Popul Res Policy Rev
DOI 10.1007/s11113-009-9152-7

The Impact of Local Labor Market Conditions on Work and Welfare Decisions: Revisiting an Old Question Using New Data

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Received: 15 October 2008 / Accepted: 15 June 2009

Abstract Using Maryland administrative data between 1996 and 2005, this paper examines the impact of local labor market conditions on work and welfare use among single mothers. Our estimates rely on the new Census Bureau Quarterly Workforce Indicators database, which provides county-level economic indicators filtered by industry, gender, and age-group. We specify a multinomial choice model to estimate the effects of these local labor market variables on the full set of work–welfare combinations. The results indicate that lower unemployment rates and increased new hires and new hires' earnings in key industries increase the likelihood that women choose alternatives that include work. African American women and those with fewer years of education respond differently to changing economic conditions. Our results are robust to controls for fixed effects, county-specific time trends, and endogenous migration.

Keywords Employment decisions · Local labor market conditions · Multinomial logit models · Welfare reform

Introduction

With the passage of the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) in 1996, the focus of state welfare offices shifted from

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providing cash assistance to helping disadvantaged individuals enter the labor force. Recent research evaluating single mothers' transition from welfare to work has stressed the role of supply-side side factors such as human capital development and marriage decisions, as well as the flurry of social policy reforms implemented throughout the 1990s. Many of these studies treat policy reforms as nuanced and multidimensional explanations of work–welfare decisions, and it is now common to include detailed measures of welfare reform, the EITC, and child care subsidies in such models.

Economic conditions are also expected to influence work and welfare use, both through the availability of jobs and the generosity of wage offers. However, little is known about the effects of demand-side factors. What we do know comes from the literature's nearly exclusive focus on the state unemployment rate. Exceptions to the use of this measure are rare, but they include aggregate employment growth rates (Schoeni and Blank 2000), median wages (Wallace and Blank 1999), 20th percentile wages (Wallace and Blank 1999; Fang and Keane 2004); average earnings in various industries (Hoynes 2000), gross job flows (Bartik and Eberts 1999) and employment-to-population ratios (Hoynes 2000).

Studies matching micro data with the unemployment rate or an employment-based alternative suffer from a number of drawbacks. Perhaps the most important constraint is the lack of sub-state-level geographic identifiers in survey datasets. Confidentiality concerns typically preclude identification of labor markets smaller than the state, thereby masking potentially large within-state variation in economic conditions. Moreover, survey data generally provide homogeneous samples of single mothers or welfare recipients, making it difficult to determine how labor market conditions impact important sub-groups. Finally, the collection of annual data on employment and welfare use prevents researchers from analyzing short-term labor demand shocks. These shortcomings mean that previous studies have tended to define local labor markets imprecisely, and it is not surprising that results from this body of work suggest a small role for the unemployment rate and many of its alternatives.

Nevertheless, it has become increasingly important to understand the relationship between labor demand conditions and work–welfare outcomes. Funding for cash assistance has changed from an open-ended entitlement to a close-ended block, which essentially means that states now face greater financial risks as labor market conditions fluctuate. Added to this are strict work requirements and time limits for welfare recipients that operate in the same manner across all economic conditions. Many states have also begun experimenting with policies that deter potential welfare recipients from receiving aid. Currently, 20 states mandate job search activities at the time of application, and 30 states operate cash diversion programs. The work-first nature of these policies relies heavily on the ability of local labor markets to absorb welfare recipients, or at least encourage such individuals to seek economic well-being outside the welfare system.

This study merges administrative data for the state of Maryland with new indicators of county-level labor market conditions to provide substantially improved estimates of the effects of economic conditions on employment and welfare use among single mothers. The micro data provide complete (quarterly) employment and (monthly) welfare histories for a sample of women over the period 1996–2005. We draw the labor market data from the Census Bureau's Quarterly Workforce Indicators (QWI) series, a partnership between the Bureau, 41 states, and the District of Columbia to disseminate detailed information on local labor market conditions. Among the distinctive features of the QWI are industry-specific measures of county labor market conditions that can be further disaggregated by gender and age-group. A key advantage of the QWI series is the opportunity to replicate local-level analyses across most states. We consider several measures of economic conditions including the unemployment rate, new hires and new hires' earnings in common destination industries, and job flows. We specify a multinomial choice model to estimate the effects of gender- and age-specific labor market variables on the full set of work–welfare combinations. Our study sheds light on how economic conditions influence two groups with important policy implications: women who are neither working nor receiving welfare and women who are working without welfare.

Results suggest that women respond to labor market conditions in ways predicted by standard economic models. Specifically, we find that increased new hires and new hires' earnings in key industries increase the likelihood that women choose alternatives that include work, even if women work in conjunction with welfare receipt. We also show that African American women and those with less than a high school degree are more sensitive to labor market opportunities when considering employment and welfare use simultaneously, while white women and those who completed high school are more responsive as they move into work without welfare. Our main results are robust to specifications that include controls for unobserved relocation choices.

The remainder of this paper proceeds as follows. The next section briefly summarizes the relevant empirical literature. The following section provides a conceptual framework highlighting the importance of labor market conditions. The administrative and QWI data sources are described next and the empirical model is specified. Results are then presented, followed by conclusions.

Previous Research on the Impact of Labor Market Conditions

There are many ways to organize the literature on the effects of economic conditions, but it is useful for the purposes of this paper to consider separately studies using state- and county-level measures of labor market opportunities.

Empirical work using state-level measures is voluminous and can be further classified by whether such measures explain aggregate welfare caseloads or individual reports of work and welfare use from survey datasets. The majority of papers examining national or state caseloads use the unemployment rate (Bartik and Eberts 1999; Blank 2001; CEA 1997, 1999; Figlio and Ziliak 1999; Levine and Whitmore 1998; Moffitt 1999; Schiller 1999; Wallace and Blank 1999; Ziliak et al. 2000). Results from these studies are remarkably consistent, finding that a one percentage point increase in the unemployment rate increases welfare caseloads by 5–7%. A noteworthy exception is Klerman and Haider's (2004) stock-flow analysis, which finds effect sizes between 12 and 47%. Overall, these studies indicate that changes in economic conditions explain 20–80% of the pre- and post-TANF caseload change throughout the 1990s. A few caseload studies have experimented with alternative measures of state labor market conditions, such as gross job flows (Bartik and Eberts 1999), minimum wages (CEA 1999), median and 20th percentile wages (Wallace and Blank 1999), and wage premiums (Bartik and Eberts 1999). There is weaker consensus about the influence of these factors, but many studies continue to find significant, albeit smaller, effects of employment-based alternatives. These early caseload studies were then modified by researchers using individual-level data from large, nationally representative surveys. The most common methodological approach is to pool time series of cross-sectional samples of single mothers and estimate a discrete choice employment or welfare participation equation as a function of individual characteristics, social policy reforms, and labor market conditions (Fang and Keane 2004; Grogger 2003, 2004a; Herbst 2008; Looney 2005; Meyer and Rosenbaum 2001; Noonan et al. 2005; O'Neil and Hill 2001; Schoeni and Blank 2000). As with caseload studies, the preponderance of this research uses the unemployment rate as the key indicator of state labor market conditions. Estimates suggest that a one percentage point increase in the unemployment rate is expected to decrease employment by 0.5–2.0 percentage points and increase welfare use by one percentage point. These point estimates imply that the unemployment rate accounts for approximately 25% of single mothers' employment growth and 20% of the decline in welfare use throughout the 1990s.[1]

Turning to the studies that examine county-level indicators of labor market opportunities, one finds that most of this research is embedded in the welfare dynamics tradition. Several studies analyze the impact of county-level unemployment rates on welfare exits (Blank 1989; Fitzgerald 1995; Harris 1993), while others investigate returns to welfare (Harris 1996; Pavetti 1993; Mueser et al. 2007). Some of these studies find small or statistically insignificant effects of local labor market conditions. A noteworthy exception is Hoynes (2000), who uses administrative data from California between 1987 and 1992 to analyze several county-level labor market variables on exits from and reentry to welfare.[2] Specifically, the author considers total UI covered employment, employment in the retail and service sectors, earnings, and employment-to-population ratios. Hoynes' results suggest that a 10% increase in employment or a 3.5% increase in employment-to-population

[1] Analyses of exits from and reentry to welfare are also relevant to this study. Early research on welfare dynamics consistently uses either the state unemployment rate or average wages as the principal controls for economic conditions (Bane and Ellwood, 1983; Blank and Ruggles, 1996; Fitzgerald, 1991; Hoynes and MaCurdy, 1994). Some of these studies find statistically insignificant effects of labor market variables. More recent work by Hofferth et al. (2002, 2005) and Grogger (2004b) indicate mixed results.

[2] It is important to mention some recent work using survey and administrative data that control for county-level measures of economic conditions (Dyke et al., 2006; Fitzgerald and Ribar, 2004a, b; Heinrich et al. 2005; Ribar 2005). The primary focus of these studies, however, is not estimating the impact of local labor markets on employment and welfare outcomes. Rather, these variables are included to control for contextual factors that might be correlated with social policy reforms, the key variables of interest. Furthermore, many of these studies rely on the unemployment rate or skill-specific measures of earnings.

ratios is associated with a 7–15% increase in the likelihood of exiting welfare. Earnings are found to increase exit probabilities by a smaller amount (5%).

By merging administrative data with a new set of county-level labor market indicators, we improve upon previous research in a number of ways. Our sample is diverse enough to facilitate comparisons across sub-groups of welfare recipients, including racial/ethnic groups and educational levels. Second, the use of administrative data allows us to more accurately measure employment and welfare participation than studies relying on self-reports, which suffer from recall bias, stigma effects, and other problems associated with misinterpreting survey questions. A related problem with survey data is the well known underreporting of program participation, which has worsened over time.[3] Third, identification of employment and welfare use on a quarterly basis is preferable to annual measures. With quarterly data we are able to capture the effects of short-term shocks to local labor markets as well as observe employment and welfare spells lasting only short periods.[4] Finally, our data cover a longer time period (1996–2005, or 40 quarters) than most longitudinal survey datasets would allow.[5] Hence, we are able to examine the effects of labor market conditions during periods of economic growth and contraction.

Conceptual Framework

Standard labor supply models suggest that the decision to work is a utility maximization problem in which individuals choose quantities of leisure and levels of government assistance (e.g., welfare, food stamps, and the Earned Income Tax Credit) subject to budget and time constraints. Arguments in the budget constraint include the expected wage offer, welfare grant amounts, and other sources of non-wage income. Therefore, under reasonable conditions, utility can be stated as a linear function of leisure time, all sources of income, and stigma associated with

[3] Underreporting program participation in the CPS is especially problematic. A recent analysis of California's Medicaid and welfare programs finds that the CPS undercounts Medicaid by 30% as compared to administrative data, while welfare undercounts are in the range of 50–60% (Klerman et al. 2005). This problem pervades other surveys and programs as well. For example, a study of the Maryland Food Stamp program finds the published estimate of food stamp receipt from the U.S. Census Bureau's American Community Survey, Supplementary Survey for 2001 (ACS/SS01) is substantially lower than the count from the State's administrative records. Most of the discrepancy of about 75,000 food stamp households is due to underreporting (Taeuber et al. 2004).

[4] The Panel Study of Income Dynamics (PSID) is the primary survey dataset used in studies of welfare dynamics. However, it captures only annual welfare spells and allows for only state-level geographic identifiers (public use file). The National Longitudinal Survey of Youth (NLSY) provides monthly data, but only for a small fraction of the sample. It does, however, allow users to append county-level variables to the dataset. The other option for analysts is the Survey of Income and Program Participation (SIPP). While the SIPP provides monthly employment and welfare data, it is hampered by the fact that sample members remain in the sample for relatively short periods, between 24 and 48 months. The SIPP does not provide access to county-level identifiers in its public use file, but some researchers have accessed these data via special arrangements with the Census Bureau (for example, see Fitzgerald and Ribar, 2004b).

[5] Our paper is not the first to use administrative data to examine the work behavior of single mothers after the implementation of welfare reform. A recent paper by Mueser et al. (2007) examines three cohorts of women in Maryland and Missouri using the county-level unemployment rate as the primary control for local economic conditions.

participation in means-tested programs. The probability that a single mother works is the probability that expected utility while working exceeds the expected utility of not working. A testable implication of this model is that local labor markets influence the relative utility from working in two ways. First, mothers operating in high employment growth areas experience fewer structural constraints on finding jobs, thereby increasing the likelihood of employment. Second, conditional on finding a job, increased wage offers make work relatively more attractive than welfare, also increasing the likelihood of employment.

Critical to our study is that employment and wages are measured for key destination industries among welfare recipients. Retail trade and the service sectors hire large numbers of Maryland welfare leavers. A recent study by Mueser et al. (2007) found that these two industries in the period following welfare reform account for 73–82% of the jobs taken by individuals in the first quarter after leaving welfare. Eating and drinking places and department and grocery stores alone account for 14–22% of the destination jobs. Therefore, it is important that labor demand in these industries is strong enough to encourage welfare recipients to enter the labor force.

Data Sources and Empirical Implementation

Administrative Micro-Data

To illustrate the usefulness of the new economic measures, we use administrative data to define a sample of young single mothers that have been the focus of recent welfare reforms. Our findings are limited to this important age group, but the data sources and methodology can be replicated for other age groups and in other states.

Our starting point was the universe of AFDC/TANF case-head recipients in Maryland. From this population we selected women born in 1977 with a valid Social Security Number issued in the State.[6] We then extracted monthly welfare histories for these women between 1996 and 2005, so that in effect we are able to follow individuals between ages 19 and 28. For reasons explained below, women in our sample must have received at least one month of welfare during the last 4 years of the 10-year observation period.[7] Our final analysis sample consists of 1,712 women, providing 68,480 person-quarter combinations.

Sample members were then matched to quarterly wage record data from Maryland's Unemployment Insurance (UI) program. Individuals in covered jobs are coded as being employed if positive earnings are reported by an employer in a given quarter. While the overwhelming majority of jobs are covered by the UI system, we

[6] The SSN 'issued in Maryland' criterion serves as a proxy for our sample women's otherwise unobserved residency in the state throughout the decade long observation period. Furthermore, we were able to calculate that 71.4% of the sample received TANF or appeared in a quarterly UI wage report, or both, in 1996 *and* 2005. Therefore, we are confident that a large majority of women with a valid SSN issued in Maryland were present in the State at the beginning *and* end of the 10-year observation period.

[7] In other words, to be included in the sample women must be observed receiving welfare for at least 1 month between January 2002 and December 2005. Of course, we retained sample members who also received welfare in any month prior to this period.

necessarily miss self-employment, informal employment, and those employed outside the State and by the Federal Government. This would be a concern for some groups of Maryland workers, given the concentration of residents living close to the District of Columbia. However, most of the State's welfare recipients live in Baltimore City and Baltimore County, so we do not anticipate that many of these individuals commute to jobs outside the State.[8]

The 1977 birth-year criterion is chosen so that the implementation of welfare reform in Maryland coincides with sample members being age-eligible to receive cash assistance as a case-head.[9] This approach is taken for several reasons. First, we are interested in the impact of labor market opportunities for women entering their prime welfare-receiving years during the transition to a work-first welfare policy. Women in this age-group comprise about 61% of Maryland's TANF caseload (DHHS 2006). Understanding the relationship between economic conditions and work–welfare outcomes for young women is therefore crucial from a policy perspective. Second, in a lifecycle context, the availability of labor market opportunities for young women could be important predictors of work and welfare use in subsequent years. Finally, restricting the observation period to the post-TANF era leads to a cleaner analysis of the effects of labor market conditions. Given that all sample members are "exposed" to welfare reform, it obviates the need to account for these policy changes in the model.

Although the micro-data are well suited for this analysis, they have some drawbacks. First, a common issue for studies using administrative data is that demographic information on welfare recipients is limited and observed only when individuals receive cash assistance.[10] We deal with this limitation by creating a sample of women who received welfare between 2002 and 2005, when full demographic information is available in Maryland's administrative file.[11] Second, since samples based on administrative data are actually samples of welfare recipients (rather than all single mothers), women are selected on the basis of previous or future interactions with the welfare system. Finally, our sample is cohort-specific and reflects the work–welfare experiences of women younger than most previous research. However, given that young women are less responsive to economic conditions, our estimates will understate the impact of labor market variables (Bound and Holzer 1995).

[8] As of 2004, 66% of Maryland's welfare recipients filed for benefits in Baltimore City or Baltimore County, while just 10% filed in Prince George's County, a jurisdiction adjacent to the District of Columbia. Furthermore, a recent report found that of Maryland's employed welfare leavers between 2003:Q1 and 2004:Q2, only 4–5% were employed in surrounding states and <1% were employed by the federal government and the military (Jacob France Institute, 2005). Therefore, the overwhelming majority of employed welfare leavers are working inside Maryland.

[9] Because women in our sample were born in 1977, they turned 19, the minimum age to receive welfare as case-heads, sometime during the first year of the study.

[10] That demographic data are only available for the months during which sample members receive cash assistance can lead to sub-optimal strategies to ameliorate the problem. For example, in an analysis of reentries to welfare using California administrative data, Hoynes (2000) is forced to use values of demographic characteristics as of the end of the previous welfare spell.

[11] Specifically, women who received cash assistance between January 2002 and December 2005 have full demographic information available. Conversely, these data are unavailable for women who received cash assistance between January 1996 and December 2001 but not at any time between 2002 and 2005.

Local Labor Market Variables: The QWI Database

Our labor market information comes from the QWI, a previously unused database available online and maintained by the Census Bureau.[12] The QWI is the culmination of the Bureau's efforts to merge administrative data on workers with business employment dynamics. Currently, 41 states and the District of Columbia participate in this data-sharing arrangement, and others have agreed to become partners. Such widespread participation is a strength of the QWI, enabling researchers to replicate analyses across a large number of states. These data provide unprecedented detail on local labor market dynamics. Among the distinctive features of the QWI are the availability of economic data on a quarterly basis, disaggregated to the county-level, and which allow researchers to organize information by industry code, gender, and age. Eight economic indicators are publicly available, but we focus on a sub-set of three gender and age-specific measures: new hires, new hires' earnings, and job flows.[13] For comparison purposes, we also estimate models using the county unemployment rate. The QWI data are used to construct a county-by-quarter data file on new hires, new hires' earnings, and job flows in retail trade, accommodation/food services.[14] We focus on these industries because, as previously stated, they are common destination jobs for individuals leaving welfare. Summary statistics for these measures are provided in Table 1.

The QWI measures of local labor market conditions are superior to the unemployment rate for a number of reasons. As others have noted, county-level unemployment rates are measured with substantial error, especially for small counties (Bartik 1996; Hoynes 2000). Furthermore, unemployment rates are blunt measures of economic conditions, since they do not provide information on the industry mix and generosity of wage offers within and across geographic areas. Unlike the QWI indicators, the unemployment rate is inadequate for describing labor market conditions among such narrowly defined sub-groups as single mothers. Finally, indicators of new hires and new hires' earnings are more accurate signals of labor market opportunities than aggregate employment and earnings. These latter measures commingle information on new workers with a large number of individuals who have been with the same employer for substantial periods. Comparisons of new

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[12] Data are available at http://lehd.did.census.gov/led/datatools/qwiapp.html. Detailed definitions of all variables are available at http://lehd.did.census.gov/led/library/techpapers/QWI_definitions.pdf.

[13] The indicators are: total employment, earnings, new hires, new hires' earnings, net job flows, job creation/separation, and turnover. The Census Bureau defines these variables in the following manner. New hires: "A worker i is defined as a new hire for employer j in t if i has positive earnings at j in t but no earnings from j in $t-1$, $t-2$, $t-3$, $t-4$." New hires earnings: "Sum of quarterly earnings at j in t for all i who are full-quarter new hires, divided by the number of full-quarter new hires at j in t, divided by three." Net job flows is "end-of-quarter employment in t minus beginning-of-quarter employment in t." The primary data sources underlying the QWI indicators are UI wage records and employer reports for the Quarterly Census of Employment and Wages (QCEW) or ES-202 series. All QWI indicators are based on the employer's address.

See http://lehd.did.census.gov/led/library/techpapers/Brookings_QWI.pdf for a technical description of QWI variables.

[14] The two-digit NAICS code for these industries are as follows: retail trade, pp. 44–45; accommodation and food services, p. 72.

Table 1 Summary statistics for the analysis sample

Variable	Mean	Standard deviation	Minimum	Maximum
Demographic variables				
Age at first quarter of welfare receipt				
18–19	0.522	0.499	0	1
20–22	0.176	0.381	0	1
23–25	0.179	0.383	0	1
26–28	0.121	0.326	0	1
Black	0.767	0.422	0	1
White	0.211	0.408	0	1
Other/missing race or ethnicity	0.021	0.143	0	1
At least high school degree	0.580	0.493	0	1
Missing information on education	0.042	0.200	0	1
Number of children at first quarter of welfare receipt				
0	0.130	0.337	0	1
1	0.665	0.471	0	1
2	0.138	0.345	0	1
3+	0.033	0.180	0	1
Additional children after first quarter of welfare receipt	0.573	0.494	0	1
Missing information on children	0.031	0.174	0	1
Labor market variables				
County unemployment rate, quarterly	6.15	1.97	1.60	18.6
New hires, quarterly, 1000s				
Retail trade	6.823	3.907	0.054	20.552
Accommodation/food services	6.837	2.980	0.043	14.817
New hires earnings, quarterly, 1000s				
Retail trade	1.402	0.211	0.654	3.259
Accommodation/food services	0.991	0.204	0.293	3.426
Job flows, quarterly, 1000s				
Retail trade	0.213	1.330	−5.165	8.471
Accommodation/food services	0.206	1.325	−6.425	7.834
Number of individuals	1,712			
Number of person-quarter combinations	68,480			

Notes: The county unemployment rate is an aggregate measure. New hires, new hires' earnings, and job flows are industry-, gender-, and age-specific

hires' earnings with aggregate earnings in a given industry reveal that the latter are significantly higher than the former. Therefore, new hires' earnings are more precise indicators of what welfare leavers can expect when they receive a job offer.

Do these labor market measures vary sufficiently across counties and over time to identify their impact on employment and welfare use? One way to assess this is by regressing each economic indicator on a vector of fixed county and time effects and then calculating the R^2. The greater the R^2 after removing these factors, the less remaining variation there is in each measure to identify the effects on employment and welfare use. When this exercise is conducted on the county unemployment rate, the R^2 is 0.87, suggesting that 13% of the total variation in this measure remains after removing area and time effects. The R^2 values for new hires in retail trade and accommodation/food services and new hires' earnings in both industries are, respectively, 0.86, 0.86, 0.50, and 0.61. It appears that labor market conditions vary sufficiently in Maryland, although the earnings measures contain significantly more variation than the employment measures.

The Model

We estimate the effects of local labor market conditions on the full set of work–welfare choices faced by single mothers using a multinomial logit model. Whereas previous studies examine any work/welfare participation or likelihood of ending a welfare spell, our analysis captures two important alternatives for disadvantaged women: work without receiving welfare and neither work nor welfare. The number of single mothers working without welfare increased dramatically throughout the 1990s, and this outcome is increasingly the focus of states' welfare reform efforts. Unemployed mothers who are not receiving cash assistance are also a concern for policymakers because these women likely have several employment barriers—including low education, health problems and a history of domestic violence—and yet their financial needs are not being met by government aid (Blank 2007).[15]

Stated formally, the probability that individual i, residing in county c in year t chooses the jth work–welfare alternative is modeled in the following manner:

$$\Pr(y = j|\mathbf{x}_{ict}) = \frac{\exp\left(\alpha + \mathbf{Z}'_{ict}\beta\right)}{\sum_{j=1}^{J}\exp\left(\alpha + \mathbf{Z}'_{ict}\beta\right)}. \tag{1}$$

Four work–welfare combinations (j) are examined: (1) no work, no welfare; (2) no work, welfare; (3) work, welfare; and (4) work, no welfare.[16,17] Estimates are

[15] It should be emphasized that women who are neither working nor receiving welfare comprise a non-trivial fraction of all mothers in our sample in any given quarter. This figure declines from about 40 to 20% during the first 12 months in the sample, and then remains in the neighborhood of 20–30% thereafter. That such a large number of single mothers are in this category is one of the key motivations to examine them in this paper.

[16] A concern with the use of multinomial logit models is the independence of irrelevant alternatives (IIA) assumption. Briefly, the IIA states that the odds of being observed in one of the alternatives does not depend on the presence or absence of other alternatives (i.e., the disturbances across the various categories are constrained to equal zero). This has implications for our dependent variable because two of the alternatives include welfare receipt and the other two include employment. To address this concern, we conduct Hausman tests of the null hypothesis of independence across the categories of the outcome. This test compares the coefficients from our full model (i.e., all outcome categories, observables, fixed effects, period effects, and time trends) with those from a series of models that exclude one of the categories. If the test statistic (distributed as a chi-squared) is significant for a given comparison, the IIA fails for that category. Results from our tests indicate that across each comparison, we cannot reject the null hypothesis of independence.

[17] It is possible that our "comparison" category (women neither working nor receiving welfare) commingles women who are disconnected from both the labor market and welfare system with those who

interpreted in relation to category (1), women who are neither working nor receiving welfare. The $\mathbf{Z}'$ is modeled in the following manner:

$$\mathbf{Z}'_{ict}\beta = \mathbf{X}'_{i}\psi + \psi \mathrm{L}_{ct} + \psi \mathrm{County}_{c} + \psi \mathrm{Period}_{t} + \psi(\mathrm{County}_{\mathrm{c}} \times \mathrm{Trend}_{t}), \qquad (2)$$

where $\mathbf{X}'$ is a vector of time invariant demographic controls including age, race, education, number of children, and the presence of additional children. As with most studies using administrative data, demographic characteristics are not observed when sample members are not receiving welfare.[18] Therefore, we fix these variables as of the first quarter of welfare receipt. The L represents controls for labor market conditions across counties and quarters, specifically the aggregate unemployment rate, log of gender- and age-specific new hires, gender- and age-specific new hires' earnings (/1,000), and gender- and age-specific gross job flows (/1,000).[19] We also include county fixed effects and period effects in all models.[20] County fixed effects control for unobserved area specific labor market shocks that influence work–welfare decisions, while period effects eliminate statewide year-to-year shocks. In other models, we incorporate county-specific time trends to account for unobserved factors that are trending within counties. Thus, labor market effects are identified using either within-county, year-to-year changes in economic conditions or using variation in the differential trends in economic opportunities across counties.[21]

Footnote 17 continued
exited welfare due to marriage. Based on figures presented in Hofferth et al. (2002), however, marriage-based exists from welfare are relatively rare. These authors find that within two years of entering welfare, 32% of women leave for work and only 15% leave for "other" reasons, one of which is marriage. According to the authors, "family changes represent only a small portion of these 'other exits'" (p. 446). Therefore, it is likely that our comparison group is comprised largely of the most disadvantaged women.

[18] Given the limited number of demographic controls available in the administrative data, we considered including county-level Census data in the models. However, these variables would also have to be fixed at a single point in time. In addition, many of these aggregate variables are highly correlated with each other, limiting their usefulness in the regressions. We decided in the end to rely on fixed effects to purge the model of individual unobserved heterogeneity.

[19] A possible objection to our parameterization of the labor market variables is that they are not adjusted for the fact that more populous counties will tend to have a comparatively large number of new hires in various industries (and so these measures should be scaled by the size of the labor force in each county). This strategy has several drawbacks, however. In particular, there are two key ways for this ratio to increase: (1) if both the numerator and denominator increase, but the numerator (new hires, for example) increases relatively more than the denominator; and (2) if the numerator remains fixed (or declines), and the denominator declines (or declines relatively more than the numerator). Scenario (2) is a concern because it indicates an improvement in economic conditions even though fewer people are being hired by the relevant industries. Given this concern and the common practice in the literature to avoid ratio measurements of labor market indicators (See Hoynes 2000 and Bartik and Eberts 1999), we decided to take the log of new hires and simply divide all earnings variables by 1,000.

[20] Maryland contains 23 counties plus Baltimore City, which is considered a separate county in the Census Bureau nomenclature.

[21] It is important to note that coefficients on the local labor market variables reflect the average effects of these factors across all counties in Maryland. Insofar as social policies vary across Maryland's counties, one might expect a differential effect of local labor markets across these county-specific policy regimes. Indeed, Herbst (2008) finds this to be the case using state-level policy and economic data. There is some evidence of policy variation across Maryland's counties: for example, Montgomery County became the nation's first county to enact an EITC. However, the degree of policy variation across the state was deemed insufficient to take advantage of this in the current analysis.

Although we include controls for unobserved geographic and time effects, at least two additional sources of bias must be considered. The first deals with the measurement of the QWI variables, which are filtered by industry, gender, and age-group. The increased precision of these variables carries the risk that mechanical rather than behavioral impacts of job openings or higher wages are estimated. This "reflection bias" arises because changes in labor demand conditions are necessarily reflected in the observed set of work–welfare choices. A related concern is the general equilibrium effects of single mothers' employment decisions. We attempt to deal with these issues by incorporating one-period lags of each QWI indicator and the aggregate county unemployment rate into the basic model. Controlling for lagged new hires and new hires' earnings mitigate feedback mechanisms that may be correlated with contemporaneous economic conditions and mothers' work–welfare decisions. Furthermore, adding the overall unemployment rate allows us to determine whether industry-specific measures of local economic activity are related to work and welfare decisions net of aggregate conditions.

The second issue deals with women's selective migration to areas based on labor market conditions. Endogenous migration is a well known form of unobserved heterogeneity in studies using low levels of geographic aggregation. Endogeneity will result, for example, if high-skilled women move to counties that implement strong work-based reforms in favorable economic conditions. It is also plausible that women with weak work preferences and multiple employment barriers live in counties with unfavorable economic conditions. Both cases will result in biased estimates of the effects of labor market variables. To deal with this issue, we create an indicator for whether women received a TANF grant through more than one county welfare office during the observation period. Eligibility and benefit determination occur at the county level in Maryland, allowing us to construct fairly complete inter-county migration histories.[22] In sensitivity analyses, we extend the basic model by adding a control for whether women received TANF through multiple local offices. Summary statistics for this variable suggest that movement between counties is a relatively rare event, and multiple moves are even rarer. Of the 1,712 women in our sample, 174 (10.1%) moved exactly one time between 1996 and 2005 and 111 (6.5%) moved multiple times. We also find mixed evidence on the relationship between mobility and several observable characteristics. Sample members are more likely to move when they are young (that is, in the early period of the panel); movers are more likely to be white and located in counties with lower unemployment rates; and they are equally likely to be employed or receiving welfare.

[22] This approach is valid only to the extent that women receive at least two months of welfare, or have at least two welfare spells, because they must show up in the data as having received welfare at two different county welfare offices. Another problem is that, while it controls for underlying preferences for welfare, we leave unobserved differential tastes for work that are correlated with economic conditions. This might not matter much either because tastes for welfare are likely to be inversely related to tastes for employment.

Results

In this section, we present multinomial logit estimates of the impact of several local labor market indicators on the four work–welfare combinations defined above. To ease interpretations, we present marginal effects, evaluated at the mean of each variable. Results are organized in the following manner. For each labor market variable, we first show estimates based on models that do not include county-specific time trends, followed by those that do. We begin with the unemployment rate (Table 2), since this is the most familiar measure, and then move on to new hires (Table 3), new hires' earnings (Table 4), and job flows (Table 5). Tables 3, 4, and 5 show results for retail trade and accommodation/food services combined and individually.[23]

Labor Market Conditions and Women's Work–Welfare Choices

Looking first at Table 2, we find that the county unemployment rate is statistically significantly associated with women's work–welfare choices. Results for the first model, which omit county time trends, suggest that increases in the unemployment rate raise the likelihood that women choose alternatives that include welfare receipt and decrease the likelihood of choosing alternatives that include employment, even in combination with welfare receipt. For example, a one point increase in the unemployment rate is associated with a 2.1 percentage point increase in the probability of not working and receiving welfare, and a three percentage point decrease in the probability of working without welfare. The effect of the unemployment rate on work and welfare receipt appears to be negative, but the coefficient is not statistically significant. The second model adds county time trends to the specification. Generally speaking, this has the effect of substantially reducing the magnitude of the unemployment effect, while leaving the standard errors unchanged. For the no work/welfare category, the coefficient declines by 43%; for the work/no welfare category, the coefficient declines by 55%. Interestingly, the magnitude of the unemployment coefficient in the work/welfare category increases and becomes statistically significant.

Coefficients on the demographic variables in Table 2 take the expected sign and for the most part are statistically significant. Older women and those with at least a high school degree are more likely to work without receiving welfare. For example, the coefficient on educational attainment indicates that more highly educated women are about 14 percentage points more likely to be employed without receiving welfare. Conversely, women with more children in the assistance unit are less likely to work without welfare. Finally, the introduction of additional children into the assistance unit is associated with a decreased likelihood of working without welfare, with reductions of about 5.5 percentage points. Parameter estimates on the

[23] We experimented with several adjustments to the standard errors, including Huber/White robust standard errors as well as corrections for clustering at the county- and individual-level. Our results are not sensitive to these corrections, and so we ultimately calculated cluster-robust standard errors at the individual-level. This decision was made because women in our sample appear multiple times (40 quarters).

Table 2 Multinomial logit estimates of work–welfare combinations: county unemployment rate

Variable	No work, welfare	Work, welfare	Work, no welfare	No work, welfare	Work, welfare	Work, no welfare
Unemployment rate	0.021***	−0.002	−0.029***	0.012***	−0.004*	−0.013***
	(0.003)	(0.002)	(0.004)	(0.003)	(0.002)	(0.003)
Age 20–22	−0.055***	−0.033***	0.001	−0.055***	−0.033***	0.001
	(0.009)	(0.005)	(0.016)	(0.009)	(0.005)	(0.016)
Age 23–25	−0.132***	−0.083***	0.086***	−0.133***	-0.083***	0.086***
	(0.008)	(0.004)	(0.018)	(0.008)	(0.004)	(0.018)
Age 26–28	−0.175***	−0.107***	0.124***	−0.175***	−0.107***	0.124***
	(0.007)	(0.003)	(0.024)	(0.007)	(0.003)	(0.024)
Black	0.017*	0.057***	−0.008	0.018*	0.057***	−0.008
	(0.010)	(0.005)	(0.016)	(0.010)	(0.005)	(0.016)
Other/missing race	0.048	0.033	−0.058	0.049	0.033	−0.058
	(0.036)	(0.021)	(0.041)	(0.036)	(0.021)	(0.042)
At least high school degree	−0.074***	−0.004	0.135***	−0.074***	−0.004	0.136***
	(0.008)	(0.004)	(0.012)	(0.008)	(0.004)	(0.012)
Missing education	−0.097***	−0.043***	0.141***	−0.097***	−0.042***	0.141***
	(0.014)	(0.007)	(0.031)	(0.014)	(0.007)	(0.031)
1 child	−0.001	0.015**	−0.0007	−0.001	0.015**	−0.0006
	(0.010)	(0.006)	(0.017)	(0.010)	(0.006)	(0.017)
2 children	0.020	0.012	−0.041*	0.020	0.012	−0.041*
	(0.015)	(0.009)	(0.022)	(0.015)	(0.009)	(0.022)
3+ children	0.035	0.006	−0.087**	0.035	0.005	−0.087**
	(0.033)	(0.016)	(0.035)	(0.033)	(0.016)	(0.036)
Additional children	0.072***	0.012**	−0.056***	0.073***	0.013**	−0.056***
	(0.008)	(0.005)	(0.013)	(0.008)	(0.005)	(0.013)

Table 2 continued

Variable	No work, welfare	Work, welfare	Work, no welfare	No work, welfare	Work, welfare	Work, no welfare
Missing children	−0.074***	−0.020	0.062	−0.074***	−0.019	0.062
	(0.024)	(0.015)	(0.040)	(0.024)	(0.015)	(0.041)
Constant	−0.287***	−0.080	0.246***	−0.767***	−0.266***	0.470***
	(0.044)	(0.078)	(0.094)	(0.151)	(0.086)	(0.167)
County fixed effects	Yes			Yes		
Period effects	Yes			Yes		
County-specific time trends	No			Yes		
Person-quarter combinations	68,480			68,480		
Log-likelihood	−82,829.233			−82,516.419		

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

Table 3 Multinomial logit estimates of work–welfare combinations: quarterly new hires in various industries

Variable (in logs)	No work, welfare	Work, welfare	Work, no welfare	No work, welfare	Work, welfare	Work, no welfare
(1) Combined new hires in retail trade and accommodation/ food services	−0.007	0.070***	−0.003	−0.041***	0.044***	0.058***
	(0.012)	(0.009)	(0.013)	(0.010)	(0.010)	(0.012)
(2) New hires in retail trade	0.009	0.050***	−0.029**	−0.023	0.018	0.035**
	(0.011)	(0.008)	(0.013)	(0.015)	(0.013)	(0.016)
(3) New hires in accommodation/ food services	−0.007	0.070***	−0.003	−0.029***	0.037***	0.039***
	(0.012)	(0.009)	(0.013)	(0.007)	(0.006)	(0.009)
County fixed effects	Yes			Yes		
Period effects	Yes			Yes		
County-specific time trends	No			Yes		
Person-quarter combinations	68,480			68,480		

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. The new hires variables are expressed in logarithms. All models include the variables presented in Table 2, except the unemployment rate. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

demographic variables do not change substantially when the linear time trends are added, so we suppress the presentation of these variables from all remaining tables.

Table 3 presents results substituting the log of new hires in retail trade and accommodation/food services for the unemployment rate. For comparison purposes, estimates are shown for models with and without county time trends. Model (1) suggests that combined new hires in both industries are strongly related to all work–welfare combinations. Specifically, an increase in new hires decreases the likelihood of choosing alternatives in which work is not one of the activities and increases the likelihood of choosing any work alternative. Interestingly, the magnitude of the positive effect is 32% greater for women working without welfare, relative to those working and receiving welfare. Such a finding indicates that young women are responsive to the level of hiring in key industries, and that a period of healthy employment growth provides a powerful incentive to rely on work alone. Models (2) and (3) examine individual measures of new hires in retail and accommodation/food services. It appears that the results in model (1) are largely driven by the strong estimates for accommodation/food services. The coefficient is statistically significant across all work–welfare states, whereas just the coefficient in the work/no welfare category is significant for retail trade. This pattern is not surprising given that nearly one-fourth of Maryland's employed welfare leavers work in the service sector in the first quarter after exiting. Finally, a close inspection of the models with

Table 4 Multinomial logit estimates of work–welfare combinations: quarterly new hires earnings in various industries

Variable (in $1000s)	No work, welfare	Work, welfare	Work, no welfare	No work, welfare	Work, welfare	Work, no welfare
(1) Combined new hires earnings in retail trade and accommodation/food services	−0.065***	0.043***	0.059***	−0.033**	0.034***	0.032**
	(0.012)	(0.010)	(0.014)	(0.012)	(0.010)	(0.014)
(2) New hires earnings in retail trade	−0.037***	0.030***	0.037***	−0.015	0.027***	0.019*
	(0.010)	(0.008)	(0.011)	(0.009)	(0.008)	(0.011)
(3) New hires earnings in accommodation/food services	−0.057***	0.033***	0.060***	−0.027**	0.022**	0.018
	(0.014)	(0.008)	(0.013)	(0.013)	(0.008)	(0.013)
County fixed effects	Yes			Yes		
Period effects	Yes			Yes		
County-specific time trends	No			Yes		
Person-quarter combinations	68,480			68,480		

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. The new hires earnings variables are expressed in thousands of dollars. All models include the variables presented in Table 2, except the unemployment rate. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

and without controls for county time trends indicates that adding trends generally strengthens the results by generating coefficients with the expected sign.

Table 4 presents estimates of new hires' earnings in retail trade and accommodation/food services. Generally speaking, these results are less sensitive to the inclusion of county time trends. A plausible explanation is that employment carries a much stronger seasonal component (e.g., temporary holiday employment) than the determinants of earnings. However, most coefficients experience fairly steep declines when the trend variables are included. Nevertheless, combined new hires' earnings are strongly related to women's decisions in all categories, and once again, increased earnings are positively associated with alternatives that include employment. A $1,000 increase in quarterly new hires earnings is associated with a 6.6 percentage point increase (3.4 + 3.2) in the probability of choosing either work state. Unlike the case of employment, the magnitude of the effect of combined earnings is similar across the work/welfare and work/no welfare alternatives. Another interesting difference between the employment and earnings results is that retail earnings are related to all work–welfare combinations. Thus, it appears that the returns to work in key industries provide important incentives to choose any work alternative, even if women are working in conjunction with receiving cash assistance.

Table 5 Multinomial logit estimates of work–welfare combinations: quarterly job flows in various industries

Variable (in 1000s)	No work, welfare	Work, welfare	Work, no welfare	No work, welfare	Work, welfare	Work, no welfare
(1) Combined job flows in retail trade and accommodation/ food services	0.0002	0.00003	0.0007	−0.0001	0.0008	0.0006
	(0.0006)	(0.0005)	(0.0007)	(0.0006)	(0.0005)	(0.0007)
(2) Job flows in retail trade	−0.001*	0.002**	0.003***	0.0004	0.0005	−0.0003
	(0.0009)	(0.0008)	(0.001)	(0.0009)	(0.0008)	(0.0010)
(3) Job flows in accommodation/ food services	0.001**	−0.001**	−0.001	−0.0007	0.001	0.002*
	(0.0008)	(0.0007)	(0.001)	(0.0009)	(0.0008)	(0.001)
County fixed effects	Yes			Yes		
Period effects	Yes			Yes		
County-specific time trends	No			Yes		
Person-quarter combinations	68,480			68,480		

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. The job flow variables are expressed in thousands. All models include the variables presented in Table 2, except the unemployment rate. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

Estimates for the final measure of labor market conditions, job flows, are presented in Table 5. Capturing the effects of job flows is complicated by the fact that the measure varies substantially across counties but also takes on negative values, which precludes us from expressing the measure in logarithmic form. None of the coefficients associated with combined job flows is statistically significant, and the magnitude of the coefficients implies economically small effects. When models are estimated separately for retail trade and accommodation/food services, most of the coefficients become significant as long as county time trends are omitted. Including the trends renders most of the estimates insignificant, suggesting that a large component of job flows is trending with unobserved factors associated with the work–welfare alternatives.

The Economic Impact of Labor Market Conditions

We assess the relative importance of labor market conditions by simulating the effects of changes in the unemployment rate, new hires, and new hires' earnings on the likelihood of choosing each work–welfare alternative. Specifically, we hold all variables at their mean values and calculate the percent change in the probability that women are observed in each category as the unemployment rate declines from

Table 6 Anticipated effect of changes in local labor market conditions and human capital characteristics on work and welfare participation

	No work, no, welfare (%)	No work, welfare (%)	Work, welfare (%)	Work, no welfare (%)
Change in labor market condition				
UR declines from 7.0% to 5.0%	−4.0	−11.2	8.0	6.7
30% increase in new hires	−6.1	−5.2	10.3	3.6
20% increase new hires' earnings	−3.2	−3.8	7.7	1.9
Change in human capital characteristic				
High school dummy variable: 0 to 1	−19.7	−28.9	−4.2	40.9

Notes: Each cell represents the percent change in the predicted probability (i.e., the change in the probability, in percent terms) that women in the sample will be observed in each work–welfare state given changes in each labor market variable. Percent changes are evaluated against a baseline of all variables, with the exception of the unemployment rate, set to their mean values. The human capital simulation holds all variables at their mean values and changes the high school dummy variable from 0 to 1. The unemployment rate is defined as the quarterly, county-level unemployment rate. New hires are quarterly new hires in retail trade and accommodation/food services industries. New hires' earnings are quarterly earnings of new hires in trade and accommodation/food services industries. Simulations are based on estimates from corresponding models in Tables 2, 3, and 4

7.0 to 5.0%, new hires increase 30%, and new hires' earnings increase 20%.[24] Changes of this magnitude are common for Maryland over the 10-year study period, and represent reasonable shifts during periods of economic expansion and contraction. The simulations are based on models estimated in Tables 2, 3, and 4 and include county and period fixed effects as well as county-specific time trends. For comparison purposes, we simulate the effects of changes in educational attainment, an important supply-side determinant of work and welfare receipt. Results from this exercise are presented in Table 6.

Generally speaking, the simulations reveal fairly consistent economic impacts across all three labor market measures. Decreases in the unemployment rate and increases in new hires and new hires' earnings increase the likelihood that single mothers choose alternatives that include work, irrespective of whether individuals also receive welfare. However, labor market improvements appear to have stronger effects on employment that coincides with welfare receipt. For example, as the unemployment rate declines from 7.0 to 5.0%, there is an 8% increase in the probability of moving into work while receiving welfare, and a 6.7% increase in the probability of working without welfare. These differences are more pronounced for both employment-based measures. An increase of 30% in the number of new hires in retail trade and accommodation/food services leads to a 10.3% increase in the fraction of women working and receiving welfare, but only a 3.6% increase in employment alone. The corresponding figures for new hires' earnings are, respectively, 7.7% and 1.9%. Overall, these results indicate that labor market

[24] Baseline simulation results for new hires and new hires' earnings are derived from both variables set to their mean values. These simulations are conducted on the combined measures for retail trade and accommodation/food services. It should be noted that the simulated change in the unemployment rate might not be applicable to the economic downturn that began in December of 2007.

improvements provide strong work incentives, but most of this employment effect is concentrated among women who continue receiving cash assistance.

One way to assess the relative importance of labor market conditions is by comparing the above simulation results with factors that can reasonably be targeted by policy interventions. We focus on increases in educational attainment, a commonly cited human capital determinant of economic independence. In particular, we simulate the impact of providing all single mothers with at least a high school education, compared to a baseline in which no mother has such a degree. As shown in Table 6, this policy experiment results in comparatively large effects on work–welfare decisions that are concentrated among women who are working without welfare. One should bear in mind, however, that this simulation is based on a large change in human capital accumulation, and that simulations based on increasing education from the sample mean (60%) yield smaller employment effects.

Extensions and Tests of Robustness

To assess the robustness of the main results, we undertake a number of sensitivity analyses. We provide only a sub-set of results and a brief discussion here, but full results from these tests are available upon request. Recall that we attempt to account for two key threats to the validity of our labor market estimates: "reflection bias" and endogenous migration. We deal with the first issue by adding to the main model controls for lagged QWI indicators and the aggregate county unemployment rate. Results from this exercise are presented in Table 7, and are limited to new hires and new hires' earnings. Generally speaking, the impact of contemporaneous labor market variables is robust to both specification checks, although in some cases the magnitude of estimated coefficients declines somewhat. That the main results are largely unchanged after accounting for overall economic conditions is important, suggesting that the filtered QWI variables are not mirroring equilibrium feedbacks or broader forces in the local economy.

To the deal with the possibility that women's unobserved characteristics lead some to migrate to areas based on labor market conditions, we estimate models that control for whether women received cash assistance through multiple county welfare offices during the 10-year observation period. The coefficient on the "mover" variable is never statistically significant, indicating that conditional on observable characteristics and fixed effects, women who moved sometime during the observation period are equally likely to be located in any of the work–welfare states. Moreover, estimates on the economic variables are unchanged. As a final check, we estimate the basic model only on women who did not move during the 10-year observation period. These results show little change in the estimated effects of labor market conditions.

To this point, we have estimated the average impact of local labor market conditions. There are good reasons to believe, however, that such factors have heterogeneous effects depending on the observable characteristics of women (Hoynes 2000). Table 8 presents results for a number of key demographic subgroups, including African American women and those with less than a high

Table 7 Sensitivity of main results to alternative specifications

Specification check	No work, welfare	Work, welfare	Work, no welfare
Panel A: Model includes one-period lag of the QWI variables			
ln(combined new hires)	−0.051***	0.038***	0.068***
	(0.015)	(0.012)	(0.016)
ln(new hires in retail trade)	−0.009	0.007	0.027
	(0.018)	(0.014)	(0.019)
ln(new hires in accommodations/food services)	−0.038***	0.031***	0.043***
	(0.008)	(0.007)	(0.010)
Combined new hires' earnings	−0.034**	0.032***	0.033**
	(0.014)	(0.012)	(0.016)
New hires' earnings in retail trade	−0.015	0.025***	0.017
	(0.010)	(0.008)	(0.011)
New hires' earnings in accommodations/food services	−0.023	0.023**	0.016
	(0.014)	(0.009)	(0.014)
Panel B: Model includes aggregate county-level unemployment rate			
ln(combined new hires)	−0.031***	0.042***	0.046***
	(0.010)	(0.010)	(0.012)
ln(new hires in retail trade)	−0.015	0.016	0.025
	(0.015)	(0.012)	(0.016)
ln(new hires in accommodations/food services)	−0.023***	0.036***	0.031***
	(0.007)	(0.006)	(0.009)
Combined new hires' earnings	−0.021	0.031***	0.020
	(0.013)	(0.011)	(0.014)
New hires' earnings in retail trade	−0.006	0.024***	0.011
	(0.010)	(0.008)	(0.011)
New hires' earnings in accommodations/food services	−0.022*	0.020**	0.011
	(0.013)	(0.008)	(0.012)

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. Models in Panel A ($N = 66{,}768$) include a one-quarter lag of a given QWI variable in addition to the contemporaneous measure. Models in Panel B ($N = 68{,}480$) include the overall county-level unemployment rate in addition to the QWI variables. For ease of interpretation, only the coefficients on the relevant QWI measures are presented. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. All models include the variables presented in Table 2, county fixed effects, controls for period effects, and county-specific time trends. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

school degree. To conserve space, we conduct the analyses on combined new hires and new hires' earnings in retail trade and accommodation/food services.

Comparing the estimates for African American and white women, it is difficult to detect a clear pattern in the results. Increases in new hires and new hires' earnings generate larger effects for African American women at the no work/welfare and work/welfare alternatives, but white women appear to be more responsive at the

Table 8 Multinomial logit estimates of work–welfare combinations: retail trade and accommodation/food services industries, by sub-group

Sub-group	Person-quarter combinations	No work, welfare	Work, welfare	Work, no welfare
Panel A: Combined new hires in retail trade and accommodation/food services (in logs)				
Black	52,560	−0.049***	0.046***	0.061***
		(0.013)	(0.013)	(0.015)
White	14,480	−0.028*	0.038***	0.070***
		(0.016)	(0.012)	(0.020)
Less than high school	25,880	−0.054	0.075***	0.038**
		(0.021)	(0.019)	(0.016)
High school+	39,720	−0.026*	0.028**	0.069***
		(0.012)	(0.012)	(0.016)
Panel B: Combined new hires earnings in retail trade and accommodation/food services (in $1000s)				
Black	52,560	−0.036**	0.032**	0.032*
		(0.015)	(0.014)	(0.017)
White	14,480	−0.022	0.028**	0.059**
		(0.020)	(0.011)	(0.028)
Less than high school	25,880	−0.048**	0.079***	0.035
		(0.024)	(0.020)	(0.022)
High school+	39,720	−0.019	0.012	0.030*
		(0.014)	(0.012)	(0.018)

Notes: Marginal effects are presented, along with standard errors (in parentheses). Standard errors are adjusted for clustering at the individual level. The dependent variable is defined as follows: 1 = no work, no welfare; 2 = no work, welfare; 3 = work, welfare; 4 = work, no welfare. The base category in all models is 1. The dependent variable in Panel A is expressed in logarithms, and the dependent variable in Panel B is expressed in thousands of dollars. All models include the variables presented in Table 2, except the unemployment rate. ***, **, * indicate statistical significance at the 1%, 5%, and 10% levels, respectively

work/no welfare alternative. Thus, it appears that improved labor market opportunities offer a strong incentive for African American women to work in conjunction with welfare receipt. White women, on the other hand, respond to labor market improvements in a way that encourages them to rely on work alone for financial well-being. A possible explanation for these differential effects is that some groups of women have greater skills and fewer work barriers that allow them to take full advantage of economic expansions.[25]

Turning to the models for educational attainment, we find similar patterns. High school drop-outs are more responsive to labor market opportunities at the no

[25] Another explanation for these results is that white women earn more than black women, and thus lose eligibility for welfare more rapidly. To explore this possibility, we compare quarterly earnings across white and black women over entire study period. Average quarterly earnings for white women are about $2,660, while earnings for black women are $2,723. Therefore, it appears that earnings disparities cannot explain the pattern of results emerging from Table 8. It should be noted that comparisons of marginal effects across black and white women should be done cautiously, given that baseline percentages of work and welfare are quite different for these groups.

work/welfare and work/welfare alternatives, while those who completed high school are more responsive at the work/no welfare alternative. The results appear to corroborate the skills argument. While a favorable economic environment is a strong incentive for low-skilled women to combine work with welfare, it cannot by itself encourage such women to move entirely off welfare and into work. Higher-skilled women operating in favorable conditions are in a better position to take advantage of labor market opportunities, and thus are more responsive to such opportunities as they move into employment without continuing on the welfare rolls.

We also examine the impact of labor market conditions on the traditional dichotomous work and welfare decisions. All employment variables, including new hires and new hires' earnings, are strongly associated with employment, but only the unemployment rate is statistically significant in the welfare use model. However, the absolute value of the unemployment coefficient is appreciably larger in the employment model than in the welfare model (coefficient value −1.7 vs. 0.8). Such results imply that individuals are more sensitive to labor market opportunities when deciding to engage in any work, even if employment coincides with welfare use. Hoynes (2000) finds similar results: economic conditions are consistently related to welfare exits but largely unrelated to reentry.

Conclusion

Merging administrative data for the state of Maryland with gender- and age-specific labor market information, this paper estimates the impact of economic conditions on the full set of work–welfare alternatives. In doing so, we make a number of contributions to the literature. We introduce the QWI database as a potentially rich resource for county-level economic indicators. New hires and new hires' earnings, in particular, are superior to measures of overall employment and earnings because they accurately reflect what recent welfare leavers are likely to encounter during their job search. Moreover, the QWI database allows users to filter economic measures by gender, age-group, industry, and region. Our administrative data also have a number of desirable qualities: diverse sample, accuracy in employment and welfare reporting, and access to quarterly observations over a 10-year period.

Our results suggest that women respond to local labor market opportunities in ways predicted by standard economic models. As the unemployment rate declines and new hires and earnings in key industries increase, women are expected to choose alternatives that include work, even if it coincides with welfare receipt. Improved economic conditions decrease the incentive to remain out of work and on welfare. These findings hold for local employment and earnings opportunities in the retail and accommodation/food services industries, as well as for the unemployment rate. Our findings also suggest that African American women and high school drop-outs respond differently to labor market conditions than white women and those with at least a high school degree. Specifically, the former groups tend to be more responsive to economic improvements when they are encouraged to work in conjunction with welfare receipt. Economic opportunities provide more powerful incentives for the latter groups as they move into employment without welfare receipt.

Two important questions are raised by this paper. First, can our results shed light on why previous studies have not consistently found significant relationships between economic conditions and work–welfare decisions? Second, to what extent can labor market opportunities, by themselves, fulfill the goal of recent welfare reforms? As to the first question, our results suggest that county as opposed to state-level labor market data are important for precisely estimating labor market effects. Disaggregating labor market measures by gender, age-group, and relevant industries is also preferable to overall measures of the unemployment rate, employment, and earnings. Still another factor is the availability of multiple observations per year, so that short-term employment shocks are accounted for and short employment/welfare spells are identified. While reducing the magnitude of estimated effects, in most cases the inclusion of fixed effects and time trends does not qualitatively alter results from simpler models. Thus, unobserved heterogeneity is important but not determinative.

As to the second question, it is important to stress that in an era of work-based social policy, strong economic growth is increasingly relied upon to move former welfare recipients toward economic independence. Our results suggest, however, that economic growth is not a panacea. Favorable labor market conditions alone cannot reliably move welfare recipients into employment. For many women, improved labor market opportunities provide strong incentives to combine work with welfare. For fewer women are these incentives strong enough to rely on work alone for financial well-being. In addition, the presence of heterogeneous labor market effects complicates our understanding of who is most likely to benefit from favorable economic conditions, or which characteristics public policy can reasonably alter so that women are in a position to take full advantage of economic expansions.

Our findings have important policy implications. Several provisions of states' welfare reforms rely extensively on the ability of local labor markets to absorb welfare recipients and provide them with incentives to offset participation in cash assistance programs. The increased use of mandatory job search programs and formal cash diversion grants are evidence of this reliance. A mixed regime of policy "carrots" and "sticks" that places equal emphasis on strong work mandates when economic conditions are favorable and softening those mandates when conditions are less favorable is necessary in an environment that stresses a "work first" philosophy. One way to achieve this balance is for national policymakers and state welfare administrators to shift focus from state-level assessments of economic conditions to one that emphasizes local employment opportunities. Anecdotal evidence from Maryland suggests that county administrators adjust policy instruments in an ad hoc manner in order to deal with changing economic conditions. Such policy adjustments are likely to decouple the relationship between work–welfare decisions and labor market opportunities during an economic downturn, followed by a strengthening of this relationship during periods of growth. In any case, an explicit policy shift that provides local offices with additional flexibility would allow caseworkers to systematically tailor TANF reforms to the economic environment in which welfare recipients operate.

A few caveats should also be mentioned. Most obviously, given that our analysis focuses on young women (ages 19–28) in one state (Maryland), generalizations of our results must be done cautiously. It remains to be seen whether our estimates apply to all working-age single mothers in states with increasingly heterogeneous economic environments. Future work in this area should therefore attempt to pool multiple states, now a strong possibility given that the Workforce Indicators data are available for most states. Another caveat is that our observation period, 1996–2005, covers years in which the U.S. economy grew at a rapid pace, notwithstanding the brief recession in 2001. Future work should also extend the analysis period to cover the downturn that began in December of 2007. A final caveat is that our sample is constructed to include women with current or future interaction with the welfare system. Future research should broaden the sample selection criteria to include all single mothers.

Acknowledgements We thank Richard Larson, Peter Mueser and Jane Staveley for useful comments on an earlier draft of this paper. Other expository editing followed ACF/OPRE Welfare Research and Evaluation conference and National Association for Welfare Research and Statistics conference presentations.

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