Economics of Transition
Volume 11 (1) 2003, 93–122

Employees and second-job holding in the Federal Republic of Yugoslavia

An empirical analysis[1]

Barry Reilly* and Gorana Krstić**
*Department of Economics, University of Sussex, Falmer, Brighton BN1 9QN, UK.
E-mail: b.m.reilly@sussex.ac.uk
**Economics Institute, Kralja Milana 16, 11000 Belgrade, FR Yugoslavia.

Abstract

This paper explores the second-job holding (or 'moonlighting') behaviour of a sample of employees using data from a unique survey conducted for the Federal Republic of Yugoslavia (FRY) in January 1998. Both participation in 'moonlighting' and the number of hours worked are examined. The participation model performs better in an econometric sense and provides the focus for our discussion. We note a strong regional dimension to 'moonlighting' in the FRY with employees in Central Serbia disproportionately represented in this activity. In addition, blue-collar workers are found to be more likely to engage in 'moonlighting' than white-collar workers. The set of labour supply variables implied by neo-classical theory exerts a strong influence and explains a significant amount of the phenomenon of interest. Our calculations suggest that if main (or regular) job earnings are restored to levels that prevailed at the time of the 'break-up' of the federation, employee second-job holding in the FRY would only fall by about one-seventh.

JEL classification: J21, J22, P2, P3.
Keywords: Second-job holding, 'moonlighting', employees, wages, Yugoslavia.

[1] The authors would like to thank two anonymous referees and a co-editor of this journal for constructive comments on an earlier draft of this paper. The comments provided by Mike Sumner are also appreciated. In all cases, the usual disclaimer applies.

Published by Blackwell Publishing, 9600 Garsington Road, Oxford OX4 2DQ, UK and 350 Main Street, Malden, MA 02148, USA.

1. Introduction

It is generally accepted that of all the erstwhile centrally-planned economies, the Socialist Federal Republic of Yugoslavia was best positioned to make a rapid transition to a market economy once political conditions permitted.[2] The secession of Slovenia, Croatia, and Macedonia from the Federation in the early 1990s and the subsequent war in Bosnia-Herzegovenia, however, precipitated the disintegration of socialist Yugoslavia. The economic situation was exacerbated by the imposition of sanctions by the United Nations Security Council on the remaining Yugoslav republics of Serbia and Montenegro in early 1992. By the mid-1990s per capita GDP fell back to a level that last prevailed in Yugoslavia in 1969.

Open unemployment and under-employment rose rapidly with the contraction of GDP. An increase in welfare payments, including generous pension provision, led to a sharp rise in the fiscal deficit (see Petrović, Bogetić, and Vujošević, 1999). An attempt by the authorities to monetize the deficit led to a protracted hyperinflation.[3] This resulted in the collapse of the country's monetary and fiscal systems, and precipitated a dramatic flight from the national currency as a means of exchange. Social and economic conditions were also exacerbated by the influx of nearly 700,000 displaced persons who chose or were forced to relocate to either Serbia or Montenegro. A stabilization programme, the Monetary Reconstruction Programme, was introduced in January 1994 to fight inflation and restore confidence in the national currency. The reforms also contained measures designed to liberalize prices and wages. The programme enjoyed modest success in that it reduced inflation and created a stable economic environment within which economic agents could operate.

The collapse in economic activity, despite the limited success of the reforms, led to an increase in the impoverishment of the Yugoslav population. Pošarac (1998) analyzed the evolution of poverty incidence in Yugoslavia over the period 1990 to 1995. Using data from the Surveys on Household Consumption conducted by the Federal Statistical Office in Belgrade, the analysis revealed that by 1995 almost three million people (just under a third of the population) were estimated to be below a recognized poverty threshold. This represented a doubling of the incidence of poverty relative to 1990 when well over one million individuals were estimated to be in poverty.

The rise in both open unemployment and under-employment provides part of the explanation for the observed rise in poverty. However, the contraction in real

[2] In contrast to most centrally-planned economies, the Yugoslav 'self-management' model relied more on market forces with worker incomes linked to enterprise performance. In addition, there was a greater degree of liberalization in both foreign trade and market prices, and agriculture was dominated by private smallholders who responded to market price signals.

[3] Petrović, Bogetić and Vujošević (1999) provide an econometric analysis of the Serbian hyperinflationary episode.

incomes – primarily labour market incomes – explains much of the poverty increase in Yugoslavia through the 1990s. The relative stability of the estimated Gini coefficients over much of the time period suggests that the drop in income is more important than its dispersion in explaining the increased incidence of poverty (see Human Development Report, FRY (1997) and Bogicević, Krstić and Mijatović (2001).

The reduction in income generating opportunities in the formal sector of the Yugoslav economy encouraged informal activity and provided for many an important safety net. Although informal or hidden activity was common in Yugoslavia and other formerly socialist economies (e.g., see Nove, 1987, pp. 213–14 and Gaddy, 1991), the confluence of political and economic events created conditions within which it flourished. Declining incomes created incentives for participation. The poor enforcement of the legal and taxation system and the mild penalties applied to participation also encouraged this activity.

A large body of the existing research on informal activity for the transitional economies has focused on calibrating the size of the unmeasured sector relative to its measured counterpart (e.g., see Johnson, Kaufmann and Schliefer, 1997; EBRD, 1997; Božović, 1993; Krstić, 1998; Lacko, 2000).[4] The definition and measurement of informal (or hidden) activity in transitional economies is clearly problematic but some innovative methods have been used to quantify its size (e.g., see Johnson, Kaufmann and Schliefer, 1997). The estimation of the size of the hidden economy in the Federal Republic of Yugoslavia (FRY), for instance, was the subject of research by Krstić (1998) who, using an aggregate labour market approach (see Contini, 1992), estimated informal activity as approximately equivalent to a third of measured GDP in 1997.

The EBRD's Transition Report (2000) argues that although informal activities have played a critical role in providing employment and earnings opportunities for many individuals in transitional economies, the motives for participation differ depending on the country's stage of reform. In economies at a mature stage of reform (e.g., central and eastern European countries), the motives tend to be more 'market-related' and guided by a desire to evade tax and avoid other bureaucratic constraints. This motive could be taken to reflect individuals responding to informal labour market opportunities. In economies at an incipient stage of reform (e.g., CIS and southern European countries), informal activity is interpreted as being driven by poor opportunities in the formal sector and is seen as providing a 'coping strategy' for survival. For those in formal employment, the 'coping strategy' is achieved through second-job holding. Although this dichotomy provides a useful classification, it is likely that both motives co-exist in a given transitional country regardless of the stage of the reform process.

The examination of informal activity is clearly not new but there is a limited empirical literature available for transitional economies at the level of the individual.

[4] Unmeasured economic activity can be taken to include a range of different activities. See Thomas (1992) for a discussion of terms and definitions.

This is mainly attributable to the absence of worthwhile data that can be usefully exploited to inform this issue.[5] This paper uses a relatively recent dataset to explore and model the 'moonlighting' or second–job holding behaviour of employees in the FRY. The labour market activity of this former socialist economy has attracted only modest empirical research since the era of reforms commenced – a fact attributable in large part to the political events that have marked the country since the early 1990s.

As noted above, the analysis of informal activity at the level of the labour force participant has been limited. Foley (1997), Kolev (1998) and Guariglia and Kim (2000), however, provide useful exceptions with their analyses of multiple-job holding in Russia using various rounds of the Russian Longitudinal Monitoring Survey (RLMS). Foley emphasized the role of wage arrears in determining participation and Kolev concluded that multiple job-holding provided an important safety valve for individuals rationed in the formal labour market. Guariglia and Kim found that 'moonlighting' behaviour was transitory in nature and generally associated with career shifts. In all three studies about one-tenth of RLMS respondents indicated involvement in informal activity and this could be viewed as an under-estimate.

The primary purpose of this paper is to ascertain the key determinants of second-job holding for a sample of employees using a unique survey on informal activity undertaken in the FRY in early 1998. It is important to make a distinction between informal job-holding and second-job holding. The former concept incorporates the activity of the employed, unemployed, retired, and other non-participants in the formal labour market. In that sense it conflates a diverse set of individuals into a single category. The use of such a broad categorization may not be all that helpful to an understanding of key aspects of transitional labour markets. In our approach we ignore the informal activity of the unemployed and non-active members of the labour force. We justify these exclusions on the grounds that the majority of labour market participants in the FRY over the late 1990s were employees (over 70 percent) and that exploring all other diverse aspects of informal activity demands a more complex theoretical and econometric approach than the one adopted here. In addition, our emphasis on employee behaviour facilitates a more direct comparison with the small volume of existing literature on multiple-job holding for Russia.

The structure of the paper can now be outlined. The next section briefly adumbrates the theoretical framework within which our empirical analysis is couched, followed by a section detailing the econometric methodology used, and a section describing the data. The penultimate section contains a discussion of the empirical

[5] For instance, estimates of the multiple-job holding rates derived from the Russian Longitudinal Monitoring Surveys (RLMS) (e.g., Kolev, 1998 or country Labour Force Surveys (e.g., Krstić, 1999 or EBRD, 2000) generally provide multiple-job holding rates that appear implausibly low and close to those obtained for more established capitalist economies (see Bell, Robert and Wright, 1997).

results and is followed by a concluding section. Throughout this paper, we use the terms 'moonlighting' and second-job holding interchangeably.

2. Theoretical issues

Our analysis is situated entirely within a static labour supply framework. We assume that the labour market is segmented into formal and informal sectors. Following Kolev (1998) we assume individuals make their formal and 'moonlighting' labour supply choices sequentially such that formal labour market characteristics are exogenous to the 'moonlighting' decision. As correctly noted by Foley (1997) for an economy in transition, an individual's job is likely to have been determined under a centrally-planned regime. The desired volume of 'moonlighting' hours worked can be solved through maximizing an individual's utility function subject to a budget constraint. This approach yields a 'moonlighting' labour supply function of the following general form:

$$H_s = f(W_s, Y_m, T - H_m, \mathbf{Z}) \tag{1}$$

where H_s = desired 'moonlighting' hours worked; W_s = the hourly 'moonlighting' wage rate; Y_m = total wages on the main (or primary) job; H_m = hours worked on the main (or primary) job; $T - H_m$ = the time available for 'moonlighting'; $\mathbf{Z}$ = a set of individual and household level characteristics taken to include non-labour market income, household structure, and other variables. In the subsequent empirical analysis lower case lettering will be used to denote the natural logarithms of these variables.

This simple neo-classical approach to 'moonlighting' labour market participation provides the theoretical framework for our empirical analysis. In this study we measure 'moonlighting' activity through both participation and the number of 'moonlighting' hours worked. There are data constraints that prevent us introducing some preferred empirical constructs for the set of variables hypothesized in $\mathbf{Z}$ and these limitations are discussed in the data section.

The predictions of this neo-classical model can now be outlined. Assuming formal and 'moonlighting' activities are sequentially decided and that non-market time is a normal good, an increase in main job earnings reduces 'moonlighting' activity through its effect on income. An increase in the number of hours worked on the main job reduces 'moonlighting' activity through the time constraint effect. The effect of the 'moonlighting' wage rate (i.e., the own wage effect) on 'moonlighting' activity is ambiguous and dependent on the relative strength of substitution and income effects.

It might be tempting to use the relative magnitude of the own wage effect and the main job earnings and other non-labour market income effects to discriminate between 'market related' and 'coping strategy' motives for participation. It is certainly the case, on the assumption that non-market time is a normal good, that negative income effects (broadly defined) are consistent with a 'coping strategy' motive. A positive own wage effect while compatible with a 'market related' motive is also consistent with a 'coping strategy'. For instance, an individual may be obliged to dispose of assets in order to 'make ends meet'. The availability of informal labour market opportunities may allow such disposals to be deferred. Under such circumstances individuals attempting to 'make ends meet' are likely to respond to informal labour market opportunities as mediated through 'moonlighting' wage offers in order to protect their stock of assets. Thus, a positive own wage effect could be construed as consistent with either an opportunities or a 'coping strategy' motive.

3. Econometric methodology

Our approach is to specify structural models for both the incidence of 'moonlighting' (i.e., participation) and its extent (i.e., the number of hours worked). A standard binomial probit is used to model the former (see Greene 1997), and a censored tobit is employed for the latter (see Amemiya, 1984).[6] Although we have observations on main job earnings and hours worked for all employees (see below), we only observe 'moonlighting' wages for those actually 'moonlighting'. Since these may represent a non-random sample of employees, the Heckman (1979) two-step procedure is used to correct hourly 'moonlighting' wages for potential selectivity bias. The unbiased estimates thus obtained are then applied to an individual's realizations for a set of explanatory variables to generate a predicted (or expected) wage offer for each individual. A key consideration in using the Heckman procedure is the identification of the selection effect. This is achieved through using variables that shift the probability of 'moonlighting' but not the level of hourly 'moonlighting' pay. The identification of the selection effect is discussed in the empirical section.

The maximum likelihood procedure assumes an exact knowledge of the probability distribution function employed up to a set of unknown parameters (i.e., the coefficient vector). The estimation procedure is thus highly sensitive to departures from the specification of the likelihood function (e.g., the assumption of normality, which is made for both models). It is of some importance to evaluate the structural models in regard to key econometric assumptions. The adequacy of the estimated models is assessed using efficient score tests of the type originally suggested in Chesher and Irish (1987) and are of the form $\mathbf{i'R(R'R)^{-1}R'i}$. In this case, $\mathbf{i}$ is an $n \times 1$ vector of ones, and $\mathbf{R}$ is an $n \times q$ matrix of score contributions computed for each of the k parameters from the original specification and the $k + 1, \ldots q$ parameters

[6] The use of the tobit is reasonably well established in the 'moonlighting' literature. For instance, see Shishko and Rostker (1976) for a relatively early application.

that capture the form of the alternative hypothesis, assumed zero under the null. The resultant test statistics are all distributed as chi-squared with $p = q - k$ degrees of freedom. The test represents the outer-product gradient (OPG) form of the score (or Lagrange Multiplier) test.

The efficient score tests undertaken are designed to assess the adequacy of the reported specifications in terms of functional form (RESET), omitted variables, homoscedastic errors, and normally distributed residuals. The assumed exogeneity of monthly earnings and hours worked in the main job is also evaluated within this testing framework. The failure of either structural model in terms of the diagnostic tests effectively vitiates any inferences that can be drawn from the parameter estimates regarding, *inter alia*, the predictions of neo-classical labour supply theory.

Orme (1990) has questioned the use of OPG-based tests and, using Monte Carlo simulations, demonstrated their poor finite sample properties. Orme's findings suggest that efficient score tests constructed using the OPG covariance matrix tended to reject the correct null hypothesis far too frequently. The poor performance of score tests in relatively small samples is readily acknowledged but analysis of our testing principle[7] suggests that some degree of confidence in the finite sample properties of the efficient score tests is justified.[8]

4. Data

Our empirical analysis uses information obtained from a unique individual-level survey designed to elicit information on informal sector activity in the FRY. The survey was conducted by the Economic Institute in Belgrade in January 1998 and relates to labour market activity, both formal and informal, undertaken in the previous calendar month. The survey provided, amongst other things, information on the labour market status of individuals and on any informal market activity undertaken. The data are based on responses to individual-level questionnaires drawn from randomly selected households across the republics of Serbia and Montenegro.

---

[7] We computed Wald tests based on the Hessian form of Fisher's information matrix for the set of parameter restrictions relating to omitted variables and RESET tests in our estimated models and compared them to the efficient score tests computed using the OPG covariance matrix reported in Table 1. We would anticipate relatively large differences between score and Wald test values, if the OPG covariance matrix provides a poor approximation to the Hessian. The scale of the differences encountered was not found to be consistent with the type of size bias suggested by the Monte Carlo simulations. It should also be borne in mind that the simulations reported in Orme (1990) are based on testing the Fisher information equality condition.

[8] Even if we take the size bias from the simulation studies at face value, it is clear we are setting our empirical models fairly stern tests in requiring them to satisfy the set of econometric assumptions. In particular, the direction of the bias would not alter an efficient score test inference that indicated non-rejection of a particular null hypothesis.

The survey obtained information on:

a) individual level characteristics including age, gender, marital status and educational attainment;
b) the labour market status of respondents including their occupation and skill level, establishment size, the industrial activity and ownership of their primary place of work, the hours worked and monthly earnings in the main job and in any secondary job held, and the opportunities for 'moonlighting' on the main job; and
c) household-level characteristics including size and composition of the household, household type and the settlement type within which it is located.

In addition to the conventional biases associated with obtaining complete and correct responses from a sample survey, an informal sector survey encounters difficulties arising from the reluctance of respondents to reveal informal activities to interviewers unknown to them. In order to attenuate the effect of this systematic bias on survey responses, the approach of interviewers was modified to minimize respondent bias. The confidential nature of the survey was emphasized and respondents were assured that only anonymized information would be used for research purposes. The sequence and wording of the questions on the questionnaire were adjusted to elicit as honest a response as possible under the circumstances. Individuals were initially asked questions on the sufficiency of their monthly income for basic needs and their standard of living. This was followed by questions on opportunities to earn some extra money and other information relating to informal activity.

The focus of this paper is the second-job holding (or 'moonlighting') behaviour of the sub-sample of employees.[9] Krstić (1999) explores in detail the relationship between the labour market outcomes of this survey and those associated with other data sources (including the Yugoslav Labour Force Surveys (YLFS) of 1995–98) and concludes that there are significant differences in regard to second-job holding. According to the YLFS for 1997, only nine percent of employees are reported as second-job holders – under a third of the rate reported for the survey used here. This may confirm the perception that individuals are reluctant to provide responses on their informal activity to an official agency.

One part of the population omitted from all official surveys in Yugoslavia in recent years are the Kosovar Albanians who, on the basis of the 1991 census estimates, comprised over 15 percent of the Yugoslav population. This ethnic group does not feature in this informal sector survey. The Kosovar responses are thus based entirely on Serbian respondents in Kosovo and Metohija.

The survey generated usable information on 1302 employed individuals. In contrast to the empirical work undertaken on multiple job holding for Russia our

[9] The analysis reported here concentrates on those employees with only one secondary job as only three percent of employees reported holding more than one secondary job in this survey. In addition, employees on unpaid leave are excluded.

data contains minimal information on non-labour market income and no information on wage arrears. In addition, the data on household structure are also limited with information relating to the number and age of dependent children or adults, or the number of other wage earners in the household unavailable. Table A1 of the appendix provides summary statistics for the sample of employees as a whole, and for sub-samples of employees with and without second jobs.

About 30 percent of all employees in our sample are engaged in 'moonlighting' activity. Female employees are less likely to 'moonlight' than their male counterparts. In terms of the raw unadjusted data, women are about 11 percentage points less likely to engage in such activity. Workers in the construction industry and those whose main jobs are blue-collar in nature are more likely to engage in second-job holding, but those with a high-school qualification are less likely participants, as are those in the senior professional occupational category. Employees residing in Serbia appear more likely to hold a second job than those in either Montenegro or Vojvodina. Those that do not hold a second job, on average, earn more and work longer hours in their main job than do those engaged in second-job activity.

Table A2 in the appendix provides more detailed information on the distribution of earnings in the main and secondary jobs. The dispersion in main job earnings in our data as measured by the Gini coefficient is comparable to estimates obtained using the YLFS data from 1997 (see Reilly and Krstić, 2002). The estimated Gini coefficients for hourly wage rates in the second job indicate considerably higher variation in this sector. As Table A2 indicates the hourly wage rate for the 'moonlighting' job in our study is, on average, about three times the hourly rate for the main job, which is comparable to findings for Russia (see Kolev, 1998 and Guariglia and Kim, 2000).

Over 85 percent of secondary job holders secure hourly wage rates that are greater than the comparable rates obtained on their main job. The large number of employees commanding higher wage rates in the second job indicates there are non-wage incentives for 'moonlighting' employees to maintain links with the main job. For instance, the main (or regular) job might be associated with a variety of non-pecuniary benefits (e.g., medical support and pension rights) that enhance the overall compensation for these jobs but might also provide access to equipment and materials that are useful in moonlighting activities. In an era of economic uncertainty, it is unsurprising that employees wish to maintain such links. However, the finding might also be consistent with the existence of an hours' constraint operating in the second job.

5. Empirical results

Table A3 of the appendix reports reduced form probit estimates for the 'moonlighting' participation model. The estimates are used to compute the inverse Mill's ratio term to correct the 'moonlighting' logarithmic hourly wage equation for potential selectivity bias. The identification of the selection parameter is achieved through including variables in the criterion equation that shift the probability of 'moonlighting' but not the level of hourly 'moonlighting' pay. A set of variables, including a number interacted with gender, is introduced to perform this role and, as indicated in the final row of Table A3, appears adequate for the task. The estimated selection coefficient itself provides guidance as to the presence of potential selection effects. The null hypothesis that the 'moonlighting' sample of employees is a random one is rejected by the data at a conventional level thus supporting use of the Heckman procedure. The predicted logarithmic hourly wage offer for the structural models is thus calculated using the selectivity corrected estimates and is defined $\hat{w}_s$ for future reference.[10]

We now examine the structural models for the 'moonlighting' decision and hours worked reported in Table 1.[11] We initially focus attention on the array of diagnostics computed for both estimated models. The probit model performs well in terms of all the reported diagnostics and the computed residuals violate only one assumption, that of a constant variance. This is not viewed as problematic since the model's estimated variance-covariance matrix is corrected for the presence of heteroscedasticity of unknown form.[12] As noted earlier, although we assume the main job attributes to be exogenous, this is a testable proposition. The tests suggest that the null hypotheses of exogeneity for both monthly earnings and hours worked on the main job are upheld by the data in the probit model. The instrument set is also found to comprise a valid set. In addition, the omission from the 'moonlighting' equation of selected sub-sets of variables relating to demographic and other main job characteristics is also upheld by the data. The percentage of correct predictions is satisfactory, and the Pseudo-R^2 and a complementary measure suggested by Cramer (1999) indicate a reasonable fit for a cross-sectional model.

In marked contrast the tobit model fares poorly in terms of the reported diagnostics. In particular, the model fails both the homoscedasticity and normality assumptions. This implies that the maximum likelihood estimation of the tobit model is inconsistent. The model also fails the RESET test and an additional specification test for the censored tobit form. On the basis of these diagnostic test results the interpretational emphasis is subsequently placed on the participation rather than the hours model.

The preferred and most data coherent model includes the standard labour supply variables implied by conventional theory, a set of age controls, a gender control,

[10] The average predicted hourly wage for the sub-sample of 'moonlighters' is approximately 10 percent higher than that of non-participants (see Table A1 of the appendix).

[11] It is worth noting that the structural model estimates are not materially altered if uncorrected rather than corrected OLS estimates are used to construct the predicted hourly wage measure.

[12] The relevant variance-covariance formula is based on Huber (1967). However, there is little difference between the adjusted and unadjusted estimated variance-covariance matrix in this case. None of the inferences offered in this paper are sensitive to the adjustment procedure used to correct for heteroscedasticity.

regional controls, a set of skill level variables as specified on the main job, a control for private sector ownership on the main job, and a set of household structure variables. A number of interaction terms were also found to merit inclusion in the empirical model.

The presence of interaction terms warrants a careful interpretation of a number of the estimated coefficients including those relating to gender and age. On average, female employees aged over 50 are, *ceteris paribus*, some 15 percentage points less likely to 'moonlight' compared to males in the same age group. Women aged between 25 and 30 are estimated to 'moonlight' more than men in that age group but the computed *t*-ratio for the point estimate is only 1.3.[13] The highest incidence of participation among male employees is in the 31 to 40 year age-category. Being present in this particular group increases the standardized probit index, *ceteris paribus*, by about one-half of a standard deviation (or 16 percentage points) relative to the base.

The estimate for the Montenegro regional control appears relatively large for a probit coefficient but an average effect must be computed using the mean value for the logarithmic monthly earnings measure – the variable with which it is

Table 1. Structural models – participation and hours worked for Yugoslav employees

	Estimated Coefficients	
Variable	**Probit**	**Censored Tobit**
Constant	1.309*	219.344***
	(0.890)	(62.807)
Age ≤ 25	–0.255	–30.122
	(0.266)	(13.709)
25 < Age ≤ 30	–0.217	–24.926*
	(0.187)	(13.709)
30 < Age ≤ 40	0.474***	34.968***
	(0.146)	(10.259)
40 < Age ≤ 50	0.211	12.143
	(0.138)	(9.863)
Age > 50	*f*	*f*

[13] These relatively weak gender effects might be surprising given that the raw data indicates women are 11 percentage points less likely to 'moonlight' than men. However, the gender effect is partly captured through the 'moonlighting' wage effect. On the basis of the estimates reported for the 'moonlighting' wage equation in Table A3 of the appendix, a sample average-aged 'moonlighting' woman earns, *ceteris paribus*, about 85 percent of what a man earns per 'moonlighting' hour. The gender pay disadvantage in the 'moonlighting' sector is thus likely to act as a disincentive for women.

Table 1 (cont). Structural models – participation and hours worked for Yugoslav employees

	Estimated Coefficients	
Variable	**Probit**	**Censored Tobit**
Female	–0.427**	–43.319***
	(0.203)	(15.071)
[Age ≤ 25] × Female	0.424	43.640
	(0.396)	(29.234)
[25 < Age ≤ 30] × Female	0.540*	56.313**
	(0.318)	(22.900)
[30 < Age ≤ 40] × Female	–0.191	–10.685
	(0.251)	(18.198)
[40 < Age ≤ 50] × Female	0.057	10.282
	(0.236)	(17.562)
***Region of Residence*:**		
Montenegro	3.332**	179.782
	(1.549)	(117.142)
Serbia	*f*	*f*
Vojvodina	–0.292***	–23.380***
	(0.096)	(6.913)
Kosovo and Metohija	–0.013	–14.480
	(0.176)	(13.378)
***Worker Skill Levels*:**		
Unskilled/Semi-Skilled	0.302**	18.933*
	(0.152)	(10.635)
Skilled	–0.452	–27.286
	(0.404)	(29.606)
High Skilled	0.241	18.344
	(0.187)	(14.107)
Lower White Collar	0.012	7.172
	(0.285)	(21.135)
Intermediate White Collar	*f*	*f*
Senior Professional	–0.154	–8.825
	(0.129)	(9.295)
Managerial	0.147	10.753
	(0.124)	(9.039)
***Primary Job Ownership*:**		
Private	1.063**	87.831**
	(0.535)	(41.115)

Table 1 (cont). Structural models – participation and hours worked for Yugoslav employees

	Estimated Coefficients	
Variable	**Probit**	**Censored Tobit**
***Household Structure*:**		
Couple with Children	*f*	*f*
Couple with Children & Relatives	0.169*	15.059**
	(0.093)	(6.687)
Couple without Children	–4.596**	–344.381**
	(2.031)	(155.567)
Single Parent with Children	0.042	4.954
	(0.158)	(11.010)
Single Person Household	0.157	9.988
	(0.154)	(10.960)
***Labour Supply Variables*:**		
$\hat{w}_s$	0.262***	13.336**
	(0.097)	(7.042)
y_m	–0.190**	–21.593***
	(0.077)	(5.467)
h_m	–0.206	–24.865**
	(0.146)	(10.726)
[$\hat{w}_s$] × [Skilled]	0.349*	20.346
	(0.192)	(14.077)
[$\hat{w}_s$] × Private Ownership]	–0.536**	–42.900**
	(0.197)	(14.446)
[y_m] × Montenegro]	–0.562*	–33.354**
	(0.220)	(16.643)
[y_m] × [Couple without Children]	0.658**	48.817**
	(0.285)	(21.787)
Observations	1302	1302
Log Likelihood Value	–721.890	–2652.388
Pseudo-R^2	0.092	†
Cramer's λ	0.108	†
$1 - e'e/s^2$	†	0.338
Proportion of Correct Predictions	0.717	†
σ	1.0 (imposed)	77.273***
		(3.185)

Table 1 (cont). Structural models – participation and hours worked for Yugoslav employees

Variable	Estimated Coefficients: Probit	Estimated Coefficients: Censored Tobit
Diagnostic Tests for Model		
Pseudo Functional Form RESET – χ^2_3	3.846	8.396**
	(0.279)	(0.038)
Homoscedasticity Test – χ^2_{23}	48.509***	78.733***
	(0.018)	(0.000)
Normality Test – χ^2_2	3.356	17.554***
	(0.187)	(0.000)
Tests for Omitted Variables		
Marital Status – χ^2_2	2.124	6.577*
	(0.346)	(0.037)
Completed Education – χ^2_4	2.228	0.966
	(0.694)	(0.915)
Settlement Type – χ^2_2	2.194	2.399
	(0.334)	(0.301)
Primary Workplace Size – χ^2_4	3.257	4.321
	(0.516)	(0.364)
Primary Job Branch Level – χ^2_9	3.224	4.015
	(0.955)	(0.910)
Tests for Excluded Interactions		
All Other Gender Interactions – χ^2_{21}	25.498	27.046
	(0.226)	(0.169)
Labour Supply Variables [$\hat{w}_s$, w_m, h_m] Interacted with:		
Age Controls – χ^2_{12}	13.313	10.015
	(0.347)	(0.615)
Region of Residence Controls – χ^2_8	6.998	7.631
	(0.537)	(0.470)
Worker Skill Level Controls – χ^2_{17}	21.277	24.452
	(0.214)	(0.108)
Household Structure Controls – χ^2_{11}	7.111	7.837
	(0.790)	(0.728)

Table 1 (concluded). Structural models – participation and hours worked for Yugoslav employees

	Estimated Coefficients	
Variable	**Probit**	**Censored Tobit**
Exogeneity Tests		
Log of Main Job Monthly Wages – χ^2_1	0.778	0.749
	(0.378)	(0.387)
Instrument Validity Test – χ^2_{31}	33.284	45.561**
	(0.357)	(0.044)
Log of Main Job Monthly Hours – χ^2_1	0.116	0.584
	(0.734)	(0.445)
Instrument Validity Test – χ^2_{27}	27.672	36.669
	(0.428)	(0.101)
Tobit Specification Test – χ^2_{23}	†	71.333***
		(0.000)

***Notes*:**

a) ***, ** and * denote statistical significance at the 0.01, 0.05 and 0.10 level respectively using two-tailed tests.

b) The variance-covariance matrix for the estimated probit model is corrected for the presence of heteroscedasticty of unknown form.

c) The diagnostics, with the exception of the tobit specification test, are computed using the efficient score test approach outlined in Chesher and Irish (1987) (see text). The RESET test uses the predicted standardized indices raised to the fourth power. The homoscedasticity test uses the set of explanatory variables from the original model. The normality diagnostic tests for zero skewness and a kurtosis value of 3 in the residuals. The tests for omitted variables relates to the specified sub-sets of variables. The exogeneity tests for the inclusion of residuals from either the reduced form main job wage or hours equations depending on the null (see Smith and Blundell, 1986). The instrument validity tests for the inclusion of the full set of identifying instruments. The significance levels for all the tests are reported in parentheses.

d) The tobit specification test is computed as $-2 \times [L^{\text{tobit}} - (L^{\text{truncated}} + L^{\text{probit}})]$ which is distributed as chi-squared with $(k^{\text{truncated}} + k^{\text{truncated}} - k^{\text{tobit}})$ degrees of freedom. The L denotes the maximized log-likelihood value and the three superscripts denote the censored tobit, the truncated tobit and the probit models, respectively, and k is the number of estimated parameters (see Fin and Schmidt, 1984).

e) The predicted values for the hourly wage in the secondary job are based on the selectivity corrected hourly wage equation estimates reported in Table A3.

f) The expected value of secondary hours worked for the sample is given by $\overline{[X_i'\gamma + \sigma\phi(X_i\gamma \div \sigma)/\Phi(X_i\gamma \div \sigma)]} = 49.6$. In addition, $\overline{\Phi(X_i\gamma \div \sigma)} = 0.306$.

g) The Pseudo-R^2 is defined as $1 - L^{\text{restricted}}/L^{\text{unrestricted}}$, where L denotes the maximized value of the log-likelihood function.

h) Cramer's $\lambda = \overline{\Phi(X_i\hat{\beta})} \mid \text{inform} - 1 - \overline{\Phi(X_i\hat{\beta})} \mid \text{inform} = 0$, (see Cramer, 1999).

i) $1 - e'e/s^2$ is a measure of goodness of fit for the tobit. $e'e$ is the sum of squared residuals for the tobit model with residuals defined as described in Chesher and Irish (1987). s^2 is the variance of the dependent variable (see Reece, 1979).

j) The standardized index estimated at the mean of the data for the probit model is –0.588. The sample average for the dependent variable implies a standardized index $\Phi^{-1}(0.299) = -0.525$.

k) † denotes not applicable in estimation and f denotes category omitted in estimation.

interacted. The ultimate effect on the standardized index is –0.65, which translates as a 24-point reduction and is statistically significant at a conventional level. Thus, the average employee resident in Montenegro is considerably less likely to participate in secondary job holding compared to a Central Serbian counterpart.

Skilled and unskilled manual workers are estimated to have a higher probability of secondary job holding relative to intermediate white-collar workers. The estimated impact effect suggests a ten point higher rate in both cases.[14] The point estimate for private sector workers, once allowance is made for the interaction term with the logarithm of monthly earnings on the main job, is not found to be statistically significant at a conventional level. The estimates for the household structure variables, in general, are poorly determined but there is marginal evidence that an employee in a household with dependent children and relatives is more likely to 'moonlight' compared to the reference household containing only dependent children.

Attention now turns to the estimates for the conventional labour supply variables in the 'moonlighting' participation equation. In order to facilitate interpretation, the probit estimates are re-expressed as marginal effects in Table 2 and represent outcomes of a 10 percent increase in the variables of interest.[15] The marginal effects for the hours equation are also reported in this table for completeness but are not discussed.

We initially focus on the effects for the predicted 'moonlighting' wage rate. The point estimate suggests that for employees other than skilled and those in the private sector, the overall effect is positive implying domination of the substitution over the income effect. In particular, a 10 percent rise in the 'moonlighting' wage rate increases participation by about one percentage point.[16] However, a more elastic response is obtained for skilled workers where a comparable increase in the wage rate induces twice this effect. The point estimate for private sector workers is negative but less precisely determined with a t-ratio of 1.2 in absolute terms.

The international literature on 'moonlighting' and informal activity offers some contrasting estimates on the effects of the informal wage rate on participant behaviour. Lemieux, Fortin and Frechette (1994), using Canadian data, detect a negative

[14] Given the interaction with the predicted 'moonlighting' wage, the overall effect for skilled workers is computed at the average value for the predicted wage and is $-0.452 + 0.349 \times 2.115 = 0.287$. This effect on the standardized index is comparable to the one reported in Table 1 for the set of unskilled/semi-skilled workers.

[15] Given the logarithmic nature of the specified regressors, a ten percent change is expressed in fractional terms as 0.1. The resultant effect of a ten percent change on the dependent variable is obtained by multiplying the marginal effect by 0.1.

[16] The estimated 'moonlighting' own wage effect can provide some further insight into the gender dimension of 'moonlighting'. On the basis of the calculations reported in footnote 13, gender 'moonlighting' pay equality could be achieved if the female 'moonlighting' hourly wage rate increased by about 17 percent. On the basis of the marginal effect this would induce a 1.5 percentage point rise in female participation. Thus, a modest one-tenth of the raw gender differential in 'moonlighting' activity can be ascribed to the 'moonlighting' gender pay differential itself.

Table 2. Marginal effects for the probit and tobit model for labour supply variables

Variable	Participation	Hours Worked
	Marginal Effects [c) d)] (Percentage Points)	Marginal Effects [c) e)] (Number of Second Job Hours per Month)
A 10% rise in:		
$\hat{w}_s$ – Log of Hourly Wages:		
Secondary Job[a)]	0.894***	0.408***
	(0.330)	(2.155)
[$\hat{w}_s$] × [Skilled]	2.086***	1.031*
	(0.624)	(0.403)
[$\hat{w}_s$] × [Private]	–0.936	–0.843
	(0.799)	(0.569)
A 10% rise in:		
y_m – Log of Monthly Wages: Main Job[b)]	–0.648**	–0.661***
	(0.263)	(0.167)
[y_m] × Montenegro	–2.568***	–1.681***
	(0.737)	(0.500)
[y_m] × [Couple without Children]	1.599*	0.833
	(0.946)	(0.661)
A 10% rise in:		
h_m – Log of Monthly Hours Worked: Main Job	–0.704	–0.761
	(0.497)	(0.328)

Notes:
a) The marginal effect for all employees other than skilled and those with primary jobs in the private sector.
b) The marginal effect for all employees other than those resident in Montenegro and those in households with a couple but without children.
c) ***, ** and * denote statistical significance at the 0.01, 0.05 and 0.1 level, respectively.
d) Asymptotic standard errors, based on heteroscedastic adjustment, are reported in parentheses.
e) Asymptotic standard errors are reported in parentheses.

relationship between hours worked in the 'underground' economy and the 'underground' wage rate. In contrast, Shishko and Rostker (1976), Krishnan (1990) and Conway and Kimmel (1998) provide evidence for the United States of a positive relationship between 'moonlighting' activity and the informal wage rate. Gaddy (1991), using data drawn from a sample of emigrants, obtains a positive correlation between informal activity and informal wage rates for the old Soviet Union, and Foley (1997), Kolev (1998) and Guariglia and Kim (2000) detect a similar relationship in terms of second-job holding using more recent data from transitional Russia for the 1990s.[17]

The logarithm of the monthly wage also achieves statistical significance in the participation model. A ten percent increase in the monthly wage on the main job reduces, *ceteris paribus*, the probability of moonlighting by close to two-thirds of a percentage point for all employees other than those residing in Montenegro and those in households without dependent children. The estimated effect for Montenegrin residents is almost four times greater than for all others and suggests that a 10 percent rise in monthly earnings on the main job reduces the probability of 'moonlighting' in Montenegro by 2.6 percentage points. Both these effects are consistent with the notion that leisure is a normal good.[18] The result for households with no dependent children is conceded as somewhat perverse given neo-classical predictions.

It is clear that some caution must be exercised in comparing the dimensions of the semi-elasticities for the 'moonlighting' own wage and monthly earnings on the main job. A 10 percent increase in the monthly pay measure represents a larger effect in terms of Dinars than a comparable increase in the hourly 'moonlighting' pay rate. The ratio of monthly earnings to the hourly 'moonlighting' wage rate is approximately 55:1 (see Table A2). Given the comparable absolute magnitude of the semi-elasticities for both measures, this clearly suggests that 'moonlighting' activity, on average, is more responsive to a one Dinar change in the hourly 'moonlighting' rate than to a similar change in the monthly pay measure on the main job. As noted earlier, however, this differential provides no basis for discrimination between an opportunity and a 'coping strategy' motive for 'moonlighting'.

Finally, the point estimate for the marginal effect on the natural logarithm of main hours suggests an inverse relationship between the number of hours worked in the main job and 'moonlighting'. However, the estimated *t*-ratio is 1.4 in absolute terms. Kolev (1998) and Guariglia and Kim (2000) also detect an inverse relationship in their analysis for Russia but their estimated effects are somewhat better determined.

---

[17] In spite of the fact that all three studies used various rounds of the RLMS, specifications and/or estimation methods differed. This could partly explain the range in the marginal effects estimated for the 'moonlighting' wage rate in these studies. A 10 percent rise in the hourly wage rate yields 0.6 of a percentage point rise in the second-job holding rate in Kolev (1998), a 1.2 point rise in Foley (1997), and a 5 percentage point rise in Guariglia and Kim (2000). This latter estimate does appear on the high side.

[18] The work of Foley (1997) and Guariglia and Kim (2000) suggest somewhat higher estimates for a comparable change in monthly earnings on the main job in Russia. Our calculations based on their work suggest marginal effects of around two percentage points – dimensionally closer to the Montenegro estimate in our study. The estimated effects reported in Kolev (1998) are statistically insignificant.

6. Conclusions

The dramatic fall in real incomes and the rise in unemployment in the FRY associated with the 'break-up' of the federation and international sanctions stimulated a rise in informal activity. It is estimated that approximately 750,000 employees were engaged in 'moonlighting' activity in the FRY in 1997.[19] Our empirical analysis attempted to shed some light on the activity of these employees at a time of comparative economic and political stability in the FRY.

We situated our analysis within a conventional static neo-classical labour supply model. The structural participation model passed a demanding array of diagnostics but the companion equation for 'moonlighting' hours was found to be mis-specified along a number of important dimensions. The participation model was used, therefore, to provide the focus for interpretation and the basis for exploring predictions implied by neo-classical theory.

The empirical estimates of the structural participation model indicate some modest differences in employee 'moonlighting' behaviour across age groups. The highest *ceteris paribus* participation rate is associated with men in their thirties. We inferred that the lower female participation in 'moonlighting' activity observed in our data is only partly driven by the lower predicted pecuniary returns to females from such activity (see footnotes 13 and 16). We also noted a strong regional dimension to 'moonlighting' activity in the FRY with employees in Central Serbia disproportionately represented in this type of activity. In addition, blue-collar workers are found to be more likely to engage in 'moonlighting' than all types of white-collar workers.

A key finding of our empirical analysis is that the 'moonlighting' wage rate exerts an important and independent influence on 'moonlighting' activity. The average participation elasticity for most employees is estimated at 0.3.[20] We did detect some variation across worker type. In particular, skilled workers are estimated to be more responsive to wage rate changes in the 'moonlighting' job than all other workers with a participation elasticity of 0.7. Our empirical analysis suggested an important role for earnings on the main job and there was evidence of an important regional variation in this effect. The computed participation elasticity with respect to main job earnings is −0.2 for all employees other than those in Montenegro and in households with no dependent children. The computed elasticity for Montenegro is −0.9. In terms of the Pseudo-R^2 measure, the array of

[19] The Federal Statistics Office in FR Yugoslavia estimated there to be approximately 2,507,000 employees in 1997 (see Statistical Yearbook of Yugoslavia, 1999). Our survey estimates suggest that about 30 percent of these are engaged in second-job holding activity.
[20] The participation elasticities are computed by dividing the original marginal effects by the sample average participation rate of 0.3. The elasticities for the hours worked equation can be obtained by dividing the original marginal effects by 17 – the average number of second-job hours worked for the full sample.

neo-classical labour supply variables was found to account for about 40 percent of what the participation model actually explains.

It is true that our empirical specifications lacked a non-labour income measure which, given the possible importance of other income sources in the composition of household income, could be viewed as a limitation. It is possible that the real decline of savings, social transfers, payments from abroad, pensions, *etc.* encouraged the participation of employees in 'moonlighting' activity but these issues cannot be explored with our current data. However, on the basis of the diagnostic tests reported, there is no evidence that the participation model is mis-specified through the omission of a non-labour income measure.[21]

Labour market earnings fell in real terms by about 50 percent in the FRY over the period 1990 to 1997. Monthly earnings, therefore, would have to double to return to the 1990 level. Our estimates suggest that a rise of this magnitude (outside Montenegro) would lead, on average, to a 4.5 percentage point fall in the probability of 'moonlighting', *ceteris paribus*.[22] Thus, the restoration of main job earnings to levels prevalent around the time of the political 'break-up' is only likely to reduce the scale of 'moonlighting' activity by about one-seventh. This crudely translates as an aggregate lower-bound reduction in 'moonlighting' activity of about 107,000 employees.

Although the Yugoslav government is committed to the development of a market-led economy, the reform process has been slow even during periods of relative political stability and no international sanctions. This has been particularly so in terms of the privatization process which, after some initial euphoria in the early 1990s, stalled in 1994 (see Djuricin, 1997). It is accepted that successful transitional economies require a stable macroeconomic framework and the development of adequate non-market institutions (see McMillan, 1997). Recent political events in the FRY augur well for the future in this regard but the conditions that stimulate 'moonlighting' activity are likely to remain in place for some time to come.

References

Amemiya, T. (1984). 'Tobit models: a survey', *Journal of Econometrics*, 24, pp. 3–61.

Bell, D., Robert, A. and Wright, R. (1997). 'Multiple job-holding as a hedge against unemployment', CEPR, Discussion Paper Series No. 1626, London: CEPR.

Bogicević, B., Krstić, G. and Mijatović, B. (2001). 'Analysis of poverty and reform of social assistance in Serbia', *Mimeo*, UNICEF.

[21] In this regard, we take some comfort from the fact that non-labour income exerted no independent influence on 'moonlighting' in the work of either Kolev (1998) or Guariglia and Kim (2000) for Russia. It was marginally significant in the work of Foley (1997) but incorrectly signed.

[22] The fractional equivalent of a doubling in monthly earnings, given the use of natural logarithms, is 0.69. Thus, the original marginal effect of –0.0648 is multiplied by 0.69 to induce the doubling effect in this case.

Božović, G. (1993). 'Grey economy in FR Yugoslavia – 1991 and 1992', *Yugoslav Survey*, No. 2, pp. 71–84.

Chesher, A. and Irish, M. (1987). 'Residual analysis in the grouped and censored normal linear model', *Journal of Econometrics*, 34, pp. 33–61.

Contini, B. (1992). *The Irregular Economy of Italy: A Survey of Contributors – Guide Book to Statistics on the Hidden Economy*, New York: United Nations.

Conway, K. S. and Kimmel, J. (1998). 'Male labor supply estimates and the decision to moonlight', *Labour Economics*, 5, pp. 135–66.

Cramer, J. (1999). 'Predictive performance of the binary logit model in unbalanced samples', *Journal of the Royal Statistical Society, Series D*, 48, pp. 85–94.

Djuricin, D. (1997). 'Privatization in FR Yugoslavia's transition', in *Challenges and Opportunities for the Economic Transition in Yugoslavia*: Proceedings from the International Conference, 5–6 November 1997, Belgrade, Yugoslavia, edited by G. Pitic, published by USAID, Economics Institute and Chesapeake Associates.

EBRD (2000). *Transition Report*, London: European Bank for Reconstruction and Development.

Foley, M. (1997). 'Multiple job holding in Russia during the transition', Yale Economic Growth Center Discussion Paper # 781, http://www.unc.edu/~foley/mjh.rtf

Fin, T. and Schmidt, P. (1984). 'A test of the Tobit specification against an alternative suggested by Cragg', *Review of Economics and Statistics*, 66, pp. 174–77.

Gaddy, C. G. (1991). 'Who works on the left? The determinants of Soviet citizens' supply of labour to the second economy', Berkeley-Duke Occasional Papers on the Second Economy in the USSR, # 24, California: Duke University.

Greene, W. H. (1997). *Econometric Analysis*, Third Edition, New York: Prentice Hall.

Guariglia, A. and Kim, B.-Y. (2000). 'The dynamics of moonlighting: What is happening in the Russian informal economy?' *Mimeo*, Department of Economics, University of Essex.

Heckman, J. (1979). 'Sample selection bias as a specification error', *Econometrica*, 47, pp. 153–61.

Huber, P. J. (1967). 'The behaviour of maximum likelihood estimates under non-standard conditions', in *Proceedings of the Fifth Berkeley Symposium on Mathematical Statistics and Probability*, Berkeley, CA: University of California Press, 1, pp. 221–223.

Human Development Report – FR Yugoslavia (1996). Economics Institute, Belgrade.

Human Development Report – FR Yugoslavia (1997). Economics Institute, Belgrade.

Johnson, S., Kaufmann, D. and Schliefer, A. (1997). 'The unofficial economy in transition', *Brookings Papers on Economic Activity*, pp. 159–221.

Kolev, A. (1998). 'Labour supply in the informal economy in Russia during transition', CEPR Discussion Paper in Transition Economics # 2024, London: CEPR.

Krishnan, P. (1990). 'The economics of moonlighting: a double self-selection model', *Review of Economics and Statistics*, 72, pp. 361–67.

Krstic, G. (1998). 'Analysis of the hidden economy in FR Yugoslavia with estimates for 1997 and recommendations for its legalization', Economics Discussion Paper, 504, Economics Institute, Belgrade.

Krstic, G. (1999). 'Comparing informal activity using the labour force surveys and the informal sector surveys', *Mimeo*, Economics Institute, Belgrade.

Lacko, M. (2000). 'Hidden Economy – an unknown quantity?' *Economics of Transition*, 8, pp. 117–49.

Lemieux, T., Fortin, B. and Frechette, P. (1994). 'The effects of taxes on labor supply in the underground economy', *American Economic Review*, 84, pp. 231–54.

McMillan, J. (1997). 'Markets in transition', in Kreps, D. and Wallis, K. F. (eds.), *Advances in Economics and Econometrics: Theory and Applications – Seventh World Congress*, Vol. II Cambridge: Cambridge University Press.

Melino, A. (1982). 'Testing for sample selection bias', *Review of Economic Studies*, 49, pp. 151–53.
Nove, A. (1987). *The Soviet Economic System*, Third Edition, London: Unwin-Hyman.
Orme, C. (1990). 'The small sample performance of the information matrix test', *Journal of Econometrics*, 46, pp. 309–31.
Petrović, P., Bogetić, Ž. and Vujošević, Z. (1999). 'The Yugoslav hyperinflation of 1992–1994: causes, dynamics and money supply process', *Journal of Comparative Economics*, 27, pp. 335–53.
Pošarac, A. (1998). 'Poverty and inequality in Yugoslavia in the 1990–1995 period', *Mimeo*, Institute for Economics Sciences, Belgrade.
Reece, W. S. (1979). 'Charitable contributions: new evidence on household behaviour', *American Economic Review*, 69, pp. 142–51.
Reilly, B. and Krstic, G. (2002). 'An anatomy of labour market earnings inequality in Serbia – 1995 to 2000', *Mimeo*, Department of Economics, University of Sussex.
Shishko, R. and Rostker, B. (1976). 'The economics of multiple job holding', *American Economic Review*, 66, pp. 298–308.
Smith, R. and Blundell, R. (1986). 'An exogeneity test for a simultaneous equation tobit model with an application to labour supply', *Econometrica*, 54, pp. 679–85.
Statistical Yearbook of Yugoslavia (1999). Federal Statistics Office, Belgrade, FR Yugoslavia.
Thomas, J. J. (1992). *Informal Economic Activity*, Harlow: Harvester Wheatsheaf.

Appendix

Table A1. Summary sample statistics for characteristics of Yugoslav employees

Variable	All employees	Main job holders only	Secondary job holders	*z*-score/ *t*-test
***Age and Gender*:**				
Age (years)	41.06	40.91	41.40	–0.84
Age ≤ 25	0.010	0.075	0.044	2.07**
25 < Age ≤ 30	0.177	0.085	0.138	2.65***
30 < Age ≤ 40	0.268	0.331	0.241	–3.33***
40 < Age ≤ 50	0.350	0.344	0.364	–0.69
Age > 50	0.195	0.202	0.174	1.11
Female	0.406	0.445	0.315	4.53***
***Marital Status*:**				
Married	0.738	0.720	0.779	–2.29**
Single	0.175	0.191	0.138	2.43**
Divorced/widowed	0.087	0.089	0.083	0.36

Table A1 (cont). Summary sample statistics for characteristics of Yugoslav employees

Variable	All employees	Main job holders only	Secondary job holders	*z*-score/ *t*-test
***Completed Education*:**				
No Education	0.013	0.013	0.013	0.00
Primary	0.102	0.093	0.123	–1.56
Secondary	0.556	0.551	0.567	–0.53
High School	0.159	0.174	0.123	2.45**
Higher/University	0.170	0.169	0.174	–0.22
***Settlement Type*:**				
Rural	0.216	0.212	0.226	–0.56
Town	0.228	0.216	0.256	–1.54
City	0.556	0.572	0.518	1.79
***Region of Residence*:**				
Montenegro	0.161	0.194	0.085	5.66***
Serbia	0.561	0.514	0.671	–5.42***
Vojvodina	0.236	0.255	0.190	2.65***
Kosovo and Metohija	0.042	0.037	0.054	–1.30
***Workplace Characteristics*:**				
Employment Size ≤ 5	0.099	0.106	0.085	1.21
6 ≤ Employment Size ≤ 25	0.160	0.171	0.133	1.79*
26 ≤ Employment Size ≤ 100	0.230	0.230	0.231	0.00
101 ≤ Employment Size ≤ 500	0.199	0.184	0.233	–1.96**
Employment Size > 500	0.312	0.309	0.318	–0.32
Ownership – Private	0.104	0.111	0.089	1.24
***Worker Skill Levels*:**				
Unskilled/Semi-Skilled	0.081	0.072	0.100	–1.60*
Skilled	0.178	0.162	0.215	–2.20**
High Skilled	0.039	0.034	0.051	–1.34
Lower White Collar	0.019	0.021	0.015	0.77
Intermediate White Collar	0.358	0.372	0.327	1.57
Senior Professional	0.149	0.168	0.105	3.17***
Managerial	0.176	0.171	0.187	–0.68
***Primary Job Branch Level*:**				
Industry and Mining	0.271	0.277	0.260	0.64
Agriculture	0.045	0.044	0.046	–0.16

Table A1 (cont). Summary sample statistics for characteristics of Yugoslav employees

Variable	All employees	Main job holders only	Secondary job holders	*z*-score/ *t*-test
Construction	0.059	0.050	0.079	–1.88**
Transport	0.082	0.075	0.100	–1.43
Trade	0.104	0.109	0.092	0.95
Catering and Tourism	0.064	0.066	0.059	0.48
Crafts	0.054	0.050	0.062	–0.84
Financial and Other Services	0.046	0.054	0.028	2.32**
Education, Culture, Health etc.	0.166	0.162	0.174	–0.53
Government	0.109	0.113	0.100	0.70
***Household Structure*:**				
Couple with Children	0.539	0.551	0.510	1.36
Couple with Children & Rel.	0.239	0.223	0.277	–2.04**
Couple without Children	0.071	0.072	0.069	0.19
Single Parent with Children	0.078	0.083	0.067	1.02
Single Person Household	0.073	0.071	0.077	–0.38
***Wages & Hours (Main Job)*:**				
y_m: Monthly Wages (natural log)	7.092	7.120	7.025	2.54**
h_m: Monthly Hours (natural log)	5.140	5.150	5.117	2.14**
$\hat{w}_s$: Predicted Hourly Wage (Secondary Job)	2.115	2.086	2.183	–3.08***
Second Job (or 'Moonlighting') Rate	0.299	0.0	1.0	†
'Moonlighting' Hours Worked	†	0.0	57.22	†
Sample Size	1302	912	390	

***Notes*:**

a) The All Employees column refers to the pooled sample. The Main Job Holders column refers to the sub-sample of employees who report holding only a main job, the Second Job Holders column refers to the sub-sample who report holding a secondary job.

b) All variables reported with the exception of Wages and Hours are binary variables. A *z*-score is used to test differences in proportions between the Main Job Holders column and the Second Job Holders column. A *t*-test, based on the assumption of common variances, is used to test differences in the continuous variables: Monthly Wages, Monthly Hours, the Predicted Hourly Wage and Age.

c) The Predicted Hourly Wage (Second or 'Moonlighting' Job) is computed using the OLS procedure corrected for selectivity bias (see Table A3).

d) The *z*-scores and *t*-tests undertake a test for the difference in sample means between column two and column three.

e) ***, ** and * denote statistical significance at the 0.01, 0.05 and 0.10 level, respectively using two-tailed tests. The corresponding critical (absolute) values are 2.576, 1.96 and 1.64, respectively.

f) † denotes not applicable.

Table A2. Summary distribution of primary and secondary job earnings by selected percentiles and summary measures of dispersion

	Mean	10th	25th	50th	75th	90th	Std. Dev.	CV	Gini
All Employees Sample Size = 1302									
Monthly Wages – Main Job	1439.5	560	850	1200	1800	2450	1047	0.727	0.318
Hourly Wages – Main Job	8.590	3.261	4.891	7.065	10.349	14.130	8.022	0.934	0.336
Non-Participants Sample Size = 912									
Monthly Wages – Main Job	1474	600	893	1243	1800	2494	1073	0.728	0.315
Hourly Wages – Main Job	8.766	3.261	5.000	7.113	10.479	14.417	8.584	0.978	0.338
Participants Sample Size = 390									
Monthly Wages – Main Job	1359	500	797	1200	1705	2390	980	0.721	0.325
Hourly Wages – Main Job	8.177	3.161	4.620	6.929	10.326	13.587	6.514	0.797	0.331
Monthly Wages – Second Job	951	250	500	700	1200	2000	810	0.852	0.405
Hourly Wages – Second Job	24.697	5.189	8.719	15.000	30.000	50.00	28.147	1.140	0.491
Ratio	3.020	1.642	1.887	2.165	2.905	3.680	4.321	1.430	†

***Notes*:**
a) All earnings measures are expressed in new Dinars.
b) Std.Dev. denotes the standard deviation; CV denotes the coefficient of variation; and Gini denotes the Gini coefficient.
c) Ratio denotes the ratio of the hourly wage rate in the secondary job to the hourly wage rate in the main job for participants.
d) † denotes not applicable.

Table A3. Secondary job hourly wage rate equation – corrected for selectivity bias

Variable	**Probit model**	**Log hourly wages**
Constant	1.177***	2.489***
	(0.874)	(0.418)
Age ≤ 25	−0.181	†
	(0.276)	
25 < Age ≤ 30	−0.158	†
	(0.197)	
30 < Age ≤ 40	0.449***	†
	(0.144)	
40 < Age ≤ 50	0.159	†
	(0.137)	
Age > 50	*f*	†
Female	−0.188	−3.511**
	(0.231)	(1.544)
[Age ≤ 25] × Female	0.602	†
	(0.392)	
[25 < Age ≤ 30] × Female	0.565*	†
	(0.319)	
[30 < Age ≤ 40] × Female	−0.086	†
	(0.256)	
[40 < Age ≤ 50] × Female	0.143	†
	(0.246)	
Age	†	−0.041
		(0.047)
$Age^2 \div 100$	†	0.068
		(0.056)
[Age] × [Female]	†	0.169**
		(0.080)
[Age^2] × [Female] ÷ 100	†	−0.212**
		(0.100)
Marital Status:		
Married	0.041	†
	(0.226)	
Single	−0.193	†
	(0.211)	
Divorced/Widowed	*f*	†

Table A3 (cont). Secondary job hourly wage rate equation – corrected for selectivity bias

Variable	Probit model	Log hourly wages
***Completed Education*:**		
No Education	−0.311	−0.756**
	(0.370)	(0.378)
Primary	−0.010	−0.444***
	(0.168)	(0.141)
Secondary	*f*	*f*
High School	−0.026	0.005
	(0.188)	(0.140)
Higher/University	−0.257	0.313**
	(0.278)	(0.124)
***Settlement Type*:**		
Rural	−0.128	−0.460**
	(0.119)	(0.129)
Town	*f*	*f*
City	−0.093	0.062
	(0.099)	(0.108)
***Region of Residence*:**		
Montenegro	−0.695***	0.025
	(0.129)	(0.209)
Serbia	*f*	*f*
Vojvodina	−0.337***	−0.057
	(0.100)	(0.124)
Kosovo and Metohija	−0.025	0.023
	(0.192)	(0.195)
[Montenegro] × [Female]	†	−0.672**
		(0.324)
***Worker Skill Levels*:**		
Unskilled/Semi-Skilled	0.286	†
	(0.195)	
Skilled	0.255**	†
	(0.119)	
High Skilled	0.226	†
	(0.207)	
Lower White Collar	0.044	†
	(0.315)	

Table A3 (cont). Secondary job hourly wage rate equation – corrected for selectivity bias

Variable	Probit model	Log hourly wages
Intermediate White Collar	*f*	†
Senior Professional	−0.145	†
	(0.203)	
Managerial	0.313	†
	(0.279)	
***Primary Job Branch Level*:**		
Industry and Mining	*f*	*f*
Agriculture	0.339	0.281
	(0.228)	(0.228)
Construction	0.421**	0.394*
	(0.195)	(0.200)
Transport	0.295*	0.617***
	(0.156)	(0.166)
Trade	0.437**	0.118
	(0.210)	(0.178)
Catering and Tourism	0.173	0.241
	(0.197)	(0.216)
Crafts	0.311	0.441**
	(0.200)	(0.229)
Financial & Other Services	−0.719*	−0.115
	(0.430)	(0.271)
Education, Culture, Health etc.	0.258*	0.099
	(0.147)	(0.178)
Government	0.002	0.462***
	(0.153)	(0.157)
[Agriculture] × [Female]	−1.120**	†
	(0.572)	
[Construction] × [Female]	−0.947**	−1.415**
	(0.461)	(0.602)
[Trade] × [Female]	−0.670**	†
	(0.278)	
[Crafts] × [Female]	†	0.671*
		(0.376)
[Financial & Other Services] × [Female]	0.867*	†
	(0.494)	

Table A3 (cont). Secondary job hourly wage rate equation – corrected for selectivity bias

Variable	Probit model	Log hourly wages
[Education, Culture, Health etc.] × [Female]	†	0.877***
		(0.224)
***Primary Job Characteristics*:**		
Private Ownership	−0.105	0.415**
	(0.172)	(0.173)
Employment Size ≤ 5	*f*	†
6 ≤ Employment Size ≤ 25	−0.145	†
	(0.188)	
26 ≤ Employment Size ≤ 100	0.122	†
	(0.209)	
101 ≤ Employment Size ≤ 500	0.087	†
	(0.196)	
Employment Size > 500	0.003	†
	(0.214)	
[26 ≤ Employment Size ≤ 100] × [Female]	−0.474**	†
	(0.207)	
[Employment Size > 500] × [Female]	−0.404**	†
	(0.206)	
***Household Structure*:**		
Couple with Children	*f*	†
Couple with Children & Relatives	0.168*	†
	(0.093)	
Couple without Children	0.043	†
	(0.158)	
Single Parent with Children	0.090	†
	(0.229)	
Single Person Household	0.267	†
	(0.198)	
h_m: Natural Logarithm of Monthly Hours Worked in the Primary Job	−0.346**	†
	(0.156)	
$\frac{\phi}{\Phi}$	†	0.579**
		(0.227)
Observations	1302	390
σ	†	0.889

Table A3 (cont). Secondary job hourly wage rate equation – corrected for selectivity bias

Variable	Probit model	Log hourly wages
ρ	†	0.651
Adjusted-R^2	†	0.247
Melino Test for Sample Selection – χ_1^2	†	6.718***
		(0.000)
Identifying Restrictions for Selection – χ_{23}^2	43.245**	24.993
	(0.025)	(0.297)

Notes:
a) ***, ** and * denote statistical significance at the 0.01, 0.05 and 0.1 level, respectively.
b) The variance-covariance matrix is corrected for heteroscedasticity and the inclusion of a generated regressor. Asymptotic standard errors are reported in parentheses.
c) $\frac{\phi}{\Phi}$ denotes the inverse of the Mills ratio term which is calculated using the reduced form probit estimates. The Melino test is the square of the t-statistic on the selection term using the uncorrected OLS standard error (see Melino, 1982).
d) ρ denotes the estimated correlation coefficient for the errors in the reduced form probit selection equation and the secondary job earnings equation. The estimated coefficient for the inverse of the Mill's ratio term is the product of the standard deviation, adjusted for selection bias, and ρ.
e) † denotes not applicable in estimation and f denotes category omitted in estimation.