I Z A Institute of Labor Economics

Initiated by Deutsche Post Foundation

DISCUSSION PAPER SERIES

IZA DP No. 12425

Is There a Business Cycle Effect on the Incidence of Dual Job Holding?

Chung Choe
Ronald L. Oaxaca
Francesco Renna

JUNE 2019

I Z A Institute of Labor Economics

Initiated by Deutsche Post Foundation

DISCUSSION PAPER SERIES

IZA DP No. 12425

Is There a Business Cycle Effect on the Incidence of Dual Job Holding?

Chung Choe
Hanyang University

Ronald L. Oaxaca
University of Arizona, LISER, IZA and PRESAGE

Francesco Renna
University of Akron

JUNE 2019

ISSN: 2365-9793

IZA – Institute of Labor Economics

Schaumburg-Lippe-Straße 5–9
53113 Bonn, Germany

Phone: +49-228-3894-0
Email: publications@iza.org

www.iza.org

ABSTRACT

Is There a Business Cycle Effect on the Incidence of Dual Job Holding?

This paper examines the extent to which the incidence of dual job holding is cyclically sensitive in the context of hours constraints on labor supply. Linear probability models of the incidence of dual job holding are estimated separately for each hours constraint regime. Selection effects are accounted for in a correlated random effects setting in which selection into overemployment, unconstrained hours, and underemployment is separately estimated each year from an ordered probit model. As measured by the local unemployment rate, transitory business cycle movements have no effect on the incidence of dual job holding. However, a sustained change in the local unemployment rate reduces the incidence of dual job holding among workers who are not hours constrained on their main jobs.

JEL Classification: J01, J22, J49

Keywords: dual job, labor supply, hours constraint, business cycle

Corresponding author:
Francesco Renna
Department of Economics
University of Akron
Akron, OH 44325-1908
USA

E-mail: frenna@uakron.edu

1 Introduction

Changes in economic conditions over the business cycles have important repercussions for the labor market. The unemployment rate is expected to fluctuate in a countercyclical fashion as the economy expands or contracts and new jobs are created or eliminated. Not clear, however, is the behavior of dual job holding over the business cycle. In fact, the U.S. multiple job rate remained relatively constant over the 2000s notwithstanding the 2001 recession and the great recession of 2008 (F. Hipple, 2010). More specifically, while the U.S. unemployment rate increased by 5.5 percentage points between December 2007 and June 2009, the rate of dual job holding increased by only 0.06 percentage points, from 5.44 to 5.50, during the same time period (Lalé, 2015). Not surprisingly the limited research on this topic has found at best weak evidence that dual job holding responds to worsening of employment prospects (Bell et al., 1997). In this study we expand on the previous literature by disaggregating dual job holders according to their motivations for holding two jobs.

On theoretical grounds, dual job rates can vary either procyclically or countercyclically. Over the business cycle, employers can adjust to the new quantity of labor demanded by either changing the number of workers or the number of hours of work per worker (or some combination of the two). If the fixed cost of hiring and laying-off workers is high enough, employers may prefer to hoard labor during a downturn of the economy and cut back on the number of hours. Under the latter scenario, workers will be more likely to move to a situation of underemployment and would try to obtain a second job to bridge the gap between the desired and the actual hours of work.[1] Thus we would expect the preference for dual job holding to increase during a recession. Conversely, during an expansion of the economy, employers may ask employees to work overtime instead of hiring more workers. This decision would lead to a decrease in the incidence of dual job holding during an expansion because the increase in the hours worked on the main job together with the increase in the effective marginal wage due to overtime regulation would decrease the propensity

[1] A similar argument would apply in the case of a type of added worker effect for joint labor supply in which if a spouse loses their job, the other spouse seeks a second job to compensate for the reduction in family labor income.

to take a second job (Shishko and Rostker, 1976). These arguments would suggest that dual job holding should be countercyclical.

However, since labor demand is also expected to decline during a recession, the fact that more workers are seeking a second job during a recession does not necessarily lead to more workers having a second job. Moreover, we would expect firms to first discharge workers with the lowest job attachments during a contraction of the economy. To the extent that secondary jobs tend to be non-traditional arrangements, it is more likely that dual job holders would lose their second job before regularly employed workers. As such the rates of dual job holding would tend to exhibit a procyclical pattern. Ultimately, whether the dual job rate is procyclical or countercyclical is an empirical question.

Interestingly, not every dual job holder seeks a second job in response to an underemployment condition on the main job. In fact, there is strong evidence that the largest proportion of workers hold two jobs for reasons other than an hours constraint (Choe et al., 2018). These alternative motives include: 1) a way to gradually transition to a new primary job or to self-employment (Panos et al., 2014); 2) seeking utility enhancing characteristics on a second job that are not available on the main job because jobs are not perfect substitutes (Kimmel and Smith Conway, 2001); or 3) two jobs may be complimentary with one another such as when the worker uses the credential earned on one job to provide consulting services on the side (Paxson and Sicherman, 1996). Typically this set of motives is grouped under the "job portfolio" term by the previous literature. Contrary to the rate of dual job holding due to underemployment conditions, which is the result of the relative shift in labor demand and supply over the business cycle, the rate of dual job holding under the job portfolio motive should be subject only to fluctuation of labor demand because there is no reason to believe that the motivation for holding a portfolio of jobs should be affected by the business cycle.[2] Consequently, the rate of dual job holding under the job portfolio regime should be clearly procyclical.

[2] A notable exception to this conclusion is the decision to hold two jobs as a form of hedging against the risk of unemployment. However, the current literature finds little evidence for such behavior (Bell et al., 1997).

2 Literature Review

The previous literature has not been able to precisely pinpoint the nature of the behavior of dual job rates over the business cycle, probably because of its theoretical ambiguity. For example, Amuedo-Dorantes and Kimmel (2009) found that the dual job rate for males does not fluctuate over the business cycle but the female dual job rate does. Surprisingly, though, they found that the female dual job rate behaved countercyclically during the 80s, it was not responsive to business cycles during most of the 90s, and it behaved procyclically afterward. This observed behavior would be consistent with a change in the composition of dual job holders over time, with constrained workers being more prevalent in the latter part of the period under analysis. However, since information about the motivation for holding two jobs is missing from their data set, this hypothesis cannot be tested and the change in the nature of cyclicality of dual job rates remains unexplained.

Using a sample of 27 European countries between 1998 and 2011, Zangelidis (2014) finds that dual job rates are negatively related to national unemployment rates, but that long-term unemployment (measured by the percentage of people who remain unemployed for more than 12 months) increases the incidence of dual job holding. Contrary to Amuedo-Dorantes and Kimmel (2009), the author finds that women are less sensitive to the worsening of economic conditions. Using CPS data from 1998 to 2013, Hirsch et al. (2016) finds that dual job rates are slightly procyclical but this result becomes insignificant after conditioning on Metropolitan Statistical Area fixed effects. This conclusion is robust to using alternative measures of the business cycle such as the local unemployment rate and local employment growth.

One possible factor that could explain this weak evidence of a statistical relationship between the business cycle and dual job holding is whether the empirical model is trying to capture the short or the long run effect of the unemployment rate on dual job holding rates. For example, using aggregate measures of dual job holding for the 48 lower states from 1994 to 1998, Partridge (2002) finds some evidence of procyclicality in the dual job holding rates in the short run, but no statistically significant relationship in the long run. This result would be consistent with the interpretation that most of the dual jobs created during the economic expansion disappear in the long run, thus suggesting that the dual job market acts as a buffer for tight markets, but eventually this effect fades out as regional labor markets adjust, possibly due to migration. Still, even this literature does not reach a consensus about the behavior of dual job holding rates over the business cycle. In fact, Panos et al. (2014), using a sample of British data from BHPS, concludes that dual job holding is countercyclical in the short run, although it becomes insignificant in the long run.

In this study, we contribute to the literature on the behavior of dual job holding over the business cycle by incorporating in our model the motivation for holding two jobs. This is an important extension because many studies have concluded that dual job holders are not a homogenous group of workers and as such they may respond differently to fluctuations in labor market conditions.

3 Conceptual Framework

We estimate the business cycle effect on the incidence of dual job holding for each hours constraint regime in a correlated random effects setting. Our estimation strategy employs a two stage procedure. The first stage sorts out the selection effects associated with hours constrained versus unconstrained employment on a worker's main job. The second stage uses a linear probability model (LPM) and conditions on the selection mechanism to estimate the effects of the business cycle on the probability of being a dual job holder within each category of hours constrained workers.

3.1 First Stage

The strategy employed to capture selection into each of the three constrained labor supply regimes is to obtain Inverse Mills Ratios from ordered probit models estimated separately for each year of the panel. Workers are classified into three mutually exclusive employment states: Underemployed, Unconstrained, Overemployed. Hours constrained workers are either working less than their desired hours on the main job (Underemployed) or working more than their desired hours on the main job (Overemployed). Those who are working their desired hours on the main job are assigned to the Unconstrained employment category.

The degree of constrained hours on the main job is defined as $C_{it}^* = h_{1it} - h_{1it}^* \gtreqless 0$, where h_{1it} is the observed hours worked on the main job and h_{1it}^* is the unobserved desired hours of work on the main job. Although we do not observe the latent variable C_{it}^*, we observe the ordered categorical variable C_{it}:

$$C_{it} = \begin{cases} 2 & \text{if } h_{1it} - h_{1it}^* > 0 \text{ (overemployed)} \\ 1 & \text{if } h_{1it} - h_{1it}^* = 0 \text{ (unconstrained)} \\ 0 & \text{if } h_{1it} - h_{1it}^* < 0 \text{ (underemployed).} \end{cases}$$

We examine the case in which the multinomial selection of C_{it} into the three mutually exclusive categories of constrained labor supply follows an ordered probit process:

$$\begin{aligned} C_{it}^* &= \psi_0 + X_{it}\beta_t + \varepsilon_{it} \\ &= I_{it} + \varepsilon_{it}, \end{aligned}$$

where I_{it} is an index function, ψ_0 is a constant term, X_{it}^* is a vector of determinants of constrained hours, and $\varepsilon_{it} \sim N(0,1)$. It follows that

$$\begin{aligned} C_{it} &= 0 \;\text{ if }\; C_{it}^* \leq 0 \\ &= 1 \;\text{ if }\; 0 < C_{it}^* \leq \mu_1 \\ &= 2 \;\text{ if }\; \mu_1 \leq C_{it}^*, \end{aligned}$$

The ordered probit model is estimated separately for each year t in order to construct the estimated Mills Ratios for the LPM's that will appear in the dual job LPM's for each of the subsamples corresponding to three constrained hours regimes.

The probabilities for each constrained hours regime are determined according to

$$
\begin{aligned}
Prob(C_{it} = 0 \mid X_{it}) &= \Phi(-I_{it}) \\
&= 1 - \Phi(I_{it}) \\
Prob(C_{it} = 1 \mid X_{it}) &= \Phi(\mu_1 - I_{it}) - \Phi(-I_{it}) \\
&= \Phi(\mu_1 - I_{it}) - [1 - \Phi(I_{it})] \\
&= \Phi(\mu_1 - I_{it}) + \Phi(I_{it}) - 1 \\
Prob(C_{it} = 2 \mid X_{it}) &= 1 - \Phi(\mu_1 - I_{it}).
\end{aligned}
$$

The corresponding densities are determined according to

$$
\frac{\partial \Phi(-I_{it})}{\partial(-I_{it})} = \phi(-I_{it}) = \phi(I_{it})
$$

$$
\begin{aligned}
\frac{\partial \left[\Phi(\mu_1 - I_{it}) + \Phi(I_{it}) - 1\right]}{\partial(-I_{it})} &= \frac{\partial \Phi(\mu_1 - I_{it})}{\partial(-I_{it})} + \frac{\partial \Phi(I_{it})}{\partial(-I_{it})} \\
&= \frac{\partial \Phi(\mu_1 - I_{it})}{\partial(\mu_1 - I_{it})} \frac{\partial(\mu_1 - I_{it})}{\partial(-I_{it})} + \frac{\partial \phi(I_{it})}{\partial(I_{it})} \frac{\partial I_{it}}{\partial(-I_{it})} \\
&= \phi(\mu_1 - I_{it}) - \phi(I_{it})
\end{aligned}
$$

$$
\begin{aligned}
\frac{\partial[1 - \Phi(\mu_1 - I_{it})]}{\partial(-I_{it})} &= -\frac{\partial \Phi(\mu_1 - I_{it})}{\partial(-I_{it})} = -\frac{\partial \Phi(\mu_1 - I_{it})}{\partial(\mu_1 - I_{it})} \frac{\partial(\mu_1 - I_{it})}{\partial(-I_{it})} \\
&= -\phi(\mu_1 - I_{it}).
\end{aligned}
$$

As shown in (Choe et al., 2018), the labor supply model for weekly hours is a function of the wage rates for job 1 (the main job) and job 2 (for dual job holders), and non-labor income. Since the ordered probit model for hours-constrained employment regimes includes both unitary and dual job holders, X_{it}^{*} in the index function (I_{it}) includes the wage rate for job 1 and non-labor income as these variables are common to all workers. Other covariates are age, education, marital status, number of dependent children, local unemployment rate, industry, and occupation.

3.2 Second Stage

The second stage of the estimation entails estimating linear probability models of dual job holding separately for each of the hours constraint regimes.[3] The linear probability models are corrected for sample selection. We construct the IMR's from the ordered probit constrained hours selection model:

$$\lambda_{0it} = \frac{\phi(I_{it})}{1 - \Phi(I_{it})} \text{ (underemployed)}$$

$$\lambda_{1it} = \frac{\phi(\mu_1 - I_{it}) - \phi(I_{it})}{\Phi(\mu_1 - I_{it}) + \Phi(I_{it}) - 1} \text{ (unconstrained)}$$

$$\lambda_{2it} = \frac{-\phi(\mu_1 - I_{it})}{1 - \Phi(\mu_1 - I_{it})} \text{ (underemployed).}$$

The selectivity corrected LPM's for dual job holding are specified by

$$D_{it} = x_{it}\alpha_j + \bar{z}_i\gamma_j + \theta_j\hat{\lambda}_{jit} + u_{jit},\ j = 0, 1, 2$$

where $D_{it} = 1$(dual job holder), x_{it} is a vector of time varying covariates, $\bar{z}_i$ is a vector of the time averaged means (Mundlak variables), λ_{jit} is the IMR for the jth constrained hours regime, and u_{jit} is an error term. The conditioning variables include the wage rate for job 1, non-labor income,

[3] An individual can of course appear in more than one category at different points in time.

age, education, marital status, number of dependent children, and the local unemployment rate. An additional set of variables are introduced to capture the effects of correlated random effects following Wooldridge(1995; 2010) and Oaxaca and Choe (2016). These additional variables are the time-averaged means of all of the exogenous variables in the model, including the exclusion variables. Also included is the Inverse Mills Ratio (IMR) interacted with year dummies, where the IMR's are obtained from the estimated first stage multinomial probit models.

Following Panos et al. (2014), we can obtain the permanent (persistent) effects as the sum of the transitory and mean (Mundlak) effects of the variables. The central variable of interest here is the local unemployment rate. Let u_{it} represent the local unemployment rate for the ith individual at time t. Let U_{mi} represent the time averaged values of the local unemployment rates for individual i, where $U_{mi} = \frac{\sum_{t=1}^{T_i} U_{it}}{T_i}$ and T_i is the number of time series observations for individual i. The transitory, mean, and persistent effects of the local unemployment rate on dual job holding in a linear probability setting are respectively given by

$$\frac{\partial D_i}{\partial U_i} = \alpha_u$$

$$\frac{\partial D_i}{\partial U_{mi}} = \gamma_{um}$$

$$\begin{aligned} \left(\frac{\partial D_i}{\partial U_i}\right)_p &= \frac{\partial D_i}{\partial U_i} + \frac{\partial D_i}{\partial U_{mi}} \frac{\partial U_{mi}}{\partial U_i} \\ &= \alpha_u + \gamma_{um}, \end{aligned}$$

since $U_{it} = U_i \ \forall t \Rightarrow U_{mi} = U_i \Rightarrow \frac{\partial U_{mi}}{\partial U_i} = 1.$

4 Data

The estimation of the model specified in the previous section is carried out using data from the British Household Panel Survey (BHPS) which was started in 1991. The last year the survey was administered was 2008. The initial sample consisted of 5,050 household representative of the British population south of the Caledonian Channel. Over time the sample was expanded to represent the entire UK population. By 2001, about 10,000 households were included in the sample. We define the local unemployment rate as the annual unemployment rate for each of the nine English regions (North East, North West, York and the Humber, East Midlands, West Midlands, East, London, South West, and South East), plus Wales, Scotland, and Northern Ireland. Using the geographical location of the male head of the household, we match the local unemployment statistics obtained from the Office of National Statistics (https://www.ons.gov.uk/timeseriestool) to each respondent in BHPS. Since the local unemployment rates are missing for 1991, our sample does not include the first wave of BHPS. The conceptual framework spelled out in the previous section abstains from the option of not holding any job at all, as such it seems to better describe the economic decisions for men, who historically have shown stronger labor force attachment. Also, self-employed individuals can freely adjust their hour of work to their desired target, thus we restrict our analysis only to male employees.

Figure 1 plots the incidence of dual job holding among the three categories of hours constrained workers along with the overall UK unemployment rate over the period of our study. The unemployment rate generally declined over the period of our study with a small rise after 2006. Dual job holding among workers unconstrained on their main job did not bear much overall relationship to movements in the unemployment rate, with frequent sign changes in movements. Among those who were underemployed on their main job, dual job holding tended to move in the opposite direction of the unemployment rate until 1996 (procyclical). After 1996 dual job holding among the underemployed generally declined along with the unemployment rate until 2005 (countercyclical) and then rose thereafter (procyclical). After some initial ups and downs prior to 1998, dual job holding among workers who reported being overemployed on their main job declined fairly steadily along with the overall unemployment rate (countercyclical).

Table 1 reports the descriptive statistics of the entire set of variables used in the analysis. All monetary values have been adjusted for inflation using 2008 as the base year. To derive the hourly wage on job 1, we divide monthly labor earning from job 1 by the sum of standard weekly hours and overtime hours worked on job 1 multiplied by 4.3. Non-labor income is computed as annual family non-labor income divided by 52. The information about dual job status pertains to the month before the interview date. Finally, the constraint regime for each individual is identified in our data by the answer to the survey question "*Thinking about the hours you work, assuming that you would be paid the same amount per hour, would you prefer to ...*", with respondents having to choose among three possible scenarios: (1) work fewer hours, (2) work more hours, and (3) continue with same hours.

When looking at Table 1, two facts that are relevant to our research question stand out. While individuals can move across hours constraint regimes over the sample period, on average unconstrained individuals who report only 1 job face the same unemployment rate as unconstrained individuals who work two jobs. This would suggest that dual-job holding is acyclical for unconstrained workers. Under and overemployed dual job holders, instead, face slightly higher unemployment rates although this difference is statistically significant only for overemployed workers, thus suggesting that overemployed dual job holding is countercyclical. Also, we found that the incidence of dual job holding declined over the sample period.

We observe substantial differences in some of the observable characteristics of individuals across labor supply regimes. Single job holders earn higher wages on job 1 irrespective of their hours constraint status. Interestingly, the hourly wage increases monotonically as people move from the underemployed to the overemployed regime, which is consistent with the presence of a backward bending supply curve. While unconstrained dual job holders have higher non-labor income than unconstrained single job holders, we find no statistical evidence of a difference in mean weekly non-labor income between constrained unitary and dual job individuals. Unitary unconstrained and overemployed individuals are typically older than their dual job holder counterparts. Consistently across all three constraint regimes, individuals who did not complete any secondary education (*None of these*) have a lower incidence of dual job holding. Among underemployed workers, dual job holders are less likely to have an *A level* diploma. Among unconstrained, dual job holders are less likely to have an undergraduate (*1st degree*) or a two year (*HND, HNC, Teaching*) college degree but more likely to have an *O level* and a *CSE* diploma. Finally, among overemployed, dual job holders are more likely to have an *O level* diploma.

Family characteristics also vary across constraint regimes and dual job status. The incidence of being married monotonically increases as we move from underemployment to unconstrained to overemployment. While there is no statistical difference in marital status between unitary and dual job unconstrained individuals, unitary unconstrained and overemployed workers are more likely to be married than their dual job holder counterpart. Unitary and dual underemployed job holders have on average the same number of children, however unconstrained and overemployed dual job holders have more children than unitary job holders.

Finally, we observe some variation in terms of the distribution of workers classified by industry and occupation of employment on their main job. For example, only 16.6 percent of the sample of unconstrained unitary job holders work in a public, health or teaching establishment while 25.3 percent of the dual job holders work in these types of establishments on their main job. We argue in this paper that industry and occupation on the main job do not directly affect the probability of holding one versus two jobs, but these do affect the odds of being in one of the three constraint regimes because different industry/occupation employments may require different contractual hours of work. Note that under this assumption, industry and occupation of the main job would still affect the probability of holding two jobs, but only indirectly through the IMR and the Mundlak variables.

5 Empirical Results

Following Wooldridge (2010, pp.832-834), we conduct tests for sample selection bias in a panel data model. In our setting, the tests involve first estimating a pooled ordered probit model and obtaining the Inverse Mills Ratio for each of the three hours-constrained regimes. We next estimate the LPM dual job choice model by fixed effects for each regime with the IMR included. Based on robust standard errors for the estimated coefficients on the IMR's, we can reject the null hypothesis of no sample selection bias only in the case of unconstrained workers. However, for ease of comparison across regimes we estimate the full sample selection model for all three regimes.

While the first stage of the two-stage estimation strategy for testing for the presence of sample selection bias uses pooled ordered probit, we also estimate a random effects ordered probit model to provide a better sense of the effects of the determinants of selection into the three constrained hours regimes over the period of our study. Table 2 reports the estimated parameters and marginal effects from the random effects ordered probit model. The first two columns of results correspond to the estimated parameters and standard errors of the model. With the exceptions of the first (underemployed) and last (overemployed) categories of constrained hours, the signs of the parameters are uninformative as to the directions of the effects of the determinants on the probability of assignment to the middle catgegory (unconstrained). Consequently, we report the estimated marginal effects and standard errors in the remaining columns.

We briefly examine a subset of the variables that are statistically significant. Non-labor income raises the probabilities of being underemployed or unconstrained while reducing the probability of being overemployed. Being married is associated with a reduction in the probabilities of being underemployed or unconstrained and an increase in the probability of being overemployed. Marginal effects associated with the local unemployment rate are particularly large. Increases in the (local) unemployment rate raise the probabilities of being underemployed or unconstrained while lowering the probability of being overemployed. To the extent that higher unemployment reduces the availability of work hours, the less likely individuals would be working more than desired hours and the more likely that they would be working either their desired hours or less than their desired hours.

Table 3 reports the estimated panel data LPM dual-job choice models corrected for sample selection. As was the case in the fixed effects LPM estimation in the test for sample selection bias, the current (transitory) effect of the local unemployment rate on the incidence of dual job holding is statistically insignificant in all three hours-constrained regimes. Of course the mean effects of the local unemployment cannot be estimated in the fixed effects setting. As estimated in the dual-job choice model corrected for sample selection, the mean (Mundlak) local unemployment rate effect is statistically insignificant for hours constrained workers but is statistically significant and negative for those who are unconstrained. The persistent local unemployment rate effect on dual job holding among hours unconstrained workers is 0.12 - 0.75 = -0.63. With an estimated asymptotic standard error of (0.22), the persistent effect is statistically significant at the 1% level. To put this in perspective, the mean of the positive changes in the historical local unemployment rate was $\Delta U^{+} = 0.0055$, and the mean historical negative change was $\Delta U^{-} = -0.0080$. So for a worker whose hours on their main job are unconstrained, a permanent increase in the local unemployment rate of 0.55 percentage points would decrease the probability of holding a second job by $(0.0055) \times (-0.63) = -0.0035$, i.e. a reduction of 0.35 percentage points. Likewise, a permanent decrease in the local unemployment rate of -0.80 percentage points would increase the probability of holding a second job by $(-0.0080) \times (-0.63) = 0.0050$, i.e. an increase of 0.5 percentage points. These procyclical findings are consistent with the increase (reduction) in the availability of second jobs for workers who are unconstrained on their main job when the local unemployment rate falls (rises).

Current real hourly earnings on the main job has a negative effect on the probability of dual job holding for all three hours constrained categories but is statistically significant only for under-employed and unconstrained workers. Thus, higher wages on one's main job reduce the incentive for holding a second job even if one reports working less than their desired hours on their main job. On the other hand, although current real non labor income and married status also exhibit negative effects on dual job holding across all three hours constrained regimes, these factors are never statistically significant. We measure child responsibilities by the number of children under the age of 6. The presence of young children is associated with a higher incidence of dual job holding across all three hours constrained regimes but is statistically significant only for those who were underemployed on their main job. This could very well reflect a targeted earnings objective for these workers. Individual year effects are generally statistically insignificant with the exception of overemployed workers. There are four years in which the year effects relative to 1992 are statistically significant: 1993, 2004, 2006, and 2007. For all four years, the year effects are negative and slightly increasing in negativity. There does not appear to be any obvious events in these years that would account for the diminishing incidence of dual job holding among those who reported that they were working more than their desired hours on their main job.

6 Summary and Conclusions

In seeking to identify any business cycle effects on the incidence of dual job holding, we first address potential selection issues pertaining to facing constrained working hours on one's main job. Our tripartite division of the hours constraints regimes consists of underemployment, unconstrained, and overemployment. The natural ordering of this division of hours constraints leads to an ordered probit model from which Inverse Mills ratios are constructed and added to linear probability models of dual job holding for each of the three hours constraints regimes.

After correction for sample selection in a correlated random effects setting, inferences about the cyclicality of dual job holding are somewhat in aggreement with what one might conclude on the basis of the unconditional patterns presented in Figure 1. In the case of overemployed workers, the transitory countercyclical pattern of dual job holding present in raw data continues to hold in the estimated model. Similarly among unconstrained workers, the lack of any systematic relationship in the raw data between dual job holding and the unemployment rate is maintained in the estimated model. On the other hand after controlling for covariates and correction for sample selection, it remains the case that there are no persistent business cycle effects on the incidence of dual job holding among underemployed workers. With respect to this latter result, the counteracting factors from the demand and supply sides of the labor market are offsetting.

Recall that our theoretical prediction was that dual job holding would be procyclical only under the job portfolio hypothesis (unconstrained dual job holding), and to some extent our results confirm that hypothesis. On the other hand for constrained workers the relationship could not be predetermined. During an economic downturn, if employers are hoarding labor by keeping workers on the payroll but reducing the number of work hours, we would expect the fallout to especially impact the overemployed. And we found evidence of that from the ordered probit results in which the relatively large negative marginal effect of the local unemployment rate on the probability of being an overemployed worker. As such, some overemployed dual job holders may transition to unconstrained dual job holders during a recession. This type of sorting process may explain why in the linear probability model for dual job holding we do not find a significant impact of our measures of the business cycle on the probability of holding two jobs for overemployed workers.

We should point out that identification of the relationship between local unemployment and the incidence of dual job holding may be hindered by the small number of observations for underemployed dual job holders, especially when the variation in local labor market rests on only 12 regional markets. From a policy point of view, our results suggest that labor policies which permanently alter the long run equilibrium unemployment rate may have important repercussion on the secondary job market of dual job seekers, at the very least for those operating under the job portfolio paradigm. As these individual maximize their utility by holding two jobs, any restriction on their ability to find a second job would be associated with a welfare loss.

References

Amuedo-Dorantes, C. and Kimmel, J. (2009). Moonlighting Over The Business Cycle. *Economic Inquiry*, 47(4):754–765.

Bell, D., Hart, R. A., and Wright, R. E. (1997). Multiple Job Holding as a 'Hedge' Against Unemployment. CEPR Discussion Papers 1626, C.E.P.R. Discussion Papers.

Choe, C., Oaxaca, R. L., and Renna, F. (2018). Constrained vs unconstrained labor supply: the economics of dual job holding. *Journal of Population Economics*, 31(4):1279–1319.

F. Hipple, S. (2010). Multiple jobholding during the 2000s. *Monthly labor review / U.S. Department of Labor, Bureau of Labor Statistics*, 133:33–34.

Hirsch, B. T., Husain, M. M., and Winters, J. V. (2016). Multiple job holding, local labor markets, and the business cycle. *IZA Journal of Labor Economics*, 5(1):1–29.

Kimmel, J. and Smith Conway, K. (2001). Who moonlights and why? evidence from the sipp. *Industrial Relations: A Journal of Economy and Society*, 40(1):89–120.

Lalé, E. (2015). Multiple jobholding over the past two decades. *Monthly Labor Review*, page 1.

Oaxaca, R. and Choe, C. (2016). Wage decompositions using panel data sample selection correction. *Korean Economic Review*, 32(2):201–218.

Panos, G. A., Pouliakas, K., and Zangelidis, A. (2014). Multiple job holding, skill diversification, and mobility. *Industrial Relations: A Journal of Economy and Society*, 53(2):223–272.

Partridge, M. (2002). Moonlighting in a high growth economy: Evidence from u.s. state-level data. *Growth and Change*, 33(4):424–452.

Paxson, C. H. and Sicherman, N. (1996). The dynamics of dual job holding and job mobility. *Journal of Labor Economics*, 14(3):357–393.

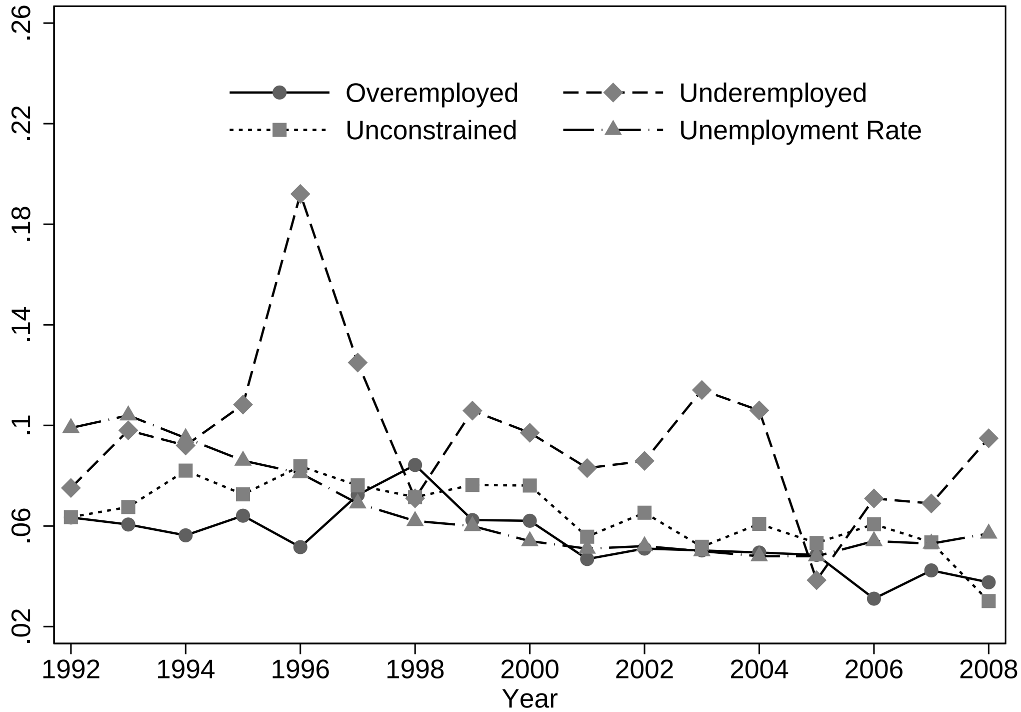
Shishko, R. and Rostker, B. (1976). The economics of multiple job holding. *The American Economic Review*, 66(3):298–308.

Wooldridge, J. (2010). *Econometric Analysis of Cross Section and Panel Data*, volume 1. The MIT Press, 2 edition.

Wooldridge, J. M. (1995). Selection corrections for panel data models under conditional mean independence assumptions. *Journal of Econometrics*, 68(1):115 – 132.

Zangelidis, A. (2014). Labour Market Insecurity and Second Job-Holding in Europe. Working Paper 1626, C.E.P.R. Discussion Papers.

Figure 1: Incidence of Dual Job Holding across Different Labor Supply Regimes



Notes: Rates based on a sample from the British Household Panel Survey (1992-2008) restricted to individuals aged 18 to 65.

Table 1: Summary Statistics: Mean and Difference

Variable	Underemployed Unitary	Underemployed Dual	Underemployed Difference	Unconstrained Unitary	Unconstrained Dual	Unconstrained Difference	Overemployed Unitary	Overemployed Dual	Overemployed Difference
Wage rate on job 1	9.455	8.418	1.037∗	11.712	10.416	1.296∗	12.492	11.571	0.922∗
Weekly non-labour income	0.074	0.083	-0.009	0.069	0.070	-0.002	0.067	0.064	0.003
Age	32.898	32.115	0.783	37.718	36.101	1.618∗	40.835	38.302	2.534∗
Higher degree (omitted)	0.021	0.031	-0.010	0.036	0.043	-0.007	0.035	0.063	-0.028∗
1st degree	0.113	0.123	-0.010	0.139	0.120	0.019†	0.149	0.146	0.003
HND, HNC, teaching	0.072	0.057	0.014	0.088	0.073	0.015†	0.099	0.090	0.010
A level	0.258	0.341	-0.083∗	0.244	0.246	-0.002	0.237	0.235	0.002
O level	0.284	0.291	-0.007	0.261	0.297	-0.036∗	0.246	0.274	-0.028‡
CSE	0.083	0.061	0.022	0.067	0.083	-0.016†	0.056	0.064	-0.008
None of these	0.169	0.096	0.073∗	0.164	0.138	0.026∗	0.178	0.128	0.050∗
Married (=1)	0.580	0.544	0.036	0.712	0.667	0.045∗	0.805	0.774	0.031†
Number of children	0.709	0.670	0.038	0.656	0.723	-0.068∗	0.689	0.818	-0.129∗
Local unemployment rates	0.074	0.076	-0.002	0.070	0.070	-0.000	0.070	0.072	-0.002‡
Year	1999.937	1999.383	0.554‡	2000.746	1999.983	0.763∗	2000.585	1999.823	0.761∗
Mining/Utility	0.017	0.008	0.009	0.023	0.011	0.012∗	0.025	0.013	0.012†
Manufacturing	0.283	0.184	0.099∗	0.285	0.203	0.082∗	0.303	0.239	0.064∗
Construction	0.075	0.046	0.029‡	0.074	0.060	0.013‡	0.068	0.082	-0.014
Wholesale/Accomodation	0.205	0.245	-0.040	0.150	0.194	-0.044∗	0.149	0.141	0.007
Trans/Comm	0.074	0.061	0.012	0.083	0.066	0.017†	0.110	0.067	0.043∗
Finance/Real estate	0.133	0.088	0.045†	0.157	0.122	0.035∗	0.156	0.119	0.037∗
Public/Education/Health	0.133	0.253	-0.120∗	0.166	0.253	-0.087∗	0.136	0.261	-0.125∗
Other not-for-profit	0.068	0.103	-0.036†	0.049	0.064	-0.014†	0.044	0.066	-0.022∗
Professional	0.072	0.084	-0.013	0.095	0.118	-0.024∗	0.106	0.134	-0.028†
Associate professional & technical	0.094	0.107	-0.013	0.117	0.107	0.010	0.102	0.121	-0.019‡
Clerical & secretarial	0.129	0.126	0.003	0.109	0.100	0.009	0.077	0.071	0.007
Craft	0.206	0.153	0.053†	0.184	0.185	-0.000	0.176	0.167	0.010
Personal & protective service	0.110	0.172	-0.062∗	0.070	0.099	-0.030∗	0.047	0.073	-0.026∗
Sales	0.058	0.088	-0.030‡	0.047	0.055	-0.008	0.042	0.044	-0.002
Machine operatives	0.151	0.111	0.040‡	0.144	0.120	0.024†	0.146	0.143	0.004
Others	0.108	0.111	-0.003	0.071	0.097	-0.026∗	0.061	0.049	0.012
Number of individuals	2482	261		22244	1511		13666	792	

Notes: Based on British Household Panel Survey (1992-2008). All income variables are gross figures with prices in 2008. ∗, † and ‡ indicate significance at 1, 5 and 10 percent levels respectively.

Table 2: Estimated Ordered Probit Model of Constrained Labor Supply Regimes

	Coef.	S.E.	$\partial P(C=0)/\partial x$	S.E.	$\partial P(C=1)/\partial x$	S.E.	$\partial P(C=2)/\partial x$	S.E.
Wage rate job 1	-0.002	(0.002)	0.000	(0.000)	0.000	(0.000)	-0.001	(0.000)
Weekly non-labour income/100	-0.195∗	(0.068)	0.019∗	(0.007)	0.037∗	(0.013)	-0.056∗	(0.019)
Age	0.021∗	(0.001)	-0.002∗	(0.000)	-0.004∗	(0.000)	0.006∗	(0.000)
1st degree	0.099	(0.066)	-0.010	(0.006)	-0.019	(0.012)	0.028	(0.019)
HND, HNC, teaching	0.047	(0.073)	-0.005	(0.007)	-0.009	(0.014)	0.013	(0.021)
A level	-0.044	(0.066)	0.004	(0.006)	0.008	(0.012)	-0.013	(0.019)
O level	-0.045	(0.067)	0.004	(0.007)	0.008	(0.013)	-0.013	(0.019)
CSE	-0.050	(0.080)	0.005	(0.008)	0.009	(0.015)	-0.014	(0.023)
None of these	-0.179†	(0.071)	0.017†	(0.007)	0.034†	(0.013)	-0.051†	(0.020)
Married	0.221∗	(0.023)	-0.022∗	(0.002)	-0.042∗	(0.004)	0.063∗	(0.007)
Number of children	-0.015	(0.010)	0.001	(0.001)	0.003	(0.002)	-0.004	(0.003)
Local unemployment rates	-2.244∗	(0.437)	0.219∗	(0.043)	0.423∗	(0.083)	-0.642∗	(0.125)
Year	-0.008∗	(0.002)	0.001∗	(0.000)	0.002∗	(0.000)	-0.002∗	(0.001)
Mining/Utility	0.121	(0.099)	-0.012	(0.010)	-0.023	(0.019)	0.035	(0.028)
Manufacturing	0.117	(0.079)	-0.011	(0.008)	-0.022	(0.015)	0.033	(0.023)
Construction	0.104	(0.084)	-0.010	(0.008)	-0.020	(0.016)	0.030	(0.024)
Wholesale/Accomodation	0.052	(0.080)	-0.005	(0.008)	-0.010	(0.015)	0.015	(0.023)
Trans/Comm	0.263∗	(0.083)	-0.026∗	(0.008)	-0.050∗	(0.016)	0.075∗	(0.024)
Finance/Real estate	0.098	(0.081)	-0.010	(0.008)	-0.018	(0.015)	0.028	(0.023)
Public/Education/Health	-0.073	(0.082)	0.007	(0.008)	0.014	(0.015)	-0.021	(0.023)
Other not-for-profit	0.010	(0.084)	-0.001	(0.008)	-0.002	(0.016)	0.003	(0.024)
Professional	-0.202∗	(0.035)	0.020∗	(0.003)	0.038∗	(0.007)	-0.058∗	(0.010)
Associate professional & technical	-0.297∗	(0.032)	0.029∗	(0.003)	0.056∗	(0.006)	-0.085∗	(0.009)
Clerical & secretarial	-0.391∗	(0.034)	0.038∗	(0.003)	0.074∗	(0.006)	-0.112∗	(0.010)
Craft	-0.261∗	(0.032)	0.026∗	(0.003)	0.049∗	(0.006)	-0.075∗	(0.009)
Personal & protective service	-0.465∗	(0.043)	0.045∗	(0.004)	0.088∗	(0.008)	-0.133∗	(0.012)
Sales	-0.305∗	(0.042)	0.030∗	(0.004)	0.057∗	(0.008)	-0.087∗	(0.012)
Machine operatives	-0.349∗	(0.034)	0.034∗	(0.003)	0.066∗	(0.006)	-0.100∗	(0.010)
Others	-0.459∗	(0.039)	0.045∗	(0.004)	0.087∗	(0.008)	-0.131∗	(0.011)
κ_1	-17.697∗	(4.989)						
κ_2	-15.226∗	(4.989)						
σ_u^2	0.660∗	(0.021)						
Log pseudo-likelihood	-3.1e+04							
N	40956							

Notes: Based on British Household Panel Survey (1992-2008). ∗, † and ‡ indicate significance at 1, 5 and 10 percent levels respectively. Robust standard errors are in parentheses. Marginal effects are estimated at the means of covariates.

Table 3: Dual Job Choice Regressions

	Underemployed		**Unconstrained**		**Overemployed**	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Wage rate job 1	-0.007∗	(0.002)	-0.001‡	(0.001)	-0.000	(0.001)
Weekly non-labour income/100	-0.048	(0.070)	-0.002	(0.024)	-0.018	(0.019)
Age	0.004	(0.019)	-0.004	(0.005)	0.012†	(0.005)
Married	-0.042	(0.029)	-0.008	(0.009)	-0.013	(0.012)
Number of children	0.027†	(0.013)	0.003	(0.004)	0.005	(0.004)
Local unemployment rates	1.003	(0.995)	0.119	(0.210)	0.168	(0.245)
Year 1993	0.048	(0.247)	-0.004	(0.020)	-0.108‡	(0.061)
Year 1994	-0.148	(0.228)	0.024	(0.022)	0.040	(0.051)
Year 1995	-0.025	(0.220)	0.010	(0.025)	0.073	(0.058)
Year 1996	0.380	(0.274)	0.013	(0.029)	0.020	(0.056)
Year 1997	0.250	(0.262)	0.036	(0.033)	-0.020	(0.064)
Year 1998	0.075	(0.280)	0.014	(0.037)	-0.033	(0.068)
Year 1999	0.104	(0.273)	0.022	(0.043)	-0.014	(0.070)
Year 2000	-0.000	(0.286)	0.039	(0.047)	-0.090	(0.073)
Year 2001	0.103	(0.288)	0.021	(0.046)	-0.036	(0.068)
Year 2002	0.094	(0.321)	0.051	(0.057)	-0.083	(0.069)
Year 2003	0.005	(0.320)	0.014	(0.060)	-0.119	(0.076)
Year 2004	0.011	(0.329)	0.034	(0.061)	-0.138‡	(0.079)
Year 2005	-0.179	(0.333)	0.042	(0.066)	-0.083	(0.082)
Year 2006	0.065	(0.354)	0.038	(0.068)	-0.153‡	(0.084)
Year 2007	-0.279	(0.358)	0.046	(0.075)	-0.140‡	(0.084)
Year 2008	0.033	(0.390)	0.033	(0.079)	-0.118	(0.097)
Wage rate job 1 (m)	0.004	(0.003)	-0.001	(0.001)	-0.002	(0.001)
Weekly non-labour income/100 (m)	0.248‡	(0.135)	0.008	(0.033)	0.036	(0.042)
Age (m)	-0.004	(0.019)	0.004	(0.005)	-0.013†	(0.005)
1st degree (m)	-0.018	(0.067)	-0.016	(0.017)	-0.046†	(0.022)
HND, HNC, teaching (m)	-0.035	(0.067)	-0.018	(0.016)	-0.044‡	(0.024)
A level (m)	-0.002	(0.064)	-0.009	(0.017)	-0.030	(0.023)
O level (m)	-0.018	(0.063)	-0.003	(0.018)	-0.023	(0.024)
CSE (m)	-0.032	(0.066)	0.004	(0.022)	-0.023	(0.027)
None of these (m)	-0.055	(0.067)	-0.019	(0.018)	-0.036	(0.024)
Married (m)	0.079†	(0.035)	0.002	(0.013)	0.011	(0.015)
Number of children (m)	-0.045∗	(0.016)	0.004	(0.005)	0.004	(0.006)
Local unemployment rates (m)	-1.572	(1.131)	-0.753†	(0.305)	-0.268	(0.326)
Mining/Utility (m)	-0.148‡	(0.087)	-0.062	(0.055)	-0.014	(0.036)
Manufacturing (m)	-0.110	(0.081)	-0.064	(0.052)	-0.010	(0.024)
Construction (m)	-0.095	(0.090)	-0.060	(0.053)	0.029	(0.026)
Wholesale/Accomodation (m)	-0.063	(0.083)	-0.029	(0.054)	0.004	(0.027)
Trans/Comm (m)	-0.098	(0.086)	-0.051	(0.053)	-0.018	(0.025)
Finance/Real estate (m)	-0.114	(0.084)	-0.047	(0.054)	-0.009	(0.026)
Public/Education/Health (m)	0.008	(0.085)	0.005	(0.054)	0.059†	(0.027)
Other not-for-profit (m)	0.035	(0.087)	-0.024	(0.056)	0.045	(0.036)
Professional (m)	0.037	(0.048)	0.022	(0.016)	0.019	(0.019)

Continued on next page

Table 3 – *Continued from previous page*

	Underemployed		**Unconstrained**		**Overemployed**	
	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Associate professional & technical (m)	0.050	(0.055)	0.007	(0.014)	0.025‡	(0.015)
Clerical & secretarial (m)	0.019	(0.044)	-0.028†	(0.013)	-0.017	(0.016)
Craft (m)	0.016	(0.036)	0.012	(0.013)	-0.009	(0.014)
Personal & protective service (m)	-0.003	(0.044)	-0.009	(0.018)	-0.012	(0.024)
Sales (m)	0.041	(0.050)	0.015	(0.021)	0.006	(0.028)
Machine operatives (m)	0.016	(0.040)	0.004	(0.014)	0.002	(0.017)
Others (m)	0.018	(0.051)	0.021	(0.018)	-0.016	(0.017)
Year 1993 (m)	0.037	(0.088)	0.041	(0.037)	0.094	(0.059)
Year 1994 (m)	-0.024	(0.084)	0.008	(0.039)	0.063	(0.061)
Year 1995 (m)	-0.034	(0.131)	-0.048	(0.048)	0.051	(0.059)
Year 1996 (m)	0.305‡	(0.166)	0.001	(0.050)	0.099	(0.063)
Year 1997 (m)	-0.149	(0.148)	-0.077	(0.050)	0.016	(0.063)
Year 1998 (m)	0.072	(0.151)	-0.036	(0.051)	0.107‡	(0.058)
Year 1999 (m)	-0.047	(0.183)	-0.018	(0.046)	0.133†	(0.055)
Year 2000 (m)	0.013	(0.177)	-0.056	(0.048)	0.095	(0.060)
Year 2001 (m)	-0.089	(0.207)	-0.058	(0.053)	0.140†	(0.058)
Year 2002 (m)	-0.003	(0.222)	-0.059	(0.058)	0.175†	(0.070)
Year 2003 (m)	-0.002	(0.250)	-0.092	(0.071)	0.079	(0.077)
Year 2004 (m)	-0.080	(0.277)	-0.082	(0.071)	0.121	(0.078)
Year 2005 (m)	0.085	(0.273)	-0.097	(0.073)	0.212†	(0.085)
Year 2006 (m)	-0.040	(0.304)	-0.060	(0.080)	0.228†	(0.094)
Year 2007 (m)	-0.044	(0.307)	-0.082	(0.085)	0.149	(0.100)
Year 2008 (m)	-0.037	(0.328)	-0.129	(0.082)	0.246∗	(0.091)
Inverse Mills Ratio	-0.011	(0.102)	0.076	(0.079)	-0.051	(0.045)
IMR × 1993	-0.020	(0.136)	-0.054	(0.118)	-0.083	(0.060)
IMR × 1994	0.095	(0.125)	0.014	(0.115)	0.071	(0.049)
IMR × 1995	0.028	(0.121)	-0.091	(0.115)	0.111†	(0.056)
IMR × 1996	-0.173	(0.140)	-0.151	(0.119)	0.078	(0.053)
IMR × 1997	-0.103	(0.129)	0.010	(0.119)	0.022	(0.058)
IMR × 1998	-0.040	(0.137)	-0.118	(0.125)	0.014	(0.063)
IMR × 1999	-0.024	(0.123)	-0.112	(0.137)	0.072	(0.061)
IMR × 2000	0.019	(0.134)	-0.061	(0.142)	-0.003	(0.061)
IMR × 2001	-0.032	(0.118)	-0.076	(0.096)	0.078	(0.053)
IMR × 2002	-0.033	(0.127)	0.019	(0.130)	0.046	(0.052)
IMR × 2003	0.030	(0.122)	-0.148	(0.118)	0.017	(0.057)
IMR × 2004	0.027	(0.117)	-0.102	(0.102)	0.012	(0.053)
IMR × 2005	0.081	(0.119)	-0.028	(0.103)	0.084	(0.054)
IMR × 2006	-0.026	(0.118)	-0.117	(0.101)	0.046	(0.052)
IMR × 2007	0.164	(0.116)	-0.059	(0.096)	0.056	(0.049)
IMR × 2008	0.002	(0.124)	-0.030	(0.107)	0.097‡	(0.053)

Notes: Based on British Household Panel Survey (1992-2008). ∗, † and ‡ indicate significance at 1, 5 and 10 percent levels respectively. Standard errors in parentheses are bootstrap estimates from 200 replications. IMR×Year indicates the interactions between lambda terms and year dummies. (m) indicates time averaged variables.