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Hiring credits and employer wage subsidies are tools that policymakers have available to attempt to improve labor market conditions for workers. This study explores how capped-wage subsidies affect firms' labor market decisions, in particular, their reliance on part-time and low-skill workers. We focus on the federal Empowerment Zone program, which offers firms in targeted areas a 20 percent wage subsidy (capped at $3,000 per year) for each employee who also resides in the Empowerment Zone. Results using different methods of identification suggest that firms respond to capped-wage subsidies by expanding their use of part-time workers, particularly where the subsidy cap is likely to bind. We also provide evidence of a shift toward lower-skill workers.

JEL: H25, J23, J48.

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1. Introduction

One set of tools that governments have for intervening in labor markets is subsidies or credits given to employers to hire workers. Such programs could theoretically be used to encourage hiring of disadvantaged workers, spur hiring during economic downturns, or encourage hiring in impoverished areas. Given these potential benefits, it is not surprising that there has been academic interest in studying these programs for several decades (see Kaldor, 1936; literature reviews in Katz, 1998 and Neumark, 2013). The existing evidence, however, does not provide strong support for the effectiveness of such programs, at least in terms of job creation. While some authors find evidence of job creation in certain programs, or for specific populations, or in the short term (for example, Katz, 1998; Jaenichen and Stephan, 2011; Boockmann et al., 2012; Faulk, 2002; Hamersma, 2008), others find little evidence of employment increases, particularly in the long run (for example, Burtless, 1985; Hamersma, 2008). Researchers have discussed and investigated when these programs could be successful, for example, during recessions (for example, Neumark, 2013; Neumark and Grijalva, 2017), or how they can be designed and implemented differently to achieve their goals (for example, Almeida, Orr, and Robalino, 2014; Bartik and Bishop, 2009).

The literature has identified a variety of reasons why this policy may not be effective at improving labor market conditions for workers. There is evidence that targeted workers may be negatively affected because of stigma effects (Burtless, 1985). These programs have also often suffered from low take-up rates by firms (for example, the Targeted Jobs Tax Credit (Bishop and Montgomery, 1993) and the Work Opportunity Tax Credit (WOTC) and Welfare-to-Work Tax Credit (WtW) (Hamersma, 2011)). This could be due to a lack of awareness of the programs or the administrative costs of identifying eligible workers. Using a survey, Hamersma and Heinrich

(2008) find that a large sample of firms were aware of the WOTC program but that it did not influence their hiring. And Hamersma (2011) finds evidence that there appear to be fixed costs of firms' participation in the WOTC/WtW programs, possibly because of minimum job duration requirements.

We aim to study the effect of another common feature of these policies: a limit on the amount of money provided to firms. We do so by looking at the federal Empowerment Zone (EZ) program, which was enacted in 1994. The EZ program provided a 20 percent wage credit, up to $3,000, for each eligible employee at eligible firms. Importantly for us to focus on the credit cap, several features of the EZ program help to avoid some of the potential problems of wage subsidies previously identified. The program chose a set of economically distressed areas within cities, and firms are eligible to receive the credit if the firm is located in the area (or relocated to the area) and the worker is a resident of the area. So the program is designed not to directly target disadvantaged individuals but to target areas with a large proportion of disadvantaged individuals. As noted previously in the literature (see, for example, Katz 1998), this location-based targeting could avoid some of the negative stigma problems that can occur with individual-targeted credits (Burtless 1985).

Another difference is that firms may be more likely to respond to the program than other policies previously studied. First, the wage credit applies to both new and existing workers, so even firms that are not interested in hiring new workers have incentives to claim the wage subsidy. Second, the program allowed firms to collect the credit for at least 10 years, with the possibility for more years if the program was renewed (the EZ program was renewed through 2019). So firms have the potential to receive money for a long time period, making it more likely that the benefits of hiring outweigh any administrative or fixed costs of firm participation in the program.[1]

Our interest, however, is in how the $3,000 cap on the wage credit affects firms' employment decisions, particularly in terms of weekly hours of work. The wage credit cap means that the marginal cost of an hour of work increases at annual earnings of $15,000. We then expect that, assuming firms hire permanent positions, some firms may respond to this change in marginal cost by either employing more workers at less than full-time hours or by hiring workers in lower-wage occupations. Our argument is that if firms face low market wages, then the wage credit cap does not bind. At higher market wages, the wage credit cap binds at less than full-time work and firms have incentives to adjust their hiring. Firms may hire additional workers only at low hours, or they may substitute workers and weekly hours, which allows them to claim more credits for the same total labor hours.

We are unaware of research that has studied the incentives created by capped-wage subsidies for firms to substitute toward part-time and low-wage work.[2] Neumark (2013) notes that targeting credits at employment could create incentives to hire at low hours. Hamersma (2011) notes that few workers in the WOTC/WtW are near the credit cap, so she cannot investigate firms' responses to the cap. Some prior research has tested for rent-seeking behavior by firms. For example, Lorenz (1995) discusses problems in the design and implementation of the Targeted Jobs Creation Tax credit program that allowed rent-seeking. Hamersma (2011) finds no evidence that firms attempt to maximize their credits in their hiring decisions, possibly

[1] As discussed later, there is an extensive literature that attempts to evaluate the success of the EZ program as policy for local economic development. Additionally, the focus of that literature is more about whether the targeted areas improved in some measurable way (for example, housing values, unemployment, poverty). While we use methodologies similar to the ones in this literature, our goal is to use the EZ program to learn more about how firms respond to the design of wage subsidy programs.

[2] Researchers have found that firms respond with part-time work in response to other policies, most notably the Affordable Care Act (for example, Even and Macpherson, 2016).

because it was administratively difficult to identify the subsidized workers or because the WOTC/WtW programs had minimum job tenure requirements that may be less easy to manipulate than weekly hours.[3] Ajilore (2012) studies the WOTC program and finds no evidence of a decrease in employment among workers who he argues are close substitutes for targeted workers. A report by the US Government Accountability Office (2001) finds no evidence that firms fired workers to claim credits on new hires in the WOTC program.

We present evidence that the wage subsidy cap in the federal EZ program is likely to bind for some firms. To investigate whether firms respond by using more part-time work, we use data from the Census Transportation Planning Packages. This is public-use data on workers aggregated to the place-of-work Census tract. It provides information about the hours usually worked by tract workers in 1990 and 2000. Combining these data with other tract information from 1980 and 1990, we employ a difference-in-differences methodology comparing the change in the proportion of workers in EZ tracts working less than 35 hours a week from 1990 to 2000, with the change in a set of comparison tracts that were eligible, but not selected for, the EZ program. We also consider triple-difference models that account for trends in the surrounding cities, and estimate the effects using both OLS regressions and propensity score matching as well as instrumental variable methods for robustness. Differencing methods are commonly used to analyze the effects of wage credits, and the specific empirical strategies we employ are common in the related literature that evaluates the effectiveness of the federal EZ program as a local economic development program (Neumark and Simpson, 2015).

[3] Hamersma (2011) does find that firms with a higher potential exposure to the wage credit, having a higher proportion of workers with job tenures exceeding the minimum thresholds, are more likely to claim the credits. Similarly, we argue that some firms are more likely to respond to the wage subsidy cap based on whether market wages cause the cap to bind.

Our results show a modest increase in part-time work among EZ workers, approximately 1 to 2 percentage points, although the estimates are not always statistically significant. These results would be consistent with a 5 to 10 percent increase off the baseline share of part-time work of approximately 19 percent. However, as discussed, only some firms are likely to face market wages at which the wage subsidy cap binds. We provide estimates showing that reliance on part-time work increases more in tracts with a higher proportion of workers earning at least $15,000 in the tract in 1990. This is consistent with a larger part-time response by firms in areas more likely to face the earnings cap of the subsidy. We also provide evidence of increase in part-time work among men, who are more likely to work part-time for economic reasons, but not among women. Finally, we provide evidence that some firms may have responded by hiring lower-wage and lower-skill workers, which is another way to avoid the subsidy cap. We find evidence that occupations associated with lower levels of education expanded as a result of the EZ program, while the share of occupations associated with higher levels of education shrank.

Overall, we see our findings as evidence that firms were responding to the cap on the wage credit in the federal EZ program. The findings provide further evidence that the implementation and design of wage credit programs are important factors in their success. In Section 2, we discuss the federal EZ program and present a more detailed discussion of potential firm responses. We discuss our data and estimation methodologies in Sections 3 and 4, respectively. In Section 5 we present the results of our estimations and we conclude with a discussion in Section 6.

2. Labor Demand Responses to the Federal Empowerment Zone Program

2.1 Federal Empowerment Zone Program

The federal Empowerment Zone program was created as part of the 1993 Omnibus Budget Reconciliation Act (OBRA 1993, P.L. 103–66) to incentivize firms to locate in and hire workers living in specified distressed areas. State and local governments were encouraged to submit EZ applications that met two criteria: at least 20 percent of the population had to live below the poverty line and the area had to have an unemployment rate of at least 6.3 percent (USGAO, 2004). A total of 78 (Wallace, 2004) applications were received and six urban areas within cities (located in Atlanta, Baltimore, Chicago, Detroit, Philadelphia/Camden, New York) and three rural areas (Kentucky Highlands, Mississippi Delta, and the Rio Grande Valley in Texas) were chosen.[4]

EZ areas received several benefits from the program. However, the most generous and utilized benefit (Hanson, 2011) was a capped refundable wage credit to businesses for operating in the EZ areas and hiring residents of the areas.[5] A business located in the EZ area would be allowed to claim a wage tax credit of 20 percent on the first $15,000 in wages paid to a qualified worker. A qualified worker was someone who lived in the EZ, performed the majority of their work in the EZ, and was employed by the firm for at least 90 days. The wage credit was available both to firms already operating in the EZ areas and to any firm that relocated to the EZ area. Unlike many policies aimed at increasing new employment by restricting the benefit to only new hires, such as the Job Creation Tax Credit, the EZ program allowed firms to apply the wage credit to all qualified workers, both existing and new employees. The wage credit,

[4] EZ areas did not have to be contiguous. See Figure 2 for maps of the New York and Chicago EZ areas.
[5] The other incentives associated with EZ designation in urban areas was a one-time $100 million Social Service Block Grant (SSBG) and a set of relatively small capital incentives. Neither of these portions of the program should induce the labor demand changes that we document in this paper. The SSBGs were spent differently in each city on a variety of projects such as transportation and consulting. The capital incentives included the ability to exclude capital gains on the sale of certain assets and increasing the amount of property that can be immediately expensed. For a complete description of the EZ incentives, see USDHUD, 2001).

therefore, was available to both new and existing employers in the area and was applicable to both existing and new qualified employees.

The fact that the federal EZ program applies not just to newly hired workers could increase take-up rates, although with the possible concern that firms will simply claim the credit and not expand hiring. While targeting job creation is typically desired in wage subsidy policies to avoid rent-capture by firms, the EZ program was designed as a local economic development tool. In that case, while new job creation would be good, wage increases, additional hours, or firm relocation could all be beneficial outcomes. In fact, a wide literature has used the EZ program to investigate the effectiveness of place-based economic development programs, including a broad range of outcomes such as firm movement, property values, and resident outcomes (for example, earnings, employment, and poverty). Neumark and Simpson (2015) review this literature and generally find that place-based policies lead to increases in local property values, while the evidence on resident outcomes is somewhat mixed.

2.2 Potential Firm Responses

For discussion purposes, assume that firms in the EZ areas employ a production function $Q = f(K, E)$ involving capital (K) and total weekly labor hours (E). For expositional ease, assume that all workers are hired for 52 weeks a year. The fact that firms are able to claim the credit for multiple years and only for workers employed at least 90 days helps justify this assumption. Firms are able to choose total weekly labor hours by combining the number of employees (N) and the hours worked (h). We will allow firms to set different hours for different workers so that $E = \sum_{i=1}^{N} h_i$. Further assume that firms operate in a competitive labor market so that the labor cost per worker is $C_i = wh_i$ where w is the market wage. Finally, assume for now that there are no employment fixed costs and no adjustment costs, that the production technology allows firms to substitute between h and N, and that the labor supply is perfectly elastic so that the wage credit does not affect workers' earnings

The wage subsidy in the EZ program changes the marginal expense to the firm of employing an hour of work. Assuming that the worker is hired for 52 weeks a year, the marginal cost of employing an eligible employee for an additional hour of work will be $0.8w$ if $52wh \leq S$, where S is the annual earnings cap for the EZ program ($15,000). If $52wh > S$, then the marginal cost of an additional hour is simply w. This change in the marginal cost of employment could cause firms to adjust their employment decisions. For example, some firms could decide to hire new employees. Other firms could decide to expand the hours of current employees (both new and current employees are eligible for the wage subsidy in the EZ program). How firms respond will depend on the level of the annual salary cap of the subsidy and the level of the wage that the firm must pay. We expect that the weekly hours chosen by the firm will decrease as market wages increase, since firms will reach the maximum wage credit at lower levels of hours worked.

Figure 1 depicts the combination of weekly hours and market wages that exactly equal the federal EZ earnings cap of $15,000 (assuming 52 weeks of work). For any wage greater than $7.21 ($= \frac{\$15{,}000}{52h}$), firms would reach the earnings cap at less than 40 hours of weekly work.[6] For comparison, the federal minimum wage in 2000 was $5.15 and the median wage was $14.12.[7] Exact wage data on the EZ areas are not available, since individual observations are not identified in public-use data at that small level of geography. However, the average wage in the area can be approximated using Census tract data by combining the aggregate earnings in a tract

[6] In our analysis, we will define part-time work as working less than 35 hours a week. The equivalent wage, assuming 35 hours of work and 52 weeks a year, is $8.24.
[7] Authors' calculations of workers ages 16 to 64 with non-zero wages in the 2000 Census, using IPUMS data (Ruggles, et al., 2015).

with the distribution of hours and weeks worked. The average of this calculation for the EZ tracts is $11.71, high enough that firms reach the earnings cap well below full-time employment.

Thus, we expect that firms in the EZ areas will have an incentive to employ a higher proportion of workers in a part-time capacity. First, since labor demand elasticities tend to be larger for the extensive margin than for the intensive margin, then we might expect that wage credits have a larger effect on hiring than on hours (Neumark, 2013). However, given the existing wages, firms responding to the wage subsidy by hiring new workers will likely hire part-time employees, unless they pay well below the area average (for example, at the federal minimum).

Second, firms whose wage and hours combination is beyond the earnings cap frontier in Figure 1 have an incentive to change their total hours composition to employ more workers, but at lower weekly hours per worker. For example, firms that face the average wage of $11.71 and employ workers at 40 hours per week for 52 weeks a year will spend (40*52*11.71) = $24,356.80 per employee. Following the EZ designation, the firm can claim the full $3,000 wage credit for each employee. However, suppose the firm splits each position in half, employing twice as many workers for 20 hours each. Now each employee earns (20*52*11.71) = $12,178.40 and generates a wage credit of (0.2*12,178.40) = $2,435.68. The total wage credits per 40 hours of work are now (2*2,435.68) = $4,871.36, an increase of $1,871.36. Thus, firms attempting to maximize their wage subsidy will have an incentive to move toward part-time work, as long as they face high enough market wages. Firms paying low wages are not able to increase their wage credit by substituting workers and hours.

This discussion leads to the main empirical question the paper is trying to address. Did the earnings cap on the EZ wage credit, in combination with predominant market wages, cause firms to respond by expanding their use of part-time work? While we have discussed this in terms of individual firm responses, it is also possible that the firms that benefit the most from the wage subsidies expand more than other firms. Since firms with more part-time workers benefit more from the subsidy, these firms' share of employment in the EZs would grow in response to the incentives. This is another way in which the capped-wage subsidy could induce an increase in part-time share. Given our data, such a composition change would be observationally equivalent to establishment-level changes. We interpret either a composition change or establishment-level changes as evidence that capped-wage subsidies encourage part-time work.

While the structure of the wage subsidy and the predominant wages in EZ areas would appear to incentivize firms to shift to more part-time work, there are reasons why such a shift may not occur. For example, if there are administrative costs to take up, as found in the WOTC/WtW program by Hamersma (2011), then there would be no overall effect. However, as discussed in the introduction, the fact that the EZ wage credit applies to both new and current eligible workers increases the incentives for firms to take up the wage credit.

Second, if firms face fixed costs for employment or some form of adjustment costs to new hires, then we may see muted responses. Similarly, indivisible labor would tend to limit firms' ability to change the composition of their workforce to maximize their wage credits, because they cannot easily substitute workers and hours. However, it is not clear that these theoretical possibilities are likely to be important in our context. For example, it seems unlikely that employment fixed costs are high enough to prevent firms from responding to the wage credit, especially since part-time workers often do not receive benefits, reducing their compensation costs. Furthermore, we will be investigating worker outcomes six years after the EZ program was created, allowing substantial time for firms to adjust their workforces. And while it may be difficult to substitute workers and hours in some occupations and industries, it may be easier in the lower-skill occupations that we later show expand in response to the EZ program.

Another reason for a muted response is that firms respond on the intensive margin by increasing the hours of current workers. Under our assumption of permanent, full-year employment, the wage credit takes a form of a kink in the marginal cost of an hour of work. A similar kink in the cost of labor is caused by overtime rules, which firms have been found to respond to on the intensive margin (for example, Costa, 2000; Hamermesh and Trejo, 2000). Again, we might not expect as much of a response of hours compared to the extensive margin, given differences in the elasticity of labor demand (Neumark, 2013). Furthermore, an intensive margin effect would be more likely in the face of adjustment or hiring costs that we have argued are unlikely to be prevalent in our application. Finally, while some firms may raise the hours of their employees, this effect is unlikely to cause an increase in full-time work because the earnings cap is achieved below 40 hours for any wage above \$7.21. Again, only firms paying wages significantly lower than the area average are likely to respond by increasing hours of current workers to full-time status.

Another reason that we may not find a shift toward part-time work is that firms will respond not by hiring at low levels of hours but by hiring at low wage levels. Assuming that the firms cannot pay below market wages for a particular job, this would mean that they hire more positions at the bottom of the wage scale, presumably lower-skill workers. These workers can be hired for more hours before hitting the subsidy cap. In addition to exploring a change in part-time work, we also explore shifts in the occupations in the EZ areas. If firms respond on this margin, we would expect to see a relative increase in the proportion of less-skilled workers in these areas. We test for this in Section 5.3.

The discussion so far has assumed that the wage subsidy has had no effect on market wages. A wage subsidy in a market can distort market wages, with the incidence dependent on the relative elasticity of the labor supply and demand. Recent evidence from outside of the United States has found that cuts in payroll taxes in Sweden (Saez, Schoefer, and Seim 2019) and hiring credits in France (Cahuc, Carcillo, and Le Barbanchon 2018) had little effect on market wages, but clear effects on employment. Hamersma (2008) provides evidence that about 38 percent of the wage credit in the WOTC/WtW programs is passed through to workers in the form of higher wages. However, the wage subsidy in the EZ program could have less of an effect on market wages. First, the wage subsidy does not apply to the entire labor market but instead only applies to firms operating within a limited geographic area. It seems unlikely that a wage credit only available to firms in small parts of larger cities would induce changes in market wages. Second, the distressed nature of the EZ areas makes it likely there is initially an excess supply of labor, reducing the likelihood that wages are bid up. Finally, to the extent that wages are pushed up by the subsidy, it would further increase the likelihood of an overall shift toward part-time work, since the earnings cap is achieved at lower hours of work as the wage rises.

3. Data

Studying firms' responses to the wage credit in the federal EZ program is difficult because data on individual firms or workers are not generally publicly available.[8] Instead, we use data from the Decennial Census aggregated to the Census tract level, chosen because the Empowerment Zone applications were defined by tract geography. Much of the existing

[8] That information is only available in the confidential, restricted-access Decennial Census Long-form data that were used by Busso, Gregory, and Kline (2013). Accessing those data is cost-prohibitive for many researchers.

literature on the EZ program uses data from the Summary File data sets of the Decennial Censuses, which are tabulations of population and housing characteristics within Census tracts. In addition to being utilized by previous researchers, these tabulations were also used in the application process for the EZ program. Importantly, however, these tabulations are of the population that *resides* in these areas. Our interest is in the response of firms that *employ* individuals in these areas. While we use some Summary File data from 1980 and 1990 to control for tract characteristics and trends, we construct our outcome measures from the 1990 and 2000 Census Transportation Planning Packages. This is the only publicly available data source for the attributes of workers by place-of-work tract and it enables us to investigate changes in the characteristics of workers who work in the designated EZ areas.[9]

The Census Transportation Planning Package (CTPP) provides counts of workers by place-of-work for a variety of traits. It can be thought of as the equivalent to the Summary Files, except tabulated by place-of-work instead of place-of-residence. The universe for the CTPP is everyone in the Decennial Census Long-form sample (one in six households) age 16 and over who worked for pay in the week before the Census. The CTPP is primarily designed to meet the needs of transportation planners, but it also contains information on workers' traits, including hours worked, annual earnings, occupation, industry, gender, ethnicity, and race.

In our primary analysis, we use the information on hours worked to investigate whether there is a shift toward more part-time work. The data have counts of workers in different ranges of hours usually worked per week. The percent of tract workers with weekly hours less than 35 hours a week will be our primary outcome of interest throughout our analysis.[10] While our

[9] Tract boundaries can change over time. We use a Graphical Information System (GIS) procedure to generate consistent geographies across 1980, 1990, and 2000.

[10] Our results are similar if we define part time as working 20 hours or less per week, but the magnitudes are smaller.

interest is in the share of part-time employment, we will also present evidence on the level of part-time and full-time workers, since the prior literature has found increases in employment in EZ areas (see Busso, Gregory, and Kline, 2013). Furthermore, as mentioned previously, firms could respond to the wage subsidy incentives by increasing hiring primarily at low-wage levels, for which the earnings cap is less likely to bind. Thus, we will use information on the types of occupations to investigate changes in the skill mix of workers that might arise due to the incentives of the capped-wage subsidy.

There are two limitations of the CTPP data that will tend to cause our estimates to attenuate toward zero. One limitation of the CTPP is that it does not allow us to separately identify information about zone employees who live in the EZ area and those zone employees who do not. Because the wage tax credit applies only to zone employees who reside in the zone, it would be ideal to be able to study their hours and earnings separately. Because the data do not allow us to separate zone residents from commuters, we cannot study the effects of the capped-wage subsidy on only the eligible workers. A second limitation is that we only observe total (usual) hours of work across all jobs for workers. If firms respond to the subsidy cap by showing a preference for part-time work, workers may respond by getting a second job. Because we do not observe hours at a particular job, we will fail to identify some shift toward part-time work if workers are using multiple jobs to reach 35 hours or more a week.

Both of these limitations result from the aggregated nature of our data. Since we cannot focus only on the qualified workers, our estimates will contain a number of zero effects from nonqualified workers, attenuating our results toward zero. Additionally, since we do not observe hours at each job, we will miss some proportion of the firms' shift toward part-time work. As discussed later, we find some evidence of the first problem as we tend to find larger effects when we weight the regressions by the share of workers in each tract who are zone residents.[11] However, we cannot solve the second limitation. Thus, our results may underestimate the true effect of changes in part-time work among the workers actually eligible for the subsidy.

4. Methodology

We utilize a difference-in-differences methodology, which compares worker outcomes in EZ areas to a set of counterfactual areas before and after the enactment of the policy. Using the federal EZ program to study how firms react to capped-wage subsidies presents three methodological issues typically found in the place-based policies literature: selection of counterfactuals, controlling for city effects, and endogeneity concerns. In order to correctly identify the effect of the wage subsidies on firms in the EZ areas we would need to know the labor decisions of firms in the EZ areas without the wage subsidy. Since this is impossible, we use as a basis for constructing our counterfactual areas the federal Enterprise Community (EC) program. These were the areas that applied to and qualified for the EZ program but were not chosen. The use of these areas as counterfactuals is common in the EZ evaluation literature (for example, Hanson, 2009, Busso, Gregory, and Kline, 2013; Reynolds and Rohlin, 2015).[12]

EC areas are used to construct counterfactuals for a number of reasons. First, EC areas received some of the benefits provided to EZ areas, but they did not receive the EZ wage subsidy.[13] Comparing EZ areas to locations that were not entitled to any of the program's

[11] The CTPP's Flows tables provide some tabulations for combinations of place-of-work tracts and place-of-residence tracts, but does not provide information on hours, earnings, or occupation. Therefore, we cannot separate zone worker-residents outcomes from zone worker-nonresident outcomes, but we can calculate the share of workers in a tract who are residents of the zone.

[12] Following much of the existing literature on the EZ program, we remove Cleveland, Los Angeles, and Washington DC from our set of EC areas. Cleveland and Los Angeles received EZ status early in the 2000s and Washington DC also qualified for a wage subsidy. However, all results are robust to inclusion of these cities in our set of counterfactuals.

[13] Enterprise Communities received $3 million in Title XX Social Services Block Grant funds and new tax-exempt bond financing. Four Enhanced Enterprise Communities each additionally received a $22 million Economic Development Initiative grant (USDHUD, 2001). We treat Enterprise Communities and Enhanced Enterprise Communities as equivalent for our analysis.

benefits would only further confound our identification of the wage subsidy effect.[14] Second, using counterfactual areas that applied to the program but did not receive the subsidy removes application bias. Third, EC areas are not located in the same cities as the areas that received EZ status. This is important because it has been shown that the EZ program had spillover effects within cities in which the EZ areas were designated (Hanson and Rohlin, 2013).[15]

While the EC and EZ areas all met the minimum qualifications for empowerment zone consideration, the EZ areas were more distressed in a variety of ways, as has been documented previously in the EZ literature (for example, Busso, Gregory, and Kline, 2013; Hanson, 2009; Reynolds and Rohlin, 2014). Table 1 presents summary statistics on characteristics of EZ and EC areas using data from 1980 and 1990. While both EZ and EC areas had high poverty rates and unemployment in 1990, the EZ areas exceed the EC areas in both categories by approximately 6 percentage points. Furthermore, the EZ areas were higher in both categories in 1980 as well, indicating that they were more distressed on average for at least 15 years prior to EZ designation. EZ areas also have lower median incomes, higher rates of receiving welfare, lower homeownership rates, and lower home values than EC areas. EZ areas also have a higher minority rate and a higher rate of female-headed households. These differences mean that while EC areas are useful for comparison, we cannot simply use them as counterfactuals. Consequently, we will include the 1980 and 1990 variables as covariates in our analysis to account for these differences.[16]

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[14] For example, residents of EZ and EC areas ages 18 to 24 were qualified for the WoTC program, so firms could receive up to $2,400 for hiring these residents. Because both EZ and EC areas had the same benefits, our methodology should eliminate any confounding effects from this feature of the WoTC program.

[15] An additional concern about using areas within cities that received EZ status is endogenous selection by local policymakers, since these alternative areas were specifically not included in the original application, despite the fact that they were similarly distressed.

[16] Note that the reliance on part-time workers in both EZ and EC areas falls from 1990 to 2000, although the change is larger for the EC areas. Thus, unconditional estimates comparing the EZ and EC areas over time would suggest a relative increase in part-time work in EZ areas.

The simplest version of our identification strategy compares the change from 1990 to 2000 (the tax incentives were provided in 1995) in outcomes in EZ areas with changes in EC areas over the same time period. By differencing within each geographic area, we remove any time-invariant factors that may affect firms' labor demand decisions, such as access or proximity to transportation hubs and other business amenities that did not change during the decade. In these regressions the dependent variable takes the form:

$$Y_i = Y_{i2000} - Y_{i1990} \tag{1}$$

where i denotes a Census tract in an EZ or EC area and Y is an outcome of interest, for example, the proportion of workers with weekly hours less than 35.

However, controlling for city-wide changes is warranted because EZ areas could be affected by overall economic trends in the cities in which they are located. Therefore, our preferred approach is a triple-difference that compares relative outcomes between EZ areas and the rest of their cities, with EC areas and their cities, before and after the policy. We isolate the effect of the EZ policy from city-specific effects by making a comparison across time. The underlying assumption is that the difference between these applicant areas and their cities would have grown the same amount without the policy. Therefore, the dependent variable used in our regressions accounting for city-wide effects takes the following form:

$$Y_i = (Y_{i2000} - \bar{Y}_{j2000}) - (Y_{i1990} - \bar{Y}_{j1990}) \tag{2}$$

where Y_{i2000} and Y_{i1990} are defined as before and $\bar{Y}_{j2000}$ and $\bar{Y}_{j1990}$ represent the average outcomes in 2000 and 1990, respectively, in the j Census tracts in the respective EZ and EC cities.

The differencing between tract and city in equation (2) eliminates any differences in outcomes due to city-specific changes between 1990 and 2000, including changes in city-wide economic growth, as well as any city-wide local government policies that influence firms' labor demand decisions. Another possible approach to controlling for city-specific changes would be to use similarly distressed areas within the EZ and EC cities. These areas more closely match the distressed EZ and EC areas within each city and, therefore, may more closely capture the local economic trends in distressed areas of the cities. However, these areas may be susceptible to contamination, as Hanson and Rohlin (2013) provide evidence that these areas experience negative spillover effects as firms relocated within cities to the EZ areas. As it is difficult to definitively argue whether these areas are better controls for local trends, we choose to present two different sets of triple-difference estimates: one using the entire remainder of the city and the other using just the economically similar areas within each city.[17]

Our base methodological approach to determine how firms respond to capped-wage subsidies is a standard OLS regression. The estimating equation we use is

$$Y_i = \alpha + \beta EZ_i + X_i'\delta + \epsilon_i \qquad (3)$$

where *i* indicates Census tracts, *EZ* represents a dummy variable for whether the EZ wage subsidy is provided (meaning the Census tract is in an EZ area). A vector of control variables is represented by *X* and includes the 1990 and 1980 variables summarized in Table 1.

We also employ propensity score methods, which are another common approach used in the place-based policy literature to address potential endogeneity concerns (Busso, Gregory, and Kline, 2013; Reynolds and Rohlin, 2015; Neumark and Young, 2019). Propensity score matching is a technique for balancing the distributions of observable covariates through the use

[17] We identify economically similar areas by estimating, within each city, the probability of being an EZ or EC tract using the 1990 and 1980 covariates in Table 1. We then use a nearest neighbor algorithm with a single match to find the tract within the city that most closely matches each EZ or EC tract.

of the propensity score, or predicted probability of treatment conditional on the observable characteristics (Rosenbaum and Rubin, 1983). In our context, we first estimate a logit of

$$EZ_i = \alpha + X_i'\delta + \epsilon_i \qquad (4)$$

and then calculate the propensity score as the predicted probabilities $\hat{P}(X)$.[18] We then use these predicted probabilities in a nearest neighbor matching algorithm, where the estimated treatment effect is the average difference between the outcome in each EZ tract and the average outcome of the *n* closest matches. Specifically, for each EZ tract we find the *n* EC tracts with the closest values of $\hat{P}(X)$. The choice of n involves a bias/variance tradeoff: higher numbers of matches produce more precise estimates but may do so at the cost of higher bias in the estimates. We choose $n = 10$ but limit matches to be within 1 percentage point of $\hat{P}(X)$. Thus, we use up to 10 possible matches but only within an acceptable range of match quality (commonly called the caliper). Our matching method does a good job of balancing covariates, reducing the difference in means for all controls variables between EZ and EC areas, and the remaining differences are usually not statistically significant (see Appendix Table A-1). For inference, we use a block-bootstrap procedure with 1000 replications. All of our main results are robust to our choice of *n* and caliper width.

5. Results

5.1 Part-time versus Full-time Employment

We begin by testing our hypothesis that the subsidy cap will encourage part-time work. Table 2 presents the baseline results of the effect of EZ designation on the proportion of workers employed less than 35 hours a week, across our various specifications. Each cell represents the results of a specific estimation strategy defined by the type of differencing (double difference,

[18] Importantly, we do not directly match on reliance of part-time workers in 1990 because that is part of our differenced dependent variable.

city-wide triple difference, or partial-city triple difference), empirical method (OLS or propensity score matching), and set of covariates (none, 1990, or 1980 and 1990). Standard errors clustered by city are presented for the OLS estimates, while block-bootstrapped standard errors are presented for the propensity score estimates.

The unconditional estimate in the first row of column (i) of the top panel indicates that EZ areas experienced a 1.6 percentage point increase in the proportion of area workers working less than 35 hours a week compared to EC areas. Table 1 indicates that part-time workers made up roughly 19 percent of the workforce in EZ and EC areas; so this is roughly an 8.4 percent increase in part-time workers relative to the baseline. Including 1990 controls in column (ii) has no effect on this estimate, and adding 1980 and 1990 controls in column (iii) slightly reduces the estimate to a 1.4 percentage point increase. These results are consistent with the expectation that the earnings cap on the wage subsidy increased the use of part-time work. The propensity score matching estimates in the second row of the top panel present similar results, although the estimates are only statistically significant when using 1990 controls.

These results present suggestive evidence of an increased use of part-time workers from our simple difference-in-difference strategy between EZ and EC areas. However, these results do not account for city-wide trends in part-time work that might be confounding our estimates. The middle panel of Table 2 presents our triple-differencing strategy that removes any city-wide trends in part-time work. These triple-difference estimates are very similar, if slightly smaller, than the simple difference estimates in the top panel. The bottom panel presents the triple-difference results that use similarly distressed areas within the cities to account for local trends. These estimates are smaller than those in the top two panels, suggesting an increase in part-time work somewhere between 0.7 and 0.8 percentage points, although the estimates are generally imprecise.

The overall pattern of modest increases in the share of part-time work is fairly robust across methods (OLS versus matching) and across differencing strategies. For additional robustness we also consider an instrumental variable approach used by Hanson (2009) and Hanson and Rohlin (2011a, 2011b). The potential worry is that EZ areas were chosen based on their future economic expectations, either positive or negative. The ideal instrument for obtaining EZ status would be correlated with receiving EZ status while being unrelated to our outcomes (the share of workers employed in part-time or low-wage jobs). This robustness check follows Hanson (2009) and uses two instruments: a dummy variable for whether the area had a representative on the United States House of Representatives Ways and Means Committee, and the number of years that the representative was in office at the time of EZ designation. These political variables have been shown to be highly correlated with EZ designation (Wallace, 2004; Hanson, 2009) and are demonstrated to be so in Appendix Table A-2. These instruments are individually significant as well as jointly significant as shown by their respective *p*-values and instrument F-test.[19] The IV estimates in Appendix Table A-3 largely produce estimates similar to the OLS and matching methods, but the estimates are imprecise and not statistically significant.

The estimates provided so far have looked at the share of workers employed part time. While the main focus of our paper, it could miss interesting changes in the number of full-time and part-time workers if firms responded to the EZ program by expanding total employment. Research on the federal EZ program has often found positive effects on total employment (for

[19] We test for over-identification using the Sargan–Hansen J-statistic and find a Hansen J-statistic of 1.547 and Chi-sq p-value of .2136.

example, Hanson, 2009; Busso, Gregory, and Kline, 2013) but recent evidence suggests that the effects are sensitive to specification (Neumark and Young, 2019).[20] To investigate, we constructed our outcomes as the natural log of employment (plus 1 to account for tracts with 0 employment), and can thus be interpreted as percentage changes. We find evidence of an increase in total employment generally between 10 and 14 percent, which is broadly in line with Busso, Gregory, and Kline (2013), although we use a different data set. However, the change is consistently larger across specifications for the number of part-time workers compared to full-time workers (see Appendix Table A-4). Thus, at least part of the shift toward a higher share of part-time work appears to be caused by a larger increase of part-time workers on the extensive margin.

Finally, while we generally find modest effects on the share of part-time work, we suspect our main results in Table 2 on the share of part-time employment could be attenuated as discussed previously. For example, the fact that our outcome potentially includes both subsidy-eligible and ineligible workers would attenuate our results. We find some evidence for this in Appendix Table A-5, where we weight our estimates based on the proportion of workers that we know live in EZ areas (based on CTPP data). Essentially, we would expect that tracts with very few eligible workers would show less response compared to tracts with many eligible workers. We generally find larger effects of EZ designation on part-time work using such weights. So there may be some attenuation due to our inability to identify the effects among eligible workers. Furthermore, we suspect that we are missing workers who have multiple part-time jobs, equating to over 35 hours of total work. This would also attenuate our results, but the data do not enable us to address this concern. The other caveat is that there are reasons to suspect that these average

[20] Studying French enterprise zones, Briant, Lafourcade, and Schmutz (2015) find that employment effects also depend on how spatially integrated the zone area is with the surrounding city.

effects mask important heterogeneity in the response of firms, which we explore in the next two sections.

5.2 Results by Subsidy Cap Constraint

Our discussion in Section 2 argued that the prevailing wages in the EZ areas would tend to incentivize the use of part-time work. However, we know that there is variation in wages within the EZ areas and we argued that the incentives to increase the use of part-time work should increase with wages. We can test the role of the prevailing wages by using information on the percent of workers in each tract whose annual earnings are at, or above, the subsidy cap of $15,000. More workers above that threshold in 1990 would suggest that firms in these areas have jobs that are more likely to be constrained by wages and the wage subsidy cap, thereby encouraging part-time work. So the distribution of the annual earnings of workers in 1990 allows us to test whether the earnings cap has an effect on the hiring decisions of firms.

To perform this test, we put tracts in three categories based on the proportion of workers in 1990 earning at least $15,000. The first group is tracts with less than 50 percent of workers above the salary cap; the second group is tracts with 50-75 percent of workers above the cap; and the last group is tracts with more than 75 percent of workers above the cap. These groups loosely correspond to tracts where few firms are constrained, more firms are constrained, and many firms are constrained, respectively. We then include dummy variables for these categories, and their interaction with the EZ dummy variable, in our OLS regression framework. Our left-out category is the first group, the least constrained tracts. For simplicity, we present only the two triple-difference models, which account for local trends, and those that include covariates (we do not report the naïve estimates, those without covariates, or the results of the double-difference models; all the results are similar and available upon request).[21]

The first two columns of Table 3 present the results from the triple-differencing strategy that accounts for city-wide trends. The coefficient on the EZ dummy variable represents the effect on the reliance on part-time workers in EZ tracts that are least constrained (have less than 50 percent of workers above the earnings cap). Regardless of controls, the coefficients suggest that these less constrained tracts experienced a decrease in part-time work among employees, approximately a 3.1 percent decline. The coefficients on the noninteracted terms suggest that, on average, part-time use increased across the EZ/EC tracts from 1990 to 2000 relative to the least constrained tracts.[22]

Importantly, the interaction terms suggest that as the earnings cap in the EZ areas begins to bind more, as more workers earn above the cap, then the use of part-time work increases among workers. The coefficient on the interaction of the EZ dummy variable with the middle category of tracts (50-75 percent of workers above the cap) is consistent with a 4.2 percentage point increase in part-time work relative to the less constrained tracts. The coefficient on the interaction term for the highest category (greater than 75 percent of workers above the cap) suggests a 6.4 percentage point increase in part-time work, relative to the least constrained tracts. So, while both EZ and EC areas saw an increase in part-time work in the higher-income tracts,

[21] Our cutpoints put most tracts in the middle category: approximately 12 percent of tracts are categorized as least constrained, 74 percent of tracts are in the middle category, and 14 percent of tracts are in the most constrained category. We chose our cutpoints both because we feel that they are easy to interpret but also because they roughly match the inflection points from a model that fully interacts a cubic function of the percent of workers making over $15,000 with the EZ indicator. The results from that specification, while somewhat less easy to interpret, match the overall results from our categorical specification.

[22] This could be due to higher-wage areas expanding more in part-time work because they face higher labor costs in general or because of the industrial mix in these areas. Estimates show that these tracts have a higher proportion of service and retail jobs, which are two sectors more likely to employ part-time workers (part-time workers are approximately 42 percent for retail workers, 31 percent for service workers, and 14 percent for all other workers, based on authors' calculations from the 1990 and 2000 Census).

the effect is roughly double for the EZ tracts. The results using the triple-differencing strategy that accounts for partial-city trends, presented in columns 3 and 4, are very similar, if a bit larger.

These results are broadly consistent with the discussion of firm responses in Section 2.2. Areas where fewer workers are above the earnings cap are areas where firms are likely to face lower market wages for the types of jobs at their firm. Such firms are unlikely to be constrained by the earnings cap when making decisions about weekly hours. Therefore, they are more likely to either expand worker hours to full time or to hire new employees at full-time hours. As the constraint begins to bind for more workers, consistent with higher hourly wages, firms are less likely to hire at full time and have a stronger incentive to substitute between hours and workers to maximize the wage credit. Thus, while Table 2 presents suggestive evidence of small increases in part-time work in EZ areas, the evidence in Table 3 demonstrates that firms adjust part-time work in both directions, depending on whether the earnings cap is binding.

5.3 The Occupation and Skill Mix of Workers

To this point our results suggest that when firms are offered wage subsidies that are capped toward the lower part of the annual earnings distribution ($15,000), they respond by increasing their reliance on part-time workers. As discussed in Section 2.2, firms could also respond to the subsidy cap by hiring more low-wage/low-skill workers. If this occurred, the share of lower-skill and less-educated workers employed in the area would correspondingly increase. [23] The CTPP data that provide place-of-work information at a small geographic scale do not provide education information, but they do provide the occupational mix of the area.

[23] In theory, we could also test whether there is an increase in workers making earnings below $15,000. However, as previously discussed, we cannot identify earnings at a particular job and workers can respond by taking on multiple jobs. The results of this test, presented in Appendix Table A-6, show some evidence of an increase in workers earning less than $15,000 but the results are somewhat sensitive to estimation strategy.

Given differences in skill requirements and wages paid across occupations, we can investigate changes in occupational mix to infer information about firm responses across skill and wage.

Table 4 presents triple-difference estimates of workers in different occupations. We have ordered the occupations from left to right in decreasing order of the nationwide proportion of workers with a BA. Overall, findings suggest that there was an increase in workers in occupations that predominately employ uneducated workers while simultaneously finding a decrease in workers in occupations that predominately employ educated workers. Specifically, estimates suggest an increase of between 0.5 and 2.2 percentage points in the "production" and "transportation" occupations; only 5 percent of US workers in either of these occupations had at least a college degree. Likewise, the most educated occupation, "professional," had a decrease of 2.0 to 2.7 percentage points. We get similar results if we create rough terciles of occupation by educational attainment and estimate the changes in the share of workers in each education group. There is a decrease in the share of the highest education occupations and an increase in the share of the lowest education occupations (see Appendix Table A-7). Additionally, the evidence on occupational shifts may also suggest a differential change in work for men versus women. For example, men are more likely to be working in production and transportation occupations, which saw increased reliance in the EZ areas. As a robustness check, Appendix Table A-8 provides evidence that the substantive change in the use of part-time workers is driven by the hours worked by men. Together, we take the evidence of changes in the skill mix and gender composition of EZ workers as further evidence that the structure of the wage subsidy induced firms to change their hiring patterns.

6. Conclusion

We discuss how a capped-wage subsidy creates an incentive for firms to use part-time or low-wage workers in order to get the maximum subsidy possible. We then use federal Empowerment Zones (EZs) to test whether firms increase their use of part-time workers in response to a capped 20 percent wage tax credit. EZs are a good way to test this because the combination of rejected applicant areas and the literature evaluating the EZ program allow us to construct good comparisons for the EZ tracts, and the EZ wage credit is relatively large, long-lived, and less prone to stigma effects than other hiring tax credits. As far as we know, this is the first paper to test the economic incentives of the wage subsidy in the EZ program, as the existing EZ literature has focused on evaluating the program as a tool for local economic development.

We find evidence of modest increases in the use of part-time workers in EZ areas from 1990 to 2000, approximately equal to 1 to 2 percentage points, or a 5 to 10 percent increase relative to the baseline. However, that average effect masks important heterogeneity. We find that EZ tracts where the $15,000 annual earnings cap is least likely to bind saw a relative decrease in part-time work, while the EZ tracts where the cap is more likely to bind have a large increase in the prevalence of part-time work. From this we conclude that, at least for the EZ program, the capped-wage subsidy led to a shift to more part-time employment. Complementary results show that firms in EZs shift toward lower-skill workers – another way to stay under the earnings cap and maximize the subsidy received.

While inducing firms to shift to using part-time workers is an economic distortion from the capped-wage subsidy, it is unclear how policymakers would view this distortion. Distributing the employment benefits over a larger number of workers may be in line with policy goals, even if it means fewer hours per employed worker. Encouraging firms to hire lower-skill workers, who often have more limited job prospects than higher-skill workers, may also be consistent with policymakers' goals.

If policymakers do not want to induce firms to shift to part-time workers, they could cap the credit by hourly wage instead of by total earnings. If structured this way, the credit would be the same for each marginal hour of employment, eliminating the incentive to use part-time workers. However, the incentive to use lower-wage workers would remain because wages beyond the hourly wage cap would not increase the amount of wage subsidy received. To give a specific example, suppose that there is a 20 percent wage credit applied to the first $10 of each eligible worker's hourly wage. Then the credit would be 20 percent for all wages up to $10 per hour and less than 20 percent for all wages beyond $10 per hour. In this way, the credit could be targeted to lower-wage workers while not inducing part-time employment.

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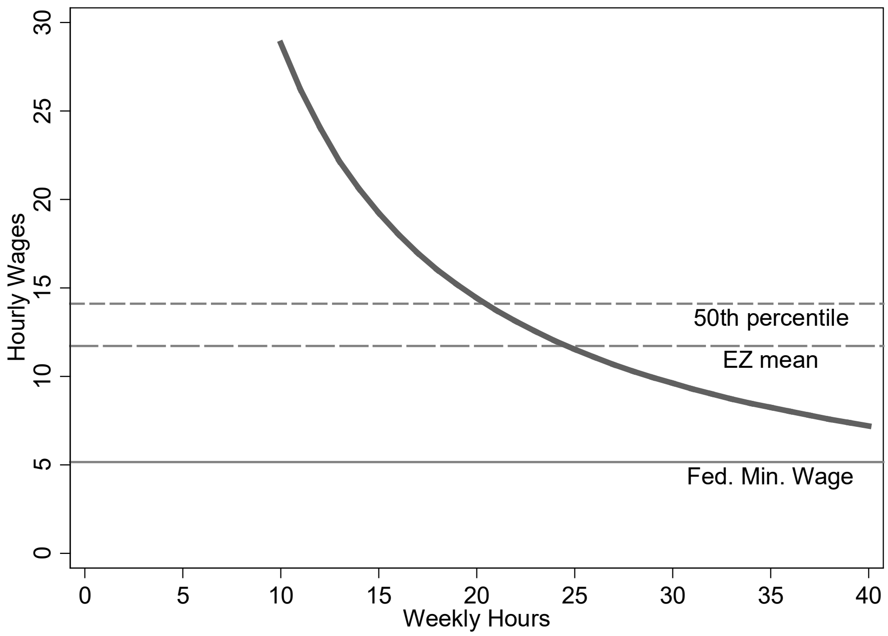
Figure 1: Earnings Cap of the Federal Empowerment Zone Wage Credit by Weekly Hours and Hourly Wage, 2000



Notes:
1) The curve represents the combination of weekly hours and hourly wages at which firms would reach the per worker earnings cap of $15,000 if they employed workers for 52 weeks a year.
2) The federal minimum wage in 2000 was $5.15. The median wage in 2000 among workers 16-64 years old was $14.12 (authors' calculations from IPUMS data). The average wage in the EZ areas is calculated from Summary 3 tract data of the aggregate earnings of area residents combined with the distribution of weeks worked and usual hours per week worked.

Figure 2: Maps of the New York (left) and Chicago (right) Empowerment Zone Program

Table 1: Summary Statistics

Variable	EZ	EC	EZ - EC
Poverty rate, 1990	0.465	0.401	0.064***
Unemployment rate, 1990	0.227	0.160	0.067***
Share nonwhite, 1990	0.843	0.647	0.196***
Share BA or higher, 1990	0.065	0.083	-0.018***
Median income ($10,000), 1990	18.748	21.034	-2.286***
Homeownership rate, 1990	0.192	0.316	-0.124***
Median house value ($10,000), 1990	4.562	6.675	-2.114***
Share female-headed households, 1990	0.637	0.533	0.104***
Share welfare, 1990	0.353	0.249	0.104***
Average housing age, 1990	38.369	36.402	1.967**
Average housing age squared, 1990	1598.226	1434.29	163.935***
Poverty rate, 1980	0.407	0.326	0.081***
Unemployment rate, 1980	0.176	0.122	0.055***
Share nonwhite, 1980	0.803	0.593	0.210***
Share BA or higher, 1980	0.041	0.065	-0.024***
Median income ($10,000), 1980	19.034	22.085	-3.051***
Homeownership rate, 1980	0.185	0.332	-0.147***
Share female-headed households, 1980	0.558	0.465	0.093***
Share welfare, 1980	0.339	0.228	0.111***
Average housing age, 1980	34.668	31.911	2.757***
Average housing age squared, 1980	1309.31	1099.079	210.231***
Worker characteristics			
Share workers with hours less than 35, 1990	0.180	0.192	-0.012**
Share workers with hours less than 35 , 2000	0.160	0.156	0.003
Share workers with earnings less than $15000, 1990	0.336	0.380	-0.044***
Share workers with earnings less than $15000, 2000	0.232	0.240	-0.008
N	273	844	

Notes:
1) All data are tabulated at the Census tract level. Area characteristics come from Summary 3 files, while worker characteristics come from Census Transportation Planning Package (CTPP) files.
2) Asterisks denote statistically significant differences at the 10% (*), 5% (**) and 1% (***) levels.

Table 2: The Effect of Capped-Wage Subsidies on Firms' Reliance on Part-Time Workers.

Dependent Variable: Percent of workers with less than 35 hours per week

		Double-difference	
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.016**	0.016***	0.014**
	(0.006)	(0.005)	(0.006)
Propensity Score Matching	-	0.022*	0.013
	-	(0.009)	(0.010)

		Triple-difference: city-wide	
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.010**	0.012**	0.010*
	(0.005)	(0.004)	(0.006)
Propensity Score Matching	-	0.016	0.010
	-	(0.009)	(0.010)

		Triple-difference: matched sub-city	
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.007*	0.008**	0.007
	(0.004)	(0.004)	(0.004)
Propensity Score Matching	-	0.012	0.007
	-	(0.009)	(0.010)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table 3: Estimates of the Effect on Firms' Reliance on Part-Time Workers by Portion of Tract Workers Earning Above the Annual Wage-Salary Cap ($15,000)

Dependent Variable: Percent of workers with less than 35 hours per week				
	Triple-difference: city-wide		Triple-difference: matched sub-city	
EZ	-0.031**	-0.033**	-0.034**	-0.034**
	(0.014)	(0.016)	(0.014)	(0.015)
EZ * I(50-75% above cap)	0.042***	0.043***	0.040***	0.040***
	(0.012)	(0.013)	(0.012)	(0.013)
EZ * I(>75% above cap)	0.064***	0.062**	0.067***	0.065***
	(0.023)	(0.022)	(0.022)	(0.022)
I(50-75% above cap)	0.034***	0.034***	0.034***	0.034***
	(0.008)	(0.008)	(0.008)	(0.008)
I(>75% above cap)	0.046**	0.047***	0.040***	0.040***
	(0.010)	(0.010)	(0.010)	(0.013)
1990 controls	Y	Y	Y	Y
1980 controls		Y		Y

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) I(50-75% above cap) is an indicator that between 50 and 75% of workers in 1990 in a tract earned more than the annual cap on the wage subsidy. Similarly, I(>75% above cap) indicates that more than 75% of workers in 1990 in a tract earned more than the annual cap on the wage subsidy. The left-out category is tracts where less than 50% of workers in 1990 earned above the annual cap on the wage subsidy.
3) All estimates are from OLS regressions and standard errors are clustered at the city level. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table 4: Estimates of the Effect of Capped-Wage Subsidies on Proportion of Workers in Selected Occupations, Including 1980 and 1990 Controls

	Triple-difference: city-wide							
	Professional	Management	Technicians	Sales	Clerical	Craftsmen	Production	Transportation
OLS	-0.024***	0.001	0.004	-0.008	0.007	-0.002	0.017	0.007
	(0.008)	(0.006)	(0.003)	(0.006)	(0.006)	(0.005)	(0.010)	(0.007)
Propensity Score Matching	-0.020	0.008	0.004	-0.012	0.004	-0.006	0.022**	0.005
	(0.016)	(0.008)	(0.005)	(0.013)	(0.010)	(0.008)	(0.0087)	(0.008)
	Triple-difference: matched sub-city							
	Professional	Management	Technicians	Sales	Clerical	Craftsmen	Production	Transportation
OLS	-0.027**	-0.008	-0.003	0.001	-0.003	-0.001	0.015*	0.017**
	(0.013)	(0.006)	(0.004)	(0.005)	(0.006)	(0.007)	(0.009)	(0.007)
Propensity Score Matching	-0.024*	-0.002	-0.003	-0.003	-0.007	-0.004	0.019**	0.015**
	(0.016)	(0.008)	(0.005)	(0.014)	(0.010)	(0.008)	(0.008)	(0.009)
% BA+	0.75	0.51	0.39	0.28	0.16	0.07	0.05	0.05
% < HS	0.09	0.22	0.21	0.42	0.48	0.66	0.78	0.76

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-1: Balancing Tests for the Propensity Score Method

	Unmatched		Matched		Matched	
Variable	EZ	EC	EZ	EC	EZ	EC
Poverty rate, 1990	0.465	0.401***	0.450	0.477**	0.453	0.449
Unemployment rate, 1990	0.227	0.160***	0.216	0.224	0.222	0.213
Share nonwhite, 1990	0.843	0.647***	0.832	0.849	0.827	0.805
Share BA or higher, 1990	0.065	0.083***	0.067	0.063	0.068	0.068
Median income ($10,000), 1990	18.748	21.034***	19.188	18.777	19.325	19.554
Homeownership rate, 1990	0.192	0.316***	0.208	0.223	0.208	0.185*
Median house value ($10,000), 1990	45.617	6.675***	4.871	4.847	4.751	4.652
Share female-headed households, 1990	0.637	0.533***	0.629	0.654	0.620	0.578**
Share welfare, 1990	0.353	0.249***	0.338	0.360*	0.343	0.330
Average house age, 1990	38.369	36.402**	38.593	39.569	38.382	37.302
Average house age squared, 1990	1598.226	1434.290***	1610.591	1660.695	1598.970	1543.405
Poverty rate, 1980	0.407	0.326***			0.397	0.393
Unemployment rate, 1980	0.176	0.122***			0.171	0.163
Share nonwhite, 1980	0.803	0.593***			0.780	0.748
Share BA or higher, 1980	0.041	0.065***			0.044	0.047
Median income ($10,000), 1980	19.034	22.085***			19.212	19.220
Homeownership rate, 1980	0.185	0.332***			0.197	0.189
Share female-headed households, 1980	0.558	0.465***			0.547	0.496**
Share welfare, 1980	0.339	0.228***			0.330	0.306*
Average house age, 1980	34.668	31.911***			34.360	33.707
Average house age squared, 1980	1309.31	1099.079***			1284.219	1257.679
Hotelling test (F-stat.)		28.396***		0.926		0.836
(p-value)		(0.000)		(0.514)		(0.676)

Notes:
1) Table presents asterisks representing statistical significance from pair-wise t-tests of differing means between EZ and EC areas for each covariate. The Hotelling test is a test of whether there is an overall difference in means across covariates between EZ and EC areas. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-2: First-Stage Regression for Instrumental Variable Estimation

	(i)	(ii)	(iii)	(iv)	(v)	(vi)
Ways and Means Member	0.123***		-0.293***	-0.147***	-0.090	-0.090
	(0.036)		(0.065)	(0.055)	(0.058)	(0.209)
Terms on Committee		0.024***	0.047***	0.029***	0.024***	0.024
		(0.003)	(0.006)	(0.005)	(0.005)	(0.020)
1990 controls				Y	Y	Y
1980 controls					Y	Y
Clustered standard errors						Y
R^2	0.013	0.042	0.059	0.272	0.342	0.375
F-stat	11.97***	41.07***	44.55***	25.83***	21.78***	1.15
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.322)

Notes:
1) Standard errors are presented in parentheses below point estimates. P-values are presented in parentheses below F-statistics. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-3: Instrumental Variables Estimates of the Effect of Capped-Wage Subsidies on Firms' Reliance on Part-Time Workers

Dependent variable: Percent of workers with less than 35 hours of work			
	Without City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
Instrumental Variables	0.034**	0.025	0.020
	(0.015)	(0.023)	(0.026)
	With City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
Instrumental Variables	0.022	0.015	0.009
	(0.015)	(0.025)	(0.028)
	With Matched Sub-city Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
Instrumental Variables	0.013*	0.008	0.010
	(0.007)	(0.014)	(0.017)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below IV estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-4: The Effect of Capped-Wage Subsidies on Employment

Dependent Variables: ln(all workers), ln(part-time workers), or ln(full-time workers)

	Double-difference		
	ln(all workers)	ln(part-time)	ln(full-time)
OLS	0.104*	0.174*	0.088
	(0.060)	(0.092)	(0.059)
Propensity Score Matching	0.107	0.195*	0.087
	(0.093)	(0.112)	(0.095)
	Triple-difference: city-wide		
	ln(all workers)	ln(part-time)	ln(full-time)
OLS	0.140**	0.207**	0.129**
	(0.059)	(0.089)	(0.059)
Propensity Score Matching	0.162*	0.245**	0.148*
	(0.093)	(0.111)	(0.095)
	Triple-difference: matched sub-city		
	ln(all workers)	ln(part-time)	ln(full-time)
OLS	0.092*	0.124*	0.084
	(0.049)	(0.063)	(0.051)
Propensity Score Matching	0.111	0.160	0.100
	(0.095)	(0.112)	(0.097)

Notes:
1) The outcome is the change in logged employment from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification. All regressions include 1980 and 1990 controls.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-5: The Effect of Capped-Wage Subsidies on Firms' Reliance on Part-Time Workers
(Weighting by share of employees who live in the EZ or EC)

Dep. Var: % of Workers with less than 35 hours per week	Without City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.014*	0.022***	0.020**
	(0.006)	(0.006)	(0.007)

Dep. Var: % of Workers with less than 35 hours per week	With City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.009*	0.018***	0.016**
	(0.005)	(0.005)	(0.007)

Dep. Var: % of Workers with less than 35 hours per week	With Matched Sub-city Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.008*	0.016***	0.015**
	(0.004)	(0.005)	(0.007)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-6: The Effect of Capped-Wage Subsidies on Firms' Reliance on Workers Earning Less Than $15,000.

Dependent Variable: Percent of workers earning less than $15,000

	Without City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.036***	0.035***	0.028*
	(0.012)	(0.013)	(0.015)
Propensity Score Matching	-	0.046***	0.023**
	-	(0.012)	(0.013)

	With City Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	-0.000	0.008	0.004
	(0.007)	(0.008)	(0.009)
Propensity Score Matching	-	0.016	-0.002
	-	(0.011)	(0.012)

	With Matched Sub-city Fixed Effects		
	Naïve	1990 Controls	1980 and 1990 Controls
OLS	0.005	0.015*	0.013
	(0.009)	(0.009)	(0.011)
Propensity Score Matching	-	0.026	0.005
	-	(0.011)	(0.012)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-7: Estimates of the Effect of Capped-Wage Subsidies on Proportion of Workers in Occupations Grouped by Education, Including 1980 and 1990 Controls

	Triple-difference: city-wide		
	High education	Middle education	Low education
OLS using Differencing	-0.019**	-0.000	0.016*
	(0.008)	(0.007)	(0.008)
Propensity Score Matching	-0.009	-0.009	0.016
	(0.019)	(0.015)	(0.016)

	Triple-difference: matched sub-city		
	High education	Middle education	Low education
OLS using Differencing	-0.038***	-0.000	0.036**
	(0.014)	(0.007)	(0.014)
Propensity Score Matching	-0.029*	-0.010	0.037**
	(0.019)	(0.015)	(0.016)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.

Table A-8: The Effect of Capped-Wage Subsidies on Firms' Reliance on Part-Time Workers by Gender

Dependent Variable: Percent of workers with less than 35 hours per week

	Triple-difference: city-wide			
	Males		Females	
	1990 Controls	1980 and 1990 Controls	1990 Controls	1980 and 1990 Controls
OLS	0.021***	0.022***	0.002	-0.003
	(0.006)	(0.005)	(0.009)	(0.010)
Propensity Score Matching	0.023*	0.023*	0.008	0.003
	(0.012)	(0.013)	(0.014)	(0.018)

	Triple-difference: matched sub-city			
	Males		Females	
	1990 Controls	1980 and 1990 Controls	1990 Controls	1980 and 1990 Controls
OLS	0.020**	0.021***	-0.003	-0.007
	(0.007)	(0.006)	(0.007)	(0.008)
Propensity Score Matching	0.021*	0.023	0.001	-0.001
	(0.012)	(0.013)	(0.013)	(0.017)

Notes:
1) The outcome is the change from 1990 to 2000 and the policy was designated in 1994.
2) Each point estimate represents the effect of the federal EZ capped-wage subsidy from a different specification.
3) Standard errors clustered at the city level are presented below OLS estimates. Standard deviations from 1000 block-bootstrapped replications are presented below propensity score matching estimates. Asterisks denote statistical significance at the 10% (*), 5% (**) and 1% (***) levels.