THE LONG-TERM EFFECTS OF UNEMPLOYMENT INSURANCE: EVIDENCE FROM NEW BRUNSWICK AND MAINE, 1940-1991

Peter Kuhn, University of California, Santa Barbara

Chris Riddell, Queen's University

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Using data spanning half a century for adjacent jurisdictions in the U.S. and Canada, we study the long-term effects of a generous unemployment insurance (UI) program on the distribution of weeks worked. We find substantial effects. For example, in 1990, about 12.6 percent of working-age men in Maine's northernmost counties worked between 1 and 39 weeks; just across the border in New Brunswick that figure was 25.6 percent. According to our estimates, New Brunswick's much more generous UI system accounts for over three fourths of this differential. In part because part-year workers are drawn from both ends of the distribution of annual weeks, UI's estimated effects on total labor supply are modest, while its effects on UI program participation and expenditures are substantial.

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Most studies of the labor market effects of income support programs share the following features: they focus on short-term responses to relatively small changes in a single program parameter, in a restricted subpopulation of workers. For example, a typical research paper on the effects of unemployment insurance (UI) might examine the impact of the benefit replacement rate on unemployment durations using historical state-by-year variation in replacement rates for a sample of persons who have entered an unemployment spell. To avoid confounding the effects of such a policy change with business cycles and other factors, attention is typically restricted to measuring effects that occur within a year or less of the implementation of the policy change.[1]

For the most part, the above limitations are not only unavoidable, but desirable: tightly-defined situations are likely to yield more precise estimates of policy effects. However, these limitations have led a number of critics (e.g. Murray 1994) to argue that existing econometric studies could significantly understate the long-term work incentives of income support programs. According to such arguments, individuals may be poorly informed about small policy changes – the incentives to acquire this information are relatively low – and may choose not to make large behavioral adjustments if they expect program changes to be temporary. Over a longer time horizon, individuals eligible for a generous programs that they expect to be permanent may make larger behavioral adjustments (including, for example, entering the labor force to qualify for it), ultimately building a 'lifestyle' around the program.[2]

We estimate the long-term labor supply effects of large-scale changes in the overall generosity of an income support program on the entire working-age population of a region. Specifically, we exploit a dramatic natural experiment resulting from a national border – between

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[1] See Krueger and Meyer (2002) for an up-to-date literature review on unemployment insurance and labor supply. Studies of other income support programs share similar features. For example, a welfare study might measure the effects of the benefit level on length of claim for a sample of single mothers.

[2] Regarding negative income tax experiments, Murray states: "Presumably people are less likely to burn bridges behind them if they know that the guaranteed income ends in three years than if it is legislated for life" (p.153).

the Canadian province of New Brunswick and the U.S. state of Maine – that divides a relatively homogeneous region. UI policy, however, changed in very different ways on either side of this border. In 1940, Maine had a modest UI system, but New Brunswick had none. By 1950, the two regions had roughly similar UI systems. Since then, New Brunswick's system experienced two major expansions, and is currently much more generous than Maine's. For instance, in 1980, 10 weeks worked per year in New Brunswick yielded an average annual income (including UI benefits) equivalent to 33 weeks worth of earnings (vs. 13 weeks in Maine). These benefits could be received year after year, without penalty. We use this dramatic policy divergence to estimate the long-term effects of UI program generosity on the distribution of weeks worked per year.[3]

We analyse annual labor supply patterns in two distinct ways. First, we examine the incidence of part-year work (defined as from 1 to 39 weeks) using cell mean data for the 1940-91 Census years. Because the available data is grouped by industry, this approach necessarily restricts attention to persons who work at least one week per year (non-employed persons are not attached to an industry). In this analysis, the entire state of Maine serves as a "control group" for New Brunswick. Second, we study the effects of UI on a more disaggregated distribution of annual work weeks using Census microdata for the period 1970-91.[4] In this case, we can restrict attention to Maine's six northernmost counties to obtain a tighter control group, and to include persons who work zero weeks per year in the analysis. The latter allows us to incorporate potential "entitlement effects" of UI (Hamermesh 1979), i.e. the possibility that UI induces non-

[3] Unfortunately, consistent data on hours worked per week are not available across the countries and years examined here. Thus our labor supply measure throughout the paper will be weeks worked per year.

[4] Aside from a desire to put the natural experiment provided by Canada's 1971 UI Act at the center of our analysis, we terminate our analysis after 1991 for two reasons. One is the likely endogeneity of Canadian UI policy changes in the 1990s, many of which were motivated by high dependency rates in regions like New Brunswick. In addition, our correction for potential post-1970 hysteresis in New Brunswick's UI generosity requires us to calculate UI benefits at unemployment rates prevailing in 1970 (see note 20). We were reluctant to extend this procedure all the way to the 2000 Census.

workers to enter the work force to accumulate enough work weeks to qualify for UI benefits, in addition to the work disincentive effects examined in most studies.

Both approaches generate substantial estimates of UI program effects. For example, in 1990, about 12.6 percent of working-age men in Maine's northernmost counties worked between 1 and 39 weeks per year; just across the border in New Brunswick that figure was 25.6 percent. According to our estimates, New Brunswick's much more generous UI system accounts for over three fourths of this differential. Overall, our findings suggest that a ten percent UI-induced increase in the income associated with working 1-39 weeks per year (holding the incomes associated with nonwork and with full-year work constant) raises the number of men (women) in that category by about 20 (10) percent, for an elasticity of two (one). Estimated elasticities of UI program participation with respect to the same policy change are .60 for men and .45 for women; the number of paid UI weeks is more responsive at 1.67 and 0.88 respectively. These large effects are, however, consistent with a small overall effect of UI on mean annual weeks worked in the region, since (a) even in the presence of generous UI, part-year workers are a minority of the work force, and (b) UI "pulls" workers from *both* tails of the weeks distribution toward the middle, part-year categories.

Methodologically, this paper may be of interest in that it incorporates some advantages of cross-national studies (large and long-lasting variation in program parameters) while avoiding some important disadvantages. The latter include large and uncontrolled differences in culture and the economic environment, major differences in data collection procedures, and a tendency to rely on one-dimensional "summary" indicators of program generosity which neglect important details of how policies work (Hamermesh 2002). Our approach examines two similar regions in different countries – where data collection procedures have been highly comparable for a long time – using a multidimensional measure of program generosity that incorporates most key aspects of the two regions' unemployment insurance systems.

Background

Maine is the northeasternmost state in the U.S.A, known for its coastal scenery, cold climate, rural character and relatively low incomes. New Brunswick is the Canadian province that borders directly on Maine, known for similar traits. Both regions' populations are overwhelmingly white and native-born. In 1990, Maine and New Brunswick had total populations of 1.2 million and 740,000 respectively. Over the preceding 50 years, these populations grew by 54 and 59 percent, well behind their respective national averages of 90 and 135 percent. Both regions' incomes were below their national averages as well: from 1945-1991, personal income per capita in Maine was on average 20% below that of the U.S. while New Brunswick was 27% below the Canadian average.

Compared to other Canadian provinces, New Brunswick is of particular interest for the study of UI for several reasons. One is geography: New Brunswick is the only Canadian province that borders a single U.S. state of roughly similar size and population.[5] Second, despite the federal nature of UI in Canada, several features of the program generate significantly higher UI replacement rates and benefit entitlement periods in New Brunswick than those facing most Canadians.[6] This generates a much more dramatic natural experiment than comparisons across

[5] Alberta is the only other Canadian province that borders a single state (Montana). New Brunswick is also the only Canadian province that shares a longer border with the U.S. than with other Canadian provinces. Both of these factors are critical because the 1971 public use Canadian Census data, which plays a central role in our analysis, does not identify geographical areas smaller than an entire province (with the exception of the Toronto and Montreal metropolitan areas). This effectively rules out other cross-border comparisons in the period before the 1971 UI reforms.

[6] The main reasons for this are (a) New Brunswick's low wage rates (UI benefits are proportional to the previous wage up to a maximum, which binds for most workers in the more prosperous regions of Canada); (b) before 1971, certain seasonal benefit programs affecting industries that are greatly over-represented in New Brunswick; and (c) after 1971, UI regulations that tied the number of benefit weeks to a jurisdiction's previous unemployment rate. The latter feature introduces an element of endogeneity to UI benefits at the economy-wide level, which we address by

other parts of the Canada-US border. Indeed, the combination of New Brunswick's small population *and* the federal financing of UI in Canada generates levels of program generosity that would likely be unsustainable if UI was self-financing within the region. Finally, because of their similar resource bases, New Brunswick and Maine share a history of seasonal employment that predates the introduction of UI to either economy, which may make these jurisdictions more sensitive to UI than other economies. Thus, the study of these jurisdictions may shed light on the process of labor force adjustment. For instance, we can examine the extent to which expansion of UI in Canada helped preserve a (seasonal) "way of life" that was forced into extinction by market forces elsewhere.

Figure 1 shows UI expenditures as a share of provincial/state GDP as well as unemployment rates. For New Brunswick, there were large increases in UI expenditures during the 1950s and early 1970s, likely reflecting two major increases in program generosity (described in the next section).[7] In contrast, UI expenditures in Maine were (a) roughly constant over the entire time period, and (b) much smaller as a share of state GDP. By the end of our sample period, the UI share of GDP in New Brunswick, at 6 percent, was about *six times* the share in Maine. As we argue in more detail below, this dominant role for UI in New Brunswick helps explain why the labor supply effects of this single program are detectable at the economy-wide level. Unemployment rates are suggestive of a UI-based story as well. Prior to 1953, Maine had a higher unemployment rate than New Brunswick. The unemployment rate in New Brunswick then rose substantially above Maine's in the 1950's, a gap which closed somewhat by 1970. Beginning in about 1976, a large and persistent gap re-emerged, with New Brunswick's rate

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calculating UI benefits at unemployment rates prevailing before the 1971 reforms.

[7] The most severe recession during this period was in the early 1980s, and its effects on UI expenditures are visible in Figure 1, but note that this cyclical effect is much smaller than the increases that occurred in the 1950s and 1970s.

consistently above 12 percent from 1982 onwards, and Maine's consistently below 8 percent from 1984 onwards.[8]

Further confirmation of the pervasive influence of UI in New Brunswick is provided by examining receipt of UI at the end of our sample period. We find that about 30 percent of New Brunswick's workers received some UI benefits in 1990 relative to about 6 percent for men in Maine, and about 3 percent for women in Maine. If we examine the entire working age population, we also find that an astonishing 23 percent of *all* New Brunswick men aged 25-59 received some UI income in 1990. This large fraction of persons receiving UI may also lessen any social stigma attached to it – a factor that might contribute to a larger labor supply response. Finally, we also compared UI participation with participation in other income support programs. Only 5 to 6 percent of New Brunswick's workers received transfers other than UI, and only 4 percent of Maine's did. Thus, the number of New Brunswick's workers receiving UI was five to six *times* the number receiving all other transfers combined, further motivating our focus on UI as the income support program that is most likely to explain weeks-worked differences between these two jurisdictions, especially among those with positive weeks worked.

Unemployment Insurance Policy

In this section we briefly summarize the evolution of the key UI parameters in the two regions, with particular attention to the census years.[9] Relative to New Brunswick, the most salient feature of Maine's UI system is its stability over the half century between 1939 and 1989.

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[8] The emergence of this gap mirrors, in a considerably magnified fashion, the well-known emergence of the Canada-US unemployment rate gap. On the national level, a number of studies have attributed a substantial share of this gap to international differences in the composition of non-work time between unemployment and nonparticipation (see for example Riddell 1999). As our own data show, however, substantial differences in annual work weeks did emerge between New Brunswick and Maine after 1970, and these differences are the focus of our study here.

[9] A more detailed discussion is available at our websites. Because Canadian Census years end in '1' while U.S. Census years end in '0', and the Census weeks-worked questions refer to the previous year, all our policy discussion and statistical analysis focuses on years ending in '9' for Maine and '0' for New Brunswick. For simplicity, we shall refer to, for example, both 1989 Maine and 1990 New Brunswick data and UI rules simply as "1990" throughout the rest of the paper.

As in all other states, this program has been state-run and state-funded since its inception. Also since inception, eligibility for Maine's UI benefits is based on total earnings during a one-year base period preceding an unemployment spell.[10] In contrast to the U.S., UI in Canada is administered and financed at the federal level. Thus, UI is not required to be self-funding within New Brunswick. Also in contrast to most states, eligibility in Canada is determined by the *number of weeks worked* rather than total earnings during a base period. Finally, the Canadian program underwent some dramatic changes during our sample period.

As noted, Canada had no operating UI system in 1940. By 1950, UI eligibility depended on the number of days worked during the two years preceding a claim and included a so-called 'ratio rule', which imposed significant limits on benefit duration for seasonal workers (i.e. those with claims in successive years). The first key legislative change was the introduction of an explicit system of Seasonal Benefits in 1955, payable only during the winter months but under quite broad conditions (including exhaustion of regular benefits). Seasonal Benefits constituted a significant enrichment of Canada's UI program for persons with short work histories. A second significant expansion of benefits occurred in 1972 where UI was massively overhauled. The new system based weeks of benefit entitlement on weeks worked in the past year, with generous benefits at low levels of weeks, especially in high unemployment regions – where it was now possible to qualify for 40 weeks of UI benefits with just 10 weeks of work. Further, workers could continue such a pattern of part-year work indefinitely, without penalty.[11]

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[10] The key changes made between 1939 and 1989 include two increases in maximum benefit duration (from 16 to 20 weeks between 1939 and 1949; to 26 weeks by 1959); a shift from calendar-year to a four quarter base period (between 1959 and 1969); extension of coverage to federal government employees by 1959; to state and local government workers by 1979; the introduction of a dependents' allowance (by 1969), and federal income taxation of UI benefits (by 1979).

[11] In some parts of the country this practice became known, unofficially, as "Lotto 10-40".

We compute a summary measure of benefit generosity that incorporates many components of each region's UI system as shown in Figure 2.[12] Our measure summarizes the weeks worth of *income* a person would receive if he/she worked w weeks per year on a regular basis and was unemployed the remainder of the year.[13] In particular, suppose that an individual earns \$$E$ for every week worked. If working w weeks qualifies the worker for c weeks worth of UI benefits worth \$$B$ per week, then the total weeks worth of *income* received for w weeks of work are $(wE + cB)/E > w$.

For 1940, this generosity measure follows a 45-degree line as a function of weeks worked in New Brunswick, because there was no UI program. To reduce clutter we omit some years, but we note that in 1950 the two UI systems were highly similar. In 1960, the Canadian UI revisions summarized above – in particular the Seasonal Benefits – led to greater benefit generosity at low weeks of work. In particular, 20 work weeks now yielded essentially 30 weeks of income, with the greatest total subsidy at about 25 work weeks. A particularly dramatic change occurred following the 1971 UI Act: by 1980, an average worker in New Brunswick could, on a regular basis, receive 33 weeks worth of income at his/her regular rate of pay by working 10 weeks per year. In most years, Maine's subsidy is biggest at around 30 weeks of work, which would yield between about 36 and 39 weeks worth of income for a worker earning the average manufacturing wage.

Conceptual Framework

As noted, a distinguishing feature of this paper is our interest in UI's effects on the distribution of annual weeks worked in a region's entire working-age population. This differs

[12] Our measure incorporates all of the following parameters: benefit durations, weekly benefit amounts, coverage, dependents rules, taxes, repeater rules, waiting periods and seasonal benefits. A much more detailed discussion of all of these parameters is available at our websites.

[13] In most cases the qualifier "on a regular basis" is irrelevant. We introduce it to allow for restrictions on repeat, seasonal unemployment – such as Canada's "ratio rule" in the 1950's—to affect benefit entitlements.

from most UI studies, which focus on a number of outcomes (such as unemployment duration) *conditional* on a worker's becoming unemployed. In these studies, the population potentially affected by UI is by definition a relatively small share of the workforce. In contrast, we are interested in estimating impacts that could operate both through the extension of existing unemployment spells, *and* through processes that lead to the initiation of unemployment spells.

With the above goal in mind, we use a simple static labor supply model as our conceptual framework. This model is in the spirit of Moffitt and Nicholson's (1982) with a (significant) change of interpretation. Moffitt and Nicholson consider a decision made by workers *at the time of job loss*. Thus, their analysis applies only to workers who have already initiated an unemployment spell, and assumes that workers can choose to end that spell at any point by taking a job. We abstract even further from the dynamics of job search, hires and layoffs by assuming that any person, whether employed or unemployed, can choose the number of weeks worked per year with the remaining non-work weeks compensated – according to benefit eligibility formulas – by unemployment insurance.

Maine and New Brunswick severely restrict UI eligibility for workers who quit their job. We ignore these restrictions in the analysis, and thus our framework implicitly assumes that workers can find seasonal or limited-term contract jobs that match their desired durations (or indeed that such jobs are supplied by the labor market in response to the UI system). Equivalently, our model describes an environment where workers can induce firms to lay them off when this is convenient to the worker, or that employers agree to "re-label" quits as layoffs.[14] Incidentally,

[14] Canada's UI system had no experience rating during this period; thus employers face no marginal tax cost when they re-label a separation. Maine's experience rating system changed very little between the mid-1960s and the end of our sample period, with a taxable wage base and contribution rates at or near the federally-mandated minimum throughout the period. Relatedly, we note that quits are a small fraction of separations leading to unemployment in both Canada and the US; thus UI disqualifications due to quitting are unlikely to be quantitatively important in either region.

there is evidence of both of these practices in Canada. For example, Green and Sargent (1998) show how jobs tend to end (via layoffs) when UI eligibility is established while Kuhn and Sweetman (1998) provide evidence of the "relabelling" of separations.[15]

We recognize that this is a highly abstract model that ignores the job search, hiring and layoff processes. To an important extent, however, worker's limited control over the length of their employment or unemployment spells (which is modelled by a stochastic wage offer process in the search framework) is incorporated in our approach by the broad weeks-worked categories used.[16] In our main analysis, which uses microdata from the 1970-1990 Censuses we can assign every working-age individual to one of five annual weeks worked options: 0, 1-13, 14-26, 27-39 and 40-52. These are the most detailed categories that are consistent across regions and the above time periods.

In general, suppose that each worker, i, faces J labor supply options; each of these options is associated with a fixed annual amount of leisure, denoted L^j, $j \in \{1,2,\ldots J\}$. Let the annual income received by worker i if he/she chooses labor supply category j be Y_i^j. These income levels vary across individuals and over time for a variety of reasons including changes in UI policy, and UI's differential impact by skill level, region, and industry. Let the utility of option j be given by the Cobb-Douglas function:

$$U(Y_i^j, L_i^j) = \alpha \ln Y_i^j + \beta \ln L^j + \theta^j X_i + \varepsilon_i^j \equiv \alpha\, y_i^j \;\; + \beta\, l^j + \theta^j X_i + \varepsilon_i^j \qquad (1)$$

[15] There is also some anecdotal evidence of job-sharing arrangements designed to maximize employees' UI eligibility in parts of Eastern Canada. For example, Harris (1998) reports that in Newfoundland "In some cases fish plants and make-work projects would hire workers for a certain number of weeks and then lay off those workers and hire others, so they'd all have qualification for unemployment insurance" (p. 176).

[16] In fact, in a world where workers have limited control over exact job duration, the search in some studies (e.g. Green and Riddell 1997) for "spikes" in very narrow weeks-worked categories (essentially single weeks of unemployment duration) could lead to significant *underestimates* of workers' responses to UI.

wher X_i is a vector of personal characteristics. If the J ε's in (1) have independent and homoskedastic Weibull distributions[17], and individuals choose the weeks-worked category yielding the highest utility, equation (1) corresponds to McFadden's conditional logit model, where the X's represent characteristics that vary across persons only, while the income (Y) variables vary both across persons and the choice categories. We note three features of the framework in (1). First, if the vector X contains a constant (as it should in the current application), the β coefficients terms in (1) and (2) are not identified—the βl^j terms will be absorbed into the category-specific intercepts in the θ^j vectors. Second, an equal increase in $\theta^j X_i$ across all choice categories leaves optimal choices unchanged; thus we follow the usual practice of choosing a base category (j=1) and interpreting estimates of the coefficients on X in the remaining categories as estimates of X's effect on the utility of category j relative to the base (i.e. of ($\theta^j - \theta^1$). Third, an equal increase in y_i^j across all choice categories also leaves optimal choices unchanged.[18] Thus, for example, in the two-outcome case considered in the next section (full-year versus part-year work), we can normalize Y_i^1 (full year work) to 1, so the UI policy variable relevant to this binary choice becomes simply the log of the income ratio between the two choice options, i.e. $\ln(Y_i^2 / Y_i^1)$. This is essentially the UI generosity measure developed in the previous section.[19]

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[17] Most other specifications for the distribution of the errors are computationally impractical. In early work on this data, we estimated a Cobb-Douglas model with only a single, normally-distributed error term (representing the person's "tastes for work") rather than four. This model consistently did a poorer job of fitting empirical weeks-worked distributions.

[18] This has some noteworthy implications for identification. For example, if there is no zero-weeks category and workers' only source of income is wages (thus $Y_i^j = w_i h^j$, where w_i is the individual's wage and h^j is the weeks of work associated with choice j), α is not identified, because y_i^j will be perfectly correlated across all the choice options. The same is also true in the presence of a fixed UI policy, if UI benefits consist of a benefit weeks entitlement b^j that depends only on weeks worked (as is true in New Brunswick) and a weekly benefit amount that is a fixed fraction, r of the weekly wage, i.e. $Y_i^j = w_i h^j + r w_i b^j = w_i(h^j + r b^j)$. Thus, changes in UI replacement rates between Census years are essential to identification in this context.

[19] For both our cell-based and microdata-based analyses, we also estimated a linear utility model (identical to (1) except that we do not take logs of income and leisure). In the two-outcome case, the policy variable here is the income difference (in inflation-adjusted dollars) between full-year and part-year work. The results were similar.

Analysis of Cell means, 1940 to 1990

Before conducting our main analysis, which focuses on microdata for 1970-1990, this section conducts the most detailed analysis possible for a longer time horizon, 1940-1990, which spans more policy changes and goes back to a period when one of these two jurisdictions had no UI at all. For this period, the only consistent measure of annual weeks across countries is the share of employed persons working fewer than 40 weeks per year (henceforth 'part year'). Further, because Census microdata is not available for Canada prior to 1970, the unit of analysis is not the individual but an industry/gender/region/year cell. This information is taken from Canadian published data before 1970, and constructed from microdata for the remaining observations.[20]

Sample sizes in the 1971 Canadian census, and limited industry disaggregation in all the post-1970 Canadian public use files, limit the number of industries for which we can calculate cell means with a reasonable degree of precision. Given these constraints, we are left with 6 industries per year for women (trade, services, finance/insurance/real estate, transportation/storage/communication/utilities, manufacturing, and public administration) and 9 for men (the above plus agriculture, construction and primary) for a total of 6 industries * 6 years * 2 regions = 72 female observations and 9 * 6 * 2 = 108 male observations. The sample is all wage earners over the age of 15 with at least one week of work in the previous calendar year.[21]

As noted in the previous section, our measure of UI policy generosity is now (the log of) a worker's relative income from working part versus full year. This measure is derived from cell-specific calculations isomorphic to those underlying Figure 2.[22] For example, suppose we use 20

[20] A detailed appendix discussing the creation of this data is available at our websites.
[21] Self-employed persons are excluded. With the exception of fishermen in Canada, the self-employed are not eligible for UI in either region. In both regions and all years, industry designations refer to the longest job held during the previous calendar year.
[22] In reality not all eligible workers apply for UI benefits. Since take-up is endogenous (Anderson and Meyer 1997), we simply use the legislated benefits as a more exogenous measure of benefit generosity. In 1980 and 1990 in New

weeks of work to represent the ‘part year’ work category (1-39 weeks). Relative income from part year work then just equals the total income associated with 20 weeks of work divided by the total income associated with 52 weeks. Table 1 shows employment-weighted means of this relative income variable, using two alternative reference points (20 and 30 weeks) to represent the part-year (1-39 weeks) option.[23] Part-year income varies cross-sectionally within a region because wage levels (and hence UI replacement rates) vary across industries, and because not all industries are covered by UI in all years. It varies over time both due to UI policy changes and to wage level changes at the industry level.

New Brunswick’s three major policy changes can be seen in Table 1. These include the introduction of UI in New Brunswick after 1940: relative income assumes its minimum possible value (20/52 = .385; 30/52 = .577) in that year. Canada’s 1971 UI Act is also clearly visible. We also see the similarity in UI generosity between regions in 1950 – especially using the 20-week parameters. It is noteworthy that – according to the 20-week specification – the largest decadal change in UI’s subsidy to part-year work occurred not in 1971, but between 1950 and 1960: a rise of (.580-.441=) .139 for men, and (.614-.455=) .159 for women, both in New Brunswick.

Table 2 shows employment-weighted means of our dependent variable by gender, region and year. In 1940, despite a modest UI program in Maine, employed men exhibited roughly similar propensities to work part-year in the two jurisdictions under study: 36 percent in New Brunswick and 33 percent in Maine. Over the next five decades, men’s part-year work continued at this high level in New Brunswick, while declining substantially (to 22 percent) in Maine. For women, the

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Brunswick, weeks of UI entitlement depended on the local unemployment rate. As already noted, we address this issue by calculating 1980 and 1990 benefits in New Brunswick at the unemployment rate prevailing in 1970.

[23] We examined the continuous distribution of weeks worked for those years where it was available; the median within the 1-39 weeks category was between 20 and 30 weeks in all cases. We also replicated the analysis using 15, 25 and 35 weeks to represent the 1-39 weeks category. There were no major differences from the results reported here. The following (microdata) section is able to improve on this approach considerably.

divergence between New Brunswick and Maine is more dramatic. In 1940, before the advent of UI in New Brunswick, working women in Maine were considerably more likely to work part-year than women in New Brunswick (37 versus 20 percent). Five decades later, part-year work had fallen to 31 percent in Maine, while increasing to 40 percent in New Brunswick. Thus, by the end of the period the rankings of the regions were reversed.

Weighted least-squares regressions of the share of workers working part year on the relative income associated with part year work are presented in Table 3.[24] Separate regressions are run for men and women, and all specifications include fixed effects for industry, region and industry*region interactions, i.e. interactions that span the cross-sectional dimension of our data. As a result identification comes only from temporal variation in part-year income within an industry/region/gender cell. In addition all specifications control for year effects as well, to net out any trends in part-year work that are common to the two regions under study, so only region-specific changes are used for identification.[25] The industry coefficients (not shown) show that part-year work is more common in agriculture, construction and primary industries relative to manufacturing (the reference category), and less common in the finance, government, and transportation industries. Time trends in part-year work implied by the estimated year effects are mostly downward, but vary across specifications. Finally, note that while the estimates using 20-

[24] Results using the log-odds ratio in the cell as the dependent variable were very similar; we show the linear results in this section for greatest ease in interpretation. The regressions are weighted by year-specific industry shares in employment, in to make the results comparable to the microdata analysis in the following section, which (implicitly) assign equal weight to each person in the data. The unweighted results are very similar; for example in the 20-week case the coefficients (*t*–statistics) on the relative income variable were .249 (3.48) for men and .348 (3.85) for women.

[25] We also estimated a version of the model that included a full set of fixed effects for region-year interactions; this allows for different overall time trends in each region and identifies UI effects only from industry-level deviations from these trends. Unfortunately, the resulting coefficients were statistically insignificant with high standard errors, specifically .064 (t=0.60) for men and -.007 (t=0.03) for women with the 20-week parameters. We interpret this as a consequence of insufficient cross-industry, within-jurisdiction variation in UI replacement rates. This is perhaps not surprising given the possibility of significant classification error in the Census's (self-reported) industry variable, which necessarily plays a key role in the cell-level analysis.

versus 30-weeks to represent part-year work are quite similar, the 20-week version of the model fits our data considerably better for women. This could be because, relative to men, a larger share of part-year women work 20 weeks rather than 30 weeks. In the discussion below, we will primarily refer to the 20-week results.

Concerning the UI policy variable, row 1 of Table 3 shows that the relative income associated with part-year work has a positive estimated effect on the share of workers working part year in all specifications; this effect is highly significant when the 20-week UI parameters are used. To interpret the size of these coefficients, recall from Table 1 that the largest decadal change in the 20-week version of our policy variable occurred between 1950 and 1960; these translate into a change in $\log(y_i^P/y_i^F)$ of log(.580/.441) = .274 for men and log(.614/.455) = .300 for women. Applying the relative income coefficients from Table 3, this translates into a predicted change in the share working part year of .081 for men and .148 for women. Using actual 1950 part-year shares as a base, this would raise the share working part-year from 26.3 to 34.4 percent for men, and from 14.0 to 28.8 percent for women.

Clearly, the above effects are large and economically meaningful, especially for women. One way to assess whether these magnitudes are plausible is to compute the labor supply elasticities they imply; when we do so (using workers' predicted responses to New Brunswick's policy change between 1950 and 1960) we calculate an elasticity of weeks worked (among workers) with respect to the weekly wage of 0.214 for men, and 0.309 for women.[26] While these are economically meaningful, at least for women, they are not implausible in magnitude. Importantly, however, even the relatively modest labor supply elasticities estimated here can

[26] Defining the wage as the average effect on income of working one additional week, the percentage wage change associated with New Brunswick's policy change 1950 and 1960 for men is 27.2 percent ((.580-.441)/(1-.5(.580+.441)). Assigning 20 weeks of work to part-year workers and 52 to full-year workers, our predicted increase of 8.1 percentage points in part-year work translates into a predicted decline in 2.6 weeks of work per year, (from 43.6 to 41.0) which yields the reported elasticity.

correspond to large responses of the incidence of part-year work to policy changes, with correspondingly dramatic implications for the number of persons participating in an income-support program and for the costs of running such programs.

To see the latter point, assume for simplicity that no workers in our “full-year” category receive UI, and suppose (as Anderson and Meyer 1997 have shown) that the UI take-up rate responds positively to benefit generosity. It follows that the elasticity of part-year work with respect to UI benefit levels provides a lower bound estimate of the elasticity of program participation with respect to the benefit level.[27] For men, we estimate that a 27.2 percent increase (from .441 to .580) in the amount of income associated with part-year work generates a 27.2 percent (from 26.3 to 34.4) increase in the share of men choosing this option, for an elasticity of one. For women, this elasticity is much higher at 2.2. Clearly, responses of this magnitude cast major doubt on forecasts of program expenditures that assume no behavioral responses by workers.

Empirical Analysis: Microdata, 1970 to 1990

As noted, public-use microdata files are only available for Canada from 1970 onwards. Thus, for the years 1970-91, we can examine the determinants of part-year work at the individual worker level.[28] To compensate for this shorter (but still long compared to most other studies) time span, the microdata have four principal advantages. First, we can exploit other sources of differential change across different types of individuals in the attractiveness of part-year work; for example, we can now allow relative income changes to vary with the respondent’s education and the presence of dependents. Second, as noted, we restrict our attention throughout this

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[27] Let $P = T * PY$, where P is UI program participation, PY is the share working part-year, and T is the take-up rate. It follows that $\mathrm{dlog}P/\mathrm{dlog}B = \mathrm{dlog}T/\mathrm{dlog}B + \mathrm{dlog}PY/\mathrm{dlog}B$, where B is an indicator of UI benefit generosity.

[28] We use the New Brunswick observations in the 1%, 2%, and 3% national Public Use microdata files for 1970, 1980 and 1990 respectively. For Maine, we use the 1% sample in 1970 and the 5% state samples for 1980-90.

section to Maine's 'northern counties', which are more similar (than Maine as a whole) to New Brunswick.[29] Third, as also noted, we can examine a more detailed distribution of weeks worked than the split between 1-39 and over 40-52 weeks. If we can show that the particular weeks-worked categories where UI subsidies increased the most also experienced the largest increases in 'popularity' among workers, our confidence that UI policy changes had causal effects on the weeks-worked distribution will be increased. Finally, the Census microdata now give us the option of including persons who work zero weeks per year in the analysis; this allows us to incorporate any entitlement effects of UI (Hamermesh 1979).[30] Because the inclusion of nonworkers, almost by definition, raises some data constraints and conceptual issues that do not apply to the analysis of workers only, we present results for both a four-outcome (1-13, 14-26, 27-39 and 40-52), and a five-outcome (the previous categories plus zero weeks) model here.

As laid out in equation (1), our analysis of workers' choices among four or five weeks-worked categories now requires measures of total income, including both earnings and UI benefits, representative of each category. Rather than choosing a single number of weeks to represent the entire category (as in the previous section) we can now calculate the sum of earnings and UI benefits for each integer number of work weeks within the category, then use a weighted average of these total income levels to represent the entire category. To avoid the possible effects of UI-induced changes in the distribution of weeks within categories on our policy measure (and because the detailed distribution of weeks is not available in all years), we

[29] The six northern counties are Aroostook, Piscatquis, Penobscot, Waldo, Hancock and Washington. Interested readers can refer to our websites for a custom map of the region.

[30] Some more minor advantages of the microdata are a more detailed set of control variables, and the ability to incorporate additional UI rules in each region. A disadvantage is that we can no longer control for industry, since non-employed persons are not attached to an industry.

use the empirical distribution of weeks within categories from a single year (1980) in these calculations.[31]

As in our analysis of the cell data, our goal in the microdata analysis is to use only time-series variation in the incomes associated with different work options to identify UI's effects on work weeks.[32] To achieve this in the absence of panel data, we calculate each individual's policy variable (total income in a weeks-worked category) using the individual's predicted rather than actual weekly earnings, using a set of wage regressions that are estimated separately by gender, region and year. These predictions are based on a vector of personal characteristics, Z, all of which appear among the covariates, X, in the conditional logits for weeks worked.[33] Because, in addition, X always includes controls for Z interacted with region, the controls in X fully span the cross-sectional dimension over which our policy measure varies. In sum, our UI policy measures give mean the income available from each weeks-worked option to a worker with characteristics Z in each year, and only intertemporal variation in these income levels is used to identify their effects.[34]

[31] Integer-level weeks distributions are not available for either region in 1970. In 1990 for New Brunswick, an unanticipated and 'accidental' suspension of the variable entrance requirement in Canada results in individuals who previously qualified with 10-13 weeks worked being ineligible for the 1990 benefit year *only* (see Green and Riddell, 1997, Baker and Rea 1998). Based on the result in both these papers that workers who otherwise would have been disqualified for UI were quite successful in finding the extra work weeks needed to qualify, we assumed that persons who worked between 10 and 13 weeks in 1980 had an 80 percent probability of qualifying for UI in 1990. Our results are not very sensitive to this fraction.

[32] Cross-sectional variation is suspect for several reasons that are well known in the labor supply literature, including ability bias and "division bias" (Borjas 1980). Although these problems take somewhat different forms in the current discrete choice framework than in the classic labor supply context, they continue to apply. See earlier versions of our paper, available from the authors, for a more detailed discussion.

[33] The vector Z comprises education (dummies for 4 levels), age and its square, plus indicators for marriage, presence of children, and school attendance. In the four-outcome analysis --which restricts the sample to workers only--, Z also includes a full set of industry effects and a control for part time work. In this context we also exclude women in the agriculture, primary, and construction industries from the sample, because very small sample sizes do not allow us to reliably impute wages in these industries. To reduce the impact of measurement error we eliminated weekly wage outliers from all the wage regressions; specifically, for each region/year/gender we dropped the top and bottom 2.5 percent of the calculated wage distribution.

[34] Intertemporal variation in these income levels stems from two sources: time trends in real wages, and in UI replacement rates. As already noted, in the four-outcome case, only the latter identify the model. This is not the

Data limitations specific to the 5-option model (which includes option to work zero weeks) are, first, the fact that nonworkers are not associated with a particular industry. Thus we cannot use industry information to calculate a worker's unemployment insurance benefits (some industries are not covered in Maine; also wage rates vary by industry), nor can we control for the sizable effects of industry on weeks worked in our analysis. In addition, of course, UI benefits depend on wages, and wage information is not available for nonworkers; we simply use the wage regressions described in the previous paragraph, estimated from workers only, to impute wages to nonworkers based on their observed characteristics.[35] Finally, income support programs other than UI, for which we have no control variables, are likely to play a larger role for persons working zero weeks than for the population working positive weeks; these programs may have uncontrolled-for effects on the propensity to work zero weeks that are more important once nonworkers are included in the analysis. For all these reasons we interpret our five-outcome results with more caution than the four-outcome results; still we find them of considerable interest.

Some basic descriptive statistics of our sample in Maine and New Brunswick are provided in Appendix 1.[36] The working age populations of Northern Maine and New Brunswick were of almost identical ages during our sample period. Both of these populations also experienced substantial increases in educational attainment between 1970 and 1990, with New Brunswick's increase starting from a considerably lower base. There was some divergence in family structure

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case in the five-outcome model, because a rise in real wages makes *all* the positive-weeks categories more attractive relative to nonwork. Thus, the five-outcome analysis implicitly uses real wage trends as well as UI replacement rate changes to identify α, effectively treating these wage trends as exogenous. If the level of the real weekly wage was affected by changing UI policy over this period, that would change the interpretation of our results.

[35] We do not "discount" the wages of nonworkers relative to those estimated for workers, because this would create a mechanical link between the choice of positive weeks and the income from working positive weeks, which would bias our analysis of the choice of weeks.

[36] For context, these Tables also show statistics for all of Maine in addition to the six northern counties. On most dimensions, Northern Maine is more similar to New Brunswick than is Maine as a whole.

between the two areas (the decline in marriage and children was more pronounced in Maine) which we expect would lead to contribute to a larger *overall* decline in men's weeks worked in Maine than New Brunswick, with the opposite expected effect for women.

Moving to the results, the estimated coefficients from conditional logit estimation of equation 1 are reported in Table 4 for both the four- and five-outcome models.[37] Clearly, the results suggest that, holding incomes in all the other weeks categories constant, an increase in the income in a weeks-worked category makes it more likely to be chosen. Although the coefficients in the 4- and 5-outcome cases are not strictly comparable, the fact that the coefficients are of the same sign and similar significance between the specifications suggests that the additional data constraints affecting the 5-outcome analysis do not dramatically affect the results.

Since conditional logit coefficients do not provide useful measures of the effects of a regressor on the expected distribution of outcomes, Table 5 reports both the actual distribution of weeks worked and the predicted distribution under two alternative scenarios. The model used to generate these predictions is the five-outcome model in the presence of controls for region*year interactions.[38] Table 5 shows, first of all, that there are substantial differences between Maine's and New Brunswick's weeks-worked distributions for both men and women. Consistent with the expected effects of UI, workers in New Brunswick were, in general, more likely to occupy the categories most highly subsidized by its UI system (1-26 weeks) and this difference seems to have grown between 1970 and 1990. For example, the share of working-age men working 1-26

[37] As noted, all the regressions in Table 4 include a vector of personal characteristics, Z and Z interacted with region. While not reported, the variables in Z have effects on the choice-specific propensities that are, in almost all cases, strongly in line with expectations (e.g. higher education reduces the attractiveness of all part-year categories relative to the baseline, full-year category).

[38] We do not report the model's baseline predictions because, by construction, the presence of a full set of region*year effects in the model allows it to fit the mean weeks distributions in each region and year perfectly.

weeks rose by 3.8 percentage points (from 2.7 to 6.5) in Maine between 1970 and 1990, compared to 8.3 percentage points (10.6 to 18.9) in New Brunswick. In Maine, the share of working age women working 1-26 weeks *fell* by four percentage points (from 16.7 to 12.7) over this period, while rising by 7.3 points (from 11.4 to 18.7) in New Brunswick.

Table 5's first counterfactual simulation, in columns 3 and 7, asks: what would have happened in New Brunswick if the 1971 UI Act had not been implemented? Specifically, we hold all variables, including the choice-specific year and region effects, at their actual levels and recompute each New Brunswicker's category-specific income variables in 1980 and 1990 as if the 1970 UI rules continued to apply.[39] For men, the simulation indicates that, instead of declining from 72.3 to 64.5 percent between 1970 and 1990, the share of New Brunswick's men working full year (40-52 weeks) would have risen slightly, to 73.1 percent, had the 1971 UI Act not been implemented. Thus the Act can account for essentially all of New Brunswick's decline in full-year work among men. Table 5 also suggests a small entitlement effect of the 1971 UI Act on New Brunswick's men: specifically the UI Act dampened the secular rise in the share of men working zero weeks (which was a national phenomenon in both Canada and the U.S. over this period), from a (11.9-7.0) 4.9 percentage point increase without the Act, to a (9.9-7.0=) 2.9 percentage point increase in its presence. Overall, as expected from a simple labor supply model, a large increase in UI generosity "pulled" men from both tails of the weeks-worked distribution (0 weeks and 40-52 weeks) which are not –or relatively lightly—subsidized by the UI system, into the middle categories (1-39 weeks), with the largest increase in the 14-26 week category where the Act's subsidy was the greatest. According to our model, by 1990, the share

[39] We adjust the nominal 1970 UI benefit levels for inflation using the consumer price index. Note also that these simulations treat the changing education mix over these two decades as exogenous. To the extent that rising UI generosity reduced the incentives to acquire additonal education, these simulations would understate the effects of UI.

of New Brunswick's men working part year (1-39 weeks), was (4.2+14.7+6.7=) 25.6 percent, compared to (0.6+8.2+6.1=) 14.9 percent had the 1971 UI Act not been implemented.

For women, we see a similar predicted effect of UI, again pulling persons from the unsubsidized tails of the weeks-worked distribution into the subsidized middle. The key difference is that this UI effect is set against an underlying time trend of rising, rather than falling overall labor force attachment. Specifically, according to our estimates, absent the UI Act the share of New Brunswick's women working zero weeks would have fallen by 24 percentage points, from 55.0 to 31.0 percent. According to our estimates, the UI Act accentuated this decline in nonwork somewhat, to its actual level of (55.0-29.0=) 26 percent. At the other end of the weeks distribution, the UI Act dampened the increase in women's full-year work, from (48.6-28.4=) 20.2 percent in its absence, to only (45.6-28.4=) 17.2 percentage points in its presence. Combining these two effects, the model predicts that the UI reforms raised the share of women working part year in 1990, from 20.4 to 25.4 percent, with the largest gains in the 14-26 week category.

Columns 4 and 8 of Table 5 examine a different policy counterfactual: what would New Brunswick's distribution of work weeks have looked like if it had Maine's UI system throughout the period from 1970 to 1990? Since Maine's UI system is much less generous than New Brunwick's and did not change much over this period, the results are quite similar. One advantage of this decomposition is that it allows us to assess the contribution of UI policy differences to the 1990 international differential in annual work patterns observed here. For example, in 1990 the share of men working between 1 and 39 weeks per year was (2.3+4.2+6.1=) 12.6 percent in Maine, versus (4.2+14.7+6.7=) 25.6 percent in New Brunswick, for an actual international difference of 13.0 percentage points. According to Table 5, this difference would have been only (0.6+8.2+6.1) -12.6 = 2.3 percentage points if New Brunswick had Maine's UI system. Thus international differences in UI policy can account for (13.0-2.3)/13.0= 82 percent of the international gap in short-year work.

As in the two-outcome model, in order to better gauge the magnitude of our estimated effects, it may be helpful to compute frequency-of-choice elasticities with respect to the total income associated with a choice, holding the incomes associated with the remaining choices constant.[40] Specifically, we conduct a simulation in which we apply New Brunswick's 1970 UI rules (adjusted for inflation) to New Brunswick's 1980 population. This effectively holds real income in both the zero and 40-52 week categories constant (because zero weeks generates no UI entitlements and the new UI rules had virtually no effect on the 40-52 week group).[41] In this example, we compute an elasticity of 2.17 for men and 0.99 for women for the middle three weeks categories combined. In other words, increasing the mean income associated with working 1-39 weeks by one percent raises the number of men (women) choosing that activity by roughly two (one) percent. While these elasticities seem high, it is worth recalling that they are calculated on a low base –the initial share working part-year--, and are consistent (see below) with very modest overall labor supply elasticities.

To assess the implications of our estimates for UI program participation, we note first that Canada's 1971 UI Act had a purely mechanical effect on this outcome: The new rules extended UI entitlements to workers with very short histories who previously did not qualify. Since our interest is in behavioral adaptation to legislation, we net out these mechanical effects by

[40] We compute more traditional labor supply elasticities (percent increase in weeks worked associated with a one percent increase in the net return to working one more week) in an earlier version of the paper. They are quite modest, similar to the estimates that emerge from our cell means analysis.

[41] Not only did UI generosity change relatively litte in this category, a large majority of workers in this category worked 52 weeks per year, and so are unaffected by UI.

computing changes in UI participation under a fixed (1980) set of UI rules.[42] Under these rules, our estimates imply an elasticity of UI program participation with respect to the average income associated with part-year work (1-39 weeks) of .60 for men, and .45 for women.[43] Elasticities of paid UI weeks (also calculated using the 1980 empirical distribution of within-cell work weeks) are 1.67 for men, and 0.88 for women. In other words, holding constant the levels of income associated with both nonwork and full-year work, a ten percent, UI-induced increase in the income associated with part-year work is predicted to raise the number of men receiving UI benefits by 6 percent, and the number of UI benefit weeks paid to men by 16.7 percent. At the same time, the elasticity of total weeks worked with respect to this policy change is quite modest: -.25 for men and -.09 for women. This contrast underscores the notion that UI, as predicted, raises some workers' labor supply while reducing others, in both cases moving people out of categories where little UI is received to categories where more is received. Thus, aggregate labor supply falls only slightly, while the distribution of weeks worked and the costs of running the UI system expand dramatically.

Alternative Explanations and Other Effects of UI

One concern that often arises in difference-in-difference estimates of policy effects is the possibility that uncontrolled-for region-specific macroeconomic shocks impacted one of the jurisdictions under study and not the other. Could more strongly-declining macroeconomic conditions in New Brunswick, rather than UI changes, account for its relative increase in part-year work over the period studied here? On this point, we note first that according to Figure 1, 1970, 1980 and 1990 were relatively similar years from a cyclicial point of view in both jurisdictions—all were relatively good years, a year or two *before* recessions that took place in

[42] As in our construction of the relative income measures associated with each choice category, we use the 1980 distribution of work weeks within our choice categories to impute UI participation.

[43] These compare to .43 (men) and .95 (women), estimated in the cell-based analysis.

the early years of those decades. Figure 1 also reveals that what happened to unemployment in New Brunswick around the two major UI expansions in our data (in the 1950s and 1970s) does not resemble a typical recession.[44] For example, it is well known that strong recessions occurred in 1983 and 1991; these correspond to increases in unemployment, but much smaller ones than those which appear to be associated with the UI policy changes studied in this paper. We also note that between 1970 and 1990, real wage growth was significantly stronger in New Brunswick than Maine, which is inconsistent with weaker macroeconomic trends in New Brunswick.[45]

That said, the two main difficulties associated with an "unobserved macro shocks" explanation of our results are (a) the fact that our microdata-based results are robust to choice-specific controls for region*year interactions and (b) the fact that, between 1970 and 1990, New Brunswick's men and women became more likely to work a small number of weeks (1-26) but *less* likely to work zero weeks.[46] These changes are more consistent with the predicted incentive effects of UI, which include both entitlement effects that draw workers into the labor force and incentive effects that reducing weeks worked among workers, than with a macroeconomic shock, which should shift the entire distribution of weeks worked either right or left.

In the remainder of this section we briefly consider other possible long-term effects of Canada's 1971 Act, beyond the effects on weeks worked on which we have focused here. We

[44] Note that changes in the unemployment rate should reflect both reduced work weeks among persons who once worked full year and increased weeks labelled as unemployment among persons who once worked zero weeks.
[45] Using our own data and country-specific CPIs, we find that real weekly wages in Maine grew by 0.9 percent and 13.0 percent for men and women respectively between 1970 and 1990. In New Brunswick those figures were 33.6 and 35.9 percent respectively. These increases become considerably smaller if one adjusts for the considerable educational upgrading in both jurisdictions, but the large international difference remains.
[46] According to Table 5, between 1970 and 1990 in our raw data the share of men working zero weeks increased less in New Brunswick than Maine (2.9 versus 4.9 percentage points), while the share working 1-26 weeks increased more (8.3 versus 3.8 percentage points). The share of women working zero weeks fell much more in New Brunswick than Maine (26.0 versus 18.9 percentage points), and the share of women working 1-26 weeks rose by 7.5 percentage points in New Brunswick while falling by 4 percentage points in Maine.

consider two outcomes: educational attainment and industry mix. Concerning the former, a naïve analysis might lead one to expect the income support aspect of UI to blunt incentives to acquire human capital. Counteracting this tendency, however, is the fact that –at least in certain ranges of the wage distribution-- UI benefits rise with wages. This, plus the fact that UI rewards labor market participation (relative to nonparticipation) could lead to a positive effect of UI on educational attainment. According to Appendix 1, what actually happened was a large increase in educational attainment in both Maine and New Brunswick. Because this upgrading started from a much lower base in New Brunswick, it is not clear how best to compare its magnitude across jurisdictions (for example, the share of high school dropouts fell by 39.2 precentage points in New Brunwick, compared to 25.5 points in Maine's Northern Counties, corresponding to percentage declines of 49.7 and 60.3 respectively). At most, we can probably say that no obvious disincentive to acquire education is either theoretically expected, or evident in the data for New Brunswick.

Did New Brunswick's generous UI raise the size of its seasonal industries relative to what would have happened otherwise? According to Appendix 1, men's employment in New Brunswick's primary industries fell from 9.7 to 8.5 percent of employment between 1970 and 1990; in Northern Maine it increased from 4.7 to 6.5 percent; both these changes are at the margin of statistical significance (women's primary employment was less than two percent at all times in both jurisdictions). The share of construction employment increased modestly in both jurisdictions. Thus, our data do not provide obvious support for the hypothesis that generous UI affected the industrial structure of New Brunswick over this period. It follows that (a) controlling for industry mix in our cell-based and four-outcome microdata analyses does not understate the long-term incentive effects of UI, and that (b) endogenous (and uncontrolled-for) changes in industry mix in the five-outcome microdata analysis are unlikely explanations for our estimated UI effects in that analysis. Instead, especially since primary industries constitute less than ten percent of men's employment (and less than 2 percent of women's), we can conclude that the adjustment of work patterns documented in this paper took place not within inherently seasonal industries, and not via the (relative) expansion of seasonal industries, but largely within the sectors that employed most men and women in New Brunswick: manufacturing, transportation, trade and services. Indeed, we estimate a large effect of UI on women's work weeks, even though about half of New Brunwick's female workers were employed in the service sector throughout our sample period.

Conclusions

In 1990, about 12.6 percent of working-age men in Maine's northernmost counties worked part year (between 1 and 39 weeks). Just across the border in New Brunswick that figure was 25.6 percent. Our estimates suggest that over three fourths of this difference can be explained by the much more generous Canadian Unemployment Insurance system. These estimates are derived from a unique and dramatic natural experiment spanning 50 years, during which we see a link between changes in these jurisdictions' relative weeks-worked distributions and New Brunswick's UI policy changes. Further, the fact that our highly parsimonious model predicts which particular weeks-worked categories (including zero weeks) became relatively more or less common after New Brunswick's 1971 UI reforms strengthens the case for a causal role of UI policy.

In more detail, we identify substantial effects of UI program generosity on the likelihood of working part year. For example, a ten percent UI-induced increase in the income associated with part-year work is estimated to raise the number of men (women) working part year by about twenty (ten) percent, for an elasticity of two (one). These UI-induced changes in the incidence of part-year work in turn imply large elasticities of program participation – and therefore program expenditures – with respect to program generosity. Our estimates however also predict quite modest effects of UI on *total* labor supply, especially for women. In part this is due to entitlement effects of UI that attract new workers into the labor market; in part it is because, even in the presence of generous UI, most workers still work full year and are unaffected by the program.

The approach taken in this paper differs from most existing studies of the labor market effects of income support programs which (for a number of good reasons) restrict their attention to the short-run effects of small policy changes on the labor supply of specific at-risk groups (e.g. former welfare recipients; workers who have just become unemployed). In contrast, we attempt to assess the long-run impact of a generous unemployment insurance program on the annual work patterns of a region's entire working-age population. Inevitably, our results are less precise and more open to debate than in shorter-term studies. That said, we are encouraged that our results are consistent with a simple economic model --which predicts opposite effects of UI on labor supply at the two ends of the distribution of annual weeks worked-- and we hope that the limitations of our approach are counterbalanced by the importance of the research question posed.

REFERENCES

Anderson, Patricia M. and Bruce D. Meyer. 1997. “Unemployment Insurance Take-up Rates and the After-tax Value of Benefits” *Quarterly Journal of Economics,* Volume 112, Number 3, pp. 913-37.

Baker, Michael and Samuel A. Rea. 1998. "Employment Spells and Unemployment Insurance Eligibility Requirements" *Review of Economics and Statistics,* Volume 80, Number 1, pp.80-94.

Borjas, George. 1980. “The Relationship between Wages and Weekly Hours of Work: The Role of Division Bias” *Journal of Human Resources,* Volume 15, Number 3, pp. 409-423.

Dingledine, Gary. 1981. *A chronology of response: The evolution of unemployment insurance from 1940 to 1980*. Ottawa: Minister of Supply and Services.

Green, David A. and W. Craig Riddell. 1997. ‘Qualifying for Unemployment Insurance: An Empirical Analysis” *Economic Journal,* Volume 107, Number 440, pp. 67-84.

Green, David A. and Timothy Sargent. 1998. “Unemployment Insurance and Job Durations: Seasonal and Non-seasonal Jobs” *Canadian Journal of Economics,* Volume 31, Number 2, pp. 247-78.

Hamermesh, Daniel S. 1979. “Entitlement Effects, Unemployment Insurance, and Employment Decisions” *Economic Inquiry* 17 (3) , 317–332.

Hamermesh, Daniel S. 2002. “International Labor Economics” *Journal of Labor Economics,* Volume 20, Number 4, pp. 709-32.

Harris, Michael. 1998. Lament for an Ocean: The Collapse of the Atlantic Cod Fishery: A True Crime Story. Toronto: McClelland & Stewart.

Kuhn, Peter and Arthur Sweetman. 1998. “Unemployment Insurance and Quits in Canada” *Canadian Journal of Economics*, Volume 31, Number 3, pp. 549-72.

Krueger, Alan and Bruce Meyer. 2002. “Labor Supply Effects of Social Insurance”, in A. Auerbach and M. Feldstein, eds. *Handbook of Public Economics*, Volume 4. Amsterdam: Elsevier, pp. 2327-2392.

Murray, Charles A. 1994. *Losing ground: American social policy, 1950-1980.* New York: Basic Books.

Riddell, Craig W. 1999. “Canadian Labour Market Performance in International Perspective”. *Canadian Journal of Economics* 32(5): 1097-1134.

Table 1
Estimated total annual income from part-year work, as a fraction of income earned from full-year work, Cell-level census data 1940-91

	1940	1950	1960	1970	1980	1990
	Part-year = 20 weeks					
MEN:						
New Brunswick	.385	.441	.580	.551	.680	.713
Maine	.454	.456	.479	.495	.509	.509
WOMEN:						
New Brunswick	.385	.455	.614	.579	.731	.731
Maine	.474	.478	.511	.509	.513	.513
	Part-year = 30 weeks					
MEN:						
New Brunswick	.577	.688	.761	.734	.774	.796
Maine	.673	.672	.680	.702	.718	.720
WOMEN:						
New Brunswick	.577	.716	.793	.760	.808	.808
Maine	.697	.693	.709	.763	.769	.769

NOTES: All means are based on 9 industries for men and 6 industries for women, and are weighted by the underlying number of observations in each industry. The sources are the 1940/50/60/70/80/90 U.S. censuses for Maine and the 1941/51/61/71/81/91 Canadian censuses for New Brunswick. See text for discussion of the computation of the relative income variable, and further discussion of the data.

Table 2
Percent of workers working part year, Census data 1940-91

	1940	1950	1960	1970	1980	1990
MEN:						
New Brunswick	.36	.27	.33	.34	.35	.36
Maine	.33	.25	.25	.24	.24	.22
WOMEN:						
New Brunswick	.20	.14	.31	.44	.47	.40
Maine	.37	.37	.50	.48	.40	.31

NOTES: Part-year work is defined as working less than 40 weeks. The sources are the 1940/50/60/70/80/90 U.S. censuses for Maine and the 1941/51/61/71/81/91 Canadian censuses for New Brunswick. See text for further discussion of the computation of part-year work, and discussion of the data.

Table 3: Regression estimates of the incidence of part year work, Cell data 1940-91

Variable	**20-week UI parameters**		**30-week UI parameters**	
	MEN	**WOMEN**	**MEN**	**WOMEN**
	(1)	**(2)**	**(3)**	**(4)**
Relative income	.295 (4.20)	.493 (5.83)	.404 (3.20)	.277 (1.24)
New Brunswick	.101 (5.57)	.038 (1.25)	.118 (6.19)	.088 (1.78)
1950	-.077 (2.88)	-.066 (1.59)	-.092 (2.89)	-.053 (0.89)
1960	-.074 (2.33)	-.004 (0.08)	-.064 (1.94)	.083 (1.25)
1970	-.062 (1.95)	.063 (1.73)	-.053 (1.63)	.129 (2.22)
1980	-.099 (2.71)	-.020 (0.47)	-.071 (2.02)	.098 (1.35)
1990	-.111 (2.85)	-.102 (2.21)	-.082 (2.17)	.016 (0.21)
R^2	.880	.812	.867	.674
N	108	72	108	72

NOTES: Coefficients are derived from a WLS regression of the fraction of workers working fewer than 40 weeks per year on the independent variables shown, plus fixed effects for industry and for industry*region interactions. Robust *t*-statistics are in parentheses (in absolute value). The 'relative income' variable is the log of the ratio of part-year income to full-year income. 1940 is the omitted year. All regressions are weighted by annual industry shares in employment.

Table 4: Estimated Effects of Category-Specific Income in McFadden choice model

	Four-outcome choice model (positive weeks only)		**Five-outcome choice model (all persons, including zero weeks)**	
Specification	**Men**	**Women**	**Men**	**Women**
	(1)	**(2)**	**(3)**	**(4)**
Without Region* Year Effects	1.087 (2.37)	1.325 (4.82)	1.146 (3.67)	.644 (4.16)
With Region* Year Effects	2.662 (2.58)	1.457 (1.83)	2.287 (4.09)	.908 (2.67)
N (number of persons in estimation sample)	13,531	10,341	17,810	18,513

NOTES: *t*-statistics are in parentheses (in absolute value). The dependent variable equals one for the weeks worked category realized, zero for each other weeks worked category. Estimation is by conditional logit. The category-specific income variable is the log of real total income (in 1983 dollars in each country) in the weeks category. Control variables in all specifications comprise a full set of region and year effects, plus a vector of personal characteristics (*X*), and *X* interacted with region. *X* in turn contains education (dummies for 4 levels), age and its square, plus indicators for marriage, presence of children and school attendance. Relative income is calculated using an individual's predicted wage from gender/region/year-specific weekly wage regressions and the UI rules prevailing in his/her country/year. The sample is restricted to individuals of age 25 to 59. In the four-outcome choice model, the sample is restricted to persons working at least one week, and a vector of industry effects is also included in *X*.

Table 5: Actual and Predicted Counterfactual Weeks-Worked Distributions, 5-outcome model

	MEN				WOMEN			
	ME Actual	NB Actual	NB without the 1971 UI Act	NB with Maine's UI system	ME Actual	NB Actual	NB without the 1971 UI Act	NB with Maine's UI system
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1970:								
0 weeks	4.0	7.0	7.0	7.1	45.2	55.0	55.0	54.6
1-13 weeks	0.9	2.0	2.0	4.5	9.4	4.4	4.4	6.6
14-26 weeks	1.8	8.6	8.6	6.0	7.3	7.0	7.0	5.8
27-39 weeks	3.5	10.1	10.1	8.5	9.8	5.2	5.2	4.7
40-52 weeks	89.8	72.3	72.3	73.8	28.3	28.4	28.4	28.3
1980:								
0 weeks	7.8	9.5	11.0	11.1	37.1	44.3	47.7	47.8
1-13 weeks	2.2	4.6	0.6	1.3	6.9	7.8	3.6	5.2
14-26 weeks	4.8	9.9	6.1	4.8	7.8	9.4	7.9	6.5
27-39 weeks	7.2	7.2	6.9	6.2	9.7	6.4	6.6	6.0
40-52 weeks	77.9	68.8	75.5	76.6	38.5	32.0	34.2	34.4
1990:								
0 weeks	8.9	9.9	11.9	12.1	26.3	29.0	31.0	31.4
1-13 weeks	2.3	4.2	0.6	1.4	5.9	5.8	3.0	4.3
14-26 weeks	4.2	14.7	8.2	6.6	6.8	13.1	10.8	9.1
27-39 weeks	6.1	6.7	6.1	5.6	8.5	6.5	6.6	6.1
40-52 weeks	78.5	64.5	73.1	74.3	52.6	45.6	48.6	49.2

Note: Predictions calculated from McFadden conditional logit model, includes controls for region*year.

Figure 1: Unemployment rates and UI expenditures in the New Brunswick and Maine

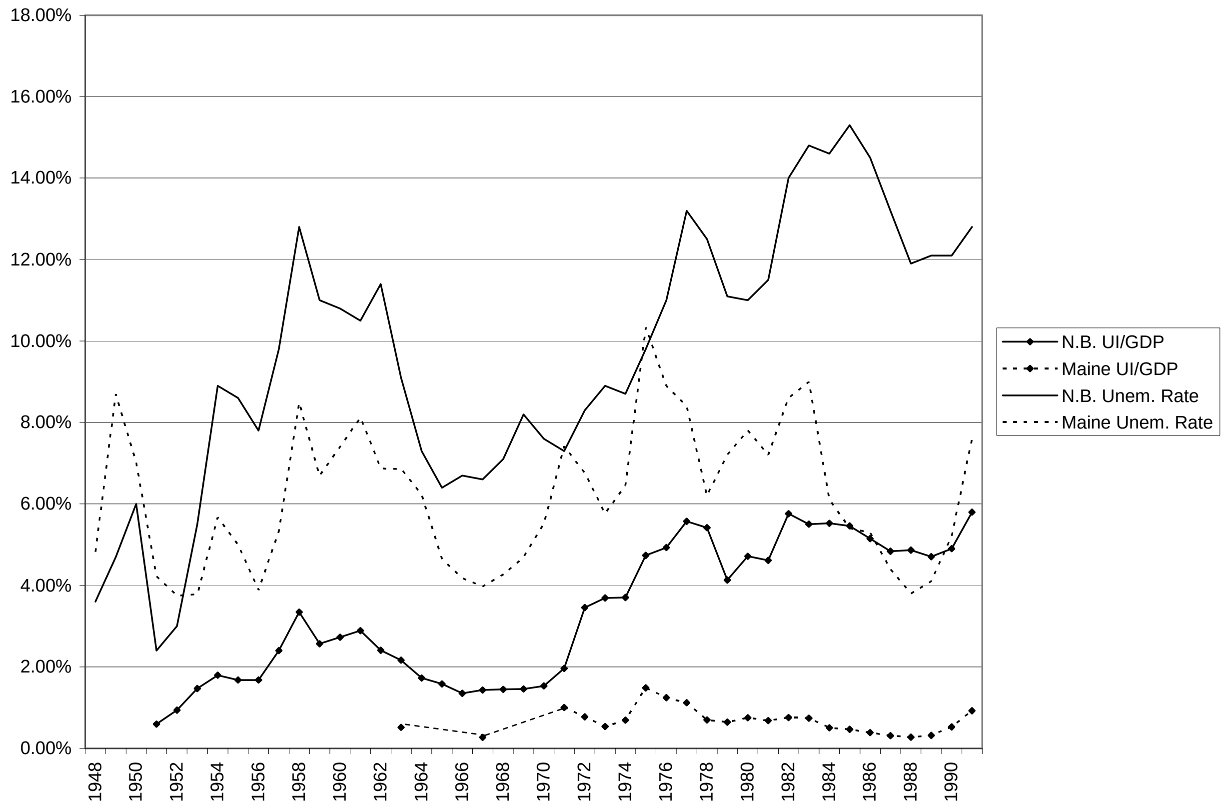
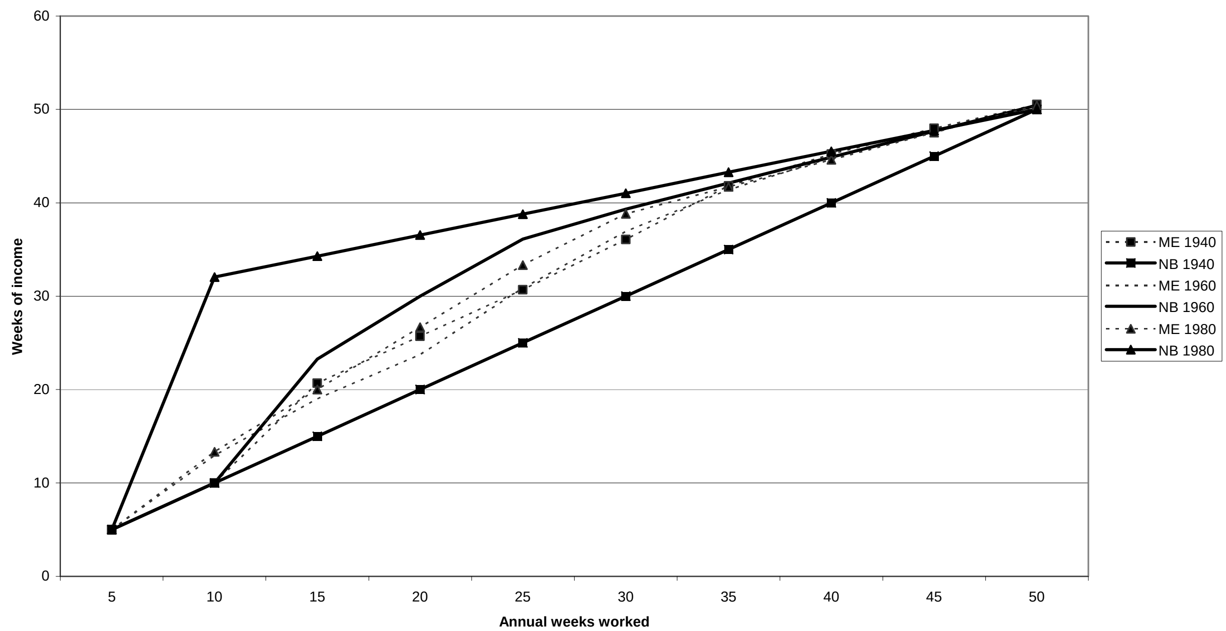


Figure 2: UI Generosity: Weeks of Income Received per Week Worked



Appendix 1: Descriptive Statistics

a) MEN

Variable	New Brunswick			Maine-Northern Counties			Maine		
	1970	1980	1990	1970	1980	1990	1970	1980	1990
DEMOGRAPHIC CHARACTERISTICS:									
Age	41.40 (.309)	39.07 (.190)	39.90 (.132)	41.58 (.416)	40.05 (.165)	39.85 (.135)	41.40 (.236)	39.96 (.096)	39.80 (.078)
Less than high school	.788 (.012)	.543 (.009)	.396 (.007)	.423 (.021)	.259 (.007)	.168 (.005)	.413 (.012)	.250 (.004)	.140 (.003)
High school degree	.104 (.009)	.279 (.008)	.404 (.007)	.397 (.020)	.422 (.008)	.436 (.007)	.376 (.011)	.396 (.005)	.422 (.004)
Some post-secondary	.042 (.006)	.075 (.005)	.077 (.005)	.089 (.012)	.148 (.006)	.235 (.006)	.102 (.007)	.165 (.003)	.241 (.004)
Four-year degree	.065 (.007)	.103 (.006)	.124 (.005)	.091 (.012)	.172 (.006)	.161 (.005)	.109 (.007)	.188 (.004)	.197 (.003)
Married	.827 (.011)	.800 (.008)	.779 (.006)	.885 (.013)	.794 (.006)	.740 (.006)	.850 (.008)	.793 (.004)	.736 (.004)
Children	.688 (.014)	.690 (.009)	.679 (.007)	.673 (.020)	.612 (.008)	.545 (.007)	.669 (.011)	.608 (.005)	.553 (.004)
Attending school	.058 (.007)	.059 (.004)	.067 (.004)	.014 (.005)	.031 (.003)	.050 (.003)	.023 (.004)	.035 (.002)	.045 (.002)
Sample Size:	1130	2822	5131	572	3910	4996	1804	11563	14488
INDUSTRY:									
Agriculture	.033 (.006)	.022 (.003)	.024 (.002)	.073 (.011)	.050 (.004)	.048 (.003)	.039 (.005)	.032 (.002)	.032 (.002)
Primary	.097 (.009)	.091 (.006)	.085 (.004)	.047 (.009)	.068 (.004)	.065 (.004)	.036 (.004)	.041 (.002)	.048 (.002)
Manufacturing	.187 (.012)	.188 (.008)	.183 (.06)	.255 (.019)	.252 (.007)	.220 (.006)	.335 (.011)	.288 (.004)	.245 (.004)
Construction	.087 (.009)	.119 (.006)	.121 (.005)	.107 (.013)	.106 (.005)	.139 (.005)	.097 (.007)	.107 (.003)	.150 (.003)
Transportation	.157 (.012)	.146 (.007)	.126 (.005)	.093 (.012)	.087 (.005)	.067 (.004)	.076 (.006)	.078 (.003)	.068 (.002)
Trade	.146 (.012)	.136 (.007)	.135 (.005)	.158 (.016)	.138 (.006)	.148 (.005)	.172 (.009)	.149 (.003)	.156 (.003)
Finance	.017 (.004)	.026 (.003)	.028 (.003)	.024 (.006)	.023 (.002)	.019 (.002)	.028 (.004)	.032 (.002)	.029 (.001)
Services	.135 (.011)	.164 (.007)	.185 (.006)	.147 (.015)	.174 (.006)	.180 (.006)	.140 (.008)	.180 (.004)	.189 (.003)
Public	.142 (.011)	.109 (.006)	.113 (.005)	.096 (.013)	.101 (.005)	.112 (.005)	.078 (.006)	.092 (.003)	.083 (.002)
Sample size	994	2495	4498	550	3614	4567	1733	10771	13487

NOTES: Standard errors are in parentheses. The sample is restricted to individuals of age 25 to 59. In addition, the sample for the industry variables is restricted to persons with at least one week of work in the calendar year preceding the Census.

b) WOMEN

Variable	New Brunswick			Maine-Northern Counties			Maine		
	1970	**1980**	**1990**	**1970**	**1980**	**1990**	**1970**	**1980**	**1990**
DEMOGRAPHIC CHARACTERISTICS:									
Age	40.99 (.296)	39.32 (.192)	39.64 (.133)	41.77 (.402)	40.21 (.165)	39.87 (.136)	41.96 (.229)	40.25 (.096)	39.77 (.078)
Less than high school	.779 (.012)	.567 (.009)	.358 (.007)	.388 (.019)	.235 (.007)	.137 (.005)	.367 (.011)	.230 (.004)	.116 (.003)
High school degree	.146 (.010)	.287 (.008)	.446 (.007)	.427 (.020)	.459 (.008)	.460 (.007)	.433 (.011)	.464 (.005)	.435 (.004)
Some post-secondary	.042 (.006)	.078 (.005)	.092 (.004)	.128 (.013)	.174 (.006)	.248 (.006)	.135 (.008)	.172 (.003)	.261 (.004)
Four-year degree	.033 (.005)	.068 (.005)	.105 (.004)	.058 (.009)	.132 (.005)	.155 (.005)	.065 (.006)	.134 (.003)	.189 (.003)
Married	.817 (.011)	.801 (.007)	.775 (.006)	.839 (.015)	.788 (.006)	.749 (.006)	.808 (.009)	.772 (.004)	.734 (.004)
Children	.817 (.011)	.654 (.009)	.699 (.006)	.660 (.019)	.672 (.007)	.627 (.007)	.672 (.011)	.659 (.004)	.619 (.004)
Attending school	.037 (.006)	.051 (.004)	.070 (.004)	.013 (.004)	.035 (.003)	.074 (.004)	.015 (.003)	.038 (.002)	.069 (.002)
Sample Size:	1138	2869	5202	626	3975	5076	1946	12007	14922
INDUSTRY:									
Agriculture	.021 (.007)	.015 (.003)	.021 (.002)	.032 (.009)	.027 (.003)	.024 (.002)	.026 (.005)	.017 (.001)	.018 (.001)
Primary	.004 (.003)	.017 (.003)	.015 (.002)	.003 (.003)	.007 (.002)	.010 (.002)	.005 (.002)	.003 (.001)	.005 (.001)
Manufacturing	.114 (.015)	.127 (.008)	.094 (.005)	.236 (.023)	.190 (.008)	.117 (.005)	.294 (.013)	.228 (.005)	.142 (.003)
Construction	.006 (.004)	.016 (.003)	.013 (.002)	.006 (.004)	.011 (.002)	.016 (.002)	.009 (.003)	.008 (.001)	.016 (.001)
Transportation	.051 (.010)	.036 (.005)	.037 (.003)	.017 (.007)	.023 (.003)	.022 (.002)	.020 (.004)	.025 (.002)	.024 (.001)
Trade	.216 (.019)	.166 (.009)	.160 (.006)	.231 (.023)	.185 (.008)	.209 (.007)	.222 (.012)	.181 (.004)	.204 (.004)
Finance	.047 (.010)	.059 (.006)	.060 (.004)	.026 (.009)	.044 (.004)	.049 (.004)	.048 (.006)	.057 (.003)	.066 (.002)
Services	.458 (.023)	.491 (.013)	.501 (.008)	.421 (.027)	.451 (.010)	.488 (.008)	.337 (.014)	.417 (.006)	.470 (.005)
Public	.083 (.013)	.074 (.007)	.099 (.005)	.029 (.009)	.063 (.005)	.064 (.004)	.038 (.006)	.063 (.003)	.055 (.002)
Sample size	472	1575	3608	347	2543	3788	1172	8030	11755

NOTES: Standard errors are in parentheses. The sample is restricted to individuals of age 25 to 59. In addition, the sample for the industry variables is restricted to persons with at least one week of work in the calendar year preceding the Census.