 2
Imports as a fraction of GDP in the United States, Canada, and France

	Imports/GDP	Percent of imports from major partner	Imports/GDP excluding major partner
A. United States			
1973	5.1	25.5	3.8
1983	7.6	20.4	6.1
1993	8.8	18.8	7.1
B. Canada			
1973	17.8	70.7	5.2
1983	18.0	72.1	5.0
1993	24.1	73.2	6.5
C. France			
1973	14.7	67.8	4.8
1983	20.0	66.3	6.7
1993	24.1	73.2	6.5

NOTES
U.S. data are taken from *Economic Report of the President*, (1994 ed.) and *Statistical Abstract of the U.S.* (1977 ed.). Canadian data are taken from *Bank of Canada Review* (Autumn 1994 ed.) and *Canada Year Book* (1980–81 ed.). French data are taken from *Annuaire Statistique de la France* (1978 and 1990 eds) and *Direction Générale des Douanes et Droits Indirects, Résultats Annuels* (1994 ed.).
a Major partner refers to Canada in the case of the United States, United States in the case of Canada, and includes ECM, Sweden, Switzerland, Norway, Finland, USSR/Russia, and Eastern Europe in the case of France. The entry for France for 1993 is based on 1989 data.

France (or indeed, with most other OECD nations), the United States is a low-trade country. Nevertheless, the *rate of growth* of imports relative to domestic output was slightly faster in the United States than in France or Canada over the past two decades. It is interesting, however, that U.S. imports expanded more rapidly during the 1970s than during the 1980s. Concentrating only on the 1980s, we see that imports grew more slowly in the United States than in the other two countries.[6]

One could argue that even though the United States is a low-trade country, it is more vulnerable to import competition from low-wage countries or from countries with technological advantages in certain products (like Japan). To address this issue, we recalculated the import-penetration ratios, excluding Canadian imports from the U.S. data, U.S. imports from the Canadian data, and European imports from the French data. The results, presented in the third column of table 2, show that even under this more restrictive notion of import penetration, the rates of growth of imports in the three countries are roughly comparable. There is no indication that Canada or France are more isolated from rising international trade than the United States is.[7]

6 This conclusion continues to hold when we use 1969, 1979, and 1989 data.

7 The large fraction of Canadian trade flows accounted for by trade with the United States suggests that shocks that affect the U.S. market will be transferred to Canada.

2.2. Sources of labour market inflexibility

A maintained assumption throughout this paper is that institutional factors that resist downward wage adjustments for lower-paid workers are relatively strong in France, somewhat weaker in Canada, and weakest in the United States. Among the most important of these features are minimum wage laws, collective bargaining arrangements, and income support programs for non-workers. In this section we briefly summarize some evidence on the importance of these institutions in the three labour markets.

The United States, Canada, and France all have minimum wage laws: the United States and France set national standards, while minimum wages in Canada are largely a provincial issue.[8] One useful summary measure of these laws is the ratio of the minimum wage to the average wage for a representative subset of workers. Figure 1 plots this ratio for the three countries over the 1966–93 period.[9] Relative minimum wages in Canada and the United States fell sharply from the late 1970s to the late 1980s, with modest rebounds in both countries in the early 1990s. In France, by comparison, minimum wages edged up slightly relative to average salaries in the early 1980s and then remained constant. The decline in the relative value of the minimum wage in the United States and Canada over the 1980s suggests that minimum wage regulations probably did not prevent relative wage declines for less-skilled workers in either country during the 1980s. The continuing high level of the relative minimum wage in France, on the other hand, suggests that minimum wages could have played an important role in reducing relative wage flexibility in the French labour market (see Katz, Loveman, and Blanchflower 1995).

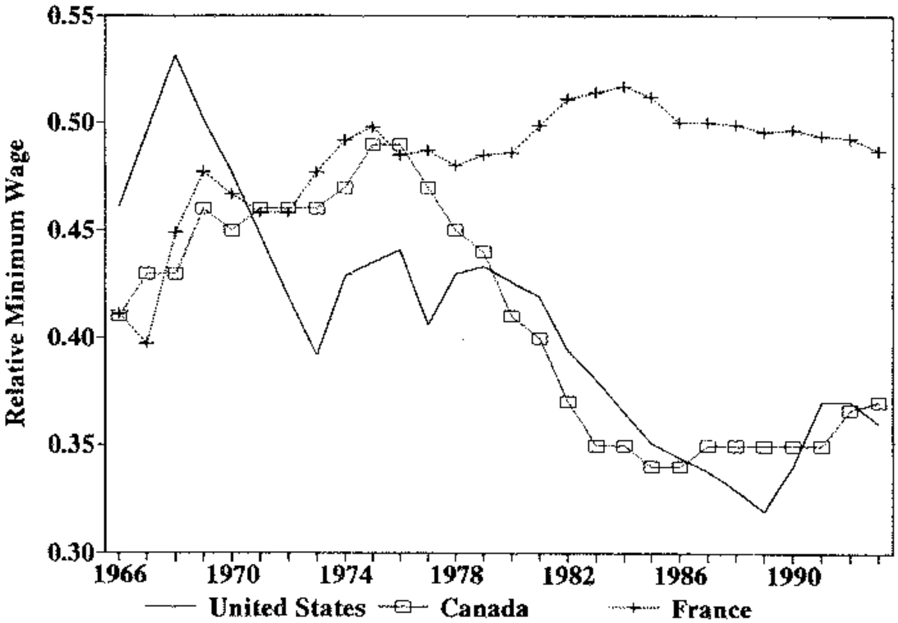
Table 3 summarizes some recent data on union membership and collective bargaining coverage in the three countries.[10] In Canada and the United States, collective bargaining is conducted at the establishment or firm level. Since about 90 per cent of workers who are covered by union contracts in both countries are union members, it is customary to use union membership rates as a measure of the extent of collective bargaining coverage.[11] The data in the first two columns of table 3 show that although unionization rates were similar in the United States and Canada in the early 1970s, by 1980 the Canadian rate was 50 per cent higher. During the

8 U.S. states can (and sometimes do) set minimum wage rates above the federal rate. For an overview of U.S. minimum wage laws, see Card and Krueger (1995). For a discussion of Canadian laws, see West and McKee (1980) and Baker, Benjamin, and Stanger (1994). For a discussion of the French minimum wage (the salaire minimum interprofessionnel de croissance, or SMIC) see Bayet (1994).

9 For the United States, we divide the federal minimum wage by the average hourly wage in manufacturing. For Canada, we use a weighted average of province-specific ratios of the minimum wage to average hourly wage of manufacturing workers. We are grateful to Dwayne Benjamin for supplying his data for this calculation. For France, we use the ratio of the average annual SMIC (net of the employee share of the payroll tax), multiplied by an estimate of average hours per year (for full time full year workers) and divided by an estimate of average net annual salaries. See Bayet (1994).

10 A closely related source of potential wage rigidity is the public sector. This may be particularly important for France.

11 In both the United States and Canada, union membership rates and union coverage rates tend to show very similar cross-sectional patterns and trends. See Lemieux (1993) and Riddell (1993).



Note: minimum wage relative to average manufacturing wage in U.S. and Canada and relative to overall average wage in France.

FIGURE 1 Minimum wage relative to average wage in the United States, Canada, and France

1980s the divergence continued, with declining unionization in the United States and roughly steady rates in Canada. To the extent that North American unions resist real wage declines and fight to maintain relative wage differentials, the lower level of unionism in the United States would be expected to enhance downward relative wage flexibility vis-à-vis Canada.

Collective bargaining institutions in most European countries are very different from those in North America. In France, industry-level bargaining sets minimum pay scales for most employees, with supplemental enterprise or establishment agreements covering about one-fifth of workers (mainly in larger companies).[12] In this setting individual union membership has no direct effect on collective bargaining coverage. As shown in the third column of table 3, union membership rates were low and falling in France throughout the 1970s and 1980s. Nevertheless, as shown in column 4, collective bargaining coverage remains very high. The broad coverage

12 See U.S. Department of Labor (1992). The baseline pay scales apply to all firms belonging to the employer's association that signed the collective agreement. In about one-half of the cases, the agreement is extended by the Minister of Labour to the rest of the industry. In 1985, 86 per cent of private sector workers employed in firms with at least ten employees were covered by an industry-level agreement, while 35 per cent of them were covered by a firm- or establishment-level agreement (Benveniste (1987)).

TABLE 3
Unionization rates in the United States, Canada, and France

	United States (membership)	Canada (membership)	France	
			Membership	Coverage[a]
1970	29.6	33.6	23.1	–
1975	28.9/22.3	35.6	22.9	–
1980–81	22.7	36.7/34.2	19.2/14.0	–
1985	17.9	35.7	10.0	–
1989–90	16.3	33.1	–	92.0

NOTES
U.S. data for 1970 and 1975 (first entry) are based on union records and are from Riddell (1993, table 4.1, column 2). Data for 1975 (second entry) and later are based on individual micro-data as reported in Curme, Hirsch, and Macpherson (1990, summary table).
Canadian data for 1970, 1975, and 1980 (first entry) are based on union records as reported in Riddell (1993, table 4.1, column 1). Data for 1981 (second entry) and later are based on individual micro-data as reported in Riddell (1993, table 4.1, column 3). French data for 1970, 1975, and 1985 (first entry) are based on union records (excluding retirees) as reported in Price (1989). Data for 1980 (second entry) and 1985 are based on individual micro-data as reported in Haeusler (1990).
a Estimated fraction of employees covered by collective bargaining contracts, taken from OECD (1994).

of industry-wide minimum pay rates suggests that collective bargaining institutions may have some effect on the French wage structure – particularly in preventing wage reductions for low-skilled workers.[13]

A final set of institutions that play a role in determining the downward-flexibility of wages is composed of unemployment insurance and income support programs for non-workers. Standard models imply that workers will not accept jobs that pay less than they can receive from unemployment insurance or welfare payments: thus higher benefit levels, or broader eligibility rules, will tend crowd out low-wage jobs.[14] France, Canada, and the United States all offer limited-duration unemployment benefits to recent job losers who meet certain eligibility requirements.[15] In the United States these benefits are paid to only a minority of unemployed individuals (30-40 per cent), whereas the Canadian and French unemployment insurance systems are far more inclusive.[16] A recent OECD study (OECD 1994) concludes that the French program is more generous than the Canadian program (especially

13 Note, however, that the median value (across industries) of the baseline pay scale negotiated at the industry level was below the SMIC during our period of analysis (Barrat, Coutrot, and Mabile 1994). This suggests that low-wage workers were better protected against a decline in their relative wages by the active SMIC policy pursued by the government than by industry level agreements.

14 In fact, both labour supply and search theoretic models suggest that the lowest wage a worker will accept is somewhat above the level of welfare or unemployment benefits.

15 The U.S. and Canadian programs are similar in broad outline, while the French system includes a second stage of extended benefits for those who exhaust the first stage.

16 During the 1980s the Canadian program typically provided benefit payments to more people than were counted as unemployed (Card and Riddell 1993), although recent reforms have scaled back coverage. The French unemployment insurance program paid benefits to about as many people as were counted as unemployed in 1990 (OECD 1994, table 8.4).

after taking account of longer-term unemployed), and that both are more generous than that of the United States.

In addition to time-limited programs for recent job losers, France, Canada, and the United States all have income support programs for certain groups of non-workers. In the United States food stamps are the only form of support widely available to able-bodied non-workers; most states limit traditional 'welfare' payments to single mothers.[17] By comparison, both Canada and France have social assistance programs that provide cash payments to non-workers regardless of family structure. All three countries also have income support programs for disabled individuals, which may be particularly important in determining the labour force behaviour of relatively unskilled older workers (see Bound and Burkhauser 1999). While the safety net systems of Canada, France, and the United States are quite complex, taken as a whole we believe that the programs are most generous and most comprehensive in France and least generous and least comprehensive in the United States.

3. Relative wages and employment patterns

3.1. Theoretical framework

In this section we present a simple theoretical model of the effects of demand and supply shocks on the relative wages and employment rates of different skill groups. We utilize a model with only one output good because previous studies (e.g., Bound and Johnson 1992; Katz and Murphy 1992) have ruled out between-industry factors as a major source of rising wage inequality in the United States. Consider a competitive economy in which firms use labour inputs $N_1, N_2, \ldots, N_J$ to produce a homogenous output Y. Suppose that output in period t is related to the inputs by a constant-returns-to-scale CES production function:

$$Y_t = f(N_{1t}, N_{2t}, \ldots, N_{Jt}) = \{\Sigma_j (c_{jt} N_{jt})^{(\sigma-1)/\sigma}\}^{\sigma/(\sigma-1)}, \tag{1}$$

where σ is the elasticity of substitution, and c_{jt} is a relative efficiency parameter for skill group j in period t.[18] If the wage rate for skill group j is denoted by w_{jt}, then the demand for labour of the jth skill group in period t satisfies the following equation:

$$\log N_{jt} = \log Y_t - \sigma \log w_{jt} + (\sigma - 1) \log c_{jt}. \tag{2}$$

Assuming that $\sigma > 1$, a rise in c_{jt} leads to an increase in the relative demand for group j. As noted by Bound and Johnson (1992), there are several possible interpretations of c_{jt}. One that we emphasize here is as a technological innovation

17 Most U.S. states and counties have very modest general assistance benefits for people who are ineligible for other forms of income assistance. See Blank and Hanratty (1993) for a detailed comparison of income support programs in the United States and Canada.

18 Any 'factor neutral' technical change results in a proportional shift in all the c_{jt}s. A model similar to the one here, but ignoring labour supply, is presented by Bound and Johnson (1992).

that alters the relative efficiency of different groups of workers. Another (that we ignore) is as a shift in the average human capital of individuals in the jth group, arising, for example, through changes in the quality of schooling.

Assume that the population in period t contains P_{jt} individuals in the jth skill class, and that each individual who chooses to work supplies a fixed number of hours per period (normalized to 1). Finally, assume that individuals within each skill group differ in their reservation wages and, that when the (real) wage rate is w_{jt}, a fraction p_{jt} choose to work, where

$$p_{jt} = N_{jt}/P_{jt} = a_j w_{jt}^{\varepsilon}, \tag{3}$$

where a_j is a taste parameter that varies across groups and ε is the elasticity of labour supply, which we assume is constant across groups. Let P_t represent the total (adult) population in period t, let $y_t = Y_t/P_t$ represent per capita output in period t, and let $f_{jt} = P_{jt}/P_t$ represent the fraction of the population in skill group j. Making use of these definitions, equations (2) and (3) can be combined to yield

$$\log w_{jt} = \frac{1}{\sigma + \varepsilon}\{\log y_t + (\sigma - 1)\log c_{jt} - \log (a_j) - \log f_{jt}\} \tag{4a}$$

$$\log p_{jt} = \frac{\sigma}{\sigma + \varepsilon} \log (a_j) + \frac{\varepsilon}{\sigma + \varepsilon}\{\log y_t + (\sigma - 1)\log c_{jt} - \log f_{jt}\}. \tag{4b}$$

These equations relate group-specific wages and employment-population rates to a year-specific aggregate term ($\log y_t$), the taste parameter a_j, and the demand and supply shift variables c_{jt} and f_{jt}.

If data are available for two periods, then the changes in wages and employment rates across different groups will satisfy

$$\Delta \log w_j = b_1 + \frac{1}{\sigma + \varepsilon}\{(\sigma - 1)\Delta \log c_{jt} - \Delta \log f_{jt}\} \tag{5a}$$

$$\Delta \log p_j = b_2 + \frac{\varepsilon}{\sigma + \varepsilon}\{(\sigma - 1)\Delta \log c_{jt} - \Delta \log f_{jt}\} \tag{5b}$$

where b_1 and b_2 are constants. These equations relate the changes in wages and employment rates to observable relative supply shocks (captured by the changes in the skill group population shares) and unobservable relative demand shifts.

We use two alternative proxies for the relative demand shifts that have affected different skill groups. The first is motivated by the observation that in the U.S. labour market, groups with higher wage levels have had faster real wage growth. For example, Juhn, Murphy and Pierce (1993) show that the real wage levels associated with different percentiles of the wage distribution 'fanned out' over the 1980s. Card and Lemieux (1996) present formal tests of the hypothesis that the real changes for different demographic groups during the 1980s are a monotonic function of their initial wage levels. Although this hypothesis is rejected, it provides a relatively accurate description of relative wage movements. Assuming that relative wage changes in the U.S. labour market in the 1980s were driven by relative demand shocks, the findings in Juhn, Murphy, and Pierce (1993) and Card and Lemieux (1996) imply that the level of wages of the jth group at the beginning of the decade ($\log w_{j0}$) is a useful proxy for the relative demand shock faced by the group over the subsequent years, although the cause of the demand shifts is not explained.

Our second proxy is motivated by recent research linking rising wage inequality in the U.S. labour market to technological changes associated with computers. Autor, Katz, and Krueger (1998) argue that changes in the utilization of computers by different groups reflect the degree of complementarity between the group and computer technologies. In the spirit of this conclusion, our second demand shift index is simply the fraction of the jth skill group who are observed using computers at work at the end of the 1980s (cu_{j1}). Assuming that relative demand shocks are driven mainly by the diffusion of computer technologies (as argued by Krueger 1993; Berman, Bound, and Griliches 1994; Autor, Katz, and Krueger 1998; Bound and Johnson 1992) the magnitude of the relative demand shock faced by a certain skill group over the 1980s is proportional to the fraction of the group who used computers on the job at the end of the decade.

Let D_j represent the proxy variable (either cu_{j1} or $\log w_{j0}$) used to predict the relative demand shift experienced by group j. We assume that the relationship between the skill-group-specific relative productivity terms and the demand proxies is approximately linear:

$$(\sigma - 1)\Delta \log c_{jt} = \alpha + \beta D_j + u_j, \tag{6}$$

where $\beta > 0$. Substitution of (6) into (5a) and (5b) leads to equations of the form:

$$\Delta \log w_j = d_1 + \frac{\beta}{\sigma + \varepsilon} D_j - \frac{1}{\sigma + \varepsilon} \Delta \log f_{jt} + e_{j1} \tag{7a}$$

and

$$\Delta \log p_j = d_2 + \frac{\beta \varepsilon}{\sigma + \varepsilon} D_j - \frac{\varepsilon}{\sigma + \varepsilon} \Delta \log f_{jt} + e_{j2} \tag{7b}$$

where d_1 and d_2 are constants, and e_{j1} and e_{j2} are stochastic terms representing a combination of the error term in the prediction equation for relative demand shifts (u_j) and any sampling errors in the observed data on wages and employment rates.

The assumptions underlying the derivation of equations (5a) and (5b) are highly restrictive. A particularly strong assumption is that there is a single elasticity of substitution between all pairs of skill groups. More realistically, one might conjecture that workers in one skill group are better substitutes for workers with 'nearly the same' characteristics. In principle, a full set of cross-elasticities of factor demand for each skill group can be introduced, but the number of parameters becomes unmanageable in the absence of some very restrictive assumptions about substitutability.

A second assumption is that workers within each skill group are equally productive and differ only in their reservation wages. This assumption implies that the mean wage for workers in a given skill group is an unbiased estimate of the wage that would be observed for the non-workers in the group if they chose to work. A (possibly) more realistic assumption is that workers in each group are perfect substitutes in production, but that different individuals in the group have different relative productivities. In this case, the potential wage of the ith person in the jth group in period t can be written as

$$\log w_{ijt} = \log w_{jt} + k_{ij}, \tag{8}$$

where w_{jt} is the wage for a 'standardized' person in group j (e.g., the person with the median productivity level). This assumption introduces two complications into the simple model of equations (5a) and (5b). First, the 'effective' labour force of the jth skill group is no longer a simple head count, but must be adjusted for the relative productivities of those who are actually observed working. A second and related issue is that the observed wage of the group may vary as the employment rate varies, depending on the relative productivities of workers and non-workers in the group.

There are a number of ways to control for changes in the observed wages of a particular skill group that are attributable to changes in the composition of employment within the group, rather than to changes in underlying supply or demand factors. One simple method that we use in the next section is to assume that non-workers would earn a relatively low wage – specifically, the minimum wage for the period of observation – and recalculate the mean wage changes accordingly. An alternative (examined in appendix B) is to use the median wage (or some other percentile), assuming that non-workers in the skill group would be paid less than the median wage for the group.

A third strong assumption is the linearity of equation (6). Note that this is testable, since it implies that the observed relative wage changes (and relative employment changes) across demographic groups will also be linear in the observed demand proxies. We present some graphical evidence on this below. In principle, however, higher-order terms of the demand shift proxies can be easily added to equation (6) and will lead to a parallel set of higher-order terms in equations (7a) and (7b).

3.2. Effect of downward wage rigidity

Equations (7a) and (7b) are derived under the assumption that wages can freely adjust to demand and supply shocks. To analyse the effects of relative wage rigidity, denote the optimal wage change specified by equation (7a) as $\Delta \log w_j^*$. Assume that institutional rigidities or other constraints lead to an *actual* wage change:

$$\Delta \log w_j = \lambda_j \, \Delta \log w_j^*, \tag{9}$$

where $0 < \lambda_j \leqq 1$ represents a flexibility parameter. If wages are more flexible upward than downward, λ_j may vary with $\Delta \log w_j^*$. A simple and plausible assumption is that institutional forces resist downward wage pressure but permit upward adjustments: in this case $\lambda_j = 1$ if $\Delta \log w_j^* > 0$, and $\lambda_j = \lambda$ if $\Delta \log w_j^* < 0$.

If wages are upward flexible but downward rigid, employment is either in equilibrium or determined by the demand side of the market. Assuming that the relative demand shocks are related to the observable proxy variables by equation (6) and that the initial wage is in equilibrium, equations (7a) and (7b) will be replaced by

$$\Delta \log w_j = d_3 + \frac{\lambda_j \beta}{\sigma + \varepsilon} D_j - \frac{\lambda_j}{\sigma + \varepsilon} \Delta \log f_{jt} + e_{j3} \tag{10a}$$

$$\Delta \log p_j = d_4 + \frac{\beta(\varepsilon + \sigma(1 - \lambda_j))}{\sigma + \varepsilon} D_j - \frac{\varepsilon + \sigma(1 - \lambda_j)}{\sigma + \varepsilon} \Delta \log f_{jt} + e_{j4}, \tag{10b}$$

where d_3 and d_4 are constant across all skill groups and e_{j3} and e_{j4} are stochastic terms. Comparisons of equations (10a) and (10b) with (7a) and (7b) show that downward rigidities dampen the responsiveness of wages to negative demand shocks and accentuate the responsiveness of employment to such shocks.

Let $\bar{\lambda}$ represent the mean of λ_j across skill groups over a particular time interval. Equations (10a) and (10b) can be rewritten as

$$\Delta \log w_j = \pi_{10} + \pi_{11} D_j + \pi_{12} \Delta \log f_{jt} + u_{1j} \tag{11a}$$

$$\Delta \log p_j = \pi_{20} + \pi_{21} D_j + \pi_{22} \Delta \log f_{jt} + u_{2j}, \tag{11b}$$

where $\pi_{11} = \bar{\lambda}\beta/(\sigma + \varepsilon)$, $\pi_{12} = -\bar{\lambda}/(\sigma + \varepsilon)$, $\pi_{21} = \beta(\varepsilon + \sigma(1 - \bar{\lambda}))/(\sigma + \varepsilon)$, $\pi_{22} = -(\varepsilon + \sigma(1 - \bar{\lambda}))/(\sigma + \varepsilon)$, $u_{1j} = e_{j3} + (\lambda_j - \bar{\lambda})/(\sigma + \varepsilon) \times \{\beta D_j - \Delta \log f_{jt}\}$, and $u_{2j} = e_{j4} - \sigma(\lambda_j - \bar{\lambda})/(\sigma + \varepsilon) \times \{\beta D_j - \Delta \log f_{jt}\}$.

Our empirical strategy is to estimate models like (11a) and (11b) using changes in wages and employment rates across skill groups in the United States, Canada, and France, and the two alternative demand proxies described above. Our primary hypothesis is that over the 1980s $\bar{\lambda}$ was smaller (wages were on average more downward rigid) in France and Canada than in the United States. According to equations (11a) and (11b), we would therefore expect that relative wage changes would be *less sensitive* to relative demand shifts in Canada and France than in the United States (i.e., π_{11} would be smaller in magnitude than it would be in the United States), while relative employment changes in Canada and France would be *more sensitive* to relative demand shifts (i.e. π_{21} would be larger in magnitude than it would be in the United States).[19]

It is worth emphasizing that the validity of any comparisons of the reduced-form coefficients from equations (11) across the three countries hinges on the assumption that the mapping between relative demand shocks and the observed demand proxies

19 Note that OLS estimation of equation (11a) will tend to yield downward-biased estimates of the coefficient π_{11}, since the demand index D_j is likely to be negatively correlated with the residual component $(\lambda_j - \bar{\lambda})/(\sigma + \varepsilon) \times \{\beta D_j - \Delta \log f_{jt}\}$ if wages are downward rigid. A similar argument applies to the estimate of π_{21}.

is similar in the three countries, that is, that the coefficient β in equation (6) is the same across countries. In the case of the computer usage proxy, we believe this is a reasonable assumption, since, as shown in table 1, overall rates of computer use are similar in the three countries and since the degree of complementarity between high-skilled workers and computer technology is presumably similar across countries.[20] In the case of the initial wage proxy the assumption may be more problematic, since wage differentials across skill groups are typically higher in the United States than in other countries (Freeman and Katz 1995). If the three economies had the same relative demand shocks across skill groups but U.S. wages varied more across skill groups, then the β coefficient that links initial wage levels to subsequent demand shocks is probably *smaller* in the United States than in Canada or France. This will make it harder for us to find relative wage compression (and relative employment divergence) in Canada and France vis-à-vis the United States when we use initial wage levels as our relative demand proxy.

3.3. Microdata on wages and employment

We have assembled labour market survey data for men and women from the beginning and end of the 1980s for the United States, Canada, and France. In order to control for the effect of cyclical factors on wages and employment, the beginning and end years are chosen to match the peaks of the business cycle in each country. The two peaks of the U.S. business cycles are 1979 and 1989, with unemployment rates of 5.8 and 5.3 per cent, respectively. In the case of France and Canada, we can closely – but not perfectly – match the timing of the business cycle because of data limitations. The beginning and end years are 1981 and 1988 in Canada, with unemployment rates of 7.5 and 7.8 per cent, respectively. The beginning and end years are 1982 and 1989 in France, with unemployment rates of 8.2 and 9.3 per cent, respectively. In appendix C, we present some evidence on how sensitive the U.S. results are to the choice of beginning and end years.[21]

Our samples include individuals age 16–65 in the United States, age 17–64 in Canada, and age 15–60 in France. We further limit our U.S. samples to whites, in order to abstract from issues of discrimination and/or differential productivity by race. A complete description of our data sources and selection criteria is presented in appendix A.

The surveys for each country measure employment status as of a particular survey week. In all cases, the wage measure is a 'point-in-time' estimate of the

20 Berman, Bound, and Machin (1998) make a similar argument.

21 Appendix C shows that the U.S. employment results are sensitive to the choice of years and to cyclical factors. For instance, the decline in the relative employment of less-skilled workers all happened between 1979 and 1981 when the unemployment rate increased from 5.8 to 7.6 per cent. Relative employment of less-skilled workers improved between 1981 and 1988 as unemployment went from 7.6 to 5.5 per cent. We think that the general improvement of the U.S. economy over this period (which tends to help less-skilled workers) offsets the impact of other sources of change in the relative demand for less-skilled workers (computer technology, etc.). In any case, since unemployment declined in the United States in our sample period but increased in both Canada and France, our choice of beginning and ending years will, if anything, bias the results in favour of the tradeoff hypothesis.

usual hourly wage for an individual's main job.[22] This measure differs from the one typically used in the literature, which is constructed from annual earnings and weeks of work over the year. Nevertheless, comparisons of alternative wage measures for the U.S. labour market reported in Card and Lemieux (1996) suggest that the two types of measures give similar estimates of the levels and changes in hourly wages. In our U.S. and Canadian data sets, wages are recorded before taxes, whereas in our French data sets, wages are recorded net of employee payroll tax contributions. Since the French payroll tax is a fixed proportion of gross earnings, this difference should not affect *relative* pay comparisons.

Tables 4 and 5 present the overall levels and changes in mean log wages and employment rates estimated from our samples. As shown in row 8 of table 4, mean log wages of men fell by 7.5 per cent in our U.S. sample, rose by 3.9 per cent in our Canadian sample, and fell by 2.7 per cent in our French sample. For women, average real wage growth was slightly positive in the United States and Canada and negative in France. The measured decline in real wages for French workers is due to the use of an after-tax wage. Between 1982 and 1989 the employees' portion of the French payroll tax increased by 5 percentage points, from 13 to 18 per cent. Thus, the measured declines in real after-tax wages between 1982 and 1989 are consistent with roughly 3 per cent increases in real before-tax wages – a trend comparable to other wage series.

The entries in row 9 of table 4 show that the standard deviation of log wages grew by 15–20 per cent between 1979 and 1989 in the United States, while this measure of wage inequality was roughly constant in Canada and rose only slightly in France. Rows 10–12 decompose the changes in the cross-sectional variances of wages in each country into a component attributable to changes in wage dispersion within narrowly defined age-education cells and a component attributable to changes in between-cell dispersion. The decompositions suggest that the sharp rises in wage inequality in the United States were about equally attributable to rising inequality within and between age-education groups. In Canada, the variances of wages between age-education cells rose only slightly less than it did in the United States. Rising wage inequality between groups was counteracted, however, by a decline in the within-cell variance of wages in Canada, leading to little net increase in overall dispersion.

Although it is possible that wage variation within age-education cells actually fell in Canada over the 1980s, an alternative explanation is that changes in survey procedures lowered the variability of survey-response errors.[23] Unlike our U.S. and French samples, our Canadian samples are based on slightly different survey in-

22 The French surveys collect usual monthly earnings and average hours per week, which we use to construct an hourly wage. The U.S. surveys collect usual hourly wages or usual weekly wages and usual hours per week, which we use to form an estimate of usual hourly pay. The two Canadian surveys ask about wage information for jobs held in the previous year, rather than currently. In all three countries, wages are unavailable for self-employed workers.

23 Since the age-education cells in Canada contain a wider range of ages and education levels than those in the United States do, one might have actually expected a bigger increase in within-cell dispersion in Canada.

TABLE 4
Changes in the level and dispersion of hourly wages of male and female workers in the United States, Canada, and France during the 1980s

	Men			Women		
	United States	Canada	France	United States	Canada	France
Base year data						
1. Mean log wage	1.857	2.130	3.398	1.467	1.867	3.246
2. Exp(mean log wage)	6.403	8.416	29.918	4.336	6.469	25.687
3. Standard deviation	0.491	0.504	0.434	0.417	0.505	0.446
End year data						
4. Mean log wage	2.317	2.534	3.711	2.024	2.256	3.553
5. Exp(mean log wage)	10.142	12.599	40.891	7.569	9.545	34.918
6. Standard deviation	0.569	0.510	0.454	0.515	0.496	0.459
Change from base to end year						
7. Mean log wage	0.460	0.403	0.312	0.556	0.389	0.307
8. Mean log real wage	−0.075	0.039	−0.027	0.021	0.025	−0.032
9. Standard deviation	0.077	0.006	0.020	0.101	−0.010	0.013
Decomposition of change in wage inequality[a]						
10. Change in variance	0.082	0.006	0.018	0.094	−0.010	0.012
11. Change in within-cell component	0.043	−0.030	0.009	0.050	−0.033	0.008
12. Change in between-cell component	0.039	0.036	0.009	0.044	0.023	0.004

NOTES
Wage measures refer to pre-tax hourly wages in the United States and Canada, (in U.S. and Canadian dollars, respectively) and the post-tax hourly wage in France (in French francs). U.S. data include whites only. Samples include individuals age 16–64 in the United States, age 17–64 in Canada, and age 15–60 in France.

a Decomposition uses 225 age-education cells in the United States, 29 cells in Canada, and 70 cells in France.

struments in 1981 and 1988. Nevertheless, we believe that changes in the Canadian survey should not have affected the estimated means by skill group, which implies that our comparisons between skill groups are valid. We conclude that the variation in wages across different age-education groups increased in the United States and Canada over the 1980s and was roughly constant in France.

Turning to employment outcomes, the first two rows of table 5 present average employment-population rates for the men and women in our samples at the beginning and end of the 1980s. As shown in row 3, male employment-population rates declined in all three countries during the decade, while female employment-population rates increased. Other things being equal, changing age and education distributions over the 1980s would have been expected to lead to rising employment rates for both genders in all three countries. When we control for the effect of these changes (based on differences in the cross-sectional patterns by age/education at the start of the 1980s) the declines in employment for men are even more pronounced,

TABLE 5
Changes in employment-population rate of males and females in the United States, Canada, and France during the 1980s

	Men			Women		
	United States	Canada	France	United States	Canada	France
Employment trends in sample[a]						
1. Employment-population rate in base year	82.8	81.9	85.3	55.0	54.9	58.5
2. Employment-population rate in end year	82.5	81.1	82.9	64.3	62.8	59.8
3. Change in employment-population rate	−0.3	−0.8	−2.4	9.3	7.9	1.3
4. Change in employment-population rate, controlling for shifts in age and education[b]	−2.6	−2.8	−3.6	7.5	5.3	−0.3
Aggregate comparisons[c]						
5. Change in aggregate employment rate for gender group	−1.3	−2.0	−4.6	6.8	5.2	−0.1

NOTES
Entries are employment-population ratios, in per cent.
a The U.S. sample includes whites age 16–65. The Canadian sample includes people age 17–64. The French sample includes people age 15–60.
b The change in employment-population rate, adjusting for expected change based on changing age-education distribution of the population and cross-sectional employment patterns by age-education group in the base year.
c Changes in overall employment rates (for gender-specific population age 16 and older in the United States age 15 and older in Canada and France). U.S. data are from *Economic Report of the President* (1994, table B-37). Canadian data are from *Quarterly Economic Review* (June 1991, table 27). French data are from *Annuaire Statistique de la France* (1990, tables C.01-1 and C.01-2).

while the increases for women are reduced – see row 4.[24] For comparative purposes, we also report the overall changes in employment-population rates by gender over the same sample periods in row 5. These tend to be fairly similar to the adjusted changes from our sample.[25]

In our empirical analysis we aggregate the micro-level data for the three countries into age-education cells and compare changes over time in the employment rates

24 A simple decomposition of the effect of changes in the age/education composition on the employment-population ratio – the 'explained' change in the employment rate – indicates that, for men, changes in age composition plays a bigger role than changes in education. In all three countries, changes in the age composition account for roughly two-thirds of the explained change in the employment rate. By contrast, for women, improvements in the level of education account for roughly 80 per cent of the explained change in employment rates in all three countries.

25 Most of the discrepancy between the two series is due to the fact that older individuals (above 60 in France, above 64 in Canada, and above 65 in the United States) are included in the calculation of the overall rates but are excluded from our samples. This is particularly important for France, where the employment rate of men age 60–64 fell drastically during the 1980s, in part because the normal retirement age was reduced from 65 to 60.

and mean wages of different cells. Each age-education cell corresponds to a skill group j in the theoretical framework of section 3.1. Note that our empirical analysis focuses only on between-group variation in wages and employment. Unlike studies such as Juhn (1992), our approach does not utilize within-group variation in wages and employment to evaluate the implicit tradeoff between these two variables. Since more than half of increase in wage inequality in the United States is due to the within-group component (table 4), we may be ignoring an important part of the changes in the relative structure of wages and employment that went on during the 1980s.[26]

Figures 2–4 illustrate some of the basic patterns in the cell-level data. The figures plot the changes in wages and employment rates for each cell against the mean wage of the cell at the beginning of the 1980s. For reference, we have added a line to each figure representing the increase in nominal wages that would have been required to maintain a constant real wage.

As shown in figure 2, the U.S. data show a strong positive correlation between the wage growth experienced by a given age-education group over the 1980s and the initial level of wages of the group in 1979. Among men (in the upper panel) only a few cells had average wage increases large enough to maintain the real value of their earnings. More of the female age-education groups had nominal wage growth in excess of inflation. On the employment side, the scatter of points in the upper panel of figure 2 shows a much weaker correlation between the initial level of wages for different groups of men and the change in their employment rate, while the correlation across different groups of women (in the lower panel) is stronger.

The data for Canadian men and women in figure 3 show some similarities with the U.S. data. For both genders, age-education groups with higher initial wages tended to have faster wage growth over the 1980s, although relative to U.S. men and women the correlation is less pronounced. The pattern of employment changes across groups is also less systematic in Canada. Indeed, the employment data in the upper panel of figure 3 suggest that groups of lower-wage men in Canada tended to have bigger employment gains than groups of higher-wage men. Notice, however, that this pattern is driven by a few low-wage cells: these are made up of younger and less-educated workers who experienced large wage losses and significant employment gains over the 1980s.

Figure 4 shows that the patterns of relative wage changes in France over the 1980s were quite different from those in the United States or Canada. Unlike their counterparts in North America, French workers with relatively low wages at the beginning of the 1980s had about the same rate of wage growth as those with relatively high wages. Assuming that French employers were affected by the same demand shocks as employers in the United States and Canada, this pattern suggests substantial relative wage inflexibility. On the other hand, the patterns of employment

26 In principle, it is possible to use both between- and within-group variation in wages and employment to evaluate the implicit tradeoff between these two variables. We focus only on between-group variation to keep the theory and the econometrics simpler.

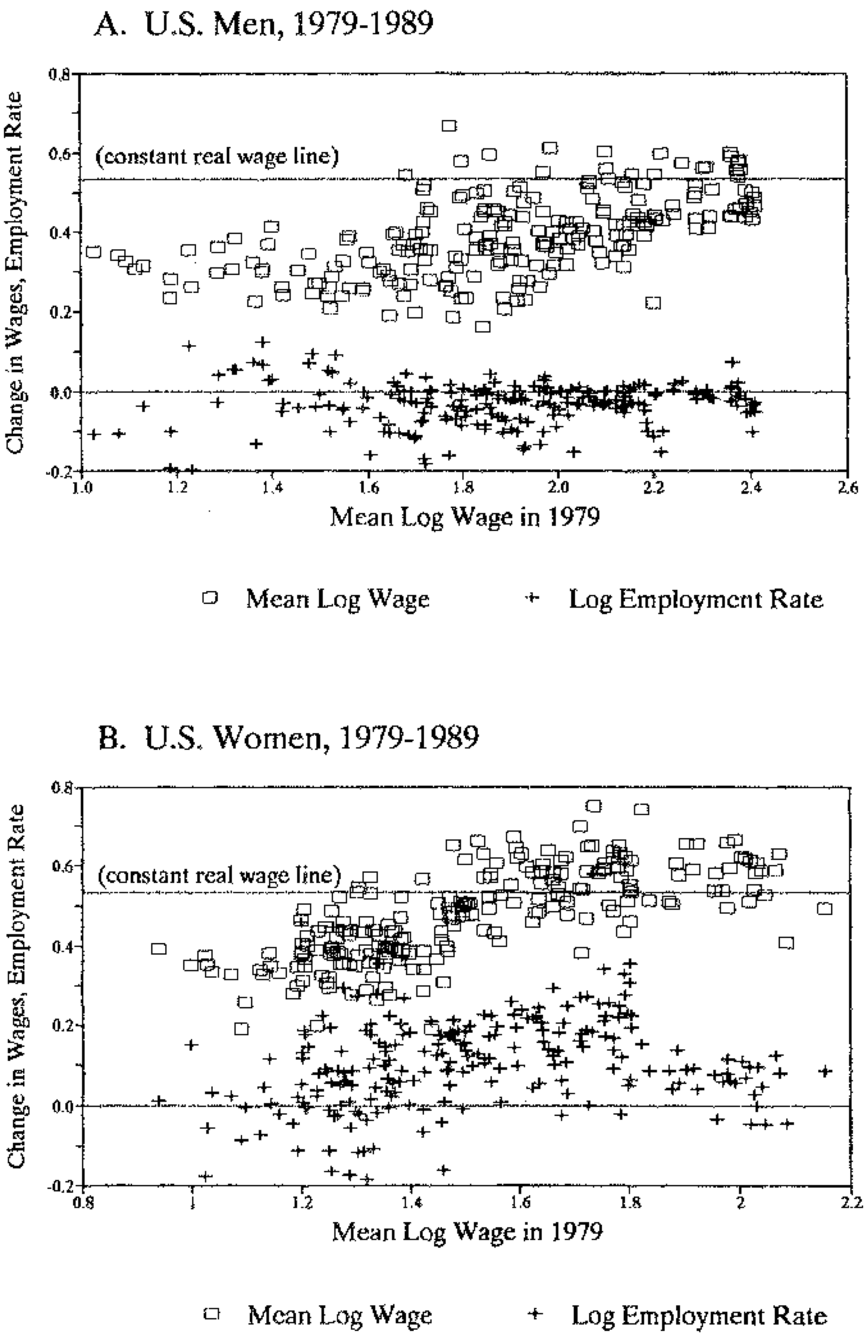


FIGURE 2 Changes in employment and wages over the 1980s: men and women in the United States

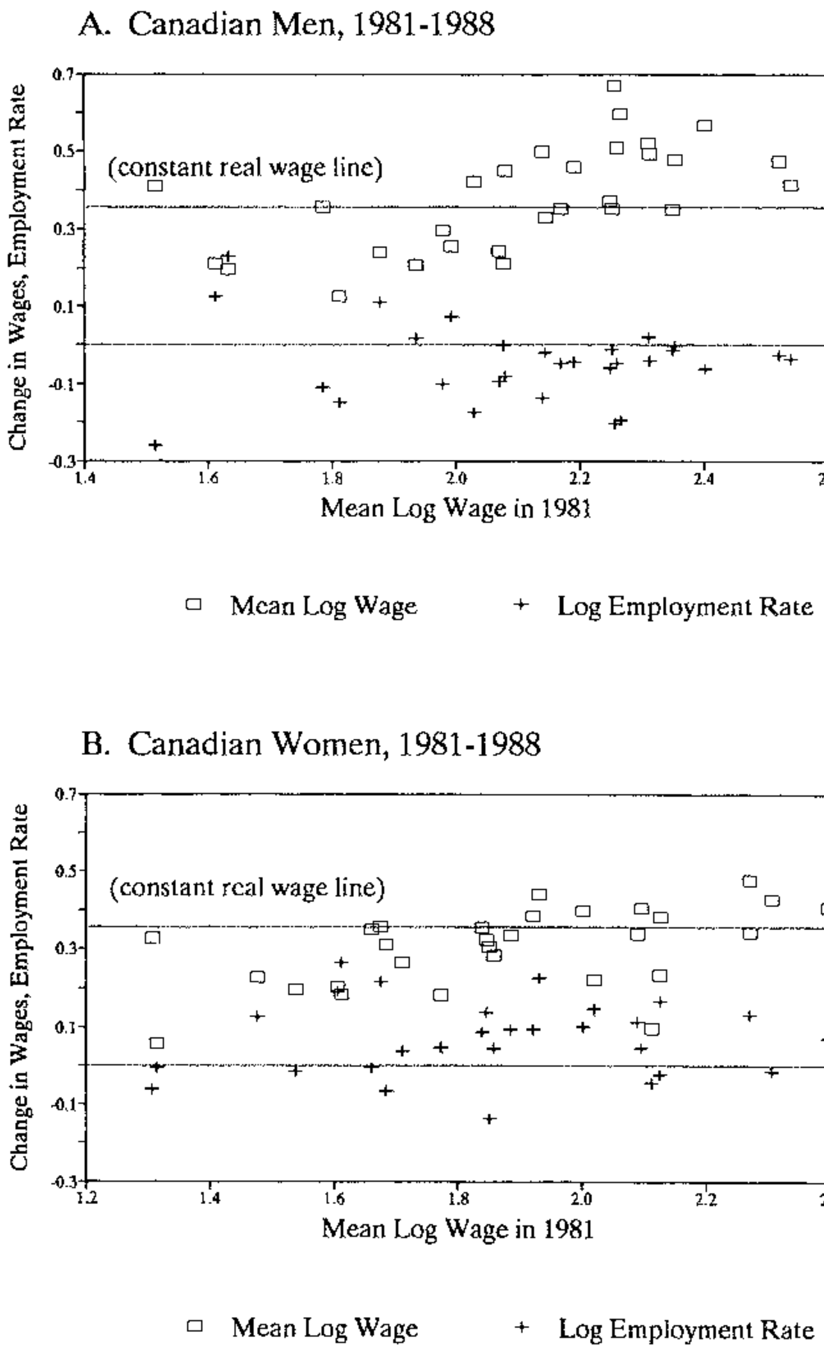


FIGURE 3 Changes in employment and wages over the 1980s: men and women in Canada

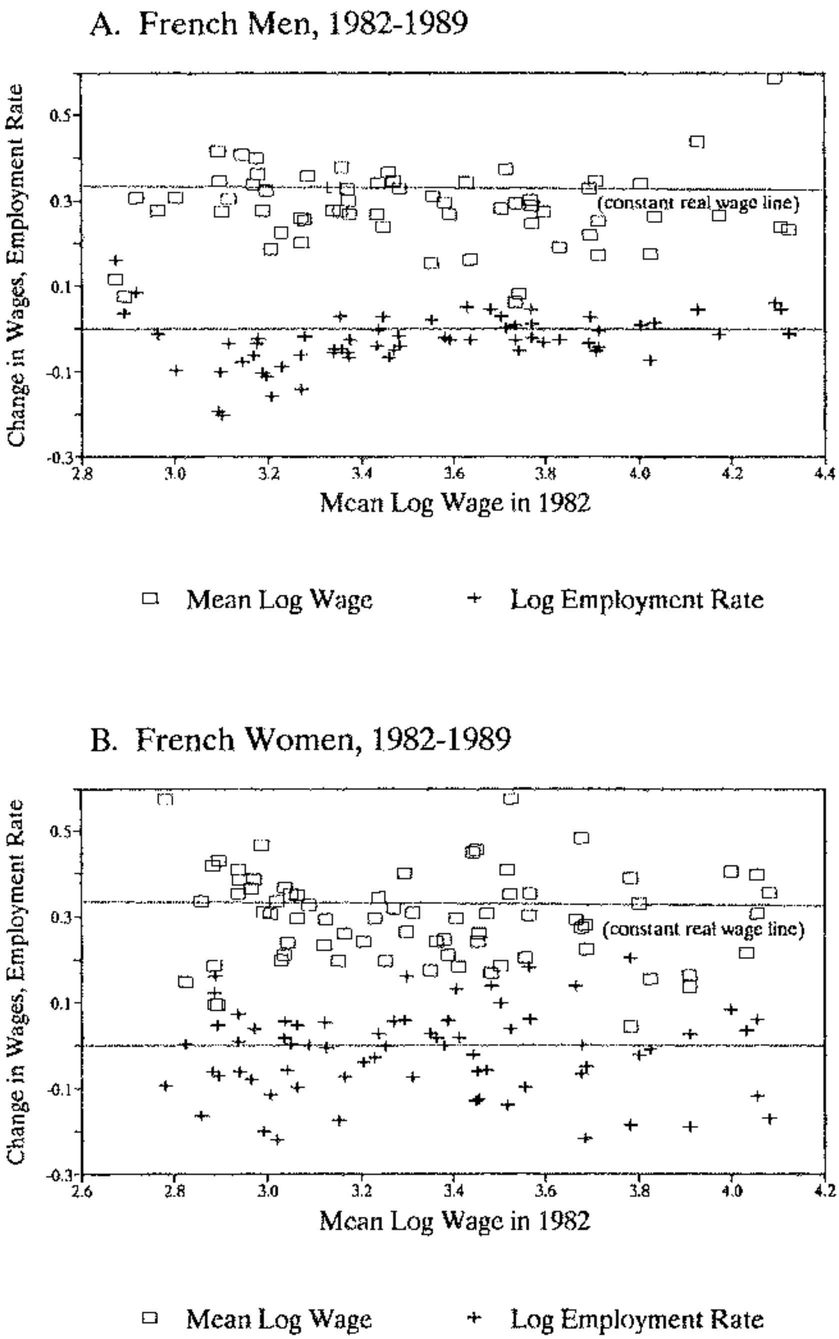


FIGURE 4 Changes in employment and wages over the 1980s: men and women in France

TABLE 6
Estimated regression models for changes in mean cell wages and employment-population rates: United States, Canada, and France

	Using mean cell wage in base year as demand shift index			Using average computer use rate in late 1980s as demand shift index		
	Mean log wage			Mean log wage		
	Raw (1)	Adjusted (2)	Log emp. rate (3)	Raw (4)	Adjusted (5)	Log emp. rate (6)
Panel A: Men						
United States 1979–89						
1. Demand shift index	0.17	0.19	0.04	0.31	0.41	0.13
	(0.02)	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)
2. Change in log population share	−0.01	0.03	0.03	−0.01	0.01	0.01
	(0.02)	(0.02)	(0.02)	(0.02)	(0.01)	(0.02)
3. R-squared	0.34	0.41	0.08	0.46	0.68	0.16
Canada 1981–88						
1. Demand shift index	0.26	0.09	−0.30	−0.14	−0.00	−0.04
	(0.10)	(0.06)	(0.08)	(0.12)	(0.07)	(0.03)
2. Change in log population share	0.06	0.13	0.16	0.26	0.18	0.05
	(0.08)	(0.05)	(0.07)	(0.08)	(0.04)	(0.09)
3. R-squared	0.43	0.55	0.34	0.31	0.51	0.01
France 1982–89						
1. Demand shift index	0.01	−0.02	0.03	−0.11	−0.10	0.11
	(0.04)	(0.03)	(0.04)	(0.06)	(0.06)	(0.06)
2. Change in log population share	−0.01	−0.00	0.05	0.01	0.01	0.04
	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)	(0.03)
3. R-squared	0.00	0.01	0.08	0.04	0.04	0.12
Panel B: Women						
United States 1979–89						
1. Demand shift index	0.30	0.38	0.12	0.18	0.38	0.13
	(0.03)	(0.03)	(0.04)	(0.03)	(0.03)	(0.04)
2. Change in log population share	0.07	0.12	0.07	0.14	0.15	0.08
	(0.02)	(0.02)	(0.03)	(0.02)	(0.02)	(0.02)
3. R-squared	0.66	0.73	0.19	0.53	0.74	0.20
Canada 1981–88						
1. Demand shift index	0.14	0.17	−0.06	−0.08	0.13	0.23
	(0.09)	(0.09)	(0.16)	(0.11)	(0.10)	(0.17)
2. Change in log population share	0.02	0.05	0.09	0.13	0.09	−0.09
	(0.06)	(0.06)	(0.10)	(0.06)	(0.05)	(0.09)
3. R-squared	0.29	0.51	0.04	0.25	0.47	0.10
France 1982–89						
1. Demand shift index	−0.12	−0.14	−0.04	−0.29	−0.29	−0.01
	(0.04)	(0.03)	(0.05)	(0.05)	(0.04)	(0.08)
2. Change in log population share	0.08	0.08	0.11	0.09	0.07	0.10
	(0.03)	(0.02)	(0.03)	(0.02)	(0.02)	(0.03)
3. R-squared	0.18	0.26	0.15	0.33	0.40	0.14

TABLE 6 *(Concluded)*

NOTES
Standard errors are in parentheses.
All models are estimated by weighted least squares on cell-level data. The U.S. sample includes 225 age-education cells; the Canadian sample includes 29 age-education cells; the French sample includes 70 age-education cells. Cell weight is the fraction of the adult population in the cell in the base year. U.S. data include only whites.
In columns 1 and 4, the dependent variable is the change in the mean cell wage between the base year and the end year. In columns 2 and 5, the dependent variable is the change in the mean cell wage, assuming that those who don't work would earn the minimum wage in the respective year. In columns 3 and 6, the dependent variable is the change in the log of the employment-population rate in the cell.

growth across different age-education cells in France are fairly similar to those in the United States or Canada.

3.4. Models of relative changes in wages and employment rates

Table 6 presents estimates of equations (11a) and (11b), fitted to the cell-level data for each country using two alternative proxies for the relative demand shocks faced by different age-education groups: the mean log wage of the group at the beginning of the 1980s and the fraction of the group who used a computer at the end of the 1980s. For each demand proxy we present two estimates of equation (11a): one using the actual change in mean log wages of workers in the cell (columns 1 and 4) and another using an adjusted wage measure that assigns the minimum wage to non-workers (columns 2 and 5).[27] All the models are fitted by weighted least squares, using as a weight the fraction of the adult population in the age-education cell in the base period.[28]

The models in column 1 of table 6 regress the changes in mean cell wages for a given country and gender group on the initial cell wage and the change in the population fraction of the group. The coefficients of the mean cell wage variable are large and positive for U.S. and Canadian men and women, but are small and positive for French men (0.01) and negative for French women (−0.12). The signs and magnitudes of the estimated coefficients are not too different when we account for potential selectivity biases by assigning the minimum wage to non-workers in each cell (column 2).

The models in column 3 regress the change in the employment-population rate on the same explanatory variables. For U.S. men the coefficient is positive (0.04) and marginally significant, while for U.S. women the coefficient is larger (0.12) and highly significant. Together with the estimates in column 1, these estimates support the conventional view that relative demand shocks reduced the demand for less-skilled workers of both genders, leading to relative declines in wages and

27 We have also estimated the wage models using the 50th and the 75th percentiles of wages for each cell (constructed over workers and non-workers, assuming that non-workers are below the median). The estimates (summarized in appendix table B1) are very similar to those in table 6.

28 The parameter estimates tend to be very similar regardless of whether the estimates are obtained by unweighted least squares or by weighted least squares with any of a variety of cell weights.

employment. According to equations (7a) and (7b), the ratio of the coefficients of the relative demand index in models for the change in log employment and the change in log wages is an estimate of the labour supply (participation) elasticity. The estimates for U.S. men imply an elasticity of 0.21–0.24, while the estimates for U.S. women imply an elasticity of 0.32–0.40.

The estimates in column 3 for Canada and France are much harder to interpret. Even though higher-wage workers in Canada had larger relative wage gains over the 1980s, the change in employment is insignificantly or even negatively related to the base wage level. In France, where relative wage growth was fairly even across skill groups, the change in employment is also only weakly related to the base wage level. There is no evidence that the rigidity of French relative wages gave rise to 'excess' employment losses for workers whose relative wages would have been predicted to fall on the basis of U.S. patterns.

The models in columns 4–6 of table 6 regress the changes in wages and employment rates on the alternative demand index variable based on observed computer use rates at the end of the 1980s. The general patterns and implications of the estimated coefficients are not too different from those in columns 1–3. Among U.S. men, for example, the coefficients relating wage and employment changes to computer use rates are slightly larger than the coefficients relating these outcomes to the base wage level, but their ratio is similar, implying a similar supply elasticity.

Experiments with other specifications revealed that the estimated demand-index coefficients for the models in table 6 are sensitive to two issues: whether or not cells of very young and very old workers are included[29] and whether or not the change in population share is included as an additional control variable. Our theoretical model suggests that age-education cells that increased their relative population shares should have had slower wage and employment growth rates. Contrary to this prediction, the estimated coefficients of the population share variables are mainly positive, and over one-half of the estimates are significant. This constitutes fairly strong evidence against our simple theoretical model.

One explanation for the 'wrong sign' on the supply-side shift variables is that the wages of a narrow skill group may depend on a broader index of supply than the relative population share of the group. In a more general model, workers in a given age-education cell may be good substitutes for workers with similar levels of age and education, in which case the relative population shares of the substitute groups must also be included in the reduced-form employment and wage models. An alternative explanation is that supply changes are an endogenous response to (unmeasured) changes in demand conditions.

Table 7 presents an alternative set of estimates of the wage and employment change models, excluding the population share variable.[30] We present estimates for

29 This sensitivity is particularly notable for the Canadian models, which are estimated on relatively few cells.

30 A closer examination of equations 10 and 11 shows that the relative bias due to the omission of the population share variable will be the same in the three countries provided that β and the coefficient of a regression of the demand shift variable on the population share variable are the

TABLE 7
Estimated models for men and women with computer use and base year wage as alternative measures of relative demand shocks

	Using mean cell wage in base year as demand shift index			Using average computer use rate in late 1980s as demand shift index		
	Mean log wage			Mean log wage		
	Raw (1)	Adjusted (2)	Log emp. rate (3)	Raw (4)	Adjusted (5)	Log emp. rate (6)
United States 1979–89						
1. All men	0.17	0.20	0.06	0.30	0.41	0.14
	(0.02)	(0.02)	(0.01)	(0.02)	(0.02)	(0.02)
2. Men age 25–54 only	0.27	0.31	0.06	0.33	0.39	0.08
	(0.03)	(0.04)	(0.01)	(0.02)	(0.02)	(0.01)
3. All women	0.37	0.50	0.19	0.31	0.52	0.21
	(0.02)	(0.02)	(0.03)	(0.03)	(0.03)	(0.03)
Canada 1981–88						
1. All men	0.31	0.20	−0.16	0.11	0.17	0.05
	(0.07)	(0.05)	(0.06)	(0.11)	(0.07)	(0.09)
2. Men age 25–54 only	0.22	0.28	0.11	0.07	0.13	0.08
	(0.13)	(0.12)	(0.05)	(0.10)	(0.09)	(0.03)
3. All women	0.16	0.24	0.05	0.12	0.27	0.16
	(0.05)	(0.05)	(0.08)	(0.07)	(0.06)	(0.10)
France 1982–89						
1. All men	0.00	−0.02	0.05	−0.10	−0.09	0.15
	(0.03)	(0.03)	(0.03)	(0.06)	(0.05)	(0.06)
2. Men age 25–54 only	−0.07	−0.03	0.09	−0.18	−0.10	0.12
	(0.03)	(0.04)	(0.02)	(0.05)	(0.06)	(0.02)
3. All women	−0.06	−0.08	0.04	−0.20	−0.22	0.09
	(0.03)	(0.03)	(0.04)	(0.05)	(0.04)	(0.07)

NOTES
Standard errors are in parentheses. See notes to table 6. In columns 1–3, the (sole) independent variable is the mean cell wage in the base year. In columns 4–6, the (sole) independent variable is the proportion of cell workers using computers at work at the end of the 1980s (1989 in the United States and Canada, 1991 in France).

both sexes using all the available age-education cells and using only the subset of men age 25–54.

Generally speaking, the results reported in table 7 follow a more stable pattern than those in table 6. As expected, demand shift variables have a positive effect on employment in all but one of the eighteen specifications reported in table 7. By contrast, the effect is negative in almost half of the models reported in table 6. This suggests that including a 'wrong signed' supply variable does more harm than good to the other coefficients of the model.

same in the three countries. If these conditions hold, excluding the population share variable should not bias the intercountry comparisons.

The results are even clearer for models that involve only prime-age males. Focusing on this group is useful, since, relative to other workers, their employment patterns are less likely to be affected by factors 'outside' the simple demand and supply model.[31] Columns 2 and 5 of table 7 show that the effect of demand shifts on wages range from large and positive for the United States to not significant for France, with Canada falling somewhere between. This suggest, once again, that wages are much more flexible in the United States than in France. If anything, wages in Canada are also less flexible than in the United States – especially when demand shocks are measured by computer usage – but more flexible than in France.

By contrast, the effect of demand shifts on employment is of the same order of magnitude in all three countries. There is, nevertheless, some weak evidence in favour of the 'tradeoff hypothesis' for these samples of prime-age males. In both column 3 and column 6, the point estimates of the employment effects are higher in France than in the United States, and this difference is marginally significant.[32]

Taken at face value, however, the tradeoff is quite small. Take, for example, the model in which the computer use rate is used as demand shift index, which is arguably our 'best' specification. The U.S.-France difference in the wage effects is 0.50 (column 5) relative to a (minus) 0.04 difference in employment effects (column 6). This means that relative wages have to decrease by more than 10 per cent for relative employment rates to increase by 1 per cent.

We reached a similar conclusion in a previous version of the paper, where we estimated the structural parameters of the model under the assumption that the flexibility parameter λ_j was the same for all age-education groups in the same country (Card, Kramarz, and Lemieux 1996). The structural estimate of the elasticity of substitution σ was equal to 0.16 (standard error of 0.06), suggesting only a small tradeoff.

The evidence for the other groups (all men and all women) in table 7 shows no evidence of a tradeoff. For the sample of all men, the effect of the demand shift variables on employment is almost identical in France and in the United States. For the sample of all women, this effect is significantly larger in the United States than in France, indicating that the relative employment rate of less-skilled women fell more in the United States than in France. Furthermore, the effect of the demand shift variables on employment is always smaller in Canada than in the United States (for all men and all women).

Overall, there is at best some weak evidence in favour of the 'tradeoff hypothesis.' Prime-age males comprise the only group for which the large drop in relative

31 Schooling and retirement decisions are important and should probably be modelled for younger and older workers, respectively. It is also hard to explain the rapid growth in female employment rates by standard supply and demand factors.

32 The effect of demand shift variables on employment is more precisely estimated for prime-age males than for other groups. It is easier, therefore, to reject the null hypothesis that parameter values are the same in all countries for this group of workers. In column 6 of table 7, for example, the critical value of the test of difference between France and the United States is 0.04 for prime-age males, compared with 0.12 for all men, which is an economically large difference in parameter values.

wages of less-skilled workers in the United States has helped to reduce the fall in their relative employment rate (relative to France or Canada). But even for this group, the tradeoff is small: the relative drop in wages is much larger than the relative gains in employment.

Finally, since downward wage rigidity is more likely to bind for low-wage than high-wage workers, we re-estimated the models for two separate samples of low- and high-wage workers in France and the United States.[33] Perhaps surprisingly, the results were qualitatively similar for the two subsamples. This suggests that U.S.-France differences in wage flexibility are not driven by few low-wage cells at the minimum wage. Factors, such as collective bargaining, that affect a much broader range of workers probably play a major role too.

4. Conclusions

The writing of this paper is motivated by a very simple observation: in labour markets with rigid relative wages, negative employment demand shocks will lead to larger employment losses for the affected groups of workers than would be expected if wages could freely adjust. We test this 'tradeoff hypothesis' by comparing changes in wages and employment rates for different age and education groups over the 1980s in the United States, Canada, and France. We argue that the same forces that are generally believed to have lowered the real wages of less-skilled workers to the U.S. labour market have affected labour markets in Canada and France. In comparison to the U.S. labour market, however, Canada and especially France have a variety of institutional features that tend to prevent relative wage adjustments. We would therefore expect the relative structure of wages to have changed less in Canada and France than in the United States during the 1980s but the relative structure of employment to have changed more.

Using comparable micro-data from the beginning and end of the 1980s for each country, we relate changes in wages and employment-population rates for narrow age and education cells to two proxies for skill-group-specific relative demand shocks: the initial level of wages for the group, and the fraction of the group who were using computers at the end of the 1980s. In the United States, wage changes over the 1980s were highly correlated with these two variables: higher-wage groups and those that were more likely to adopt computer technologies enjoyed bigger wage gains, while lower wage groups and those that were less likely to adopt computers suffered significant real wage declines. We find a slightly weaker relationship between these two variables and wage growth in Canada and virtually no relation in France. Thus, relative wages appear to be slightly less flexible in Canada than in the United States and to be completely inflexible in France.

The patterns of relative employment changes provide, at best, some weak support to the view that wage flexibility moderates the employment losses for groups

33 High-wage cells are those whose base period wage is above the median base wage and vice versa. The analysis was not performed for Canada because of the small number of cells available.

affected by negative demand shocks. In the United States, the relative employment rates of lower-wage groups declined over the 1980s. As has been noted in previous work (Juhn 1992), the parallel trends in wages and employment rates for less-skilled workers are consistent with movements along an upward-sloping supply schedule. In Canada and in France, where one might have expected relative wage rigidity to lead to even bigger employment losses for low-wage workers, the patterns of relative employment growth over the 1980s are not systematically different from those in the United States. Prime-age males are the only group for which relative employment losses are (slightly) larger in France and Canada than in the United States. We conclude that there is at best some weak support for the 'tradeoff hypothesis.' Explanations for the poor employment performance of Western Europe during the 1980s have to be found elsewhere.

Appendix A: Data description

A.1. Wage and employment data

The U.S., Canadian, and French micro-data used in this paper are derived from supplements to the regular labour force surveys in each country. The surveys from all three countries provide a measure of usual hourly earnings for the respondent's main job last week (last month in France) for individuals who were working for pay and were not self-employed.

Our U.S. analysis is based on merged monthly files for the outgoing rotation groups of the 1979 and 1989 Current Population Surveys (CPS). These files contain approximately 275,000 observations per year. We include in our samples men and women age 16–65 whose potential labour market experience (age-education-6) is positive. The employment rate is based on labour force activity in the survey week of the CPS. Our wage measure is based on the reported hourly wage (for hourly rated workers) or the ratio of reported weekly earnings to average weekly hours (for others). To limit the influence of outliers we have deleted all wage observations below \$2.01 or above \$60.00 in constant 1989 dollars in the wage analysis. We have also excluded from our wage samples any allocated wage observations.

To implement the estimation methods described in section 2 we divided individuals of each gender into 225 individual age and education cells. The cells are based on single years of education (with $\leqq$ 8 years in the lowest cell and $\geqq$ 18 years in the highest cell) and one- two- or three-year age ranges (single-year age ranges for ages up to 23, two-year age ranges for age 24–43, and three-year age ranges for age 44 and older). We then computed the mean log wage, the employment rate, and other statistics for each cell.

Our analysis for France is based on data from the 1982 and 1989 Labour Force Survey (Enquête Emploi, EE) which is conducted every year by the French National Statistical Institute (see INSEE 1989 for technical details on the survey). The EE is a 1/300 sample of the French population, based on a three-year rotating panel. Seven panels of approximately 30,000 individuals age 15–60 are available over the 1981–89 period. We use a sample of 60,723 observations for 1982 (based on all individuals in their first or second interview year in 1982) and 29,198 observations for 1989 (based on all individuals in their third interview year in 1989).

The EE collects information on usual monthly earnings net of payroll taxes for the respondent's main job last month. The variable is divided into the nineteen categories: less than 1,000FF; from 1,000 to 1,500FF; from 1,500 to 2,000FF;...; from 4,500 to 5,000FF; from 5,000 to 6,000FF,...; from 9,000 to 10,000FF; from 10,000 to 15,000FF;...; from 25,000 to 30,000FF; above 30,000FF. Individuals who refuse to answer the earnings question are excluded from the wage samples.

Usual hourly earnings are obtained by dividing the midpoint of the earnings category by usual weekly hours times 4.33 (weeks per month). For both men and women, we divide the sample into seventy individual age and education cells, based on eight level-of-education qualifications (no diploma, primary school (CEP), junior high school (BEPC), vocational or technical school (CAP), academic high school (baccalauréat), technical baccalauréat, undergraduate studies (academic or professional), and graduate studies or grandes écoles) and five-year age ranges. We use coarser age-education cells in France than in the United States to allow a reasonable number of observations per cell.

Our Canadian analysis is based on Statistics Canada's 1981 Survey of Work History (SWH) and 1988 Labour Market Activity Survey (LMAS). The two surveys are retrospective supplements to the regular Labour Force Survey that were conducted in January and February 1982 (SWH) and 1989 (LMAS). The 1981 SWH and the 1988 LMAS yield samples of 61,066 and 58,860 individuals aged 16–64, respectively.

The SWH and the LMAS collect information on hourly or weekly wages and usual hours of work for each job in the previous year. We have edited the Canadian samples to obtain samples comparable to those collected in the CPS and EE. For each month (in the SWH) or week (LMAS) of the year we determined the main job (if any), based on reported hours for all jobs held in that month or week. Each main job held at some time in the previous year was then assigned a weight based on the fraction of the year that this job was the respondent's main job. The final sample is a sample of main jobs with weights that represent the probability that the job would appear in a sample of ongoing jobs conducted at regular intervals over the year. This 'weighted sample' is thus comparable to the samples obtained in CPS or the EE. We use a similar procedure to compute an employment rate comparable to the employment rate obtained in the CPS or in the EE.

A major limitation of the Canadian samples is that only broad age and education categories are available in the public use versions of the 1981 SWH and 1988 LMAS: five education categories (primary education or less; some or completed high school; some post-secondary education; post-secondary diploma or certificate; university degree); and six age ranges (17–19, 20–24, 25–34, 35–44, 45–54, and 55–64). Excluding university graduates age 17–19, we thus have twenty-nine age-education cells. Note, finally, that there is no earnings allocation flag in the public use samples. We thus keep allocated wage information in Canadian samples.

A.2. Data on computer use at work

The proportion of workers using computers at work in each age-education cell is obtained using U.S. data from a supplement to the April 1989 CPS, French data from the 1991 Enquête-Emploi, and Canadian data from the 1989 General Social Survey (GSS). Details on the CPS and the Enquête-Emploi are provided above. The GSS is a survey of approximately 10,000 individuals conducted by Statistics Canada. The 1989 GSS included several questions on technology at work, including the use of a computer. The age categories in the GSS, the SWH, and LMAS are identical, while education categories are more detailed in the 1989 GSS. Workers in the GSS thus can be divided into the same twenty-nine age-education cells used to divide the SWH and LMAS samples.

Appendix B

TABLE B1
Estimated regression models for changes in various selection adjusted and unadjusted measures of wages for men and women: United States, Canada, and France

	Models for changes in mean log wages:		Models for changes in 50th percentile:			Models for changes in 75th percentile:		
	Unadj.	Adj.	Unadj.	Adj.	Adj. and excl. (emp-pop)<0.5	Unadj.	Adj.	Adj. and excl. (emp-pop)<0.75
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
United States 1979–89								
1. All men	0.17	0.20	0.18	0.26	0.26	0.17	0.26	0.26
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
2. Men age 25–54 only	0.27	0.31	0.29	0.38	0.38	0.29	0.34	0.34
	(0.03)	(0.04)	(0.03)	(0.04)	(0.04)	(0.03)	(0.04)	(0.04)
3. All women	0.17	0.20	0.18	0.26	0.26	0.17	0.26	0.26
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Canada 1981–88								
1. All men	0.31	0.20	0.26	0.24	0.24	0.31	0.24	0.24
	(0.07)	(0.05)	(0.07)	(0.07)	(0.07)	(0.06)	(0.06)	(0.06)
2. Men age 25–54 only	0.22	0.28	0.13	0.20	0.20	0.04	0.06	0.06
	(0.13)	(0.12)	(0.12)	(0.13)	(0.13)	(0.09)	(0.11)	(0.11)
3. All women	0.31	0.20	0.26	0.24	0.24	0.31	0.24	0.24
	(0.07)	(0.05)	(0.07)	(0.07)	(0.07)	(0.06)	(0.06)	(0.06)

TABLE B1 *(Concluded)*

	Models for changes in mean log wages:		Models for changes in 50th percentile:			Models for changes in 75th percentile:		
	Unadj.	Adj.	Unadj.	Adj.	Adj. and excl. (emp-pop)<0.5	Unadj.	Adj.	Adj. and excl. (emp-pop)<0.75
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
France 1982–89								
1. All men	0.00	−0.02	0.01	−0.10	−0.10	0.00	−0.04	−0.04
	(0.03)	(0.03)	(0.03)	(0.05)	(0.05)	(0.03)	(0.04)	(0.04)
2. Men age 25–54 only	−0.07	−0.03	−0.02	−0.12	−0.12	−0.02	−0.02	−0.02
	(0.03)	(0.04)	(0.04)	(0.07)	(0.07)	(0.04)	(0.04)	(0.04)
3. All women	−0.06	−0.08	−0.07	−0.17	−0.17	−0.04	−0.15	−0.15
	(0.03)	(0.03)	(0.04)	(0.05)	(0.05)	(0.04)	(0.05)	(0.05)

NOTES

Standard errors are in parentheses. See notes to table 7. The independent variable in all models is the base-year mean log wage for the age-education cell.

In column 1, the dependent variable is the change in the mean cell wage between the base year and the end year. In column 2, the dependent variable is the change in the mean cell wage, assuming that those who don't work earn the minimum wage in the respective year. In column 3, the dependent variable is the change in the median wage among workers. In column 4, the dependent variable is the change in the median wage in the population, assuming that those who don't work earn the minimum wage. The same dependent variable is used in column 5, but cells in which less than 50 per cent of the population is employed in either the base or the end year are excluded. In columns 6–8 the dependent variable is as it is in columns 3–5, except that the median is replaced by the 75th percentile.

Appendix C

TABLE C1
Estimated models for U.S. men for alternative time periods

	Using mean cell wage in base year as demand shift index			Using average computer use rate in late 1980s as demand shift index		
	Mean log wage			Mean log wage		
	Raw (1)	Adjusted (2)	Log emp. rate (3)	Raw (4)	Adjusted (5)	Log emp. rate (6)
1. 1979–89	0.167	0.204	0.055	0.300	0.411	0.137
	(0.015)	(0.016)	(0.014)	(0.022)	(0.019)	(0.021)
2. 1979–88	0.197	0.213	0.065	0.314	0.416	0.158
	(0.016)	(0.016)	(0.014)	(0.024)	(0.019)	(0.021)
3. 1979–81	0.033	0.024	0.091	0.047	0.067	0.131
	(0.008)	(0.008)	(0.008)	(0.013)	(0.012)	(0.014)
4. 1981–88	0.163	0.189	−0.026	0.267	0.349	0.027
	(0.013)	(0.015)	(0.011)	(0.020)	(0.019)	(0.017)

NOTES
Standard errors in parentheses. See notes to table 7. In columns 1–3, the (sole) independent variable is the mean cell wage in 1979. In columns 4–6, the (sole) independent variable is the proportion of cell workers using computers at work in 1989.

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