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‘ATYPICAL WORK’ AND COMPENSATION

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Abstract
Atypical work, or alternative work arrangements in U.S. parlance, has long been criticized in popular debate as providing poorly-compensated employment. Although the early U.S. literature seemed to confirm this perception, more recent cet. par. analysis has offered a partial but somewhat more optimistic evaluation. The present paper builds on the latter body of research with a view to providing improved estimates of the effect of the full range alternative work arrangements on worker compensation. The improvements are basically two-fold. First, we account for the skewness in atypical worker earnings while retaining the Mincerian human capital earnings function. Second, we deploy additional waves of the main dataset on atypical workers (the CAEAS), while supplementing this cross-section analysis with longitudinal data from the NLSY. Our analysis covers earnings and (access to) health benefits. We report that although one group of atypical workers (contractors) seems to enjoy a wage premium, cross-section results from the CPS and NLSY for the better-known category of temporary workers point to a negative wage differential of some 6-15 percent. It emerges that much of the disparity stems from unobserved worker heterogeneity, accounting for which still supports a wage advantage for contracting work. As far as fringes are concerned, the appearance in cross section of a potentially large deficit in access to health benefits is again reduced after accounting for the permanent unobserved individual heterogeneity, although in this case the attenuation is much more modest.

JEL Classification: J31, J33, J4

Keywords: atypical/contingent work, alternative work arrangements, wage differentials, employer-related health insurance

I. Introduction

The frequency of alternative work arrangements such as consulting, contract, and temporary work – more familiarly referred to as "atypical work" outside the United States – has steadily increased in recent decades (see, for example, Segal and Sullivan, 1997). Research on the phenomenon has tended to focus on the nature and extent of such arrangements and their impact on a worker's employment history (e.g. Addison and Surfield, 2005). The earnings, still less the compensation packages, associated with these arrangements have been accorded less attention in the literature. But the thrust of the earnings literature is frankly pessimistic: compared with regular or open-ended employment, alternative work arrangements often appear to offer inferior compensation. Alternative work arrangements might therefore seem to warrant their atypical worker tag.

There are various reasons why atypical workers might earn less than regular workers. The *standard* model sees employer use of, say, agency or contract workers as a means of covering leave or as providing a flexible buffer stock allowing the firm to meet uncertain or fluctuating demand. In the voluntary sorting process that results, there will be a match between worker preferences and job characteristics. Agency jobs offering no firm-specific training will be filled by workers who, for a variety of reasons (Booth et al., 2002), do not wish to take advantage of the training option. Groups with marginal attachment to the workforce together with younger and older workers may be expected to be over-represented among the ranks of agency workers and to receive lower wages by virtue of differences in their human capital. (Note that this buffering story – the strategic use of outside contractors to smooth the work load of the regular labor force – hinges on the exogeneity of the temporal flow of work. Flexibility in the timing of the work to be performed might otherwise allow it to be carried out by the regular workforce during off-peak periods; see Abraham and Taylor, 1996.) Observe that heterogeneous preferences among workers are insufficient to produce an atypical worker pay penalty – what Hirsch (2005, p. 526) terms an "equilibrium wage gap" – that remains after factoring in human capital differences. Such a gap requires heterogeneous (non-fungible) workers, and employers who are not indifferent as between the agency and regular workers.

The firm's use of atypical workers, and agency workers in particular, may be a response to informational problems in the hiring process and lead to wage gaps. Firms may be wary of offering permanent contracts to workers who may turn out to be lemons. Agency workers can reduce the costs to firms of hiring 'risky workers,' namely, those with poor work histories, and may be a more efficient solution to the informational problem than the probationary contract. For example, if agencies are more adept at screening and matching workers, their use raises the probability that the worker will prove suitable – and cheaper to hire and fire if unsuitable (Booth et al., 2002). The wage implications are transparent, since workers are in effect pricing themselves into regular employment (without upsetting the morale/productivity of incumbents) and the agency is selling information on worker quality to firms. The groups of workers in question will include displaced workers (see Farber, 1999) in addition to labor market entrants and younger workers. A formal model of the use of agency workers for screening purposes is offered by Houseman et al. (2003), who argue that this process will be strongly pro-cyclical.[1] The authors' also argue that the relief offered firms through the use of agency temporaries in tight labor markets can also result in a positive wage gap for workers in high-skill occupations (see below).

Even so, negative wage gaps may be more apparent than real. One possibility here is that (unmeasured) worker ability and employment in an alternative work arrangement might be negatively correlated. In this case, the unfavorable wage gap will be attributable not to the employment arrangement but instead to lesser ability. In other words, faulty inferences between earnings and type of work arrangement may be drawn if low-ability workers sort themselves into alternative work arrangements.

A *positive* wage gap might apply to atypical work if the aggregate demand for these alternative work arrangements is greater than the number of individuals seeking such employment. Thus, to borrow Hirsch's (2005, p. 527) argument in respect of part-timers, if firms use atypical work as low cost means of adjusting to variable and uncertain demand the relatively large supply of atypical workers required may run up against mobility or substitution constraints – across labor markets delineated by geography, occupation, and industry – and give rise to positive equilibrium wage gaps for atypical workers.

Some atypical workers in the 'knowledge economy' will enjoy high wages because of their skill endowments. Workers in high-skill professional, managerial, scientific, and technical occupations may turn to alternative work arrangements – contract working and use of employment agencies – if firms are unwilling or unable to satisfy knowledge workers. Forde and Slater (2005, p. 254) sketch the informational problems that arise in markets for such workers, noting that temporary employment agencies play "an important role in matching 'expert' knowledge workers with a series of short-term appointments." Abraham and Taylor (1996, p. 399) emphasize technological considerations, noting that contractors may be able to realize economies of scale that are unavailable in-house.

High wages can translate into positive wage gaps for the reasons given by Houseman et al. (2003, p. 109): the temporary use of higher-priced temporary labor may give the firm additional time to fill permanent vacancies. As before, the argument is that tight labor markets reduce the cost of agency temporaries relative to direct hires, and that it may pay to wage discriminate between direct hires and temporary agency workers in favor of the latter. Note the argument applies to high-skill occupations. The distinction with the low-skill occupations case, discussed above, is that the employment of temporary workers is now just that: it is temporary and does not perform a screening function leading to their induction of agency workers into the regular workforce, even if the goal as before is to obtain more labor without the need to raise the wages of incumbents/permanent employees.

The bottom line is that there a distinction has to be drawn between earnings differences and wage gaps. Our task will focus on explaining those differences in earnings between alternative work arrangements and regular employment that are attributable to observed *and* unobserved worker characteristics. The residual wage gaps that are left are the starting point for separate lines of inquiry not pursued here. Thus, for example, are negative gaps stepping stones to permanent employment for key groups in the labor market? Even zero gaps may be checked for their consistency with voluntarism in temporary working. Positive wage gaps are of interest precisely because of the knowledge economy and the theoretical implication that alternative work arrangements are not monolithic. We justify our more restricted focus on the grounds that there has been inadequate analysis of earnings differences between atypical and regular workers – and especially of differences in compensation – as was also the case until very recently for part-time workers (see Hirsch, 2005). Thus, our information on atypical worker compensation is either dated (in cross section) or overly restricted in the range of alternative wage arrangements examined (in longitudinal data). Accordingly, we seek to update previous findings using subsequent waves of the main dataset (the CAEAS) available to researchers in this area, while using a different longitudinal dataset (the NLSY) to correct for permanent unobserved differences between individuals for a wider range of alternative work arrangements than hitherto.

Our analysis proceeds as follows. To establish the innovations of our own empirical approach, we first briefly review the extant evidence on the relative compensation of atypical workers. We then introduce the empirical models for earnings determination and employee access to health benefits, and also provide information on the two datasets used here, focusing on the two compensation indicators and types of atypical work. Next, our findings are presented, sequenced by dataset. A short summary concludes.

II. Existing Research

Research into alternative work arrangements and their implications for the labor market took off in mid 1990s. This increased attention paid atypical work reflects two main developments. First, it was not until the 1980s that structural developments in the labor market – such as the attenuation of the common law at-will principle (see Autor, 2003) – favored the accelerated growth of atypical work. Second, there was understandably little quality data on work arrangements that had up to that point played a marginal employment role.

More recently, analysts have been able to make progress in identifying the impact of these arrangements on earnings, inter al., through the identification of workers engaged in the Temporary Help Services (THS) industry. And, since 1995, investigation of other types of atypical work has been facilitated by the publication of a Contingent and Alternative Employment Arrangement Supplement (CAEAS) to the Current Population Survey (CPS). This new supplement was administered biennially with the February CPS (in odd years), although there was no survey in 2003 and the most recent survey (for 2005) is not yet available.

(Table 1 near here)

A chronological review of the literature on atypical work and its remuneration is provided in Table 1. Many of the studies have a narrow reach, cross-tabulating employment in the various work arrangements with hourly earnings and group health insurance coverage. The study by Cohany (1998) in row 6 of the table is representative. It charts little change in the characteristics of those employed in such activities over time (specifically, between the first and the second CAEAS), and in comparing their median weekly earnings with those of workers in open-ended employment reports that agency temporaries and oncall workers fare particularly poorly. For example, the median weekly earnings of agency temporary workers are only two-thirds of those of regular workers. That said, and again consistent with the findings from the first

CAEAS, independent contractors and contract workers enjoy a wage premium of between fifteen and twenty-one percent.

Cohany's cross tabulations also uncover a sharp shortfall in atypical worker fringes. Health insurance coverage rates range from a low of seven percent (agency temporary workers) to a high of fifty percent (contract workers). And in terms of eligibility to participate in employer pension plans, temporaries are once again at the low end (eleven percent) and contractors at the high end (forty-six percent) of the scale. By contrast, more than sixty percent of regular workers had employer-related health insurance coverage, and more than fifty percent were eligible for a pension plan.

Differences in the compensation packages associated with the various employment arrangements may of course reflect differences in worker characteristics. Thus, for example, Cohany reports that atypical workers are younger, have lower educational attainments, and are more concentrated in the lower-paid industries and occupations than are regular workers. Using two of the same compensation measures as Cohany (viz. wages and health insurance coverage), the study by Polivka et al. (2000) in row 8 adopts a multivariate analysis to control for such differences in worker characteristics. The authors' OLS wage regressions put the wage disadvantage of agency temporary work at (negative) five to nine percent.[2] In the light of Cohany's much higher estimates, these cet. par. results suggest that much of the difference in wages between temporary agency work and regular work is explained by differences in worker characteristics. Yet Polivka et al. find no such attenuation for those workers at the opposite end of the earnings spectrum: contractors are still estimated to enjoy a wage *premium* of twenty-three percent, at least in 1997, on par with the results obtained by Cohany.

In their separate probit analysis of health insurance, Polivka et al. report a shortfall in atypical worker health benefits. In a specification that also includes the worker's hourly wage, it is further reported that its coefficient estimate is positive and well determined, suggesting that workers do not trade health insurance access for wages. Expressed differently, poorly-paid workers seemingly receive a poor benefits package. That said, the deficit in access to an employer's health insurance was halved once coverage through other means (such as a spouse or private insurance) was accounted for.

Polivka et al. (2000, p. 77) recognize the limitations of the dataset, noting that compensation is influenced by a number of variables not contained in the CAEAS, such as "firm-specific factors, personal tastes, and other unobserved differences that might influence who is in these arrangements."[3] The analyses of Segal and Sullivan (1997, 1998), summarized in rows 4 and 7 of Table 1, in part anticipated these concerns. Although limited to a comparison of agency temporary workers with their counterparts in regular employment, and lacking data on fringes, these two studies provide were the first to control for individual unobserved heterogeneity in atypical worker earnings determination. Thus, the study in row 7 of the table exploits administrative data (rather than the CAEAS), extracting a ten-year sample panel from the 1984-94 quarterly records contained in the Washington State Unemployment Insurance system and identifying temporary employees through their industry affiliation.[4] A comparison of the pooled and fixed effects regression estimates suggests that that accounting for permanent unobserved individual heterogeneity reduces the earnings deficit of temporary workers by between ten and fifteen percent.

Finally, in a labor market analysis of single-parent female welfare recipients initially obtaining atypical work in the temporary help service sector versus other industries, Heinrich, Mueser, and Troske (2005) report that this choice does not prejudice their future earnings development or continued employment – or for that matter welfare recidivism. Welfare recipients beginning work in this sector do earn substantially less than their counterparts in other sectors, but this difference does not seem to be the result of unmeasured characteristics. Moreover, the low earnings are not permanent: after two years the differences between those initially in atypical work are virtually the same as their counterparts who had jobs in other industries. This faster earnings growth is shown to be partly the result of atypical workers moving to other higher-paying industries. And there is no difference in the proportions of workers who do not have a job one year later across industries, including temporary help. The bottom line from this study is that welfare recipients obtain opportunities for future advancement by working in the temporary help service sector.

In the present study, we seek to build and improve upon the existing literature in a number of ways. First, unlike the cet. par. studies of Segal and Sullivan and Heinrich, Mueser, and Troske (in rows 4, 7, and 11 of Table 1) each of our two datasets allows us to examine a wider set of alternative work arrangements than agency temporary employment alone. Moreover, our one truly longitudinal dataset offers a wider array of controls than available in the Current

Population Survey, with more human capital variables as well as proxies for worker ability and screening.

Our second contribution to the analysis of the effect of alternative work arrangement on pay is technical in nature. Polivka et al. elect not to use the natural log of a worker's hourly earnings as their dependent variable given certain distributional aspects of the data. More precisely, the right-skewness in the earnings of independent contractors prevented them from directly estimating the Mincerian semi-logarithmic earnings function that has found such strong support in the literature. We adopt the [simplest] technique outlined in Blackburn (2005) that takes the variances of the earnings attaching to the various work forms into consideration, without abandoning the conventional Mincerian earnings function.

Finally, we use additional rounds of the CAEAS to test the robustness of the findings of Polivka et al. (row 8). In particular, we will seek to establish whether or not their more optimistic findings still hold using updated information from the 1999 and 2001 CAEAS.

III. The Empirical Models

To assess the impact that atypical work has on the compensation package, we conduct separate analyses of (hourly) earnings and (access to) employer-provided health benefits. Ideally, we should like to estimate a compensation model in which the dependent variable combines the *dollar value* of the health insurance coverage offered with the worker's wage. Unfortunately, this approach requires data on such things as the quality of the coverage offered and its cost to the worker, information not contained in either dataset used here. For its part, a simultaneous equations approach was ruled out because we were unable convincingly to identify a variable influencing the wage rate but not the probability of being offered health insurance (and vice versa). In line with Polivka et al. (2000), however, we will present results for a specification of the health-benefits equation that includes a wage argument.

Wage Determination

Consider the underlying wage determination model that includes worker ability

$$E(w_{i,t} \mid x_{i,t}, AWA_{i,t}, c_i) = \beta' x_{i,t} + \delta AWA_{i,t} + c_i, \qquad \text{t=1,2, … T} \qquad (1)$$

where $w_{i,t}$ is the (log) wage earned by worker i at time t, $x_{i,t}$ are the corresponding observed worker characteristics, $AWA_{i,t}$ is a dummy variable equal to one if a worker i is engaged in an alternative work arrangement at time t (zero otherwise), and c_i is worker ability. The parameter δ is the wage differential that is associated with employment in an AWA.

When equation (1) is estimated by OLS, we have

$$w_{i,t} = \beta' x_{i,t} + \delta AWA_{i,t} + c_i + u_{i,t}. \tag{2}$$

Absent controls for worker ability, OLS will estimate

$$w_{i,t} = \beta' x_{i,t} + \delta AWA_{i,t} + v_{i,t}, \tag{3}$$

where $v_{i,t} = (c_i + u_{i,t})$. Equation (3) will still provide unbiased estimates of δ provided $E[v_{i,t} \mid x_{i,t}] = 0.$ However, we will still have to correct the estimates obtained from equation (3) to allow for differences in the wage variances observed across the various alternative work arrangements. We remit discussion of this issue to the Data Appendix.

Related to equation (3) is the random effects linear estimator. It allows us to aggregate our three waves of longitudinal data into one pooled sample, assuming the cross-sectional differential has not changed over time. Although this estimator does not take unobserved individual heterogeneity into account, it will produce unbiased and consistent coefficient estimates by allowing for the cross-correlation in the error term that arises when we have repeated observations on the same individual.

If, as may be hypothesized, ability and employment arrangements are negatively correlated, however, the estimate of the wage differential will be biased downward. We can remove worker ability from the model using a fixed effects specification that will also be estimated alongside equation (3). The fixed effects specification allows for not only individual-specific intercepts but also year-specific intercepts, as follows

$$w_{i,t} = \alpha_t + \phi_i + \beta' x_{i,t} + \delta AWA_{i,t} + u_{i,t}, \tag{4}$$

where α_t captures the impact if any that time has on worker earnings and where the individual-specific intercept ϕ_i which controls for any time-invariant unobserved worker characteristics such as ability. Any elements of $x_{i,t}$ that are unchanging over time are omitted from (4).

Employer-Related Health Insurance

In analyzing the question of access to group health insurance (HI), the correlation between ability and employment in an AWA may again bias the estimate of δ. Since the dependent variable is now dichotomous, we use the logit model

$$HI_{i,t}^{*} = \beta' x_{i,t} + \delta AWA_{i,t} + c_i + u_{i,t}, \tag{5}$$

where *HI* is observed to be one if *HI**>0, zero otherwise. Ability is again represented by c_i.

$$\Pr(HI_{i,t} = 1) = \frac{e^{\beta' x_{i,t} + \delta AWA_{i,t}}}{1 + e^{\beta' x_{i,t} + \delta AWA_{i,t}}}. \tag{6}$$

If c_i is omitted, the fitted model, will yield an estimate of δ that may not be the true differential. To handle this concern, we also estimate the following fixed effects logit model

$$\Pr(HI_{i,t} = 1) = \frac{e^{\phi_i + \beta' x_{i,t} + \delta AWA_{i,t}}}{1 + e^{\phi_i + \beta' x_{i,t} + \delta AWA_{i,t}}}.$$

Familiarly, equation (7) differs from (6) in including a worker-specific intercept, allowing for a consistent estimate of the true value of δ.[5]

IV. The Data

We use two main datasets to estimate the differential attaching to atypical work: the Contingent and Alternative Employment Arrangement Supplement (CAEAS) to the Current Population Survey (CPS) – as well as the CPS itself – and the National Longitudinal Survey of Youth, 1979 Cohort (NLSY79). The advantage of the former dataset is its size, given the relatively small proportion of workers in certain of the alternative work arrangements. The disadvantage is that there is no overlap of households across the supplements, ruling out panel estimation methods. Each attribute is reversed in the case of the latter dataset.

The CAEAS/CPS Wage Data

We extract one sample from each of the four biennial CAEAS supplements. To allow for the inclusion of the 2001 CAEAS[6] and to increase the number of regular workers included in our analysis, we link each February CAEAS to the subsequent March basic CPS survey using the matching algorithm outlined in Madrian and Lefgren (1999). The CPS collects information on industry, occupation, and wages only for those regular workers who are in an outgoing rotation group. The CAEAS collects these data for atypical workers without regard to their rotational status. We link the two surveys to allow for the inclusion of regular workers for whom the crucial wage data is elicited in either February or March. For those regular workers who are observed as being in the March outgoing rotation group, we include only those who are observed in the February CAEAS/CPS data, who report having the same employer (and being employed) in both months, and who have the same activities and occupation in each month. The four cross sections were also pooled to obtain more precise estimates of the wage differentials attaching to the different work arrangements, while accounting for year effects. Note that the four samples included only those individuals who were employed in the week prior to their February interview. This restriction was imposed because those recorded as unemployed, or as non-participants, would not report any labor force or wage data.

The wage variable is the hourly wage rate. It is constructed following the detailed procedure outlined in Polivka (1999). Since for some workers the calculation of this hourly wage variable involves dividing total weekly earnings by usual hours worked weekly, we omitted multiple job-holders. Further, workers either reporting or having an imputed hourly wage rate of less than two dollars an hour and more than one hundred and fifty dollars an hour were also excluded from the samples. Two other restrictions were the inclusion of only those individuals aged twenty-five to sixty-five years at the time of the February survey (to avoid compounding different supply responses) and, familiarly, the exclusion of individuals with incomplete demographic, industrial, and occupational data.

Workers are classified according to seven mutually exclusive work arrangements. The first two categories pertain to open-ended employment and comprise regular workers and screened workers. Following the convention established in the literature, we initially distinguish between five types of atypical workers: agency temporary workers, direct-hire temporary workers, oncall workers, contract workers, and independent contractors.

Regular workers are those individuals who are directly hired into an open-ended employment arrangement using standard interviewing methods, rather than having first been screened through an alternative work arrangement. Individuals are classified as *screened workers* if they are currently engaged in open-ended employment *and* they indicate that immediately prior to being employed in this capacity they were employed by the firm in an alternative work arrangement (i.e. without any break in employment continuity). We distinguish between the two types of open-ended employment to allow for the possibility that initially serving an employer as an atypical worker strengthens the bond between employer and employee, and thereby influences the wage paid to such workers.

Returning to the five atypical work categories, *agency temporary workers* are those workers who rely on a third-party, the temporary help service, to secure their employment or who receive their paycheck from a temporary help service.[7] *Direct-hire temporary workers* are those individuals who are in a job temporarily due to an economic reason and who are hired directly by a company rather than through a staffing intermediary. (Note that there is no specific question in the supplement relating to direct-hire temporary work; rather, as described by Polivka et al. (2000, p.p. 42-43), this synthetic category is constructed from a series of questions in the CAEAS.) As a practical matter, we will subsequently aggregate these two temporary categories into a single *temporary worker* composite.

Oncall workers work for a firm on a per-diem or as-needed basis (day laborers are also folded into this classification). For their part, *contract workers* differ from independent contractors in that they, like their agency temporary brethren, rely on a third-party to provide them with the necessary clients or projects. Following the convention in the literature, we also impose the following restrictions for this category: a contract worker needs to have only one client and usually work at that client's workplace. Finally, those we describe as *independent contractors* are self-employed consultants and contractors, responsible for the acquisition of their own clients or projects.

In addition to these five categories we shall also construct some (other) composites, either to test the hypothesis that alternative employment hold uniform implications for a worker's earnings or to facilitate comparison with the NLSY79 (see below).

The NLSY79 Wage Data

The National Longitudinal Survey of Youth, 1979 Cohort (NLSY79) is the product of repeated interviews with individuals aged 14-21 years at the time of the initial interview in 1979 (and therefore just beginning to enter the labor market).

Four different samples were extracted from the NLSY79 for the wage analysis. First, three different cross-sections were created to run the standard OLS regressions, covering employees in employment in 1994, 1996, and 1998. (In 2000, there was no atypical worker question.) The restrictions imposed were the same as those applied to the CAEAS/CPS. There was one exception: we exclude from the NLSY79 samples those workers who had accumulated less than nine weeks' tenure with their current employer. This restriction is imposed because the survey does not report employer-specific information (e.g. the firm's industry affiliation or the worker's occupational classification) from the respondent unless he or she meets this particular service threshold. The wage rate is measured on an hourly basis, but is constructed from information on earnings and usual hours worked relevant to the frequency at which the individual is actually paid.

Second, for the random and fixed effects estimates, an unbalanced panel was constructed, covering all three years used for the cross-sectional analysis and applying same restrictions in respect of missing data and the truncation of the wage distribution. This unbalanced panel construct does not require that workers be recorded as employed for all three years, only for at least two of the three years. It was designed to permit a more precise estimate of δ by including as many observations on individuals as possible.

Individuals were initially grouped into six possible job categories rather than the eight constructed from the CAEAS/CPS. The first two categories of *regular workers* and *screened workers* are identical and, as before, the distinction is based on the notion that previously screened workers may be in a better job match. The same is true for agency temporaries and direct-hire temporaries that are subsequently combined into a single *temporary worker* category. But we are unable to distinguish between the two types of contract work, so we will here speak of *contractors/consultants*. The remaining category is the catch-all of *other work types*, of which the most numerous subgroup is self-employment. Other aggregations are discussed in the next section.

Data on Health Insurance

In discussing compensation, health insurance *coverage* is more central than the offer of *access to* employer-provided benefits. However, data limitations (not least for identifying instruments) lead us here to focus on the latter indicator. Whether or not the employee picks up coverage hinges on a number of factors such as it being offered in the first place, the quality of the coverage provided, the fraction of the premium paid for by the employer, and the availability of other sources of coverage. While extant research does suggest that atypical workers are systematically less likely to be covered by health insurance than regular workers (see Table 1), we are on safer ground in focusing on the question of whether or not they are offered coverage by their employer. Answers to this question can offer guidance as to whether employers are using these alternative work arrangements to reduce costs. To the extent that workers are trading off health insurance for wages, however, there may no savings to the employer purse. We will therefore also need to include the hourly wage as a regressor in modeling access to benefits. If its coefficient estimate is negative, this is prima facie evidence of such tradeoffs are being made by workers. If the point estimate is negative, on the other hand, the suggestion would be that high wages and fringes go together, thereby underscoring any finding of negative wage differentials.

Using a sequence of questions in the CAEAS, we can identify a worker's eligibility to participate in their employer's health insurance plan. These questions, however, are asked only of workers identified as wage and salaried workers, a restriction that limits our ability to include the many independent contractors that are self-employed but which does not otherwise much denude the ranks of the remaining alternative work categories vis-à-vis the samples used in the wage analysis. In short, we are confident in relying on the CAEAS to produce results that are reasonable base estimates of access rates to employer-related health insurance plans for these atypicals.

The health insurance question in the NLSY79 is more direct. Specifically, the worker respondent is simply asked whether or not the employer offered access to health insurance benefits. The terms of the offer are not identified. In investigating the effects of atypical work on employee access to benefits, we will use much the same samples as are employed in our cross-sectional and fixed effects earnings analyses.

V. Findings

Descriptive wage data for the four CAEAS/CPS cross sections and the three waves of NLSY79 data are given in Tables 2 and 3, respectively. Beginning with the larger CAEAS/CPS samples, we see that oncall and temporary workers earn lower wages than do regular workers. For oncall workers, the raw wage differential is between (negative) nine and eleven percent, while temporary workers earn between ninety and ninety-seven cents on the regular worker dollar. At the other end of the wage spectrum are contract workers and independent contractors. Contract workers enjoy wages that are seven to fourteen percent higher than open-ended employment. The wage premium attaching to independent contracting ranges from nineteen to twenty-eight percent.

(Tables 2 and 3 near here)

The earnings of atypical workers vis-à-vis regular workers are directionally the same in the NLSY79 data, but as shown in Table 3 the differences are systematically sharper. For example, the negative wage differential associated with temporary employment now ranges between thirty-five and forty-three percent, whereas workers in contracting/consulting earn a premium of between twenty-two and forty-four percent. Across both datasets, however, there is only a weak suggestion of a wage premium accruing to those who first worked for their employer as an atypical worker. Relative to non-screened regular counterparts, screened workers earn at best four percent more (Table 2) and, at worst, four percent less (Table 3).

Work diaries maintained by the NLSY79 respondents provide us with some additional human capital controls not found in the CAEAS/CPS data. Specifically, we have data on a worker's (cumulative) general labor market experience as well as his/her tenure on the current job.[8] Moreover, we can estimate a standardized measure of the number of jobs held by individuals by dividing the reported total number of jobs held by (cumulative) general labor market experience. This standardized number of *jobs* variable can be also viewed as an inverse proxy for the attractiveness of the worker to an employer. Descriptive information on each of these variables is provided in Table 4 for the 1994 wave of the NLSY79. Not surprisingly workers in atypical work arrangements have substantially less tenure with their employers than do regular workers. In terms of general labor market experience, it would appear that contractors/consultants have been employed slightly longer than regular workers. And, despite their having spent fewer years in employment, temporary workers have held more jobs on average than those engaged in open-ended employment.

We next turn to the multivariate cross-section results, beginning with those for the larger CAEAS/CPS samples. We report (albeit without further ado) the results of the semi-logarithmic model in Table 5.[9] As was flagged earlier, the coefficients produced by this model may not be consistent estimates of the actual wage differential. For dichotomous variables, such as the alternative work arrangements, differences in the dispersion of the dependent variable (in this case the natural log of the wage rate) may correlate with the model's error term. Given this error dependence, it would be improper to rely on the often used convention of subtracting one from the exponentiation of the (biased) coefficient estimate to provide estimates of the cet. par. effect of atypical employment on earnings. (For the formal argument and evidence on the need for this correction, see the Data Appendix). Accordingly, in Table 6 we report estimates of the wage differential that have been corrected for this error dependence using the procedure outlined in Blackburn (2005). To obtain the differential in percentage terms, one now simply has to multiply the estimates in Table 6 by one hundred.

Note further that for each of the CAEAS/CPS cross sections, we estimate two specifications. The first specification – reported in the odd-numbered columns – uses the full sample of workers in each cross section. The second specification – reported in the even-numbered columns – excludes those workers who had a wage component, such as hourly or weekly earnings, allocated to them by the CPS. These allocations occur where a worker does not provide his or her wage information. As shown by Hirsch and Schumacher (2004), the use of allocated workers may cause the estimates of the wage differential to be biased toward zero. When the CPS allocates an earnings component to a worker, it does so using four characteristics (age, gender, education, and occupation) to impute that individual's earnings. Although these variables are good predictors of an individual's wage rate, Hirsch and Schumacher show that coefficients estimated for those characteristics upon which the CPS does not rely to allocate earnings, such as atypical work, will be biased when analyzing workers with allocated wage components. For example, given the fairly small number of workers engaged in *oncall* work, we will understand the (negative) effect that this work form has on earnings given the high probability that the CPS will allocate an oncall worker the wage rate earned by a *regular* worker by the CPS were we to include allocated workers in the sample.

(Tables 5 and 6 near here)

Focusing therefore on the differentials reported in Table 6, we see that controlling for worker characteristics slightly increases the wage penalty attached to temporary employment. For the pooled sample in columns (9) and (10), we see that the linear estimator places the negative wage differential attaching to temporary employment at around six percent. The F-tests at the foot of Table 5 reject the possibility that agency temping and direct-hire temping have the same implications for workers earnings. When we estimate separate coefficients for the two types of temporary employment, for the pooled sample (and again correcting for error dependence) we find that agency temping reduces a worker's earnings by fully sixteen percent. Employment as a direct-hire temporary worker, on the other hand, serves to reduce wages by only four percent vis-à-vis open-ended employment. Yet we choose to aggregate these two work forms for two reasons. We do so first to maintain consistency with the results for the NLSY79 where aggregation of these two categories was not contraindicated. Second, estimates obtained from the CAEAS/CPS for a sample designed to match the NLSY79 age cohort (i.e. including only those workers aged 30 to 35) failed to reject the hypothesis that employment as an agency temporary played the same role in earnings determination as employment as a direct-hire temporary. (Results available from the authors upon request).

Returning to the results for the other alternative work arrangements in Table 6, we see that the negative implications of oncall work for a worker's hourly earnings disappear in all but one cross section. Only in 1997, then, do we see a statistically significant negative wage differential – of around three to four percent. Recall that our simple cross tabulations in Table 2 suggested that engaging in oncall work rather than open-ended employment reduced the individual's wages by nine to eleven percent. In other words, more than one-half of the raw wage difference is explained by differences in the observed characteristics of workers selecting oncall work.

We observe a similar attenuation of the simple *positive* wage premiums associated with both types of contract work (cf. Table 2). Again focusing on the results obtained from the pooled sample in columns (9) and (10), we see that electing to work as a contract worker serves to increase the worker's earnings in the range ten to thirteen percent. The corresponding differential for independent contractors is fifteen to sixteen percent. Across all cross sections, we fail to find any evidence that being screened initially as an atypical worker before being offered regular employment has any effect on the worker's wage.

Two final observations should be made. First, the notion of there being a composite atypical work category is contradicted. As shown in the hypothesis tests at the foot of Table 5, different work arrangements play distinct roles in earnings determination. Second, note that excluding those workers with allocated wage components does not materially affect the estimated wage differentials.

(Tables 7 and 8 near here)

Turning now to the analysis of the NLSY79 earnings data, we again report two sets of results – for the semi-logarithmic specification in Table 7,[10] and for the correction suggested by Blackburn (2005) in Table 8. Again focusing on the adjusted differentials, we find that the negative wage differential associated with temporary employment is generally well-determined and between five and fifteen percent. Taking into consideration the characteristics of those engaging in temporary employment nearly halves the raw estimates of the differential (cf. Table 3). The reduction in the wage premium of contracting/consulting workers is more modest. Across the three NLSY79 samples, the linear estimator places the higher wages enjoyed by contractors and consultants at between twenty-one and thirty-six percent. There is now some slight indication that screened workers earn more than their counterparts hired directly into regular employment. At best, they enjoy a four percent wage premium and, at worse, suffer a three percent wage penalty.

As was the case in the CAEAS/CPS data, we find no support for the use of a composite atypical work argument. The results contained at the foot of Table 7 clearly show that temporary employment (agency and direct hire temps) has a significantly different effect on worker earnings from contracting/consulting and other work types, Even if the two categories of temporary employment can be aggregated. It would be inappropriate, therefore, to speak of an aggregate across all these categories.

Except for tenure, the additional human capital controls found in the NLSY79 data explain very little of the variation in wages across workers. We report these coefficient estimates in Table 7 because the continuous variables are unaffected by the error dependence correction. Greater firm-specific human capital, as measured by tenure, significantly increases earnings. Each year spent with an employer adds to a worker's earnings in the range of 3.5 to 4.1 percent. Although of the expected signs, the coefficients estimated for both general labor market experience and the number of jobs held by an individual are generally poorly determined.

Unique to the NLSY79 is the availability of information on the respondent's Armed Forces Qualification Test (AFQT) test scores. The AFQT is a set of standardized tests used by the military to assess the abilities and knowledge of recruits in the following areas: general science, arithmetic reasoning, word knowledge, paragraph comprehension, numerical operations, coding ability, auto/shop knowledge, knowledge of mathematics, mechanical comprehension, and electronics information. We formed four separate categories for worker ability. Thus, for example, we combined the scores on the word knowledge and paragraph comprehension tests. The resulting aggregate score was then regressed on a set of age and education dummies at the time the test was administered, which were allowed to have nonlinear effects. The residuals from this regression were used as a proxy for a worker's *verbal ability* in the OLS wage regression. Measures of *mathematical ability*, *coding ability*, and the catch-all of *practical ability* were derived in a similar manner.[11]

(Table 9 near here)

The results of the OLS regressions including these observed ability measures are given in Tables 9. For expositional convenience, we report only the corrected estimates. Although coding and math ability are generally significant in explaining the variation in wages across workers, they do not materially alter the point estimates attaching to the various work forms. In line with the results produced in Table 8, we still arrive at an estimate of the wage penalty associated with temporary employment of between six to fourteen percent when we include these ability proxies, while the contracting/consulting wage premium is twenty-one to thirty-four percent (Table 10). The virtual absence of any role for measured ability in influencing differentials for atypical work may well mean that standardized testing does not control for certain key aspects of worker heterogeneity (such as motivation and labor force attachment) – or, in the limit, do other than confirm the individual's ability to take the relevant tests. An alternative possibility is that ability and employment in an atypical work arrangement are not negatively correlated.

We next turn to panel estimators for further consideration of whether employment in an alternative work arrangement and unobserved factors (including ability) are correlated. As noted earlier, two such panel data estimation techniques are employed. As a baseline, the random effects linear estimator provides us with more precise estimates of the wage differential attaching to atypical work in the cross section by pooling the three waves of NLSY79 data into a single sample. Since this specification does not take into consideration any bias that is induced by unobserved individual heterogeneity, we also employ the fixed effects linear specification. We present results correcting for error-dependence and also allow the effect of atypical employment on earnings to differ by gender.

(Table 10 near here)

The summary random effects least squares results obtained from the full sample of workers given in the first column of Table 10 indicate that temporary workers have wages that are seventeen percent lower than those of regular workers, while contracting/consulting workers enjoy a premium of thirty-one percent. As can be seen from the fourth column of the table, however, the wage penalty (premium) attaching to temporary work (contracting/consulting) worker are sharply reduced in magnitude after we control for unobserved individual heterogeneity. Now the wage penalty for associated with temporary employment is only some five percent – a reduction of more than twelve percentage points. The wage premium for contracting/consulting work falls by eleven percentage points to 20 percent.

The corresponding results for male workers are provided in the second and fifth columns of Table 10. In line with the results obtained for the full sample, the wage penalty for temporary employment is reduced after unobserved heterogeneity is taken into consideration: from about (negative) twenty-seven percent to eighteen percent. Interestingly, there is almost no attenuation in the wage premium for contracting/consulting work over regular employment, which is steady at around thirty percent. However, the results for females in the third and final columns of the table are very different from those of males. The negative and positive differentials found for female temporary workers and female contractors/consultants, respectively, in the baseline random effects specification are reversed in the fixed effects estimation. It seems that female temporary workers have unobserved characteristics that lead to lower earnings, and conversely for female contractors/consultants. In the former case, these characteristics most likely include a preference for flexible employment by reason of family or household demands as well as marginal attachment to the labor force. The fixed effects results would suggest that this group of females actually fare better when employed in temporary jobs than regular work. As far as female contractors/consultants are concerned, the modest negative differential observed in fixed effects can be taken as more indicative of their being of higher ability. If higher ability females are disproportionately opting for contracting/consulting work, then one would expect to find an upward bias in the cross-section estimate of the wage effect.

(Tables 11 and 12 near here)

We turn in conclusion to the issue of access to health insurance on the job, which may clearly benefit the worker even if the employer does not directly (or indirectly) contribute to its cost by virtue of the lower premiums charged for group insurance. Descriptive information on employer-related health insurance access rates from the CAEAS/CPS and the NLSY79 are reported in Tables 11 and 12. Across both datasets, it is evident that those in alternative work arrangements are less likely to be eligible for health insurance than are regular workers. From the CAEAS/CPS, we see that forty-five to sixty-one percent of temporary workers have access to employer-related health insurance. Independent contractors have the lowest access rates of twenty-one to thirty-four percent. Of the four types of atypical work identified in the CAEAS, contract workers seem to fare the best, with fifty-three to sixty-percent of them being eligible for employer-related health insurance. Note, however, that more than eighty percent of regular workers have access to these benefits. Similar but again sharper disparities between atypical workers and those in regular employment are found in the NLSY79.

(Table 13 near here)

We now turn to multivariate analysis to disentangle the impact of worker characteristics and work arrangement on the probability being offered health insurance coverage. Two sets of logit estimates are provided, with and without the worker's hourly wage. As noted earlier, we include the wage variable to test for the possibility that workers may be trading access to health insurance for higher wages. The results for the CAEAS/CPS are reported in Table 13. As can be seen, the point estimates for the four types of atypical work are generally well-determined and negatively signed. Focusing on the results obtained from the pooled sample, we see that temporary workers are twelve to eighteen percentage points less likely than are regular workers to be eligible to participate in an employer's insurance plan. Independent contractors have the largest deficit vis-à-vis open-ended employment. The reduction in their likelihood of being extended access to benefits ranges between thirty-one and thirty-nine percentage points. On net, controlling for observed characteristics almost halves the raw deficits in access as between open-ended employment and the various types of atypical work. And consistent with the results obtained by Polivka et al. (2000), we find no evidence that workers are trading health insurance access for higher earnings. Rather, we find that the probability of access to employer-related health insurance benefits is increasing in the wage. In other words, the suggestion is that low-paying jobs also attract lower fringes.

(Table 14 near here)

A similar attenuation of the differences in access to benefits as between atypical workers on the one hand and those in open-ended employment on the other is found for the NLSY79, once we control for the characteristics of workers selecting into these alternative work arrangements. The logit estimates are given in Table 14. It can be seen that, relative to regular workers, temporary workers are seventeen to twenty-one percentage points less likely to be eligible to participate in an employer's health insurance plan. For their part, contractors/consultants are twenty-five to thirty-seven percentage points less likely to have access than regular workers.

Across both datasets, there is now some evidence to favor the argument that initially serving an employer in an atypical arrangement prior to being hired as a regular worker may result in higher compensation. That is to say, relative to their non-screened counterparts, screened workers are significantly more likely to have access to employer-provided health insurance benefits: four to ten percentage points more in the CAEAS/CPS, and six percentage points more in the NLSY79 (see the first row entries in Tables 13 and 14, respectively).[12]

(Table 15 near here)

Finally, we turn to our panel estimates. As was case with the random effects linear estimator, the random effects logit model allows us to aggregate the three NLSY79 waves to provide more robust estimates. And we again rely on the fixed effects specification to statistically control for unobserved heterogeneity. We also run separate equations for male and female workers to examine whether the effects of atypical work on access to health insurance benefits differ by gender. These panel estimates are reported in Table 15. As can be seen, the samples used to estimate the fixed effects logit are considerably smaller than those for the random effects specification. This is because identification of the effect of atypical work on access to benefits using fixed effects requires not only a change in eligibility but also in atypical work status over the three year sample period. Comparing the all-worker results, fixed effects estimation produces smaller effects throughout than does random effects. Before taking account of unobserved heterogeneity, temporary workers (contractors/consultants) are a little more than thirty-five (fifty) percentage points less likely to be eligible to participate in an employer's health insurance plan. After controlling for unobserved heterogeneity, the estimate of the reduced access of temporaries (contractors/consultants) is twenty (twenty-three) percentage points. On this occasion, however, disaggregating by gender yields few surprises. Generally, although women are somewhat less likely to be eligible than males, there is the same hierarchy in the effect of alternative work arrangement on access to benefit and much the same reduction in the disparity by atypical work arrangement once account is taken of unobserved individual heterogeneity.

VI. Conclusions

Our analysis extends prior research into the effects of atypical employment on compensation. Thus, we have provided both cross section and panel estimates for a full array of atypical work forms, corrected the earnings function estimates for skewness in the earnings distribution, and incorporated two additional rounds of data of the CAEAS. With these improvements we can confirm the *broad* findings of Polivka et al. (1999) and Segal and Sullivan (1997, 1998). That is to say, much of the differences in compensation packages associated with atypical work and open-ended employment can be explained by taking into consideration both observed and unobserved differences in the types of workers filling these alternative work arrangements.

The wage results obtained from the CAEAS/CPS data would imply that atypical work is not a monolithic entity. Temporary employment negatively influences a worker's wage to the tune of six percent, while independent contractors enjoy a fifteen to sixteen percent wage premium. Using pooled data from the NLSY79 we find that the negative (positive) wage differential attaching to temporary employment (contract work) is seventeen (thirty-one) percent. But there is considerable diminution in the absolute magnitude of each differential once we control for unobserved characteristics as ability, motivation, or lifestyle preferences: the negative differential for temporary work falls to 5 percent and the advantage associated with contract work declines to twenty percent.

When we allow for atypical work to differ in its implications by gender, we find only a mild attenuation in the wage effects of engaging in temporary and contracting/consulting work in the case of males: for example, both the random and fixed effects linear models produce a wage premium for contracting/consulting of about thirty percent. However, we see a clear overstatement of the wage penalty (premium) attaching to temporary (contracting/consulting) work in the case of females. Our fixed effects estimates suggest that females are actually positively rewarded for engaging in temporary work to the tune of thirteen percent while yet confronting a modest wage penalty (of around three percent) if employed as a contractor or consultant.

Balancing these generally more optimistic results, however, is our finding of relatively large disparities in the access of atypical workers to employer-related health insurance benefits. While we find that the magnitude of the deficit is nearly halved when we control for worker differences, even the smallest estimate – that obtained from a fixed effects model – estimates that temporary and contract workers are more than twenty percentage points less likely to be have access to this fringe benefit than are regular workers. Taken in conjunction with the results from an earlier literature, this reduced access to health benefits flags a potential source of concern and merits further study.

Endnotes

[1]We do not consider here wage gaps associated with the operation of segmented labor markets because of the absence of an anchor, or permanent worker analogue.

[2]Noting a large degree of skewness in earnings, particularly for contractors, Polivka, Cohany, and Hipple (2000) use hourly wages rather than the natural log of the hourly wage in their OLS regressions. In our calculations above, we estimate the percentage differential by taking the dollar estimate from the authors' OLS regressions and then expressing this value as a percentage of mean of reported regular worker earnings in either 1995 or 1997.

[3]On the issue of firm-specific factors in benefits determination, see Lautsch (2003).

[4]As noted by Segal and Sullivan (1998), one key advantage in using the unemployment insurance administrative data over the CPS is that the source of the information on industrial affiliation is the paycheck-issuing entity (for temporary workers, this would be the THS agency). A concern with worker-reported data (as with the outgoing rotations of the CPS) is that agency temporary workers may cite the industrial affiliation of their client firm rather than that of their true employer – the temporary agency. In such cases, researchers will fail to identify temporary workers.

[5]One technical note concerning the fixed effects logit is warranted. To identify the effect that atypical work has on access to health insurance, an observed change is required in not only the dependent variable, but also in work arrangement. This restriction will limit the number of useable observations for this analysis.

[6]In 2001, due to a CPS programming error, the survey was not administered to the outgoing rotation group. For the 2001 sample, we include only those regular workers who are identified in both the February CAEAS and the outgoing rotation group in March.

[7]This last condition led to the inclusion of the miniscule fraction of the agency's workers who are engaged in open-ended employment and are paid by the temporary help service. As noted by Houseman and Polivka (2000), a 1989 Industry Wage Survey indicates that these workers comprise only 3.2 percent of an agency's total employment.

[8]The NLSY79 gives the actual number of weeks that a respondent has been employed since entering the survey, as well as the actual number of weeks employed with the current employer. This allows us to control for general human capital using actual work experience, and firm-specific training using a worker's tenure with the employer.

[9]Full OLS results for the pooled CAEAS sample are provided in Appendix Table 1.

[10]Full OLS results for the 1998 NLSY79 sample are provided in Appendix Table 2.

[11]Construction of the mathematical ability measure first required that we sum across the scores for the respondent's arithmetic reasoning, knowledge of mathematics, and numerical operations tests. In similar vein, practical ability was derived from adding the component scores for general

science, auto/shop knowledge, mechanical comprehension, and electronics information. Only the coding measure involved no initial summation.

[12]As was the case for wage compensation, the inclusion of proxies for a worker's ability does little to affect the point estimates of the probability that an atypical worker will be eligible to access an employer's insurance plan. The proxies themselves are imprecisely estimated and fail to explain the differences in worker access rates.

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DATA APPENDIX

The standard convention in calculating the cet. par. wage differential involves the regression of the log wage on a vector of worker characteristics, *x*. In using the log wage, rather than the wage in level terms, we are assuming that the underlying wage determination model is

$$E[w_i \mid x_i, AWA_i] = e^{x_i'\beta + \phi AWA_i}, \quad (1)$$

where w_i is the wage rate earned by worker *i*, x_i is a vector of worker *i*'s (observed) characteristics, and AWA_i is a variable equal to one if the worker is engaged in an alternative work arrangement (zero otherwise). The estimated model is, therefore,

$$E[\log(w_i) \mid X_i, AWA_i] = x_i'\beta + \phi AWA_i + \varepsilon_i, \quad (2)$$

where ε_i is a standard error term. This model has strong theoretical support and provides consistent estimates of β and ϕ if $E[w_i \mid \varepsilon_i] = 0$. However, it may not yield consistent estimates of the differential attached to atypical work.

After estimating (2), the differential attaching to, say, atypical employment would be obtained via

$$\delta_{AWA} = e^{\phi} - 1. \quad (3)$$

As shown in Blackburn (2005), this approach is flawed if

$$E[e^{\varepsilon_i} \mid AWA_i = 1] \neq E[e^{\varepsilon_i} \mid AWA_i = 0]. \quad (4)$$

This possibility can be tested by taking the residuals from the estimation of equation (2), squaring them, and then regressing the squared terms on the same elements of X_i and AWA_i as follows

$$E[\varepsilon_i^2] = x_i'\lambda_1 + \lambda_2 AWA. \quad (5)$$

If the estimation of equation (5) yields statistically significant estimates of λ_2, then (3) will not yield a consistent estimate of δ_{AWA}; specifically, a positive (negative) coefficient estimate of λ_2 would imply that any estimates of δ_{AWA} using (3) are over- (under-) stated.

Blackburn shows that a more appropriate estimate of δ_{AWA} is

$$\delta_{AWA} = e^{\phi} e^{\frac{\lambda_2}{2}} - 1, \tag{6}$$

where the variance of δ_{AWA} is computed as

$$\text{var}(\delta_{AWA}) = [(\delta_{AWA} + 1)^2 * \text{var}(\phi)] + \frac{1}{2}[(\delta_{AWA} + 1)^2 * \text{var}(\lambda_{AWA})]. \tag{7}$$

The table below shows the results of fitting equation (5) for both datasets. It indicates that the use of (6) is required:

OLS Estimates of Error Dependence (Dependent Variable: Squared Error Term)

(a) Pooled CAEAS/CPS data			
Screened workers	-0.036*** (0.010)	Temporary workers	0.075*** (0.006)
Oncall workers	0.140*** (0.010)	Contract workers	0.025 (0.018)
Independent contractors	0.267*** (0.005)		
(b)1998 NLSY79 data			
Screened workers	-0.034* (0.020)	Temporary workers	-0.021 (0.042)
Contractors/ consultants	0.225*** (0.038)	Other work types	-0.088 (0.071)

Table 1: Selected U.S. Studies Examining the Effect of Alternative Employment on Worker Compensation

Study	Data	Methodology	Compensation measure(s)	Work form(s) examined	Findings
1. Hipple and Stewart (1996)	1995 Contingent and Alternative Employment Arrangement Supplement to the Current Population Survey (CAEAS/CPS).	Cross tabulations alone.	Median weekly earnings, health insurance coverage from employer, and eligibility for employer-related pensions.	Agency temporary, oncall, and contract workers, independent contractors and regular/traditional workers.	Median weekly earnings of agency temporary workers ($290) and oncall workers ($386) are lower than are those of regular workers ($480). Contract employment ($512) and independent contracting ($518) carry a premium. All forms of non-traditional employment offer lower benefits in terms of employer-related health insurance and pension coverage. Among agency temporary workers only 6% have health insurance from the employer and just 7% have pension coverage. The corresponding values for regular workers are 61% and 49%, respectively.
2. Rothstein (1996)	1994 National Longitudinal Survey of Youth (1979 Cohort).	Cross tabulations.	Average hourly earnings, and average weekly hours.	Agency and direct-hire temporary workers, contractors/consultants, and regular workers.	Atypical work is associated with lesser experience: temporary workers spent a lower fraction of the weeks prior to their job in employment (46%) as compared with regular workers (76%). The wages of temporary workers are between 65 and 74% percent of those earned by regular workers.
3. Nollen (1997)	1984 Industry Wage Survey, and 1994 Occupation Compensation Survey of Temporary Help Supply Services.	Cross-tabulations.	Average hourly wages.	Agency temporary employment and non-temporary (i.e. regular) employment.	In 1994 agency temporaries earned $7.74 per hour, versus $11.94 for non-temporaries. In both 1984 and 1994, the differential attaching to temporary work approximated 34%, some of which is allied to the disproportionate number of temporaries concentrated in low-paying occupations such as clerical or laborer positions.
4. Segal and Sullivan (1997)	1983-1994 CPS outgoing rotations.	Cross-tabulations plus cross-section and fixed effects OLS estimations.	Average hourly wages.	Agency temporary employment and non-temporary (permanent) employment.	Agency temporaries earned 78 cents on the permanent-worker dollar in 1993. Cross-section OLS estimates controlling for demographic, industrial, and occupation differences are in the range -8 to -14%. Fixed effects regressions point to a negative differential of just 3%, suggesting that much of the wage difference is attributable to unobserved worker heterogeneity.

Study	Data	Methodology	Compensation measure(s)	Work form(s) examined	Findings
5. Blank (1998)	1995 and 1996 March Current Population Surveys.	Cross tabulations.	Hourly earnings, employer-provided health insurance coverage, access to pension coverage, and average weekly hours.	Agency temporary workers, regular workers.	On average agency temporaries work 36 hours per week, as compared with 43 hours for regular workers, and receive hourly wages that are 70% of those paid to regular workers. Some 24% of temporary workers have health insurance from their employer as compared with 67% in the case of regular workers; and less than 10% of temporaries have pensions compared to 54% of regular workers.
6. Cohany (1998)	1997 CAEAS/CPS.	Cross tabulations.	Median weekly earnings, health insurance coverage from employer, and eligibility for employer-related pensions.	Agency temporary, oncall and contract workers, independent contractors, and regular/traditional workers.	Findings similar to Hipple and Stewart (1996) in row 1. Agency temporary workers earn the least ($329 per week) when compared with regular workers ($510), and oncall work pays 85 cents on the regular-employee dollar. Both groups are younger, and more likely to be female or a minority member than are regular workers. For their part, contract workers and independent contractors have median weekly earnings of $619 and $587, although neither achieves the same rate of employer-related health insurance or pension coverage as do regular workers.
7. Segal and Sullivan (1998)	1984-1994 Washington State unemployment insurance administrative data.	OLS estimation using cross-section data, plus worker fixed effects estimation.	Quarterly earnings and quarterly hours worked used in calculation of average hourly earnings.	Agency temporary employment, and non-temporary (regular) employment.	The cross-section findings point to a well-determined negative earnings differential for agency temporary work in excess of 30%. But controlling for worker-specific fixed effects reduces this to between 10 and 15%.
8. Polivka, Cohany, and Hipple (2000)	1995 and 1997 CAEAS/CPS.	OLS estimation and probit analysis using cross-section data.	Hourly earnings, access to employer-related health insurance, and health insurance coverage	Agency temporary, direct-hire temporary, oncall, contract workers, independent contractors, and regular/traditional workers.	Cross tabulations of hourly wages yield results similar to Cohany (1998) in row 6. After taking into consideration differences in worker characteristics, the differential attached to agency temporary (contracting) work is estimated to be -5% (23%) in 1997. Fewer than 1 in 5 agency temporaries have access to employer's health insurance plan, compared with more than 4 out of 5 regular workers. Benchmarked to regular workers, the probit analysis indicates no attenuation for this group.

Study	Data	Methodology	Compensation measure(s)	Work form(s) examined	Findings
9. Autor (2001)	1994 Occupation Compensation Survey of Temporary Help Supply Services.	OLS estimation using cross-section data, firm fixed effects OLS estimation.	Hourly wages; training in word processing, data entry, computer programming.	Agency temporary employment only.	Temporary workers offered training were paid 2 to 3% lower wages than those not given training. Firm fixed effects results are of a negative differential in excess of 5%.
10. Houseman (2001)	1999 CAEAS/CPS.	Cross tabulations.	Employer-provided health insurance coverage, and access to pension coverage from employer.	Agency temporary, direct-hire temporary, oncall and contract workers, independent contractors, and regular workers.	Of those who have health insurance, agency temporaries are the least likely to have obtained this coverage from their employer (9%). Some 60% of insured contract workers acquired that insurance from their employer, roughly on a par with regular workers (64%). Across all types of atypical work, pension coverage is sharply lower than in regular employment; for example, only 7% of agency temporary workers have this benefit compared with 58% of regular workers
11. Heinrich, Mueser, and Troske (2005)	Administrative data for female recipients of Aid to Families with Dependent Children and Temporary Assistance for Needy Families in Missouri (1993-97) and North Carolina (1997).	Cross tabulations plus OLS earnings estimates, with selection argument for job choice (derived from a multinomial logit) to control for unobserved individual heterogeneity.	Quarterly earnings (the current and subsequent earnings of welfare recipients in the temporary help sector are compared with other employed welfare recipients in other industries). The study also examines employment and welfare dynamics.	Welfare recipients in temporary help service industry and other industries.	Welfare recipients in temporary jobs (defined by sector) receive lower earnings than their counterparts in other jobs initially, but that after 2 years their earnings are virtually identical to those received in other jobs. This implied faster earnings growth arises in part because workers in the temporary help industry are more likely to move into higher paying industries over time. Accordingly, temporary workers have appreciably better prospects than those who are not holding jobs initially (and vis-à-vis those who held jobs in other sectors they are no more likely to be unemployed or have appreciably higher rates of welfare recidivism).

Table 2: Mean Hourly Wage Rates by Employment Arrangement, CAEAS/CPS Data

	1995	1997	1999	2001
Regular workers	13.49 (8.14)	14.04 (8.00)	15.57 (9.67)	17.37 (10.80)
Screened workers	13.46 (7.01)	13.92 (6.46)	16.25 (9.37)	17.49 (10.53)
Temporary workers	12.09 (9.24)	13.58 (10.57)	14.47 (11.95)	16.42 (12.74)
Oncall workers	12.14 (11.50)	12.46 (10.89)	14.24 (12.89)	15.80 (15.29)
Contract workers	14.39 (10.09)	15.36 (9.42)	17.15 (9.03)	19.85 (12.01)
Independent contractors	16.12 (12.79)	18.00 (15.22)	19.53 (16.67)	21.08 (17.74)
n	23,655	22,286	23,627	9,486

Note: Standard deviations in parentheses.

Table 3: Mean Hourly Wage Rates by Employment Arrangement, NLSY79 Data

	1994	1996	1998
Regular workers	13.57 (9.43)	14.94 (11.22)	16.34 (11.80)
Screened workers	13.01 (7.56)	15.29 (9.73)	15.67 (9.92)
Temporary workers	8.85 (4.50)	9.59 (5.36)	9.58 (4.89)
Contractors/ consultants	16.49 (20.38)	21.29 (19.03)	23.53 (23.12)
Other work types	19.79 (15.85)	19.92 (17.82)	17.78 (6.91)
n	5,551	5,702	5,666

Note: See note to Table 2.

Table 4: Labor Market Experience by Employment Arrangement, 1994 NLSY79 Data

	Regular workers	Screened workers	Temporary workers	Contractors/ consultants	Other work types
Experience	13.03	12.56	10.27	12.69	13.11
(in years)	(3.33)	(3.24)	(3.95)	(3.77)	(3.58)
Tenure	5.24	5.01	1.31	3.96	2.11
(in years)	(4.52)	(4.13)	(2.22)	(3.73)	(2.36)
Jobs	0.78	0.85	1.58	0.97	1.08
(standardized) no. of jobs)	(0.56)	(0.57)	(1.26)	(0.55)	(0.54)
n	4,910	427	98	86	30

Note: See note to Table 2.

Table 5: OLS Cross-Section and Pooled Regression Estimates of Atypical Worker Wage Differentials, CAEAS/CPS Data
(dependent variable: log hourly wage)

	1995		1997		1999		2001		Pooled	
Variable	(1)	(2)[1]	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Screened workers	0.009 (0.016)		0.018 (0.015)	0.023 (0.017)	0.003 (0.017)	-0.003 (0.018)	-0.011 (0.030)	-0.008 (0.033)	0.008 (0.009)	0.008 (0.011)
Temporary workers	-0.119*** (0.013)		-0.080*** (0.012)	-0.088*** (0.012)	-0.107*** (0.013)	-0.124*** (0.013)	-0.099*** (0.026)	-0.107*** (0.026)	-0.101*** (0.007)	-0.106*** (0.009)
Oncall workers	-0.070*** (0.025)		-0.076*** (0.024)	-0.088*** (0.025)	-0.075*** (0.027)	-0.074*** (0.028)	-0.107* (0.058)	-0.106* (0.057)	-0.077*** (0.014)	-0.084*** (0.018)
Contract workers	0.047 (0.033)		0.112*** (0.042)	0.107** (0.044)	0.118*** (0.037)	0.113*** (0.039)	0.033 (0.085)	0.027 (0.087)	0.084*** (0.021)	0.101*** (0.028)
Independent contractors	-0.021 (0.014)		0.038** (0.015)	0.037** (0.016)	0.002 (0.015)	-0.008 (0.015)	-0.019 (0.029)	-0.027 (0.029)	0.002 (0.008)	0.007 (0.010)
Allocated?	Y		Y	N	Y	N	Y	N	Y	N
n	23,655		22,286	18,521	23,627	18,813	9,486	7,398	79,054	44,732
Adjusted R^2	0.31		0.31	0.32	0.32	0.32	0.33	0.34	0.32	0.33

Notes: Huber-White standard errors to correct for heteroskedasticity in parentheses.
***, **, * denote significance at the .01, .05, and .10 levels, respectively.
[1]In 1995, the CPS does not provide us with a means of distinguishing between, allocated and nonallocated earnings.
Additional controls are age (and age^2), gender and ethnicity, a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, six educational dummies (the omitted category is no high school diploma), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (the omitted category is residing in the Northeast), ten industry dummies (the omitted category is agriculture/fishing/forestry), and six occupational dummies (the omitted category is manager).

Using the results obtained from the full samples, F-tests were conducted to determine the following restrictions, with test statistic ρ and p-value (in parenthesis) reported below:

	1995	1997	1999	2001	Pooled
$\beta_{AGENCY\ TEMP} = \beta_{DIRECT\text{-}HIRE\ TEMP}$:	ρ=20.89 (p=0.000)	ρ= 7.62 (p=0.006)	ρ= 6.80 (p=0.009)	ρ=0.09 (p=0.760)	ρ=33.60 (p=0.000)
$\beta_{CW} = \beta_{IC}$:	ρ= 3.64 (p=0.057)	ρ= 2.83 (p=0.092)	ρ= 8.88 (p=0.003)	ρ=0.36 (p=0.561)	ρ=14.21 (p=0.000)
$\beta_{TEMP} = \beta_{OC} = \beta_{CW} = \beta_{IC}$:	ρ=13.63 (p=0.000)	ρ=18.31 (p=0.000)	ρ=18.31 (p=0.000)	ρ=2.03 (p=0.108)	ρ=48.54 (p=0.000)

Table 6: Corrected OLS Cross-Section and Pooled Regression Estimates of Atypical Worker Wage Differentials, CAEAS/CPS Data

	1995		1997		1999		2001		Pooled	
Variable	(1)	(2)[1]	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Screened workers	-0.015 (0.013)		-0.005 (0.012)	0.004 (0.014)	-0.013 (0.013)	-0.020 (0.015)	-0.009 (0.021)	-0.007 (0.024)	-0.010 (0.007)	-0.007 (0.009)
Temporary workers	-0.089*** (0.007)		-0.043*** (0.007)	-0.046*** (0.007)	-0.059*** (0.007)	-0.074*** (0.007)	-0.047*** (0.014)	-0.052*** (0.014)	-0.062*** (0.004)	-0.059*** (0.005)
Oncall workers	0.001 (0.014)		-0.029*** (0.013)	-0.039*** (0.014)	0.007 (0.014)	0.014 (0.015)	-0.008 (0.029)	-0.007 (0.030)	-0.007 (0.007)	-0.011 (0.009)
Contract workers	0.047* (0.024)		0.171*** (0.029)	0.177*** (0.031)	0.117*** (0.029)	0.116*** (0.032)	0.039 (0.061)	0.040 (0.063)	0.101*** (0.015)	0.134*** (0.020)
Independent contractors	0.106*** (0.008)		0.198*** (0.008)	0.200*** (0.009)	0.150*** (0.008)	0.140*** (0.009)	0.121*** (0.015)	0.116*** (0.015)	0.145*** (0.004)	0.160*** (0.006)

Notes: See Notes and text to Table 5.

Table 7: OLS Cross-Section Estimates of Atypical Worker Wage Differentials, NLSY79 Data
(dependent variable: log hourly wage)

Variable	1994	1996	1998
Screened workers	-0.011 (0.047)	0.039* (0.021)	0.014 (0.019)
Temporary workers	-0.075 (0.047)	-0.060 (0.046)	-0.149*** (0.043)
Contractors/ consultants	0.052 (0.080)	0.161** (0.075)	0.075 (0.066)
Other work types	0.133 (0.093)	0.233* (0.138)	0.130** (0.058)
Jobs (standardized)	-0.028** (0.013)	-0.014 (0.013)	-0.020 (0.015)
Experience	0.001 (0.009)	0.003 (0.009)	0.004 (0.007)
Experience^2	0.001*** (0.000)	0.001*** (0.000)	0.001*** (0.000)
Tenure	0.040*** (0.005)	0.038*** (0.004)	0.034*** (0.004)
Tenure^2	-0.002*** (0.000)	-0.002*** (0.000)	-0.001*** (0.000)
n	5,551	5,702	5,666
Adjusted R^2	0.40	0.42	0.44

Notes: Huber-White standard errors to correct for heteroskedasticity in parentheses.
***, **, * denote significance at the .01, .05, and .10 levels, respectively.
Additional controls are age (and age^2), gender and ethnicity, a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, education (in years), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (the omitted category is residing in the Northeast), ten industry dummies (the omitted category is agriculture/fishing/forestry), and six occupational dummies (the omitted category is manager).

F-tests were conducted to determine the following restrictions, with test statistic ρ and p-value (in parentheses) reported below:

	1994	1996	1998
$\beta_{\text{AGENCY TEMP}}=\beta_{\text{DIRECT-HIRE TEMP}}$	ρ=3.60 (p=0.058)	ρ=0.00 (p=0.999)	ρ=2.60 (p=0.107)
$\beta_{\text{TEMP}}=\beta_{\text{C//C}}=\beta_{\text{OTHER}}$	ρ=2.90 (p=0.033)	ρ=3.40 (p=0.028)	ρ=6.82 (p=0.000)

Table 8: Corrected OLS Cross-Section Estimates of Atypical Worker Wage Differentials, NLSY79 Data

Variable	1994	1996	1998
Screened workers	-0.025*	0.044***	-0.003
	(0.015)	(0.016)	(0.003)
Temporary workers	-0.058*	-0.052*	-0.147***
	(0.031)	(0.031)	(0.027)
Contractors/ consultants	0.243***	0.359***	0.206***
	(0.050)	(0.051)	(0.039)
Other work types	0.167**	0.391***	0.090
	(0.075)	(0.124)	(0.059)

Notes: See text and notes to Table 7.

Table 9: Corrected OLS Cross-Section Estimates of Atypical Worker Wage Differentials Using Ability Proxies, NLSY79 Data

Variable	1994	1996	1998
Screened workers	-0.030**	0.036	-0.001
	(0.015)	(0.016)	(0.014)
Temporary workers	-0.064**	-0.070**	-0.140***
	(0.031)	(0.032)	(0.027)
Contractors/ consultants	0.213***	0.337***	0.267***
	(0.049)	(0.032)	(0.027)
Other work types	0.190**	0.404***	0.083
	(0.084)	(0.130)	(0.059)
Coding ability	0.022***	0.006	0.003*
	(0.007)	(0.007)	(0.007)
Math ability	0.045***	0.056***	0.049***
	(0.011)	(0.011)	(0.011)
Practical ability	0.013	0.017	0.005
	(0.013)	(0.013)	(0.013)
Verbal ability	0.005	0.008	0.018*
	(0.011)	(0.011)	(0.010)
n	5,280	5,435	5,391
Adjusted R^2	0.41	0.43	0.46

Notes: See text and Notes to Table 7

Table 10: Corrected Panel Estimates of Atypical Worker Wage Differentials, NLSY79 Data

	Random Effects OLS			Fixed Effects OLS		
Variable	Full Sample	Males Only	Females Only	Full Sample	Males Only	Females Only
Screened workers	0.032*** (0.002)	-0.009*** (0.003)	0.087*** (0.003)	0.046*** (0.002)	0.014*** (0.004)	0.087*** (0.003)
Temporary workers	-0.170*** (0.006)	-0.265*** (0.008)	-0.075*** (0.009)	-0.048*** (0.007)	-0.180*** (0.009)	0.131*** (0.013)
Contractors/ consultants	0.313*** (0.010)	0.304*** (0.012)	0.304*** (0.025)	0.196*** (0.012)	0.308*** (0.023)	-0.032* (0.019)
Other work types	0.482*** (0.020)	0.374*** (0.021)	0.785*** (0.057)	0.394*** (0.028)	0.160*** (0.026)	1.222*** (0.173)
Jobs	-0.027*** (0.009)	-0.040*** (0.015)	-0.020* (0.012)	-0.019 (0.029)	-0.001 (0.040)	-0.033 (0.041)
Experience	0.015*** (0.005)	0.020*** (0.008)	0.008 (0.007)	0.087*** (0.029)	0.145*** (0.032)	0.038 (0.049)
Experience^2	0.001*** (0.000)	0.000 (0.000)	0.001*** (0.000)	-0.000 (0.000)	-0.001 (0.001)	0.000 (0.000)
Tenure	0.027*** (0.002)	0.023*** (0.003)	0.034*** (0.003)	0.017*** (0.004)	0.011*** (0.005)	0.023*** (0.006)
Tenure^2	-0.001*** (0.000)	-0.001*** (0.000)	-0.001*** (0.000)	-0.001*** (0.000)	-0.001* (0.000)	-0.001*** (0.000)
n	16,919	9,084	7,835	16,919	9,084	7,835
Adjusted R^2	0.41	0.37	0.42	0.24	0.16	0.28

Notes: ***, **, * denote significance at the .01, .05, and .10 levels, respectively.
Additional controls are age (and age^2), a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, education (in years), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (omitted category is residing in the Northeast), ten industrial dummies (omitted category is working in agriculture/fishing/forestry), and six occupational dummies (omitted category is employment as a manager).

Table 11: Availability of Employer-Related Health Insurance by Employment Arrangement, CAEAS/CPS Data (in percent)

	1995	1997	1999	2001
Regular workers	83.3%	83.1%	84.3%	85.7%
Screened workers	90.0	91.5	93.8	91.4
Temporary workers	45.1	54.8	52.2	60.8
Oncall workers	30.4	36.4	37.0	43.7
Contract workers	64.5	65.6	68.4	53.0
Independent contractors	34.1	30.6	32.5	22.1
n	18,518	18,447	18,718	7,988

Table 12: Availability of Employer-Related Health Insurance by Employment Arrangement, NLSY79 Data (in percent)

	1994	1996	1998
Regular workers	80.3%	81.1%	83.3%
Screened workers	90.4	87.3	88.1
Temporary workers	35.4	35.6	39.9
Contractors/consultants	19.3	29.7	30.4
Other work types	39.7	44.5	80.5
n	5,377	5,433	5,289

Table 13: Logit Estimates of Access to Employer-Related Health Insurance Benefits, CAEAS/CPS Data

Variable	1995		1997		1999		2001		Pooled	
Screened workers	0.554***	0.556***	0.762***	0.738***	0.960***	0.952***	0.439*	0.442*	0.707***	0.698***
	(0.140)	(0.141)	(0.148)	(0.149)	(0.172)	(0.172)	(0.243)	(0.244)	(0.082)	(0.082)
	[0.077]	[0.074]	[0.107]	[0.099]	[0.124]	[0.118]	[0.047]	[0.045]	[0.094]	[0.082]
Temporary workers	-1.633***	-1.599***	-1.346***	-1.320***	-1.298***	-1.274***	-1.315***	-1.309***	-1.415***	-1.389***
	(0.059)	(0.060)	(0.055)	(0.056)	(0.056)	(0.057)	(0.114)	(0.116)	(0.031)	(0.032)
	[-0.228]	[-0.212]	[-0.189]	[-0.177]	[-0.168]	[-0.158]	[-0.141]	[-0.133]	[-0.118]	[-0.176]
Oncall workers	-1.813***	-1.845***	-1.583***	-1.561***	-1.656***	-1.659***	-1.616***	-1.593***	-1.673***	-1.672***
	(0.103)	(0.106)	(0.101)	(0.103)	(0.099)	(0.101)	(0.209)	(0.211)	(0.056)	(0.057)
	[-0.253]	[-0.245]	[-0.223]	[-0.209]	[-0.215]	[-0.206]	[-0.173]	[-0.161]	[-0.223]	[-0.212]
Contract workers	-0.198	-0.220	-0.090	-0.222	-0.402**	-0.460**	-1.221***	-1.367***	-0.293***	-0.359***
	(0.172)	(0.175)	(0.193)	(0.193)	(0.194)	(0.198)	(0.402)	(0.413)	(0.103)	(0.104)
	[-0.028]	[-0.029]	[-0.013]	[-0.030]	[-0.052]	[-0.057]	[-0.131]	[-0.139]	[-0.039]	[-0.046]
Independent contractors	-2.226***	-2.407***	-2.308***	-2.516***	-2.462***	-2.621***	-2.902***	-2.792***	-2.367***	-2.518***
	(0.128)	(0.136)	(0.141)	(0.148)	(0.149)	(0.155)	(0.295)	(0.297)	(0.077)	(0.080)
	[-0.311]	[-0.320]	[-0.325]	[-0.338]	[-0.319]	[-0.326]	[-0.310]	[-0.283]	[-0.315]	[-0.391]
Hourly wage		0.074***		0.073***		0.055***		0.052***		0.064***
		(0.004)		(0.004)		(0.004)		(0.005)		(0.002)
		[0.010]		[0.010]		[0.007]		[0.005]		[0.008]
n	18,518		18,447		18,718		7,988		63,688	
Log L	-7,647.72	-7,448.44	-7,860.00	-7,666.96	-7,655.91	-7,513.90	-2,959.66	-2,904.18	-26,212.52	-25,631.03

Notes: Standard errors in parentheses, marginal effects in brackets.
***, **, * denote significance at the .01, .05, and .10 levels, respectively.
Additional controls are age (and age^2), gender and ethnicity, a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, six educational dummies (the omitted category is no high school diploma), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (the omitted category is residing in the Northeast), ten industry dummies (the omitted category is agriculture/fishing/forestry), and six occupational dummies (the omitted category is manager).

Using the results obtained from specifications omitting wage controls, likelihood ratio tests were conducted to determine the following restrictions, with test statistic ρ and p-value (in parenthesis) reported below:

	1995	1997	1999	2001	Pooled
$\beta_{\text{AGENCY TEMP}}=\beta_{\text{DIRECT-HIRE TEMP}}$:	ρ= 69.61 (p=0.000)	ρ= 90.15 (p=0.000)	ρ= 79.77 (p=0.000)	ρ=22.38 (p=0.000)	ρ=268.43 (p=0.000)
$\beta_{\text{CiW}}=\beta_{\text{IC}}$:	ρ= 96.51 (p=0.000)	ρ= 91.71 (p=0.000)	ρ= 74.19 (p=0.000)	ρ=11.29 (p=0.001)	ρ=277.85 (p=0.000)
$\beta_{\text{TEMP}}=\beta_{\text{OC}}=\beta_{\text{CW}}=\beta_{\text{IC}}$:	ρ=171.41 (p=0.000)	ρ=188.06 (p=0.000)	ρ=170.97 (p=0.000)	ρ=46.38 (p-0.000)	ρ=567.43 (p=0.000)

Table 14: Logit Estimates of Access to Employer-Related Health Insurance Benefits, NLSY79 Data

Variable	1994		1996		1998	
Screened workers	0.469***	0.482***	0.458***	0.447***	0.597***	0.609***
	(0.167)	(0.167)	(0.162)	(0.162)	(0.190)	(0.191)
	[0.059]	[0.060]	[0.056]	[0.055]	[0.064]	[0.062]
Temporary workers	-1.679***	-1.642***	-1.597***	-1.583***	-1.814***	-1.695***
	(0.276)	(0.277)	(0.290)	(0.291)	(0.308)	(0.309)
	[-0.211]	[-0.204]	[-0.196]	[-0.194]	[-0.196]	[-0.173]
Contractors/ consultants	-2.838***	-2.944***	-2.327***	-2.432***	-2.340***	-2.569***
	(0.324)	(0.332)	(0.322)	(0.332)	(0.366)	(0.390)
	[-0.357]	[-0.365]	[-0.286]	[-0.297]	[-0.253]	[-0.262]
Jobs (standardized)	-0.126	-0.119	-0.108	-0.109	-0.260***	-0.241***
	(0.081)	(0.081)	(0.080)	(0.080)	(0.092)	(0.092)
	[-0.016]	[-0.015]	[-0.013]	[-0.013]	[-0.028]	[-0.025]
Experience	-0.090	-0.085	-0.043	-0.040	0.003	0.017
	(0.060)	(0.059)	(0.051)	(0.051)	(0.047)	(0.047)
	[-0.011]	[-0.011]	[-0.005]	[-0.005]	[0.000]	[0.002]
Experience^2	0.007***	0.006**	0.005**	0.005**	0.002	0.001
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
	[0.001]	[0.001]	[0.001]	[0.001]	[0.000]	[0.000]
Tenure	0.299***	0.289***	0.148***	0.143***	0.132***	0.113***
	(0.032)	(0.032)	(0.029)	(0.029)	(0.029)	(0.029
	[0.038]	[0.036]	[0.018]	[0.017]	[0.014]	[0.012]
Tenure^2	-0.013***	-0.013***	-0.005**	-0.005**	-0.003*	-0.003
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
	[-0.002]	[-0.002]	[-0.001]	[-0.001]	[-0.003]	[-0.000]
Hourly wage		0.032***		0.014**		0.058***
		(0.007)		(0.006)		(0.008)
		[0.004]		[0.002]		[0.006]
n	5,377		5,433		5,289	
Log L	-2,131.28	-2,118.86	-2,181.77	-2,178.33	-1,995.46	-1,964.94

Notes: Standard errors are given in parentheses and marginal effects in brackets.
***, **, * denote significance at the .01, .05, and .10 levels, respectively.
Additional controls are age (and age^2), gender and ethnicity, a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, education (in years), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (the omitted category is residing in the Northeast), ten industry dummies (the omitted category is agriculture/fishing/forestry), and six occupational dummies (the omitted category is manager).

Using the specifications that excluded wage controls, likelihood ratio tests were conducted to determine the following restrictions, with test statistic ρ and p-value (in parentheses) reported below:

	1994	1996	1998
$\beta_{AGENCY\ TEMP}=\beta_{DIRECT\text{-}HIRE\ TEMP}$	ρ= 0.03 (p=0.866)	ρ=0.32 (p=0.574)	ρ= 0.00 (p=0.948)
$\beta_{TEMP}=\beta_{C//C}=\beta_{OTHER}$	ρ=11.33 (p=0.010]	ρ=6.25 (p=0.100)	ρ=16.53 (p=0.001)

Table 15: Panel Estimates of Access to Employer-Related Health Insurance Benefits (NLSY79 data)

Variable	Random Effects Logit Full Sample		Random Effects Logit Males Only		Random Effects Logit Females Only		Fixed Effects Logit Full Sample		Fixed Effects Logit Males Only		Fixed Effects Logit Females Only	
Screened workers	0.682***	0.671***	0.554***	0.560***	0.792***	0.791***	0.608***	0.590***	0.751**	0.750***	0.506	0.440
	(0.137)	(0.134)	(0.184)	(0.184)	(0.204)	(0.205)	(0.203)	(0.202)	(0.276)	(0.276)	(0.325)	(0.323)
	[0.109]	[0.106]	[0.087]	[0.088]	[0.127]	[0.127]	[0.096]	[0.094]	[0.118]	[0.118]	[0.081]	[0.071]
Temporary workers	-2.322***	-2.237***	-1.944	-1.879***	-2.662***	-2.645***	-1.285***	-1.288***	-1.024**	-1.025**	-1.501***	-1.509***
	(0.240)	(0.234)	(0.335)	(0.334)	(0.338)	(0.337)	(0.331)	(0.331)	(0.441)	(0.448)	(0.500)	(0.503)
	[-0.368]	[-0.355]	[-0.305]	[-0.295]	[-0.427]	[-0.424]	[-0.204]	[-0.204]	[-0.161]	[-0.161]	[-0.241]	[-0.242]
Contractors/ consultants	-3.235***	-3.281***	-3.093***	-3.160***	-3.376***	-3.564***	-1.417***	-1.414***	-1.422***	-1.424***	-1.802	-1.699
	(0.280)	(0.275)	(0.333)	(0.335)	(0.511)	(0.514)	(0.393)	(0.396)	(0.437)	(0.438)	(1.095)	(1.089)
	[-0.513]	[-0.520]	[-0.486]	[-0.496]	[-0.542]	[-0.572]	[-0.225]	[-0.224]	[-0.223]	[-0.224]	[-0.289]	[-0.273]
Other work types	-0.998**	-1.039***	-0.754	-0.816*	-1.415*	-1.496**	-0.830	-0.787	0.201	0.213	[1]	[1]
	(0.406)	(0.399)	(0.492)	(0.495)	(0.738)	(0.740)	(0.678)	(0.679)	(0.885)	(0.886)		
	[-0.158]	[-0.165]	[-0.118]	[-0.128]	[-0.227]	[-0.240]	[-0.132]	[-0.125]	[0.032]	[0.033]		
Jobs	-0.279***	-0.268***	-0.356***	-0.354***	-0.151	-0.144	-0.011	-0.019	0.566	0.555	0.012	0.002
	(0.079)	(0.076)	(0.115)	(0.115)	(0.107)	(0.107)	(0.420)	(0.422)	(0.766)	(0.766)	(0.585)	(0.589)
	[-0.044]	[-0.043]	[-0.056]	[-0.056]	[-0.024]	[-0.023]	[-0.002]	[-0.003]	[0.089]	[0.087]	[0.002]	[0.000]
Experience	-0.050	-0.047	-0.293***	-0.297***	0.107*	0.115**	0.242	0.251	-0.027	-0.025	0.487	0.503
	(0.044)	(0.043)	(0.074)	(0.074)	(0.059)	(0.059)	(0.222)	(0.222)	(0.347)	(0.347)	(0.315)	(0.316)
	[-0.008]	[-0.007]	[-0.046]	[-0.047]	[0.017]	[0.018]	[0.038]	[0.040]	[-0.004]	[-0.004]	[0.078]	[0.081]
Experience^2	0.006***	0.005***	0.015***	0.014***	0.000	-0.001	0.002	0.002	0.012*	0.012*	-0.005	-0.004
	(0.002)	(0.002)	(0.003)	(0.003)	(0.002)	(0.002)	(0.005)	(0.005)	(0.007)	(0.007)	(0.007)	(0.007)
	[0.001]	[0.001]	[0.002]	[0.002]	[0.000]	[-0.000]	[0.000]	[0.000]	[0.002]	[0.002]	[-0.001]	[-0.001]
Tenure	0.287***	0.271***	0.251***	0.244***	0.320***	0.308***	0.380***	0.387***	0.303***	0.304***	0.477***	0.491***
	(0.025)	(0.024)	(0.033)	(0.037)	(0.037)	(0.037)	(0.047)	(0.047)	(0.065)	(0.065)	(0.073)	(0.074)
	[0.046]	[0.043]	[0.039]	[0.038]	[0.051]	[0.049]	[0.060]	[0.061]	[0.048]	[0.048]	[0.077]	[0.079]
Tenure^2	-0.011***	-0.010***	-0.010***	-0.009***	-0.012***	-0.012***	-0.024***	-0.025***	-0.018***	-0.018***	-0.032***	-0.033***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.004)	(0.004)	(0.005)	(0.005)	(0.005)	(0.006)
	[-0.002]	[-0.002]	[-0.002]	[-0.001]	[-0.002]	[-0.002]	[-0.004]	[-0.004]	[-0.003]	[-0.003]	[-0.005]	[-0.005]
Hourly wage		0.026***		0.026***		0.030***		-0.012*		-0.003		-0.020*
		(0.005)		(0.006)		(0.008)		(0.006)		(0.008)		(0.011)
		[0.004]		[0.004]		[0.005]		[-0.002]		[-0.000]		[-0.003]
n	16,099		8,740		7,359		2,746		1,458		1,288	
Log L	-5,893.27	-5,889.15	-3,126.55	-3,116.40	-2,711.30	-2,703.89	-847.90	-846.23	-459.42	-459.37	-364.52	-362.82

Notes: Standard errors in parentheses, marginal effects in brackets. [1]Omitted due to collinearity issues. ***, **, * denote significance at the .01, .05, and .10 levels, respectively.
Additional controls are age (and age^2), a dummy variable equal to one if married (zero otherwise), an interaction term between gender (being female) and marital status, education (in years), a dummy variable equal to one if residing in an urban area (zero otherwise), four regional dummies (omitted category is residing in the Northeast), ten industrial dummies (omitted category is working in agriculture/fishing/forestry), and six occupational dummies (omitted category is employment as a manager.

Appendix Table 1: Full OLS Regression Estimates of Atypical Worker Wage Differentials, Pooled CAEAS/CPS Data (dependent variable: log hourly wage)

Screened workers	0.009 (0.009)	Temporary workers	-0.102*** (0.007)
Oncall workers	-0.077 (0.014)	Contract workers	0.084*** (0.021)
Independent contractors	0.002 (0.008)	Age	0.038*** (0.001)
Age^2	-0.000*** (0.000)	Female	-0.139*** (0.006)
Black	-0.085*** (0.006)	Other ethnicity`	-0.051*** (0.008)
Married	0.089*** (0.005)	Married female	-0.096*** (0.007)
High school diploma	0.176*** (0.006)	Some college	0.245*** (0.007)
Associates degree	0.292*** (0.008)	Bachelors degree	0.411*** (0.008)
Masters degree	0.507*** (0.010)	JD/MD/PhD	0.548*** (0.014)
Urban	0.507*** (0.004)	Northcentral	-0.042*** (0.005)
South	-0.085*** (0.005)	West	-0.019*** (0.005)
Construction/ mining	0.135 (0.017)	Manufacture	0.147*** (0.016)
Transportation, comm.. and utilities	0.199*** (0.017)	Retail/wholesale trade	-0.078*** (0.016)
Finance, insurance and real estate	0.125*** (0.017)	Business services	0.044*** (0.017)
Personal services	-0.041** (0.018)	Professional services	0.022 (0.016)
Public administration	0.201*** (0.017)	Technical/ sales workers	-0.124*** (0.007)
Clerical workers	-0.252*** (0.006)	Operators/ laborers	-0.333*** (0.007)
Service workers	-0.418*** (0.007)	Skilled labor	-0.154*** (0.007)
1997	0.047*** (0.005)	1999	0.123*** (0.004)
2001	0.201*** (0.006)		
n	78990		
Adjusted R^2	0.32		

Appendix Table 2: Full OLS Cross Section Regression Estimates of Atypical Worker Wage Differentials, 1998 Cross Section NLSY79 Data
(dependent variable: log hourly wage)

Screened workers	0.014 (0.019)	Temporary workers	-0.149*** (0.043)
Contractors/ consultants	0.075 (0.066)	Other work types	0.130** (0.058)
Experience	0.004 (0.007)	Experience^2	0.001*** (0.000)
Tenure	0.034*** (0.004)	Tenure^2	-0.001*** (0.000)
Jobs (standardized)	-0.020 (0.015)	Age	-0.028 (0.084)
Age^2	0.000 (0.001)	Female	-0.109*** (0.018)
Black	-0.051*** (0.014)	Hispanic	-0.015 (0.017)
Married	0.119*** (0.017)	Married female	-0.128*** (0.023)
Education	0.065*** (0.003)	Urban	0.005 (0.013)
Northcentral	-0.103*** (0.019)	South	-0.145*** (0.017)
West	0.001 (0.021)	Construction/	0.300*** (0.056)
Manufacture	0.266*** (0.053)	Transportation, comm. and utilities	0.307*** (0.055)
Retail/wholesale trade	-0.016 (0.054)	Finance, insurance and real estate	0.301*** (0.058)
Business services	0.232*** (0.056)	Personal services	0.026 (0.060)
Professional services	0.082 (0.054)	Public administration	0.267*** (0.055)
Technical/ sales workers	-0.053** (0.025)	Clerical workers	-0.243*** (0.017)
Operators/ laborers	-0.325*** (0.021)	Service workers	-0.254*** (0.018)
Skilled labor	-0.236*** (0.022)		
n	5,666		
Adjusted R^2	0.44		

Appendix Table 3: Full Logit Regression Estimates of the Determinants of Access to Employer-Related Health Insurance Benefits, Pooled CAEAS/CPS Data

Screened workers	0.707*** (0.082) [0.094]	Temporary workers	-1.415*** (0.031) [-0.188]
Oncall workers	-1.673*** (0.056) [-0.223]	Contract workers	-0.293*** (0.103) [-0.039]
Independent contractors	-2.367*** (0.077) [-0.315]	Age	0.078*** (0.009) [0.010]
Age^2	-0.001 (0.000)	Female	-0.197*** (0.040)
Black	0.209*** (0.040) [0.028]	Other ethnicities	-0.031 (0.051) [-0.004]
Married	0.334*** (0.036) [0.044]	Married female	-0.652*** (0.047) [-0.087]
High school diploma	0.496*** (0.038) [0.066]	Some college	0.595*** (0.042) [0.079]
Associates degree	0.759*** (0.052) [0.101]	Bachelors degree	0.842*** (0.048) [0.112]
Masters degree	1.082*** (0.067) [0.144]	JD/MD/PhD	1.264*** (0.104) [0.168]
Urban	0.178*** (0.026) [0.024]	Northcentral	0.056* (0.034) [0.007]
South	-0.043 (0.032) [-0.006]	West	-0.116*** (0.033) [-0.015]
Construction/	0.171* (0.094) [0.023]	Manufacture	1.687*** (0.092) [0.225]
Transportation, comm. and utilities	1.295*** (0.095) [0.172]	Retail/wholesale trade	0.398*** (0.087) [0.053]
Finance, insurance and real estate	1.064*** (0.096) [0.142]	Business services	0.293*** (0.093) [0.039]
Personal services	-0.069 (0.096) [-0.009]	Professional services	0.970*** (0.087) [0.129]
Public administration	2.261*** (0.108) [0.301]	Technical/ sales workers	-0.369*** (0.041) [-0.049]

Appendix Table 3, Continued

Clerical workers	-0.192 (0.040) [-0.026]	Operators/ laborers	-0.674*** (0.046) [-0.090]
Service workers	-1.103*** (0.041) [-0.147]	Skilled labor	-0.462*** (0.049) [-0.061]
1997	0.009 (0.029) [0.001]	1999	0.097*** (0.029) [0.013]
2001	0.198*** (0.039) [0.026]		
n	63,688		
log L	-26,212.52		

Appendix Table 4: Full Logit Regression Estimates of the Determinants of Access to Employer-Related Health Insurance Benefits (1998 NLSY79 data)

Screened workers	0.597*** (0.190) [0.064]	Temporary workers	-1.814*** (0.308) [-0.196]
Contractors/ consultants	-2.340*** (0.366) [-0.253]	Other work types	-0.102 (0.473) [-0.011]
Experience	0.003 (0.047) [0.000]	$Experience^2$	0.002 (0.002) [0.000]
Tenure	0.132*** (0.029) [0.014]	$Tenure^2$	-0.003* (0.002) [-0.000]
Jobs (standardized)	-0.260*** (0.092) [-0.028]	Age	0.692 (0.600) [0.075]
Age^2	-0.010 (0.008) [-0.001]	Female	0.088 (0.127) [0.009]
Black	0.314*** (0.105) [0.034]	Hispanic	0.025 (0.119) [0.003]
Married	0.523*** (0.117) [0.056]	Married female	-0.607*** (0.163) [-0.066]
Education	0.088*** (0.021) [0.010]	Urban	0.090 (0.092) [0.010]
Northcentral	-0.023 (0.135) [-0.003]	South	0.038 (0.125) [0.004]
West	0.277* (0.146) [0.030]	Construction/ mining	0.534** (0.258) [0.058]
Manufacture	2.174*** (0.257) [0.235]	Transportation, comm. and utilities	1.675*** (0.274) [0.181]
Retail/wholesale trade	0.716*** (0.242) [0.077]	Finance, insurance and real estate	1.357*** (0.290) [0.146]
Business services	0.936*** (0.258) [0.101]	Personal services	0.107 (0.274) [0.012]
Professional services	1.386*** (0.249) [0.150]	Public administration	2.456*** (0.343) [0.265]

Appendix Table 4, Continued

Technical/ sales workers	-0.023 (0.178) [-0.002]	Clerical workers	-0.054 (0.138) [-0.006]
Operators/ laborers	-0.584*** (0.151) [-0.063]	Service workers	-0.582*** (0.128) [-0.063]
Skilled labor	-0.612*** (0.172) [-0.066]		
n	5,289		
log L	-1,994.46		

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