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# FEMALE DUAL LABOUR MARKETS AND EMPLOYEE BENEFITS

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## Abstract

*The American workforce and the role of employee benefits have changed dramatically since the 1980s when economists seriously considered dual labour market models to describe pay and employment patterns. Then, dual labour market models described men's labour markets, but not women's and the tests applied to wages and salaries, not total compensation including employee benefits. Applying a switching regression technique using the 2000 Current Population Survey and including women workers and employee benefits, we find that the dual labour market hypothesis is consistent with both female and male labour market structures, especially when total compensation is considered.*

## I Introduction

This study reviews two decades of developments in dual labour market theory and contributes to tests of the theory by including women and employer-provided health and pension plans. We use the switching regression technique, which is an econometric procedure that determines whether two regression equations fit the data better than one. If the data are grouped into two sectors, the technique estimates which sector a worker is most likely to be in using maximum likelihood techniques. Human capital theory predicts that workers' pay depends on differences in individual productivity and preferences for cash, employee benefits, and job characteristics and not on a sector location. Conversely, the dual labour market model predicts that a return to productivity characteristics depends on what sector one is in. Based on 1980 data, Dickens *et al.* (1988) rejected the dual labour market model for women because it seemed most were located in one sector. In contrast, 20 years later and using total compensation as a measure of pay we find that the dual market hypothesis explains women's labour markets. Our results confirm the earlier findings that the dual market hypothesis explains men's pay when employee benefits are and are not included. This study advances previous dual labour market studies by

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including women and using total compensation – wages, health insurance, and pensions.

American scholars' interest in dual labour market theory waned in the 1990s, perhaps because involuntary unemployment seemed to diminish. In addition, complements to human capital models – e.g. efficiency wage theory – also predicted involuntary unemployment and pay inequality for similar workers. Human resource scholars and the popular press described a new restructured labour market (Levine *et al*., 2002). Support for the dual labour market thesis has been found more recently in non-American labour markets (Blackaby *et al*., 1995).

Developments in the American labour market outcomes are not inconsistent with dual labour market predictions; in 1979, 11.5% of the white male workforce had jobs at poverty-level wages; in 1999, the 16% were so employed (Mishel *et al*., 2001, p. 133). Indicators of job security are higher and growing for top income earners while falling for black males and the lowest earners (Neumark, 2000), and, important for this study, the inequality in access to employee – provided health insurance and pension coverage is growing significantly (Calabrese and Medoff, 2001). Both blue- and white-collar workers lost employee benefit coverage between 1979 and 1998, but blue-collar workers lost more. Blue-collar health insurance coverage rates fell from 66% to 49% in these 20 years, while white-collar rates fell from 69.3% in 1979 to 60.3% in 1998. Blue-collar workers also lost more pension coverage, from 46.4% in 1979 to 33.6% in 1998; white-collar coverage rate was maintained at over 49%.

Understanding eroding job quality and falling coverage in employee benefits requires accurate theories of labour markets. The persistent segmentation of the labour market could be one explanation, but no one has tested the segmentation hypothesis for the US in over 10 years.

Following the above introduction, the second part of the paper contrasts the dual labour market theory and human capital theory with special attention to various tests of the hypothesis, especially Dickens and Lang (1985); the third section describes our methodology and testing strategy, the fourth section explains how we use the Current Population Survey data, and the fifth section describes the results and the conclusion suggests further research and policy implications.

## II Dual Labour Market Theory

In the 1970s and 1980s, labour economics renewed interest in Cairnes (1874) and Kerr's (1954) contention that employers and workers do not compete fully in one labour market, but in several segmented markets. One focus was on theories of economic discrimination and African Americans' and women's poor labour market outcomes: low wages, job insecurity, and lower returns for skill acquisition and education. Radical economists used historical data and case studies to argue that employers actively engage in 'divide and conquer strategies' by manipulating racial and gender divisions to lower labour costs (Edwards *et al*., 1982).

Non-Marxist theories argued that dual economies were created by technological differences, causing some firms to establish internal labour markets and primary sector jobs. Piore (1971) and Berger and Piore (1980) added product market instability to the theory and argued that core firms stabilized production and shifted business cycle risk to peripheral firms. Core firms have stable employment and internal labour markets; peripheral firms face intense competition and offer low wages and insecure employment.

Bulow and Summers (1986) and Albrecht and Vorman (1992) offer technology-based explanations for non-competing groups in labour markets by extending Shapiro and Stiglitz's (1984) efficiency wage model, that employers purposefully create queues by paying jobs above market rates because certain technologies cause high monitoring costs. Furthermore, because firms have different time horizons, firms with high costs of detection or low discount rates offer secure employment and training, while others hire on-the-spot market (Arai, 1997). Other models of optimization behaviour that cause persistent disequilibrium and queues in the primary sector include: wage rigidity and insider/outsider and efficiency wage models that emphasize employers' retention motives for maintaining internal labour markets. Other reasons firms pay more than the minimum required to attract qualified workers are the following: to discourage unionization (the threat effect), to recruit good employees, and to maintain industry and workplace norms (Akerlof, 1984).

Cain (1987, p. 226) argues that modern neoclassical economics reconciles six sources of wage inequality identified as far back as Adam Smith and that new theories, such as dual labour market theories, are unnecessary: compensating wage differentials; human capital investment differences; transitory differences because of lags and demand changes; mobility barriers such as geography (Heckman and Holtz, 1986), unions and other forms of regulation; class and group deprivations (due in part to cultural bias towards ethnicity, skin colour, and sex); and legitimate differences in productivity that are difficult to measure. Not that the dual labour market theory is not strictly referring to two sectors, as Cain (1987, p. 225) observed the term, dual labour market, is a 'metaphor for unnumbered (but few) segments.' Therefore, we, like many others, use the terms, 'dual labour market' and 'segmentation,' interchangeably.

The last point bedevils labour economists and econometricians. If both productivity and choices are not controlled for, it is possible that what looks like persistent undervaluing of productivity-enhancing characteristics are actually two serious measurement issues. First, individuals are compensated with pay, fringe benefits, and non-pecuniary attributes that are often difficult to value. Moreover, even if these job characteristics could be measured, workers value the same non-pay attributes differently, so they demand different levels of compensation (Taubman and Wachter, 1986). Second, not all the characteristics that make workers productive can be identified. Recent work on earnings determination suggests that non-cognitive skills – such as reliability, persistence, and honesty, which are caused by habits and personality traits – do affect pay (Bowles *et al*., 2001; Dunifron *et al*., 2001; Heckman and Rubenstein, 2001). Not measuring these may lead an econometrician to mistakenly attribute a low

return on formal education to a worker being in the secondary sector or subject to discrimination and not to the absence of one of these non-cognitive, but relevant skills. Also, management might create firm-specific effects, which, in return, affect productivity (Abowd *et al.*, 2001).

*Race and sex in labour markets*

For the most part, women's labour markets have been excluded in the debate about dual labour markets. Dickens and Lang (1992) justify leaving out women in subsequent analyses because 'pink-collar jobs exhibit characteristics of primary and secondary jobs.' Although meant as an aside, this comment may belie a serious failing in dual labour market theory. Did Dickens and Lang mean human capital theory, and a single labour market best explains the variation in women's earnings? Does this mean that as more women enter the labour force duality will disappear? Moreover, leaving women out of the tests seems *ad hoc*, since racial identification is examined and the theory suggests that women and non-whites are more likely to be excluded from the primary sector. If a significant part of the labour market cannot be characterized as segmented, then the theory is weak. An adequate test of the dual labour market hypothesis must include all workers. Because women pose a problem, interest in dual labour market theory fell off.

Friedberg *et al.* (1988) found that, in 1980, women consistently earned less than men, but did not receive lower returns for education when occupation and industry were controlled for. Yet, simple cross tabs on earnings and education by race and sex support dual labour market theory in 2000. Educated white men earn more than every group, but black women with college degrees earn about the same as white men with high school diplomas (Bureau of Labor Statistics, 2002). Women may be in jobs that do not pay for education – the secondary sector – or in jobs that pay for men's education, but not theirs, because of discrimination. Friedberg *et al.* (1988) argue that women are paid less than men with the same education because they are consistently employed in lower-paid occupations and industries than men. However, they argue, this is not evidence that women are in dual labour markets; their switching regression technique sorts almost all women (over 30 years) into one sector and women were paid positive returns for their education. They concluded that the dual labour market model was not an appropriate way to describe women's employment patterns. Yet, they admittedly did not control for women's labour market experience.

Craypo and Cormier (2000) findings that four out of ten workers in a small sample of working poor households were in service industries and one-third were high school dropouts, could lend support to the human capital hypothesis. However, black women's wages were well below the average, even though they had the highest level of education among all race and gender groupings. White men had the least education and were paid the highest wages. These data are suggestive of some segmentation – there are queues for jobs that pay for skills – however, the data could be merely revealing discrimination. Boston (1990, p. 101) argues that market segments exhibit unexplained earnings differentials

and occupational mobility barriers for white men and women and black men and women, so that discrimination is an insufficient explanation of market segmentation.

Boston (1990) aims to untangle outcomes caused by discrimination and those caused by segmentation. We pursue Boston's ideas by checking for the effect of segmentation on women's earnings (cash and employee benefits) by testing the dual labour market hypothesis separately from men using an endogeneity test. Boston uses an exogeneity test by defining jobs that require more specific skills and training (these data are available in descriptions of occupations) as primary sector jobs. He divides the sectors with these criteria before testing for differences in rates of return on ability by segment. He performs separate tests on white men and women and on black men and women. He finds support for the hypothesis that segmented labour markets explain unexplained earnings differences for each race and sex group.

*Review of technical strategies to distinguish between theories*

Tests for differences between hypothesized market segments are 'exogenous,' sorting jobs according to preconceived characteristics of each sector and proceeding to test for sectoral mobility and differences in rates of return for productivity traits. Only two studies use econometric techniques to make sorting 'endogenous – letting the procedure define the two sectors.' These models have come to different conclusions (Dickens and Lang, 1985; Anderson *et al*., 1987). See Dickens and Lang (1992) and Boston (1990) for reviews of these approaches.

The older endogenous study's (Dickens and Lang, 1985) switching regression test specifies that there are two distinct wage-setting mechanisms. In one, firms pay for productivity characteristics of individuals and, in the other, firms generally do not. The technique assumes, as does human capital theory, that individuals want the highest-valued compensation package. The switching regression technique estimates a probability function that sorts individuals into two sectors. If the coefficients of the regressors are different in each sector, then the hypothesis that two segments exist is not rejected. The technique's major weaknesses are: (1) it assumes only two sectors exist; (2) all ability and preference characteristics are controlled for (Heckman and Holtz, 1986, p. 328; Dickens and Lang, 1992, p. 58); and (3) non-competitive institutions, such as unions, significantly affect earnings. (In subsequent studies, Dickens and Lang, 1988 argue that segmentation, not unionization, explains primary and secondary jobs.)

The second endogenous approach (Anderson *et al*., 1987) clusters jobs using eight factors in order to minimize 'within-cluster differences': (1) low wages and flat age-earnings profiles; (2) little on-the-job training; (3) little general training required; (4) low turnover; (5) low job satisfaction; (6) high absenteeism; (7) more layoffs; and (8) physically riskier jobs. They find no evidence of clustering along these dimensions that would have been in accordance with the predictions of segmentation theory. However, their criteria ignore a major principle in

labour economics. The presence of adverse job characteristics – such as risk, being unsatisfying, or unstable – do not necessarily identify a 'secondary' job.

The (inexplicable) absence of compensating wage differentials can and does reveal the existence of non-competing groups. In other words, low-paying jobs are not necessarily secondary sector jobs if the level of required skill and education is low. Jobs that do not pay for skills, education, and disamenities are secondary jobs. Jobs that require few overt skills and do not reward for loyalty or effort are also bad jobs. Any compensation equation that does not include a proxy for intangible, but important, productivity attributes – like the shuttle driver's attendance record – will underestimate the dualism. Blackaby *et al.* (1995) also suggest that dualism can be underestimated because the returns to the productivity of the unemployed are not included in analysis of all workers. They find that the unemployed are more likely to be in the secondary market.

Our definition does not differ from Osterman's (1975) and Piore's (1971) descriptions of secondary jobs, which emphasize that secondary jobs do not penalize for turnover and thus have 'no' future. These are jobs in which valuable skills, some intangible, like loyalty, attentiveness, cheerfulness, are neither accounted for nor compensated. In our framework, jobs that do not pay a return for education and do not pay for disamenities are secondary jobs.

## III Methodology and Econometric Model

Showing that dual labour market theory can reasonably explain the data does not falsify another theory, such as human capital theory (Dickens and Lang, 1992, p. 2–6). We examine whether dual labour market theory – the theory that there is a queue of qualified individuals for primary sector jobs – is a reasonable explanation for a more completely and accurately described labour market, a labour market that provides compensation in wages, pensions, and health insurance and that includes both men and women.[1]

We apply a switching regression using the recent Current Population Survey and correcting for heteroskedasticity. Next, we expand the dependent variable to include total compensation. Finally, we examine the theory for its relevance to female labour markets.

To review, the dual labour market hypothesis proffers that there are, for simplicity, two persistently non-competing markets. The best jobs are rationed so that all qualified workers, who maximize the net present value of their total compensation, cannot obtain these primary jobs.

[1] Some primary jobs can be devalued over time and become secondary sector jobs. The labour market changes in meatpacking are one example of these transformations. In the 1970s, hourly pay in meatpacking was higher than in most industries and higher than the community average. As a result of changes in industrial organization, product, technology, immigration, and global competition, hourly jobs in meatpacking pay fell below similar jobs in most industries although the relative level of danger, training, and education required remained the same. Therefore, in the 1990s, meatpacking jobs were relatively lower-paid factory jobs, more disagreeable and dangerous, and held by immigrant (often illegal) workers. They had also become mostly non-union. Pittsburgh steel industry jobs demonstrated a similar dramatic transformation between 1900 and the post-WWII period. However, the transformation was in the other direction – secondary jobs became primary jobs.

We will assume that the labour market is segmented into two markets: primary and secondary. Each segmented market has its own wage equation, and certain worker characteristics determine which sector the individual is in. We do not assume anything about which occupations or industries are in the primary and secondary sectors and we do not determine what worker characteristics determine which sector a worker is in.

The hypothesized wage equation for each market is

$$LnW_{p,i} = X_i\beta_{p,i} + \varepsilon_{p,i}, \tag{1}$$

$$LnW_{s,i} = X_i\beta_{s,i} + \varepsilon_{s,i}. \tag{2}$$

The third equation determines each worker endogenously to either the primary market or the secondary market:

$$Y_i^* = Z_i\gamma + \varepsilon_{w,i}, \tag{3}$$

where $Y_i^*$ is the unobservable; the worker belongs to the primary market if $Z_i\gamma + \varepsilon_{w,i} \geq 0$, and to the secondary market, otherwise.

Therefore, a worker is either in one or the other sector depending on the parameters.

The log-likelihood function is derived from the probability function. The likelihood function is given by

$$\Pr(\varepsilon_{w,i} > -Z_i\gamma | Z_i, X_i, \varepsilon_{p,i}) f(\varepsilon_{p,i}) + \Pr(\varepsilon_{w,i} \leq -Z_i\gamma | Z_i, X_i, \varepsilon_{s,i}) f(\varepsilon_{s,i}), \tag{4}$$

where $f(\cdot)$ is the density function of the error $\varepsilon_p$ or $\varepsilon_s$.

If we assume that $\varepsilon_p$, $\varepsilon_s$, and $\varepsilon_w$ are normally distributed, the log-likelihood function is therefore

$$\sum_{i=1}^{N} \ln\left\{ \left[ 1 - \Phi\left( \frac{-Z_i\gamma - (\sigma_{pw}/\sigma_{pp})\varepsilon_{p,i}}{\left(1 - \sigma_{pw}^2/\sigma_{pp}\right)^{0.5}} \right) \right] \phi(\varepsilon_{p,i}, \sigma_{pp}) + \Phi\left( \frac{-Z_i\gamma - (\sigma_{sw}/\sigma_{ss})\varepsilon_{s,i}}{\left(1 - \sigma_{sw}^2/\sigma_{ss}\right)^{0.5}} \right) \phi(\varepsilon_{s,i}, \sigma_{ss}) \right\}, \tag{5}$$

where $\phi(\cdot)$ is the normal density function and $\Phi(\cdot)$ is the cumulative distribution function. $\sigma_{pw}$ is the covariance of $\varepsilon_{p,i}$ and $\varepsilon_{w,i}\sigma_{ww}$ is normalized to equal one. Maximum likelihood estimates are obtained using the standard search algorithms.

A switching regression is an econometric technique that makes sorting workers into sectors endogenous. The procedure determines which sector the worker is most likely to be in using maximum likelihood techniques. The switching regression model that we adopted is one extension of the basic switching regression model. The basic model has been extended to the situation in which switching is related to a given vector of explanatory variables, $Z_i$. According to the switching model theory, this $Z_i$ in the switching equation may or may not be equal to a vector of explanatory variables, $X_i$ in the primary and

secondary equations. The worker belongs to the primary market if the value of $Z_i$ exceeds a certain quantity.

This endogenous approach is better than the exogenous approach since we do not predetermine what worker and employer characteristics are most likely to be associated with either sector.

We predict that the sector that workers are employed in greatly affect how they will be compensated in terms of cash income and access to health insurance and retirement plan coverage. Therefore, we examine how well the standard model explains the determinants of cash income and total compensation and compare the results with how well a switching regression explains compensation in primary and secondary sectors.

## IV Data

Using a March supplement of the 2000 Current Population Survey, our original sample size of 133,710 is reduced to 12,747, because some explanatory variables had many missing values.

The independent variable denoting the highest level of education attained is divided into three groups: less than high school, less than college, and college graduates and over. The variable indicating race denotes the individual as white and non-white. The 12,747 workers are, on average, 39.47 years old; 48% are women, 15% are non-white; 11% have less than a high school education, while 28% have some college or a college degree. Their average earnings are $630 per week. Fifty-eight percent have health insurance provided by their employer and 62% are covered by a pension plan. Twenty-five percent have never been married.

The men and women are similar in our sample except that men have higher pay and more health insurance, are less likely to be non-white and have high school degrees, and are more likely to have never been married. Table A1 in the Appendix display the means and ranges for the variables used in the regressions.

To explain wage levels we use a dummy variable for marital status, which equals one if the worker was ever married and zero otherwise. This variable controls for many hypothesized factors affecting pay (ranging from the individual's 'quality' to discrimination against single males). Race and age (a rough proxy for experience) are also used as controls. The dual labour market theory hypothesizes that race and sex determine access to the secondary market, but do not significantly affect compensation in the segment.

Dual labour market theory suggests that both experience and education are rewarded in the primary sector, but not in the secondary sector. Therefore, the results on the education and experience variables are our main interest. Dickens and Lang (1985, 1988, 1992) use 'age minus 5' to indicate experience and 'years of schooling' in the same regression with education. We use 'age' as a proxy for experience to avoid this multicollinearity, and because no better experience variable exists in the CPS. The Panel Study of Income Dynamics (PSID) has more experience-related variables, but the larger CPS sample size helps obtain

more reliable estimation results. The PSID, in 2001, samples approximately 7000 individuals.

The available data restrict us to measuring the sectors along the education dimension and requires the education variable to proxy for all skills. The skills rewarded at a firm are very firm specific and, sometimes, so idiosyncratic that most data sets – the CPS included – do not measure them. Formal schooling, which education attainment measures, has become a standard and acceptable proxy for skill as is age and age squared for potential experience.

We impute the value of the employee benefits using employers' cost of an employee benefit as a proxy for the benefit's value to the worker. We discuss pensions first and then health insurance. The average pension benefit cost for employers, in 2000, was 60 cents per hour, or $1100 on average, per year (Employment Cost Index, 2000). The CPS records whether workers are covered by a pension plan and what type of coverage is available; therefore, we impute the average cost of coverage per worker that year. We use the average figure in order not to bias our results towards dualism and have assumed that pension coverage is equally expensive. This assumption will bias the results away from finding compensation inequality and the different rates of coverage will be captured in the results. Since workers report the amount their employers pay for health insurance, the calculation is much more straightforward.

Approximately two-thirds of both men and women are covered by a pension plan: men's pension coverage rate is 62.25% and women's coverage rate is 62.29%. Men are more likely to have health insurance; their coverage rate is 63.51%, while women's coverage rate is 52.52%. We use the actual employer contribution of the health insurance reported in CPS to obtain the amount employers spend on health insurance. Table A2 in the Appendix describes the variables.

## V Results

We find that the dual labour hypothesis was supported for both dependent variables, wages and total compensation, and when men and women were combined in one sample and when men and women were analysed separately. The switching equation technique assigns each worker, endogenously, to either the primary market or the secondary market and estimates the determinants of pay in each market; therefore, we estimate all three equations simultaneously. A likelihood ratio test supports the hypothesis that two equations fit the data better than one equation. The likelihood ratio statistic for the sample including men and women and for earnings as the dependent variable is 508.698; for the expanded earnings-dependent variable, it is 514.816. For the men's subsample, using cash earnings and total compensation, the likelihood ratio statistics are 260.928 and 263.578, respectively. Similarly, for women, they are 313.278 and 228.59. The critical values were far smaller than our computed likelihood ratio test statistic, which we interpret as a rejection of the single labour market (OLS) model. The results support the dual labour market theory.

Table 1
*Earnings determinates assuming one market and dual markets: OLS and switching for both men and women (CPS)*

| | | Switching model | | |
|---|---|---|---|---|
| | One market: Ln weekly wage | Ln weekly wage primary | Ln weekly wage secondary | Switching |
| Intercept | 4.2874* | 4.2685* | 5.4506* | − 1.7869* |
| | (0.0529) | (0.0630) | (0.4182) | (0.4322) |
| Some college | 0.3530* | 0.3768* | − 0.1865 | 0.9683* |
| | (0.0138) | (0.0199) | (0.3799) | (0.4053) |
| College degree | 0.8030* | 0.9055* | − 0.1865 | 1.3070* |
| | (0.0162) | (0.0234) | (0.3799) | (0.4160) |
| Female | − 0.3576* | − 0.2514* | − 0.5749* | − 1.4376* |
| | (− 0.0095) | (0.0227) | (0.0459) | (0.1648) |
| Non-white | − 0.0611* | − 0.0337* | − 0.0987 | − 1.1511* |
| | (− 0.1293) | (0.0167) | (0.0830) | (0.3083) |
| Never married | − 0.0899* | − 0.0303** | − 0.2929* | − 0.6737* |
| | (0.0130) | (0.0166) | (0.0542) | (0.1896) |
| Age | 0.0823* | 0.0758* | 0.0909* | 0.0216* |
| | (0.0027) | (0.0033) | (0.0086) | (0.0036) |
| Age squared | − 0.0009* | − 0.0008* | − 0.0011* | |
| | (− 0.0000) | (0.0000) | (0.0000) | |
| *N* | 12,713 | 12,713 | 12,713 | 12,713 |
| $R^2$ | 0.3577 | | | |
| Log likelihood | − 10098.703 | | − 9844.354 | |

*Notes:*
*Statistically significant at 5%; **statistically significant at 10%.
One market regression is OLS with heteroskedasticity consistent standard errors.

We display the results of the 'one market' OLS regressions and the 'dual market' switching regressions for both men and women in Table 1 (the regressions with the earnings-dependent variable) and Table 2 (the regressions with total compensation). We used two dependent variables for all the regressions: cash earnings and total compensation (cash earnings plus the average employer contribution for employer cost of health insurance and pensions if a person is covered). Table 2 displays the results for the total compensation-dependent variable for the sample that combines men and women.

The results in both regressions for the combined sample support the dual labour market hypothesis. Education has a positive rate of return only in the primary sector, as expected, and not in the secondary sector. This is also true in the regression when total compensation is the dependent variable, with one difference – the small, but positive and weakly significant coefficients on the education attainment measures in the secondary sector (this is caused by different effects for men and women described below) Including fringe benefits seems to make a difference. In addition to finding positive and significant coefficients for education, experience is also rewarded, but if you are a man, never have been married lowers pay, as does being a woman or/and non-white.

Table 2
*Expanded earnings determinates assuming one market and dual markets: OLS and switching for both men and women (CPS)*

| | | Switching model | | |
|---|---|---|---|---|
| | One market: Ln weekly wage+fringes | LN weekly wage+fringes primary | Ln weekly wage+fringes secondary | Switching |
| Intercept | 4.3097* | 4.2454* | 3.6492* | − 0.8733* |
| | (0.0523) | (0.0663) | (0.2136) | (0.2285) |
| Some college | 0.3643* | 0.4037* | 0.3568* | 0.5671* |
| | (0.0139) | (0.0213) | (0.0722) | (0.1810) |
| College degree | 0.8028* | 0.9221* | 0.6446* | 0.5844* |
| | (0.0162) | (0.0246) | (0.0865) | (0.2052) |
| Female | − 0.3567* | − 0.2198* | − 1.1106* | − 0.9241* |
| | (0.0094) | (0.0246) | (0.1092) | (0.1466) |
| Non-white | − 0.0589* | − 0.0143 | − 0.6453* | − 0.8404* |
| | (01280) | (0.0167) | (0.1754) | (0.2083) |
| Never married | − 0.1018* | − 0.0221 | − 0.6724* | − 0.6295* |
| | (0.1290) | (0.0184) | (0.1447) | (0.1786) |
| Age | 0.0843* | 0.0758* | 0.0998* | 0.0121* |
| | (0.0027) | (0.0033) | (0.0073) | (0.0034) |
| Age squared | − 0.0009* | − 0.0008* | − 0.0011* | |
| | (0.0000) | (0.0000) | (0.0001) | |
| $N$ | 12,713 | 12,713 | 12,713 | 12,713 |
| $R^2$ | 0.3685 | | | |
| Log likelihood | − 9931.846 | | − 9674.438 | |

*Notes:*
*Statistically significant at 5%.

### *Male labour markets*

We conclude that male labour markets in 2000 are segmented in much the same way they were in 1970. A likelihood ratio test shows that the 'one market' hypothesis can be rejected at any conventional level. Having more education and being older are associated with higher wages and total compensation only in the primary market, and have no effect in the secondary sector, results that support the segmentation hypothesis. Tables 3 and 4 displays results for men's cash compensation and total compensation regressions, respectively, and indicate that a college degree (compared with having a high school degree) in the primary sector raises men's cash compensation by 75.41% and men's total compensation by about the same amount, 75.33%. Age has a positive effect on both types of compensation in the primary and secondary sector for men, which is not what segmentation theory predicts if age is a good proxy for experience and higher skills. But the effects on compensation are rather small at approximately 10%.

Being non-white and never having been married lowers men's cash and total compensation by approximately 13–17%. (There is no such penalty in the sector in the secondary sector. Having never been married could proxy for 'non-marriageability' that could be a proxy for low on-the-job productivity because of

Table 3
*Earnings determinates assuming one market and dual markets: OLS and switching for men*

| | | Men switching model | | |
|---|---|---|---|---|
| | Men: one market: Ln weekly wage | Ln weekly wage primary | Ln weekly wage secondary | Switching |
| Intercept | 4.1300* | 3.6175* | 5.0353* | − 5.4659* |
| | (0.0743) | (0.1227) | (0.5356) | (0.5302) |
| Some college | 0.3380* | 0.3137v* | 0.0265 | 0.6882* |
| | (0.0187) | (0.0227) | (0.2328) | (0.2777) |
| College degree | 0.7541* | 0.7533* | 0.1765 | 1.3008* |
| | (0.0222) | (0.0300) | (0.2623) | (0.2883) |
| Non-white | − 0.1296* | − 0.1010* | − 0.1346 | − 0.3425* |
| | (0.0189) | (0.0231) | (0.0966) | (0.1698) |
| Never married | − 0.1741* | − 0.1395* | − 0.4092* | 0.3560 |
| | (0.0181) | (0.0208) | (0.0818) | (0.2255) |
| Age | 0.0928* | 0.1279v* | 0.1096* | 0.0923* |
| | (0.0038) | (0.0078) | (0.0164) | (0.0101) |
| Age squared | − 0.0010* | − 0.0016* | − 0.0014* | |
| | (0.0000) | (0.0001) | (0.0002) | |
| $N$ | 6600 | 6600 | 6600 | 6600 |
| $R^2$ | 0.3556 | | | |
| Log likelihood | − 5188.836 | | − 5058.372 | |

*Notes:*
*Statistically significant at 5%.

prison time and irresponsibility, etc.) Our results are consistent with previous research that shows that non-white men face discrimination in the primary market, but not in the secondary market, because race operates more as a sector locator. On balance, these results confirm the dual labour market hypothesis.

### *Female labour markets*

This is the first study that finds support for the existence of non-competing markets for women because times have changed and because we include employee benefits in the measure of compensation. In the 1980s, when the older studies were conducted (Dickens and Lang, 1985; Friedberg *et al*., 1988), over 78% of female workers were in service, clerical, and technical occupations. By the late 1990s, women workers are employed in many more occupations; less than 58% of women are in service, clerical, and technical occupations (US Statistical Abstract, various years) as more women integrate into male occupations. The question is whether or not their markets are acting more like male labour markets.

Just looking at the results of the cash compensation, women like men, have no rates of return for education. Yet, unlike men, women obtain some returns in total compensation for education in the secondary sector. Having a college degree, rather than no college education, raises women's total compensation (wages and employee benefits) by 41.3% in the secondary sector (there was no

Table 4
*Expanded earnings determinates assuming one market and dual markets: OLS and switching for men*

| | | Men switching model | | |
|---|---|---|---|---|
| | Men: one market: Ln weekly wage+fringes | Ln weekly wage+fringes primary | Ln weekly wage+fringes secondary | Switching |
| Intercept | 4.0259* | 3.5828* | 4.8302* | − 5.4304* |
| | (0.0735) | (0.1199) | (0.4541) | (0.5349) |
| Some college | 0.3503* | 0.3292* | 0.1494 | 0.5695* |
| | (0.0188) | (0.0227) | (0.1474) | (0.2226) |
| College degree | 0.7569* | 0.7540* | 0.3668 | 1.0805* |
| | (0.0221) | (0.0288) | (0.1650) | (0.2328) |
| Non-white | − 0.1279* | − 0.0956* | − 0.1678* | − 0.2952** |
| | (0.0187) | (0.0229) | (0.0787) | (0.1619) |
| Never married | − 0.1911* | − 0.1562* | − 0.3981* | 0.4139** |
| | (0.0179) | (0.0206) | (0.0720) | (0.2336) |
| Age | 0.0940* | 0.1337* | 0.1081* | 0.0964* |
| | (0.0038) | (0.0076) | (0.0155) | (0.0100) |
| Age squared | − 0.0010* | − 0.0017* | − 0.0013* | |
| | (0.0000) | (0.0001) | (0.0002) | |
| *N* | 6600 | 6600 | 6600 | 6600 |
| $R^2$ | 0.3708 | | | |
| Log likelihood | − 5083.758 | | − 4951.969 | |

*Notes:*
*Statistically significant at 5%.

effect for men – the coefficients for a college degree for men in the cash and total compensation regressions in the secondary markets were insignificant; see Tables 3 and 4). We conclude that employee benefits in women's labour markets operate to segment the markets for women and also help women get paid more for their education in the secondary market.

Another contrasting result is that never having been married lowers men's cash and total compensation. Having never been married does not affect women's cash compensation in either the primary or secondary markets, but is an advantage for women in the secondary sector when total compensation is considered. In fact, never married women's total compensation is 30.8% higher than married, divorced, or widowed women. We can speculate that never married women behave in such a way that makes them more likely to obtain health and pension benefits, for instance working more hours in the secondary sector. Married women may forgo health insurance because their husband's employer offers dependent coverage.

Race discrimination seems to have different effects on women's labour market outcomes than it does for men. Although it may seem odd that being a non-white woman does not adversely affect cash compensation, these women do have lower total compensation, 7.6% less, as one would expect, in the primary sector (see Tables 5 and 6). The secondary sector results are noteworthy as well.

Table 5
*Earnings determinates assuming one market and dual markets: OLS and switching for women*

| | | Women: switching model | | |
|---|---|---|---|---|
| | Women: one market: Ln weekly wage | Ln weekly wage primary | Ln weekly wage secondary | Switching |
| Intercept | 4.0796* | 4.0558* | − 0.5639 | 2.8301* |
| | (0.0723) | (0.0782) | (1.6800) | (1.0540) |
| Some college | 0.3717* | 0.4064* | 1.7870 | − 1.1874 |
| | (0.2000) | (0.0334) | (1.4998) | (0.9892) |
| College degree | 0.8544* | 0.9435* | 1.6803 | − 1.0431 |
| | (0.0232) | (0.0310) | (1.5172) | (0.9925) |
| Non-white | − 0.0027 | − 0.0240 | − 0.8726* | 0.5344* |
| | (0.0175) | (0.0229) | (0.3610) | (0.2080) |
| Never married | 0.0070 | − 0.0155 | 0.2210 | − 0.0561 |
| | (0.0183) | (0.0193) | (0.3182) | (0.1778) |
| Age | 0.0718* | 0.0732* | 0.0774* | − 0.0182* |
| | (0.0037) | (0.0040) | (0.0187) | (0.0052) |
| Age squared | − 0.0008* | − 0.0008* | − 0.0005* | |
| | (0.0000) | (0.0000) | (0.0002) | |
| *N* | 6113 | 6113 | 6113 | 6113 |
| $R^2$ | 0.2849 | | | |
| Log likelihood | − 4830.107 | | − 4673.468 | |

*Notes:*
*Statistically significant at 5%.

Non-white women receive higher total compensation in the secondary sector (also in Table 6), suggesting either race discrimination does not exist or black women work longer hours than whites. Race discrimination could lower wages even though black women's cash earnings and total compensation are higher, because full-time work increases the probability of qualifying for employee benefit coverage. As discussed above, in the primary market – where salary is the major form of compensation and differences in hours worked do not matter as much – the coefficient on being non-white is negative and significant, suggesting, on balance, the results of all four regressions suggest that race discrimination exists against non-white women in the primary and secondary sectors, although employee benefits help close the gap in the secondary sector.

One of the most interesting outcomes of the switching regression is that fewer people are likely to be in the secondary market when total compensation is considered rather than cash compensation. As we explained above, the technique assigns workers to each sector during the regression. When we did not divide the sample by sex, 22.69% of workers were assigned to the primary market when cash wage was the dependent variable. When we used total compensation more workers, 28.47%, were assigned to the primary sector, meaning workers were more likely to be paid for their education and experience. When we consider only men and wages, only 21.74% were assigned to the primary sector, but when total compensation is used more were assigned,

Table 6
*Expanded earnings determinates assuming one market and dual markets: OLS and switching for women (CPS)*

| | | Women: switching model | | |
|---|---|---|---|---|
| | Women: one market: Ln weekly wage +fringes | Ln weekly wage+fringes primary | Ln weekly wage+fringes secondary | Switching |
| Intercept | 4.0839* | 3.6149* | 6.0091* | − 1.1603* |
| | (0.0712) | (0.1526) | (0.3674) | (0.3865) |
| Some college | 0.3817* | 0.3731* | 0.1725 | 0.4791** |
| | (0.0202) | (0.0422) | (0.1612) | (0.2731) |
| College degree | 0.8506* | 0.9535* | 0.4131* | 0.8008* |
| | (0.0233) | (0.0623) | (0.1810) | (0.2658) |
| Non-white | − 0.0005 | − 0.0765* | 0.1601** | − 0.1464 |
| | (0.0173) | (0.0370) | (0.0897) | (0.1433) |
| Never married | 0.0009 | − 0.0637 | 0.3082* | − 0.4217* |
| | (0.0181) | (0.0423) | (0.1087) | (0.1592) |
| Age | 0.0748* | 0.1064* | 0.0205 | 0.0150* |
| | (0.0038) | (0.0085) | (0.0129) | (0.0051) |
| Age squared | − 0.0008* | − 0.0012* | − 0.0002 | |
| | (0.0000) | (0.0001) | (0.0002) | |
| *N* | 6113 | 6113 | 6113 | 6113 |
| $R^2$ | 0.2927 | | | |
| Log likelihood | − 4764.293 | | − 4649.998 | |

*Notes:*
*Statistically significant at 5%; **statistically significant at 10%.

23.72%, to the primary sector. The exercise for women yields intriguing results; when cash was used as the dependent variable almost all women, 99.73%, were assigned to one sector, which implies the dual labour market hypothesis does not explain women's labour markets and confirms Friedberg, Dickens, and Lang. But when we consider pension and health insurance coverage as a form of compensation, we find that most women are in the secondary sector, 61.99%, and 38.01% were assigned to the primary sector.

## VI Conclusions

We find, in contrast to Friedberg *et al.* (1988) that the dual labour market hypothesis does explain women's labour markets. Although it may seem that that our results could be explained by the considerable changes in women's labour market experiences over the past 20 years, we also find weak support for the dual labour market hypothesis for women when only wages are used as a measure of compensation. In contrast, we find strong support for the hypothesis when cash income and health and pensions are considered in women's labour markets.

In summary, modifications of human capital theory have reconciled observations by studies of segmented labour market theories. Yet, there are

still sharp distinctions between the two frameworks for understanding labour markets, which produces sharply different policy responses. Human capital theorists presume that, over time, workers' productivity choices determine their returns to work; therefore, policies that enhance workers' decision to be more productive will solve problems of low pay and the appearance of race and sex discrimination. On the other hand, segmentationists hypothesize that 'the existence of sectors with distinct wage-setting mechanisms' (Dickens and Lang, 1985, p. 796) is rooted in the dynamics of market economies so that inter-industry and inter-firm differentials will persist, causing stubborn, involuntary unemployment. Policies, thus oriented, focus on improving secondary jobs with minimum wage laws, activist training programmes, and dedicated enforcement of employment rights and collective bargaining. In short, human capital theorists call for improving the people; segmentation theorists call for improving the jobs.

Further explorations into the complex factors affecting women's labour markets are necessary, as well as identification of the occupations and industries that are more likely to have secondary jobs. There are over two-thirds of workers who are in sectors that do not pay expected rates of return for education and experience attainment. Policies aimed at creating internal labour markets and other mechanisms that reward skill and other productivity characteristics in retail, and other fast-growing, low-paid industries, may help boost earnings, encourage skill formation, and promote employee benefit coverage in the US

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## Appendix

Table A1
*Means of variables used in regressions*

| Variable | Observed | Mean | Std. dev. | Min | Max |
|---|---|---|---|---|---|
| Age | 12,713 | 39.4744 | 12.32787 | 15 | 75 |
| Education | 12,747 | 40.26642 | 2.576238 | 31 | 46 |
| Sex | 12,747 | 1.480505 | 0.499639 | 1 | 2 |
| Health insurance paid by employer (yes, no) | 12,747 | 0.582333 | 0.493194 | 0 | 1 |
| Employer's annual health insurance contribution | 12,747 | $1624.48 | $1791.71 | 0 | $9999 |
| Pension annual contribution (imputed) | 12,747 | $684.75 | $533.26 | 0 | $1100 |
| Pension plan covered (no, do not know = 0) | 12,747 | 0.6225 | 0.540616 | 0 | 1 |
| Female | 12,747 | 0.480505 | 0.499639 | 0 | 1 |
| Non-white | 12,747 | 0.150702 | 0.357772 | 0 | 1 |
| Less than high school | 12,747 | 0.113517 | 0.317236 | 0 | 1 |
| Less than college | 12,747 | 0.606417 | 0.488563 | 0 | 1 |
| College degree | 12,747 | 0.280066 | 0.449049 | 0 | 1 |
| Never married (1, 0) | 12,747 | 0.254256 | 0.435459 | 0 | 1 |
| Weekly wage | 12,747 | $630.56 | $455.49 | $100 | 2884.61 |
| Ln weekly wage | 12,747 | $6.23 | $0.67 | $4.61 | 7.97 |
| Total compensation weekly | 12,747 | $674.96 | $475.82 | $100 | $3098.05 |
| Ln total compensation weekly | 12,747 | $6.30 | $0.67 | $4.61 | $8.04 |
| *(A) Men* | | | | | |
| Age | 6600 | 39.31636 | 12.36135 | 15 | 75 |
| Education | 6622 | 40.13787 | 2.709215 | 31 | 46 |
| Sex | 6622 | 1.000000 | 0 | 1 | 1 |
| Health insurance paid by employer (yes, no) | 6622 | 0.635156 | 0.481423 | 0 | 1 |
| Employer's annual health insurance contribution | 6622 | $1979.60 | $1968.67 | 0 | $9999 |
| Pension annual contribution (imputed) | 6622 | $684.39 | $533.37 | 0 | $1100 |
| Pension plan covered (no, do not know = 0) | 6622 | 0.622169 | 0.484882 | 0 | 1 |
| Non-white | 6622 | 0.137119 | 0.344000 | 0 | 1 |
| Less than high school | 6622 | 0.132588 | 0.339155 | 0 | 1 |
| Less than college | 6622 | 0.590154 | 0.491842 | 0 | 1 |
| College degree | 6622 | 0.277258 | 0.447679 | 0 | 1 |
| Never married (1, 0) | 6622 | 0.265630 | 0.441701 | 0 | 1 |
| Weekly wage* | 6622 | $734.49 | $509.26 | $100 | $2884.61 |
| Ln weekly wage | 6622 | $6.39 | $0.66 | $4.61 | $7.97 |
| Total compensation weekly | 6622 | $785.72 | $531.80 | $100 | $3098.05 |
| Ln total compensation weekly | 6622 | $6.46 | $0.66 | $4.61 | $8.04 |
| *(B) Women* | | | | | |
| Age | 6113 | 39.64502 | 12.29035 | 15 | 75 |
| Education | 6125 | 40.40539 | 2.416821 | 31 | 46 |

Table A1 (Contd.)

| | | | | | |
|---|---|---|---|---|---|
| Sex | 6125 | 1.000000 | 0 | 1 | 1 |
| Health insurance paid by employer (yes, no) | 6125 | 0.525225 | 0.499404 | 0 | 1 |
| Employer's annual health insurance contribution | 6125 | $1240.55 | $1485.77 | 0 | $9999 |
| Pension annual contribution (imputed) | 6125 | $685.14 | $533.188 | 0 | $1100 |
| Pension plan covered (no, do not know = 0) | 6125 | 0.622857 | 0.484711 | 0 | 1 |
| Non-white | 6125 | 0.165388 | 0.371561 | 0 | 1 |
| Less than high school | 6125 | 0.092899 | 0.290313 | 0 | 1 |
| Less than college | 6125 | 0.624000 | 0.484420 | 0 | 1 |
| College degree | 6125 | 0.283102 | 0.450542 | 0 | 1 |
| Never married (1, 0) | 6125 | 0.241960 | 0.428305 | 0 | 1 |
| Weekly wage | 6125 | $518.18 | $356.51 | $100 | $2884.61 |
| Ln weekly wage | 6125 | $6.052 | $0.63 | $4.61 | $7.97 |
| Total compensation weekly | 6125 | $555.21 | $371.28 | $100 | $3098.05 |
| Ln total compensation weekly | 6125 | $6.13 | $0.62 | $4.61 | $8.04 |

Table A2
*Source for Variables*

| | | CPS variable used | | |
|---|---|---|---|---|
| Use in regressions | Variable in regressions | CPS name | CPS ID | CPS data set |
| Expanded earnings definition for dependent variable #2 | Pension paid by employer (Y/N) | PENPLAN | 3005 | March supplement |
| Dependent variable #2 | Imputed value of $1100 if the employer paid if the answer was yes to PENPLAN | | | March supplement |
| Dependent variable #2 | Health insurance contribution by employer | EMCONTRB | 1424 | March supplement |
| Dependent variable #2 | (Health insurance paid by employer) Y/N used to confirm above | HIPAID | 4791 | March supplement |
| Dependent 1 | Earnings weekly amount | | | |
| | PTERNWA | 11472 | March supplement | |
| Independent | Age (used for experience) | AGE | 2568 | March supplement |
| Independent | Female | SEX | 2605 | March supplement |
| Independent | Non-white | RACE | 6038 | March supplement |
| Independent | Schooling | PEEDUCA | 11206 | March supplement |

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