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Local Labour Markets, Migration and Wage Determination: Theory and Evidence for the Wage Curve in New Zealand

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LOCAL LABOUR MARKETS, MIGRATION AND WAGE DETERMINATION: THEORY AND EVIDENCE FOR THE WAGE CURVE IN NEW ZEALAND*

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Abstract

Blanchflower and Oswald (1994) claim that there is a stable inverse relationship between the wages paid to workers and the local unemployment rate. Using micro data from a range of countries they find an unemployment elasticity of pay of around –0.1. We use a dynamic efficiency model based on Phelps (1994) to produce an upward-sloping locus of steady-state wage and employment rate outcomes. In our model, points along this locus are identified by local labour market shocks. In the long run, inter-regional equilibrium ensures that regions lie on a so-called Harris-Todaro condition, which is consistent with no net internal migration. Using a synthetic micro sample of 20,302 observations from the 1986, 1991 and 1996 New Zealand Censuses of Population and Dwellings, we then find support for this specification of the earnings equation. As predicted, the coefficient on the employment rate is positive, while the gross migration rate has a positive effect on wages. In contrast, Blanchflower and Oswald's specification is not robust to controlling for the possible presence of simultaneity bias, due to the endogeneity of unemployment. Separate gender regressions reveal that males, but not females, exhibit a positive employment elasticity of pay, possibly due to the greater labour supply elasticity of the latter group. Evidence that the wages of less geographically mobile groups of workers are also more responsive to changes in the local unemployment rate is found when occupations of different skill levels are analysed. Finally, we use a procedure developed by Lang and Gottschalk (1996) to show that our data set has permitted a 41% gain in efficiency of the estimation of the employment elasticity of pay over previous New Zealand studies that used aggregate data.

Keywords: earnings functions, unemployment, local labour markets, bargaining

JEL Classification Numbers: E24, J31, R23

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LOCAL LABOUR MARKETS, MIGRATION AND WAGE DETERMINATION: THEORY AND EVIDENCE FOR THE WAGE CURVE IN NEW ZEALAND

1. Introduction

During the 1990s there has been a rapidly expanding body of research on the relationship between wages and unemployment in local labour markets. This research is one response to the general challenge that modern microeconomic foundations for macroeconomic theories would need to stand up to scrutiny with emerging rich microeconomic data sets in order to be convincing. Blanchflower and Oswald (1990), using U.S. and British micro data, found evidence for an inverse relationship between the level of pay of individuals and the local unemployment rate and labelled this relationship "the wage curve". Subsequently (1994), they reported on additional evidence for the wage curve in their 1994 book using data on individuals from a wide range of countries.[1] The robustness of their finding has been confirmed by other investigators using similar data (Blackaby and Hunt 1992; Winter-Ebmer 1996; Bratsberg and Turunen 1996; Kennedy and Borland 1997; Janssens and Konings 1998; Baltagi and Blien 1998). Several time-series studies (Johansen 1997; Chiarini and Piselli 1997) also suggest a long-run inverse relationship between the wage level and unemployment.

Particularly striking in this research is Blanchflower and Oswald's (1994) finding that the elasticity of the responsiveness of pay to the local unemployment rate appears to be very similar across countries and time periods, namely about -0.1. Elasticities obtained by others have generally been of a similar magnitude, although significant estimates have been found in the range of -0.3 and -0.05 (see, for example, Table 4 in Card's 1995 review of the evidence). It is evidence of this nature that led Blanchflower and Oswald to conclude that "Every country seems to have a "wage curve"" (1994, p. 12).

Some empirical studies reject this conclusion, but they form a small minority. For example, Albaek *et al.* (1999) find no stable negative relation between wages and unemployment across regions in the Nordic labour markets once regional fixed effects are accounted for. Partridge and Rickman (1997) found evidence of an upward sloping

[1] Specifically, they estimated wage curves with data from 12 countries: USA, Britain, Canada, South Korea, Austria, Italy, Holland, Switzerland, Norway, Ireland, Australia and Germany.

wage curve, which is consistent with the view that higher wages compensate for expected higher unemployment, as advocated during the 1970s by Hall (1970, 1972), Harris and Todaro (1970) and Reza (1978), among others. In any case, the unemployment elasticity of pay may vary across different groups. For example, Baltagi and Blien (1998) found with German data that the wage curve is more elastic for unskilled workers, for younger workers and for males. Janssens and Konings (1998) find a wage curve for males but none for females in Belgium, whereas Kennedy and Borland (1998) reported that in Australia female earnings were more responsive to the unemployment rate than male earnings.

Accepting, as Card (1995) does in his review of the literature, that the wage curve may be close to an "empirical law of economics" (p. 798), the question of what causes this law remains nonetheless unsettled. Blanchflower and Oswald (1994) posit three different theories that would each predict an inverse relationship between wages and the unemployment rate. They are: a model of regionally-based implicit contracts, an efficiency wage model and a bargaining model. Card (1995) argues that the second of these theories, the efficiency wage model, is the most plausible. Despite Blanchflower and Oswald's (1994) attempt to list a range of testable implications from each of the theories (pp. 95-97), the available evidence to date may be consistent with more than one theory. For example, the higher wage curve elasticity commonly observed for "weak" bargaining groups, such as the relatively young or non-union workers, is consistent with both a bargaining model and an efficiency wage model. The latter model would point to the lesser opportunities these groups have to enhance pay through interregional migration, because they suffer from financial or other constraints. In any case, there may be alternative theoretical explanations for the wage curve, in addition to the three posited by Blanchflower and Oswald.

To date, there has been relatively little research on the determinants of pay in local labour markets in New Zealand. Morrison (1997) found a remarkable diversity in labour market outcomes across New Zealand regions, but also a positive correlation - albeit weak - between the regional employment rate and average annual income of wage and salary earners in these regions. Given that employment and unemployment rates sum to one, Morrison's correlation is consistent with the wage curve. Subsequently, Morrison and Poot (1999) estimated a wage curve with 1996 cross-sectional data for 39 regions and four occupational groups (i.e. a total of 156 observations). The data related to male salary and wage earners employed full-time. They found a statistically significant elasticity between -0.12 and -0.07 across a range of specifications. These results are similar to those obtained for other countries.

Thickett and Poot (1999) estimated a wage curve based on an alternative data set derived from the 1981, 1986, 1991 and 1996 Censuses with one observation for each of 1081 area units, making a total of 4324 observations. The panel nature of this data set was exploited. Here the data related to all gainfully employed males. The wage curve elasticity was significant and varied between -0.181 and -0.024. However, besides a measure of the level of education of the male workers, there were few other determinants of wages available and the results may be sensitive to specification bias.

In the present paper we aim to improve on this earlier research by studying the relationship between wages and local unemployment at a considerably less aggregated level. New Zealand research to date has been hampered by the absence of the type of individual level data used in much overseas research. While analysis of unit record data is possible through the so-called Statistics Laboratory of Statistics New Zealand, the laboratory has until recently operated on a full cost recovery basis that made its use out of reach for many academic researchers. In this paper, we seek a balance between obtaining sufficiently rich data to undertake the type of panel data analysis carried out elsewhere and obtaining information at relatively low cost. Our compromise involved the purchase of a synthetic micro level sample generated from the three most recent population censuses (1986, 1991 and 1996). Each observation consists of a group of wage and salary earners at the local labour market level (30 urban areas), defined according to their employment status, ethnicity, gender, age, education and occupation. For each group, an estimate of median annual income and mean weekly hours worked was obtained. We ended up with 20,302 useful observations, with group sizes varying between 3 and 12,719. In addition, several local labour market characteristics were obtained for each urban area. Some of these were obtained from census data, some from other sources.

While the relative loss in efficiency resulting from having to estimate the wage curve from a synthetic micro sample rather than from individual unit records cannot be assessed, we can estimate the gain in efficiency from using our disaggregated data set as compared with the regional macro data of Morrison and Poot (1999) and Thickett and

Poot (1999). Based on a procedure suggested by Lang and Gottschalk (1996), we show that our study produces coefficient standard errors that are 41% lower than studies using aggregate New Zealand data.

A weakness of wage curve estimation in much of the literature has been the weak link between theory and empirical estimation. Most researchers estimate wage equations in the tradition of Mincer (1974) with the local unemployment rate added as an additional regressor. In this paper we develop an extension of a simple dynamic efficiency model of the wage curve proposed by Campbell and Orszag (1998) and maintain a tight link between theory and estimation. One consequence is that we allow for a varying unemployment elasticity of pay.

The next section sets out the theoretical model. Section 3 discusses the nature of the available data and provides some descriptive statistics. Estimates of the wage curve by means of our synthetic micro sample are discussed in Section 4. This is followed in Section 5 by attention to some special issues, with Section 6 summing up.

2. The theoretical model

As noted above, Blanchflower and Oswald (1994) offer three possible explanations for the wage curve. A negative relationship between unemployment and wages could, they argue, be supported by a labour contract model, an efficiency wage model or a bargaining model.

The labour contract model makes the crucial assumption that regions differ in amenity values but that the "outside option" which laid-off workers face (the unemployment benefit or the reservation wage) is equal across regions. Firms and workers agree on a state-contingent wage level and a state-contingent employment level along the lines of the standard Azariadis (1975) and Baily (1974) implicit contracts model. Higher wages will then coincide with a higher level of contractual employment to compensate for the higher income risk. Attractive regions will be bunched at outcomes characterised by low long-run wages with high long-run unemployment. However, as noted by Card (1995) and by Blanchflower and Oswald (1995), the empirical evidence is not consistent with some of the predictions of this theory.

The contracting model would not appear to be very plausible in the New Zealand context. Poot (1986) showed that climate is a location-fixed amenity to which internal migrants respond. They are attracted by the warmer weather in the north. However, the long-term "drift north" in New Zealand coincides with higher wages on average in the north, rather than in the south. The contract model would predict wages to be higher in the colder south, *ceteris paribus*.

A more promising alternative is a union bargaining model. This model, which originated with De Menil (1971), generates a wage equation of the form $w = a + s\, \pi/n$. Here w is the negotiated wage available to union workers, a is the expected "alternative" wage in the non-union sector, π/n is the level of profits per worker and s is a relative bargaining power parameter. Because a will decrease with increasing rates of unemployment, a wage curve results. Blanchflower and Oswald (1994) provide some supporting micro-level evidence for this theory. However, the wage curve appears less elastic for union workers than for non-union workers and the curve is also less elastic in highly unionised countries (Card 1995; Albaek *et al.* 1999). Both facts contradict the union bargaining model. Moreover, the bargaining model would seem inappropriate to the decentralised wage setting structure that has existed in New Zealand since the introduction of the Employment Contracts Act (1991).

The third wage curve theory builds on the efficiency wage model of Shapiro and Stiglitz (1984). Employers, who can imperfectly monitor workers' productivity, will offer a wage that will discourage workers from shirking. Because the expected penalty for shirking, when detected, is greater when it becomes harder to find a job, firms can offer a lower wage premium during times of high unemployment.

This shirking model has, as noted by Card (1995), various advantages over the two other models. Firstly, it suggests that a short-run inverse correlation between wages and unemployment rates is not inconsistent with a long-run positive cross-sectional association between expected regional wages and unemployment rates, as suggested by Harris and Todaro (1970). An additional advantage of this theory is that it leads to the testable hypothesis that a group-specific unemployment rate should be a better predictor of group-specific wages than the average regional unemployment rate. This hypothesis can be tested to the extent that group-specific regional unemployment rates can be observed. Thirdly, since the shirking model is likely to be more relevant in relatively non-unionised economies, the model predicts that a decline in unionisation should lead to a more elastic wage curve. As noted earlier, this is consistent with evidence reported in the literature.

In the remainder of this section a wage curve model is developed which we would expect to be appropriate in a New Zealand context. Given the discussion above, we concur with Card (1995) that the efficiency wage model is the "leading contender" for this empirical regularity. In contrast to the models of Shapiro and Stiglitz (1984) and Blanchflower and Oswald (1994), wherein firms attempt to minimise the costs attributable to shirking workers, we follow here Campbell and Orszag (1998) and adopt a turnover cost-type efficiency wage model based on Salop (1979) and Phelps (1994). In this model, firms in low unemployment regions economise on the costs associated with hiring new workers by paying higher wages in order to discourage existing workers from quitting. In addition, local authorities may engage in territorial competition by offering firms training subsidies, which are funded from a local tax on wage income.

Consider an economy with R regions, indexed by $j = 1, 2, \ldots, R$. It is assumed that there are a relatively large number of regions and that no region dominates the national economy. There are N identical firms in each region, indexed by $i = 1, 2, \ldots, N$, with N being a large number. All firms produce the same homogeneous good, sold at a unit price in all markets. Following Oi (1962), labour is considered a quasi-fixed factor of production, as the recruitment of new employees involves one-off training costs. Firms face a level of staff turnover that depends on the wage they offer relative to the alternative income workers have available to them. Firms behave in a Nash fashion and treat economy-wide average wages and employment levels as given. Given the assumption of identical firms, wages and employment are equated across firms within regions. However, due to location-fixed amenities and imperfect geographical factor mobility or trade, wages may vary between regions.

Firm i in region j aims to maximise discounted profits at time t_0:

$$\Pi_{ijt_0} = \int_{t=t_0}^{\infty} e^{-\rho t} \left[\left(f(E_{ijt}) - (1+\tau_{jt}) w_{ijt} E_{ijt} - (1-\sigma_{jt}) T(h_{ijt}) E_{ijt} \right] dt , \qquad (1)$$

subject to the dynamic constraint:

$$\frac{\dot{E}_{ijt}}{E_{ijt}} = h_{ijt} - q_{ijt} . \tag{2}$$

Here E represents the firm's employment level, $f(E)$ a production function, w the wage, $T(h)$ training costs, h the hiring rate and q the quit rate. The local authority pays a training subsidy of σ to be funded out of a tax on wage income τ. The training cost $T(h)$ is measured in terms of the cost to the firm per existing employee of training new recruits arriving at rate h. The price of output has been set as the numeraire, therefore wages can be interpreted as real values.

The current value Hamiltonian for firm i in region j is:

$$H = f(E_{ijt}) - (1+\tau_{jt})w_{ijt}E_{ijt} - (1-\sigma_{jt})T(h_{ijt})E_{ijt} + \lambda_{ijt}(h_{ijt} - q_{ijt})E_{ijt} . \tag{3}$$

where λ_{ijt} is the Lagrange multiplier. The first conditions for this maximisation problem are:

$$(1-\sigma_{jt})T'(h_{ijt}) = \lambda_{ijt} ; \tag{4}$$

$$-\lambda_{ijt}\frac{\mathrm{d}q_{ijt}}{\mathrm{d}w_{ijt}} = 1+\tau_{jt} ; \tag{5}$$

$$-\frac{\partial H}{\partial E_{ijt}} = \dot{\lambda}_{ijt} - \rho\lambda_{ijt} . \tag{6}$$

The interpretation of these first-order conditions is straightforward. Equation (4) equates the marginal benefit from recruiting an additional worker, λ_{ijt}, to the marginal training cost. Equation (5) says that the marginal benefit from increasing the wage, in terms of reduced quits, should equal the marginal cost of the rising wage bill. Equation (6) expresses the marginal benefit of recruiting an additional worker in terms of the change in the present value of discounted future profits. The training subsidy σ_{jt} and tax rate τ_{jt} are assumed to be set such that the budget constraint $\tau_{jt} w_{.jt} = \sigma_{jt} T(h_{.jt})$ is satisfied.

Following Campbell and Orszag (1998), the training cost function is assumed to be a quadratic:

$$T(h_{ijt}) = \frac{A}{2}(h_{ijt})^2. \tag{7}$$

With respect to the quit rate, it shall be assumed, along the lines of Phelps (1994), that quits are a function of the ratio of the wage offered by the firm over the alternative wage offered elsewhere in the local economy. Specifically, quits are inversely related to $w_{ijt} / [w_{.jt}(1-\pi(u_{jt}))]$, where $w_{.jt}$ is the average wage in the local labour market and $\pi(u_{jt})$ is the instantaneous probability of becoming unemployed after quitting. The latter is unobservable but is assumed to be increasing in the observed unemployment rate in the local labour market u_{jt}.

This model differs from earlier formulations in that quitting is also assumed to be negatively related to the ratio of wage to non-wage income w_{ijt} / b_{jt} and that migration is possible, so that quits are affected by the level of wages available in other regions. The latter implies that quits are inversely related to $w_{ijt} / [\overline{w}_{..t}(1-\pi(\overline{u}_{.t}))]$, where $\overline{w}_{..t}$ is the economy-wide average expected wage at time t and $\pi(\overline{u}_{.t})$ is the economy-wide expected probability of being unemployed after quitting.[2] In accordance with Friedman's (1968) famous concept, $\overline{w}_{..t}$ and $\overline{u}_{.t}$ correspond to the "natural" rates of wages and unemployment in the economy, respectively.[3]

[2] Migration is driven by long-run expectations of income. Hence, workers respond to the underlying national values of wages and unemployment, rather than the contemporaneous values. $\overline{w}_{..t}$ and $\overline{u}_{.t}$ are assumed to be constructed such that each region's wage or unemployment rate is weighted by its labour force and are therefore the same everywhere. In reality, one might also expect "closer" regions to exert a greater influence than more distant or otherwise less accessible locations.

[3] Following Friedman, these are assumed to change gradually over time, hence the subscript t.

The quit rate function is assumed to be of constant elasticity and therefore can be written as:[4]

$$q(w_{ijt}, b_{jt}, w_{\cdot jt}(1-\pi(u_{jt})))$$

$$= B\left(\frac{w_{ijt}}{w_{\cdot jt}(1-\pi(u_{jt}))}\right)^{-\eta}\left(\frac{w_{ijt}}{b_{jt}}\right)^{-v}\left(\frac{w_{ijt}}{\overline{w}_{\cdot\cdot t}(1-\pi(\overline{u}_{\cdot t}))}\right)^{-\mu}. \tag{8}$$

Substituting the training cost function (7) and the quit rate function (8) into the first order conditions (4) and (5) yields the following equations:

$$A(1-\sigma_{jt})h_{ijt} = \lambda_{ijt}; \tag{9}$$

$$\lambda_{ijt}B(\eta+v+\mu)w_{ijt}^{-\eta-v-\mu-1}[w_{\cdot jt}(1-\pi(u_{jt}))]^{\eta}b_{jt}^{v}[\overline{w}_{\cdot\cdot t}(1-\pi(\overline{u}_{\cdot t}))]^{\mu} = 1+\tau_{jt}. \tag{10}$$

The model shall focus on the steady-state equilibrium path of each local labour market. This requires that the same wage be offered by all firms in a region and that the total employment level is constant. Hence, the following conditions must hold:

$$w_{ijt} = w_{\cdot jt},\ i = 1,2,\ldots,N,\ j = 1,2,\ldots,R; \tag{11}$$

$$h_{ijt} = q_{ijt},\ i = 1,2,\ldots,N,\ j = 1,2,\ldots,R. \tag{12}$$

In the long run, all regions must offer a given worker the same level of expected utility. If this were not the case, net migration would take place until either some regions were completely depopulated of workers or all inter-regional utility differences were

[4] This is consistent with Phelps' (1994) assumption of a quit rate function that is homogeneous of degree zero in all arguments and follows Campbell and Orszag's (1998) suggestion of a way to augment q to include non-wage income.

eliminated. Hence, it shall be assumed that the following spatial equilibrium condition is satisfied:[5]

$$(1-\pi(\overline{u}_{jt}))\overline{w}_{\cdot jt} = C_{jt}(1-\pi(\overline{u}_{\cdot t}))\overline{w}_{\cdot t}, \; j = 1,2,\ldots,R. \tag{13}$$

The parameter C_{jt} allows for deviations from the assumptions of perfect mobility and identical regional attributes. If $C_{jt} = 1$ for all j, condition (13) is equivalent to the standard Harris and Todaro (1970) condition that expected wages are equalised across regions. However, there are two reasons why C_{jt} may differ from unity: the existence of location-fixed amenities that affect the utility of workers or imperfect arbitrage due to some form of spatial friction, for example costly migration.[6] Consequently, the following functional form for C_{jt} is posited:

$$C_{jt} = \zeta_j e^{\theta c_{jt}}. \tag{14}$$

ζ_j measures the relative level of amenities available in a region. Regions that are more attractive than average have values of ζ_j that are less than unity, and vice versa. In the absence of regional differences in amenities, $\zeta_j = 1$. c_{jt} represents a measure of the cost of migration. The case of costless mobility, when $c_{jt} = 0$, is consistent with Blanchflower and Oswald's efficiency wage model, where all migrants from region a to region b enjoy an increase in non-pecuniary benefits regardless of their ensuing labour force status. C_{jt} is monotonically increasing in c_{jt}.

$\overline{w}_{\cdot jt}$ and $\overline{u}_{jt}$ denote regions' natural rates of wages and unemployment. In the short run these are unknown to agents, therefore the time-averaged values $w_{\cdot j}$ and u_j are used as approximations, respectively. Combining equations (9) to (14) then yields the

[5] This is equivalent to the zero-migration condition in Blanchflower and Oswald's shirking model, except that this specification assumes no uncertainty regarding labour demand shocks. However, unlike the former, it does allow for more than two regions and positive migration costs. Equation (3.28) casts a new perspective on Friedman's claim that the natural rate of unemployment represents a rate "which has the property that it is consistent with equilibrium in the structure of *real* wages" (p. 8).

[6] See Chapter 9 of Barro and Sala-i-Martin (1995) for models of long-run spatial equilibrium in growing economies with costly migration.

following relationship involving the steady-state wage level w_{jt} and unemployment rate u_{jt} in region j:

$$A(1-\sigma_{jt})B^2(\eta+\nu+\mu)(1-\pi(u_{jt}))^{2\eta} w_{.jt}^{-2\nu-2\mu-1} b_{jt}^{2\nu} \zeta_j^{-2\mu} e^{-2\mu\theta c_{jt}} w_{.j}^{2\mu}(1-\pi(u_j))^{2\mu}$$
$$=1+\tau_{jt}. \qquad (15)$$

Taking the natural logarithm of both sides and simplifying through the introduction of some new parameters allows this expression to be rewritten as follows:

$$\ln w_{.jt} = \kappa + \psi_j + \alpha \ln \frac{1-\sigma_{jt}}{1-\tau_{jt}} + \xi \ln b_{jt} - \varphi c_{jt} + \delta \ln(1-\pi(u_{jt})), \qquad (16)$$

where

$$\kappa = \frac{\ln(AB^2(\eta+\nu+\mu))}{2\nu+2\mu+1}; \qquad (17)$$

$$\psi_j = \frac{2\mu(\ln w_{.j} + \ln(1-\pi(u_j)) - \ln \zeta_j)}{2\nu+2\mu+1}; \qquad (18)$$

$$\alpha = \frac{1}{2\nu+2\mu+1}; \qquad (19)$$

$$\xi = \frac{2\nu}{2\nu+2\mu+1}; \qquad (20)$$

$$\varphi = \frac{2\theta\mu}{2\nu+2\mu+1}; \qquad (21)$$

$$\delta = \frac{2\eta}{2\nu+2\mu+1}. \qquad (22)$$

The main conclusions from this model are that there is a positive relationship between $w_{.jt}$ and $1-\pi(u_{jt})$ and that the elasticity of this relationship, δ, is constant across regions. The elasticity is solely a function of behavioural parameters, which are assumed constant across the homogeneous and mobile labour force. Regional fiscal policies, measured by the parameters σ_{jt} and τ_{jt}, and the regional social security benefit b_{jt} have no impact on the elasticity.

The probability of unemployment will be proxied by the observed unemployment rate u_{jt}, following the approach of a number of previous authors.[7] Even so, equation (16) does not represent Blanchflower and Oswald's wage curve *per se*, as it relates the regional wage level to the probability a worker is re-employed after quitting, rather than the probability he/she becomes unemployed. The equilibrium wage locus presented here will form the basis for empirical estimation. It is preferable to some arbitrary linear model with the natural logarithm of u_{jt} on the right hand side, as it follows directly from the theoretical model. As with Campbell and Orszag's (1998) steady-state wage equation, equation (16) gives no reason to expect the relationship between wages and unemployment to be one of constant elasticity. Rather, the unemployment elasticity of pay is determined by the parameter δ and the level of the unemployment rate as follows:

$$\frac{\partial \ln w_{.jt}}{\partial \ln u_{jt}} = \frac{u_{jt}}{u_{jt}-1}\delta . \tag{23}$$

As with that of Salop (1979), the model presented here is capable of explaining a labour market equilibrium, in which both wage and employment levels are endogenously determined. Combining equations (13) and (14), and taking logarithms, yields:

$$\ln \overline{w}_{.jt} = \chi_t + \ln \zeta_j + c_j - \ln(1-\pi(\overline{u}_{jt})), \tag{24}$$

[7] Blanchflower and Oswald (1994) show that the probability of being unemployed is related to the unemployment rate in the steady state as follows: $\pi(u) = 1 - \frac{h(1-u)}{u}$, where h is the hiring rate.

where

$$\chi_t = \ln(1-\pi(\overline{u}_{.t})) + \ln \overline{w}_{..t}. \qquad (25)$$

Equation (24), which shall henceforth be referred to as the Harris-Todaro condition, sets the expected wage in a given region that is necessary for a zero-migration equilibrium to exist. Every region must therefore satisfy two independent conditions in the long run and, as a consequence, a regional equilibrium is uniquely defined by the dynamic efficiency wage condition (16) and the Harris-Todaro condition (24), for any given state of the national economy. The equilibrium wage-employment locus entails a positive relationship between earnings and the employment probability in each region and dictates the wage offer a profit-maximising firm should post, given local labour market conditions. Along the locus, however, only one combination of wages and employment levels will give workers a utility level that ensures there is neither net migration into or out of the region. To the extent that equation (24) is an analogue for the wage curve, this is a similar situation to that of Blanchflower and Oswald's shirking model, wherein the relative positions of regions along a common wage curve are determined in the long run by differences in non-pecuniary benefits.

The Harris-Todaro condition implies a negatively sloped curve in wage-employment probability space. If this curve is assumed to be fixed in the long-run, a positive relationship between permanent wages and unemployment will be predicted. In the short-run, regional shocks to wages and unemployment rates will trace out points along the wage locus.[8] Hence, the model concurs with Blanchflower and Oswald's suggestion that Harris and Todaro's (1970) compensating differentials argument might sit comfortably alongside the wage curve so long as it is clear to what time frame is being referred. Figure 1 depicts the wage-employment locus (*WE*) and the Harris-Todaro condition (*HT*).

3. A description of the data used

The bulk of the data used in this study was derived from the 1986, 1991 and 1996 New Zealand Censuses of Population of Dwelling. Definitions of all the variables referred to

[8] That is, variations in the pair $(w_{.jt}, \pi(u_{jt}))$ will enable the parameter δ to be identified.

in this section are found in the appendix. Observations on annual income Y and average weekly hours worked H were obtained for 34,560 population subgroups, each containing n wage and salary earners.[9,10] The subgroups form a mutually exclusive and exhaustive set of permutations of the following qualitative characteristic categories: year t (3 values); urban area i (30 values); age (3 values); educational attainment (2 values); ethnicity (2 values); gender (2 values); full-time/part-time status (2 values); occupation (8 values). As would be expected, a large number of observations reported a value of zero for n and, hence, Y and H. These observations were deleted, leaving 20,302 observations in total. Collectively, around half a million workers were covered in each year, or approximately 56% of all wage and salary earners.[11]

The wage level w of a group in 1986 dollars was calculated as follows, using national consumer price index values for each year P:[12,13]

$$w = \frac{Y}{(52H)P} \tag{26}$$

A variety of aggregate statistics for each urban area were also collected from the 1981, 1986, 1991 and 1996 Censuses. These included the working age population N, the number of unemployed males U_M and females U_F and the number of each gender in the labour force, L_M and L_F, respectively. u_T, u_M and u_F refer to the total, male and female unemployment rates, respectively, and were constructed by dividing the relevant number of unemployed people by the relevant labour force. M represents the number of working

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[9] Non-wage and salary earners (for example, self-employed individuals) were excluded as they do not conform to the model presented in Section 2 and have values of Y and H with different interpretations to those of wage and salary earners.

[10] For confidentiality reasons, Statistics New Zealand randomly rounds all values of n to base 3. A reported count of 6 could, therefore, imply an actual count of anywhere between 4 and 8. Since the long run expected value is designed to equal the original count, the impact of this on the results of this paper should be insignificant.

[11] 532,755 wage and salary earners are included in the 1986 sample, 497,136 in 1991 and 565,788 in 1996.

[12] This assumes that H is the average hours worked per week over the previous 52 weeks, which differs from the census definition, meaning the wages of those who have not worked the entire year will be under-estimated. In addition, since Y includes non-wage income, the wages of those who receive a large amount of interest, dividends, etc. will be over-estimated. See the appendix for details.

[13] P took the value 1 for 1986, 1.499 for 1991 and 1.643 for 1996.

age persons in an urban area whose address was outside that urban area five years previously. The gross migration rate m was then determined as follows:[14]

$$m = \frac{M}{N} \tag{27}$$

One potentially important variable that was not obtained from the census was the average house sale price in each urban area, expressed as a proportion of the corresponding national value. Instead, this was drawn from half-yearly urban property sales statistics and denoted p.[15]

Table 1 gives the mean and standard deviation of each individual level variable by census year. Observations are weighted by n, so that the values reflect the average characteristics of all workers in the sample. Table 2 gives the corresponding values for each aggregate variable. Here observations are weighted by the population of each urban area.

As one would expect, nominal annual income Y increased over the sample period, albeit less rapidly between 1991 and 1996. Average hours worked declined, particularly after 1991, partly reflecting a corresponding fall in the proportion of full-time workers in the labour force. Although not reported in Table 2, the data also show that the average hours worked by part-time workers fell over the sample period, while among full-time workers H rose on average. Both changes were particularly noticeable after 1991, coinciding with the introduction of the Employment Contracts Act. Average real wages w rose between 1986 and 1991, but then fell, so that those workers in our sample earned less in 1996 than they had ten years earlier. There has been no significant difference in the wages paid to full-time or part-time workers. The coefficient of variation of H has increased by 30% over the past decade, while the value for w has remained constant.[16,17]

[14] In-migration was deemed a better value to use in m than out-migration, as individuals who leave New Zealand are not included in the census.

[15] The average values obtained were for the territorial authorities corresponding most closely to each urban area. The boundaries of these underwent substantial changes in 1989.

[16] The coefficient of variation is the standard deviation of a variable expressed as a proportion of the mean.

[17] The values for 1986 rather than 1991 were compared with those for 1996 in order to control for changes in the composition of the workforce over the business cycle (see Winkelmann (1999) for an explanation).

Unemployment follows an expected cyclical pattern, peaking in 1991. The 1991 sample is also the "oldest", with only 29% in the 15-25 age group, and the most educated, with 44% having a post secondary school qualification. The absence of younger and less skilled individuals from this sample would appear to support the notion that workers with less job-specific training are more likely to be fired during periods of recession.[18]

In general, Table 1 provides evidence of a shift from low-skilled to high-skilled occupations. The proportion of the sample employed as teachers (*OC*2), salespeople (*OC*4) or service workers (*OC*5) rises from 1986 to 1996, while the proportion of health professionals (*OC*1) remains stable. On the other hand, the proportions of production workers (*OC*6) and construction workers (*OC*7) in the sample fall over time. Exceptions to this pattern are office clerks (*OC*3), who comprise a falling share of the sample despite being considered "high-skilled", and labourers (*OC*8), who make up a greater proportion of the 1996 sample than that of ten years earlier.

Finally, a set of variables was collected, in addition to the above, to be used as instruments for the endogenous determinants of wages. The proportion of private dwellings that were rented in each urban area r was obtained from each census. Data on annual rainfall and sunshine hours in each urban area and year were collected. These were expressed as a percentage of the long run, or normal, values for the particular location and were denoted W_R and W_S, respectively. Lastly, an industry mix variable was constructed along the lines of Bartik (1991). This was a weighted average of the growth rates of national industry employment over the year prior to each census, with the weights calculated as the share of a local labour market's employment in each industry at the time of the census.[19] This variable was denoted I. Means and standard deviations for these four variables are reported in Table 2.

4. New Zealand evidence for the wage curve

Blanchflower and Oswald (1994) settled on a standard specification in which both the wage and the unemployment rate were in natural logarithm form. This implies a constant unemployment elasticity of pay. In order to produce comparable results for New Zealand

[18] See, for example, Mincer (1962).

[19] National employment changes were taken from the Quarterly Employment Survey. Differences in industry definition between this survey and the census means that some growth rates are not strictly correct. For example, the Q.E.S. does not cover farmers, whereas the census does.

the same specification was initially chosen. The relevant unemployment rate u was constructed so as to be gender-specific:

$$u = \begin{cases} u_M \text{ if male} \\ u_F \text{ if female} \end{cases} \tag{28}$$

The logarithm of real wages ln$w$ and logarithm of the unemployment rate ln$u$ would initially appear to exhibit a weak inverse relationship, as indicated by the correlation coefficient of –0.049. However, it is essential to control for the effect of other relevant determinants of workers' earnings before any conclusion about the role of the local unemployment rate can be made.

Human capital theory points to the importance of workers' personal characteristics in determining wages (Becker 1971; Mincer 1974; Mincer and Polachek 1974). In addition, the requirements of different jobs dictate different levels of pay, perhaps via the existence of compensating wage differentials or industry effects. The first column of Table 3 reports the results of a regression attempting to explain ln$w$ as a function of workers' gender, age, educational attainment, ethnicity and full-time/part-time status as well as their occupation. The wage paid to employees should compensate them for differences in the cost of living across local labour markets. A significant component of this is likely to be related to housing costs. Therefore, the natural logarithm of the median house price in each urban area and period as a fraction of the national value ln$p$ was included in the list of regressors in an attempt to capture this effect.[20] Other important determinants of ln$w$ may be nationwide changes that affect all locations equally. With panel data these can be controlled for by including a full set of year dummy variables as regressors and employing fixed time effects estimation. As with all regressions in this section, observations are weighted by the size of the subgroup, so that values of ln$w$ representing the median income of a small number of workers are given less weight. The relevant earnings equation can be written as follows, with $\mathbf{x}_{gh}$ representing a vector of control variables of gender h in group g and ι_t denoting the presence of time period effects:

[20] Winter-Ebmer (1996) and Kennedy and Borland (1997) also use house prices as a proxy for a regional price index.

$$\ln w_{ghjt} = \gamma_0 + \mathbf{x}'_{gh}\boldsymbol{\gamma}_1 + \gamma_2 \ln p_{jt} + \iota_t + \varepsilon_{ghjt} \tag{29}$$

From Table 3 it can be seen that workers aged 26-40 appear to earn around 35% more than 15-25 year olds, while those aged 41-60 are paid 39% more than the youngest group. This finding is broadly consistent with a concave age-earnings profile and captures the effect of increasing labour market experience on pay. Also consistent with Mincer's (1974) earnings function is a positive coefficient on the education dummy variable, which suggests that *ceteris paribus* the attainment of a post secondary school qualification increases earnings by 13%. Europeans earn 5% more than non-Europeans, while males earn 20% more than females. These findings might indicate the presence of discrimination in the labour market or perhaps just the omission of relevant human capital variables that are correlated with gender and ethnicity.[21] Full-time workers appear to earn 4% less per hour than part-time workers. This is also not surprising, given the extra hours many full-time salary earners work for no marginal increase in pay. The estimated time effects $\hat{\iota}_t$ are found to be significant. All else being equal, a worker earned 3% less in 1991 than in 1986. In 1996 this value was 5%, suggesting that there has been a trend towards lower real wages over the past decade. ln$p$ reports a significant positive effect, suggesting that cross-sectional variation in the cost of living *is* important.

The gender-specific unemployment rate in natural log form ln$u$ is added to the set of explanatory variables in the second column of Table 3. An estimate of the unemployment elasticity of pay of –0.098 is obtained, remarkably close to the "magic" value of –0.1 propounded by Blanchflower and Oswald. This result implies that a 10% increase in an urban area's unemployment rate will induce a 0.98% fall in the wages of workers in the urban area.

The level of pay received by workers in a particular locality may be influenced by other relevant variables that are specific to the local labour market but are constant over time, such as amenities and spatial location. To capture the effect of these "permanent" urban area attributes, a full set of 29 urban area dummies ω_j was included. Our specification is now:

[21] Besides vertical and horizontal occupational segregation, another factor contributing to men earning higher wages is that many females leave the labour market to raise children, hence their age is a less

$$\ln w_{ghjt} = \gamma_0 + \mathbf{x}'_{gh}\gamma_1 + \gamma_2 \ln p_{jt} + \beta \ln u_{hjt} + \omega_j + \iota_t + \varepsilon_{ghjt} \tag{30}$$

Table 4 presents the results of F-tests for the joint significance of both urban area and time effects, the presence of location dummy variables given the existence of time effects and the presence of time dummies given the existence of urban area effects. Strong evidence in support is found in all cases, justifying the two-way fixed effects model. As the third column of Table 3 indicates, the estimated coefficient on unemployment now increases in magnitude to –0.133. The exclusion of urban area effects clearly resulted in an under-estimated value of $\hat{\beta}$. Figure 2 plots the estimated values of the urban area effects $\hat{\omega}_j$ against the natural log of the time-averaged unemployment rate for each region. The two exhibit a positive relationship, indicating that urban areas with permanently higher levels of ln$w$ also tend to have permanently higher rates of ln$u$. The existence of this relationship alongside a negative unemployment elasticity of pay lends tentative support to Blanchflower and Oswald's (1995) claim that "movements in actual wages can be negatively correlated with movements in actual unemployment, while at the same time "permanent" unemployment... is positively related to "permanent" wages" (p. 160). Thus, the wage curve is able to be reconciled with the Harris-Todaro concept of compensating differentials across regions.[22] To provide more formal evidence than Figure 2, a Hausman test was conducted to test whether either ω_j or ι_t can be represented as random effects. Both test statistics were significant at the 1% level, indicating that both types of effect should be captured as dummy variables and that a switch to random effects estimation would introduce bias.

Given the impact different time periods have on the intercept of the earnings equation, it would seem logical to investigate the possibility that they also affect the unemployment elasticity of pay. The fourth column of Table 3 includes two interactive terms designed to capture any differences in β over time. Unemployment appeared to have a weaker effect on wages in 1986 and 1996. The fact that $\hat{\beta}$ is most negative in

suitable proxy for labour market experience that is the case with males. This issue is explored further in the next section.

[22] As Blanchflower and Oswald note, migration is most likely driven by the permanent values of earnings and unemployment.

1991 suggests that, to some extent, the *level* of ln$u$ determines the impact it has on ln$w$. Blanchflower and Oswald's assumption of a constant elasticity wage curve may then be inappropriate.

Even if the assumption of constant elasticity is inappropriate, it is unclear why unemployment had less of an influence on earnings in 1996 than a decade earlier, a time when unemployment was at approximately the same level. One possible explanation is the introduction of the Employment Contracts Act in 1991, which heralded a movement away from collective bargaining and weakened the power of unions. However, the discovery of a more elastic wage curve in the presence of a centralised system of industrial relations contradicts the empirical findings of Blanchflower and Oswald (1994), who reported that unemployment has a smaller effect on wages in unionised sectors of the United States than in non-unionised sectors.

One response is that the Employment Contracts Act can be interpreted as having increased the elasticity of labour supply in New Zealand. In this case, firms have had less need to alter wages in response to changes in product demand since 1991, as turnover rates have become more elastic. In other words the Employment Contracts Act has been responsible for an increase in workers' average values of v in equation (8). In addition, the "flattening" of the wage curve observed cannot be attributed solely to labour market reform. The ι_t dummies capture the effects of *all* region-invariant shocks and it is impossible to disentangle the effect of each.

Since it is determined together with wages in the labour market, unemployment may not just affect pay but itself be a function of the wage rate. If this is the case, the estimates of β obtained so far will be affected by simultaneity bias. Hausman's specification error test was conducted by regressing ln$u$ on all available exogenous variables. The residuals from this auxiliary regression were found to have a significant effect on wages, therefore it was concluded that ln$u$ is endogenous. To eliminate this problem, instrumental variables estimation was employed. This requires one or more valid instruments, that is, a set of variables that affect unemployment in a region but may be omitted from the earnings equation. Following Blanchflower and Oswald's (1994) approach, the following instrumental variables were chosen: annual rainfall deviation from normal W_R, annual sunshine deviation from normal W_S and an industry mix variable I.[23] Levels of rainfall and sunshine may be thought of as being key factors determining the amount of agricultural production, which, in turn, affects the economy-wide unemployment rate due to New Zealand's reliance on primary industries. I captures the effect of shocks to labour demand by treating national rates of change in industry employment as given within each urban area.[24] Oswald (1997) claimed that the proportion of houses in a local labour market that are rented r is an important determinant of unemployment.[25] Since there is no obvious direct relationship between the renting proportion and $\ln w$, this suggests that r may another plausible instrument for local unemployment in the earnings equation.

Table 5 presents the results of using the lagged value of $\ln u$ plus, in turn, W_R, W_S, I, r and all four variables as instruments for $\ln u$.[26] In contrast to Blanchflower and Oswald's results, we find that once endogeneity of the unemployment rate is controlled for evidence of a wage curve vanishes. Regardless of what combination of instruments is used, a positive and significant value of $\hat{\beta}$ is obtained. This may suggest the refection of the wage curve notion in favour of a short-run Harris-Todaro relationship, however it may also imply some form of mis-specification of the wage function. We investigate the latter with reference to our model of wage determination.

The earnings equation implied by the theoretical model presented in Section 2 differs from Blanchflower and Oswald's standard specification, used so far in this section. In particular, the former suggests that wages are determined by the probability of *employment* $1-\pi$, rather than the probability of unemployment π, and gives no reason to expect a constant value of β. To evaluate our model, an earnings equation was estimated with ordinary least squares, whereby wages were explained by the group-specific variables in $\mathbf{x}$, house prices, region and time fixed effects and the natural logarithm of the employment rate, which acts as a proxy for $\ln(1-\pi)$.[27] As the first column of Table 6

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[23] Blanchflower and Oswald also used military spending, however this is not relevant in a New Zealand setting. Local government spending is an obvious substitute, however this information is not easily obtainable for all urban areas and each year.

[24] Use of I as an instrument requires the assumption that each urban area comprises a relatively small proportion of national employment in each industry group.

[25] Oswald tentatively concluded that a 10% increase in the renting proportion decreases the unemployment rate by 2%.

[26] The lagged value of $\ln u$ refers to the corresponding value from the previous census.

[27] The employment rate $1-u$ is simply the ratio of the number of employed workers and the number in the labour force.

reports, the employment rate has the expected positive impact on wages. Our estimate of the employment elasticity of pay δ is 1.550, meaning a 10% increase in the employment rate leads to a 15.5% increase in wages.

The theoretical model from Section 2 also implies that wages are affected by the cost of migration c. This is unobserved, however it can be proxied by the migration rate m. Migration is expected to have an equilibrating effect on wages, hence it should be inversely related to migration costs. As a result, m^{-1} was added to the list of regressors in the second column of Table 6. A significant negative coefficient was found, supporting the model's predictions and entailing a value of 0.015 for φ.[28] When the migration rate is controlled for, $\hat{\delta}$ increases slightly to 1.578.

Since the employment rate, along with the unemployment rate, is determined simultaneously with wages, our estimates of δ may be biased and require instrumentation. Following Johnes and Hyclak (1999), house prices and the migration rate may also be thought of as endogenous variables in this system. To allow for this, the last column of Table 6 treats $\ln(1-u)$, $\ln p$ and m^{-1} as endogenous and reports the results of instrumental variables estimation when lags of all three plus I and r are used as instruments.[29] $\hat{\delta}$ now remains significant after controlling for simultaneity bias and increases to 2.119. The migration rate also enters with a significant coefficient, however house prices no longer appear to be a significant determinant of wages. The robustness of the earnings equation to corrections for endogeneity when the employment rate is included may indicate that the our specification of wage function is superior to that of Blanchflower and Oswald.

As Campbell and Orszag (1998) noted, when the employment rate, rather than the unemployment rate, is included in the earnings equation, a constant elasticity wage curve is no longer assumed. Rather, the unemployment elasticity of pay $\frac{\partial \ln w}{\partial \ln u} = \beta$ can be derived from the employment elasticity δ as in equation (23). At the mean, the preceding

[28] The elasticity of earnings with respect to migration is not constant, so that a 1% increase in the gross migration rate incurs a $\frac{0.015}{m}$ % increase in wages, or 0.06% at the mean.

[29] Lags of $\ln(1-u)$ and m^{-1} refer to values at the previous census, while the lag of $\ln p$ refers to the observation for the previous half-year. W_R and W_S were excluded as they were found to have significant correlations with the residuals from the weighted least squares regression and, hence, are no longer valid instruments.

regressions imply a value of $\hat{\beta}$ that ranges from –0.178, with weighted least squares, to –0.238, with instrumental variables.[30] Thus, the evidence suggests that the effect of unemployment on workers' earnings is of a similar magnitude to that uncovered with Blanchflower and Oswald's specification of the wage curve.

5. Some Additional Issues

It is widely accepted that there are considerable differences in the labour supply behaviour of men and women. Fertility decisions affect the likelihood of women participating in the labour force. In a New Zealand study, Harris (1992) found that doubling the number of pre-schoolers in a household would decrease labour force participation of women of European ethnicity by over 33%, while the presence of children tends to increase the participation of males.[31] Among those who did participate, Harris reported that the presence of children in a household has a substantial negative effect on the hours worked by females. As noted earlier, time spent out of the labour market raising children also leads to a depreciation of human capital among women. This implies the existence of a different shaped age-earnings profile. To allow for these differences, separate estimates of the earnings equation may be made for males and females. In their search for a German wage curve, Baltagi and Blien (1998) found that the unemployment elasticity of pay was greater for men than for women, as did Janssens and Konings (1998) for Belgium. In contrast, Kennedy and Borland (1997) reported that in Australia the wage curve is stronger for females than males and concluded that "the significance of the wage curve relation for males appears to be driven largely by a small group of part-time workers" (p. 24).

Although a gender-specific employment rate has been used as an explanatory variable, thus far our specifications have imposed that all coefficients are the same for both male and female workers. To examine this assumption, the earnings equation implied by our model was estimated separately for each gender. Seniority, as measured by age, seems to increase the earnings of men more than it does of women, whose estimates are reported in the first and second columns of Table 7, respectively. Men aged 41-60 earn 53% more than those aged 15-25, while for women the corresponding value is

[30] The weighted mean value of u was 0.101.

32%. This is consistent with the explanation given above. In addition, women who switch from part-time to full-time work can expect a 4% drop in hourly wages, compared to 13% for men. Neither gender provides support for the wage curve, with the coefficient on the employment rate insignificant in both cases. However, once endogeneity in the employment rate, house prices and migration rate is controlled for, a positive value of $\hat{\delta}$ is found for men but not women. The results of instrumental variables estimation for men and women are listed in the third and fourth columns of Table 7, respectively.[32]

The conclusion that the wage curve is more elastic for men than women contradicts the finding of Kennedy and Borland. In addition, our wage curve result is not driven by part-time workers. Full-time and part-time males both report a positive coefficient on the employment rate, while $\ln(1-u)$ has an insignificant effect for females, regardless of employment status. It is possible that our inability to find conclusive evidence for the wage curve in the total sample is due to the confusion of the different effects of the two gender groups. The wages of female workers do not appear to be affected by the local rate of unemployment among women, however females make up over half of our sample.

One explanation for the apparent irrelevance of the employment rate in the determination of female wages is that the supply of women in the labour force is more flexible than that of men. If a fall in the demand for labour leads to many female workers withdrawing completely from the labour force, firms will be less capable of reducing the wages of women in times of tight labour markets than men. In terms of the model presented in Section 2, women have a higher value of v in equation (8). It then follows from equation (22) that they will have a lower employment elasticity of pay.

Having investigated the possibility of differences in the wage curve between males and females, it is instructive to examine how the wage curve varies across occupation groups. Using aggregate New Zealand data, Morrison and Poot (1999) found tentative evidence that the wage curve is more elastic for relatively less-skilled occupations. However, neither Blanchflower and Oswald (1994) nor Kennedy and Borland (1997) were able to report any systematic pattern in the unemployment elasticity of pay among different occupations.

[31] Harris' study was based on data from the 1986 Census.

[32] The full set of four instruments was used. Based on a mean value of u of 0.087, our estimates imply that $\hat{\beta}$ is –0.100 for men.

In contrast to the case of differences between genders, there is no compelling theoretical reason to allow all coefficients in the earnings equation to differ with occupation. Accordingly, we allow only the employment elasticity of pay δ to vary by introducing seven interactive terms to the list of regressors. The results, which are reported in the fifth column of Table 7, suggest that four of the seven occupations have values of δ that differ from that of our reference occupation group, namely, health professionals ($OC1$). The interactive terms for teachers ($OC2$), office clerks ($OC3$) and salespeople ($OC4$) are all found to be significant and negative. The interactive term for production workers ($OC6$) is significantly positive, while those for service workers ($OC5$), construction workers ($OC7$) and labourers ($OC8$) are all insignificant.

Although there is no clear pattern, it would appear that the first four, high-skilled occupations tend to have less elastic wage curves than the other four, low-skilled occupations.[33] This is consistent with Topel's (1986) prediction that the wages of more geographically mobile groups of workers tend to be less affected by the local unemployment rate. This is seen in our model by more mobile groups having a higher value of μ in equation (8).

One important question is how much this study has gained by using group level data compared to previous research in New Zealand, which has relied on aggregated data. This can be addressed with reference to Lang and Gottschalk (1996), who derived the efficiency loss from using grouped data to estimate the coefficients of variables that vary across groups but not individuals within groups when individual data is not available on the dependent variable. They show that use of disaggregated data only increases the precision of coefficient estimates for group-specific variables if there is correlation between the individual and group level regressors. In a wage curve context, Lang and Gottschalk's approach allows us to calculate the gain in efficiency from using our "synthetic" micro data set compared to the aggregated data used by Morrison and Poot (1999).

The first step requires the employment rate $\ln(1-u)$ to be regressed on all individual level regressors, namely the age, education, gender, ethnicity employment status,

[33] We are unable to use instrumental variables estimation now as the presence of the seven interactive terms means ten instruments would be required.

occupation, urban area and year dummies. The R^2 from this regression is then compared with the R^2 from a similar regression where the individual level variables are averaged over the group, in essence making them of the same level of aggregation as $\ln(1-u)$. This procedure was applied twice. Firstly, we aggregated observations within urban areas, years and gender, reducing the number of observations from 20,302 to 180. This resulted in the R^2 of interest rising from 0.871 to 0.937. Following Lang and Gottschalk, this implies that the standard error of $\hat{\delta}$ increases by 43% when one switches to aggregate data.[34] Next, we used $\ln(1-u_T)$ rather than $\ln(1-u)$ and aggregated within urban areas and years only. This left 90 observations for the group level regression and an R^2 that rose from 0.952 to 0.983. This implies that when observations by urban area and year only are available the standard error of $\hat{\delta}$ is 69% higher than in our data set, or, alternatively, that our estimated standard error is 41% lower than with aggregate data.

6. Conclusion

This study examines the role that local labour markets play in wage determination in New Zealand. Based on the dynamic efficiency wage model of Phelps (1994), a positively sloped steady state relationship between the wages of workers in a region and their probability of finding employment was derived. The location of a particular region on this locus in the long run is determined by the position of a so-called Harris-Todaro condition, which ensures an inter-regional migration equilibrium. It is suggested that Blanchflower and Oswald's (1994) wage curve involves a mis-specification of the relationship between wages and local labour market conditions.

Using disaggregated census data we have documented for the three different time periods 1986, 1991 and 1996 how urban areas with tight labour markets experienced greater pressure on local wages. This data set affords a substantial gain in the efficiency of estimates of the wage curve over previous New Zealand studies that used aggregate data. After controlling for the endogeneity of the employment rate, prices and migration rate of an urban area, evidence, albeit weak, of a positive relationship between wages and

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[34] The increase in the standard error of δ is equal to $\left(\frac{1-R_{gr}^2}{1-R_m^2}\right)^{-\frac{1}{2}}$, where R_{gr}^2 and R_m^2 denote the R^2s of the group level and individual level regressions, respectively.

the employment rate is reported for males only. This would seem to support the general predictions of our theoretical model. Our finding of an insignificant employment elasticity of pay among women may be explained by the more elastic labour supply of this group.

We also find some evidence that the wages of workers in less-skilled occupations are most responsive to the local employment rate. This concurs with the prediction of our model that geographically mobile groups will be less vulnerable to local labour market conditions. It appears that the less migration acts as an adjustment mechanism for any subgroup of the population, the stronger statistically is its wage curve. Since these groups are disproportionately dependent on job prospects within the urban areas in which they live, their quit rates are less responsive to changes in their wage levels.

In contrast to the internationally established result that "weak" bargaining groups have relatively high unemployment elasticities of pay, our results reveal that the unemployment elasticity of pay fell with the introduction of the Employment Contracts Act in 1991. One explanation is that the Act has brought about an increase in the elasticity of labour supply in New Zealand. In addition, our finding may be influenced by an inability to control for other national shocks that have affected earnings.

Blanchflower and Oswald (1994) describe their empirical findings on wages and unemployment as featuring "two of the variables that most interest policymakers" (p. 1). It seems clear that hourly earnings are higher in New Zealand the tighter the local labour market, at least among males. In addition, our results suggest that the wages of those in less skilled occupations are more responsive to changes in employment prospects. Any attempt to reduce this vulnerability of workers in weaker bargaining groups will then require increasing the mobility of these workers. The findings of this study confirm that an increase in the "openness" to migration of an urban area leads to higher wages. An important issue for future research to address is the effect on wages of gross migration rates for different subgroups of workers.

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Figure 1
The wage-employment locus (*WE*) and Harris-Todaro condition (*HT*)

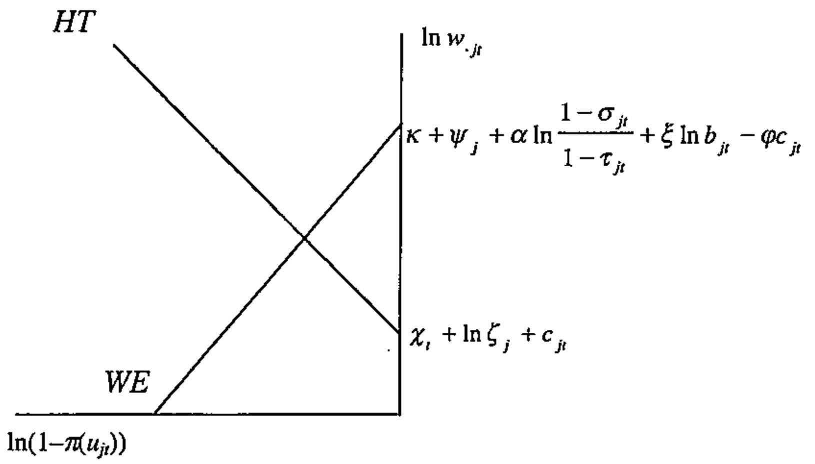
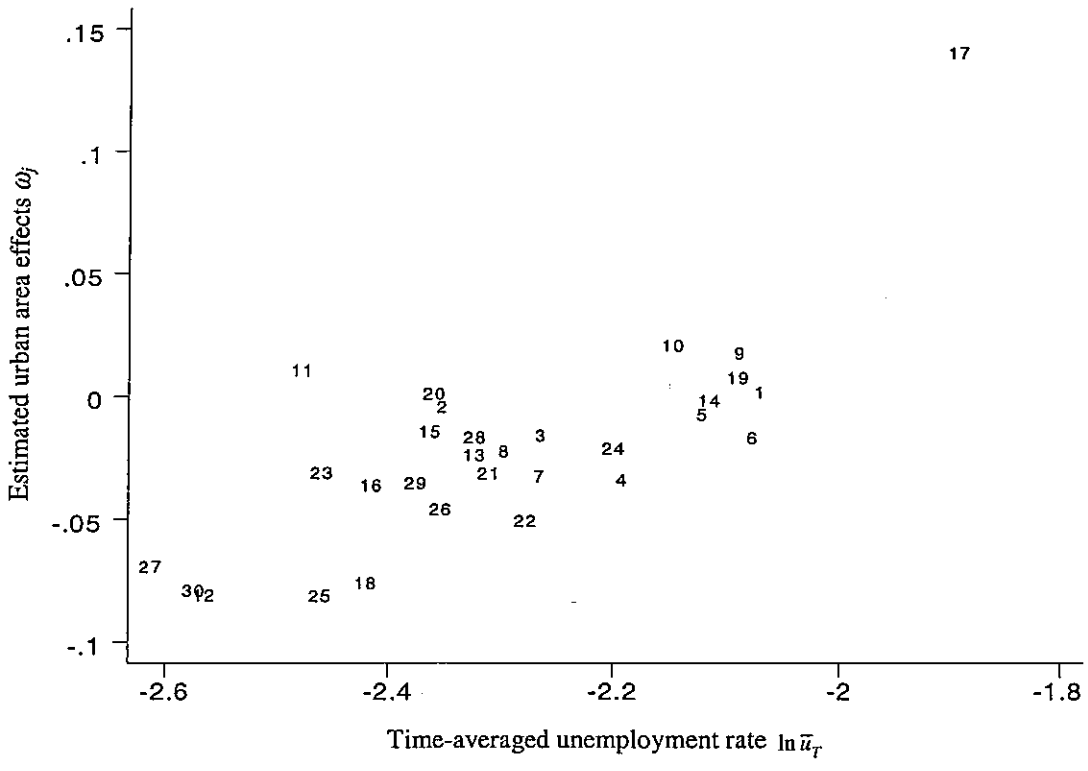


Figure 2
Scatter plot of estimated urban area effects and the time-averaged unemployment rate for New Zealand in 1986, 1991 and 1996



Note: Labels refer to urban area numbers listed in the appendix.

Table 1
Descriptive statistics for the subgroup level variables

Variable	1986		1991		1996	
	mean	standard deviation	mean	standard deviation	mean	standard deviation
Y	14070	6201	21122	9747	21582	11072
H	36.63	10.87	36.03	11.70	34.75	13.48
full-time	0.797	0.402	0.765	0.424	0.702	0.458
non-Maori	0.869	0.338	0.864	0.343	0.858	0.349
male	0.386	0.487	0.370	0.483	0.367	0.482
aged 26-40	0.334	0.472	0.352	0.478	0.340	0.474
aged 41-60	0.311	0.463	0.356	0.479	0.355	0.479
qualified	0.343	0.475	0.437	0.496	0.373	0.484
$OC1$	0.070	0.255	0.077	0.267	0.070	0.255
$OC2$	0.079	0.270	0.095	0.293	0.096	0.295
$OC3$	0.371	0.483	0.363	0.481	0.332	0.471
$OC4$	0.092	0.289	0.105	0.306	0.124	0.330
$OC5$	0.171	0.377	0.183	0.387	0.205	0.404
$OC6$	0.136	0.343	0.097	0.296	0.093	0.290
$OC7$	0.035	0.184	0.032	0.175	0.020	0.139
$OC8$	0.045	0.207	0.048	0.213	0.060	0.238
w	7.403	2.837	7.448	2.455	7.175	2.887

Table 2
Descriptive statistics for the urban area level variables

Variable	1986		1991		1996	
	mean	standard deviation	mean	standard deviation	mean	standard deviation
u_T	0.070	0.011	0.144	0.016	0.083	0.014
u_M	0.054	0.009	0.136	0.018	0.078	0.015
u_F	0.092	0.016	0.153	0.018	0.089	0.015
m	0.240	0.073	0.261	0.058	0.308	0.039
r	0.243	0.033	0.244	0.035	0.273	0.029
W_R	102.768	17.389	102.232	10.231	110.970	9.880
W_S	106.473	4.837	100.468	4.865	102.350	4.290
I	0.007	0.005	-0.016	0.003	0.032	0.002

Table 3
Results from weighted least squares estimation of the earnings equation for New Zealand in 1986, 1991 and 1996

Regressor	Specification [a]			
	(1) *WLS*	(2) *WLS*	(3) *WLS*	(4) *WLS*
constant	1.791 [c]	1.551 [c]	1.471 [c]	1.447 [c]
	(0.007)	(0.020)	(0.026)	(0.027)
ln*u*		-0.098 [c]	-0.133 [c]	-0.142 [c]
		(0.008)	(0.010)	(0.010)
aged 26-40	0.353 [c]	0.353 [c]	0.353 [c]	0.353 [c]
	(0.003)	(0.003)	(0.003)	(0.003)
aged 41-60	0.393 [c]	0.392 [c]	0.393 [c]	0.393 [c]
	(0.003)	(0.003)	(0.003)	(0.003)
qualified	0.129 [c]	0.129 [c]	0.128 [c]	0.128 [c]
	(0.003)	(0.003)	(0.003)	(0.003)
non-Maori	0.054 [c]	0.052 [c]	0.057 [c]	0.057 [c]
	(0.004)	(0.004)	(0.004)	(0.004)
male	0.200 [c]	0.174 [c]	0.163 [c]	0.162 [c]
	(0.003)	(0.004)	(0.004)	(0.004)
full-time	-0.044 [c]	-0.044 [c]	-0.044 [c]	-0.044 [c]
	(0.003)	(0.003)	(0.003)	(0.003)
occupation dummies	Yes	Yes	Yes	Yes
ln*p*	0.059 [c]	0.040 [c]	0.020	0.024
	(0.004)	(0.005)	(0.017)	(0.017)
$\ln u_{1991}$				-0.045 [b]
				(0.021)
$\ln u_{1996}$				0.063 [c]
				(0.016)
urban area dummies	No	No	Yes	Yes
time dummies	Yes	Yes	Yes	Yes
number of observations	20,302	20,302	20,302	20,302
$\overline{R}^2$	0.707	0.710	0.714	0.714

[a] Significance at the 1%, 5% and 10% level is denoted by [c], [b] and [a], respectively. Standard errors are in parentheses.

Table 4
Testing for the significance of urban area and time effects

Null hypothesis	*F*-statistic [b]
$H_0: \mu = 0$ given $\lambda \neq 0$	10.76 [c]
$H_0': \lambda = 0$ given $\mu \neq 0$	149.88 [c]
$H_0'': \mu = \lambda = 0$	20.25 [c]

[b] Critical values are, at the 1% level, 1.71 for H_0, 4.60 for H_0' and 1.68 for H_0'', respectively.

Table 5
Results from instrumental variables estimation of the earnings equation for New Zealand in 1986, 1991 and 1996

Regressor	Specification				
	(1) *IV*	(2) *IV*	(3) *IV*	(4) *IV*	(5) *IV*
constant	2.220 [c]	2.671 [c]	2.657 [c]	2.623 [c]	2.080 [c]
	(0.099)	(0.136)	(0.120)	(0.141)	(0.077)
ln$u$	0.193 [c]	0.390 [c]	0.383 [c]	0.368 [c]	0.132 [c]
	(0.043)	(0.059)	(0.052)	(0.061)	(0.033)
aged 26-40	0.354 [c]	0.354 [c]	0.354 [c]	0.354 [c]	0.354 [c]
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
aged 41-60	0.394 [c]	0.396 [c]	0.395 [c]	0.395 [c]	0.394 [c]
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
qualified	0.128 [c]	0.129 [c]	0.129 [c]	0.128 [c]	0.128 [c]
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
non-Maori	0.058 [c]	0.058 [c]	0.058 [c]	0.058 [c]	0.058 [c]
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
male	0.250 [c]	0.302 [c]	0.300 [c]	0.296 [c]	0.234 [c]
	(0.012)	(0.016)	(0.014)	(0.016)	(0.009)
full-time	-0.044 [c]	-0.044 [c]	-0.044 [c]	-0.044 [c]	-0.044 [c]
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
occupation dummies	Yes	Yes	Yes	Yes	Yes
ln$p$	0.058 [c]	0.082 [c]	0.066 [c]	0.079 [c]	0.051 [c]
	(0.018)	(0.019)	(0.019)	(0.019)	(0.018)
urban area dummies	Yes	Yes	Yes	Yes	Yes
time dummies	Yes	Yes	Yes	Yes	Yes
number of observations	20,302	20,302	20,302	20,302	20,302
$\bar{R}^2$	0.699	0.676	0.677	0.679	0.704

Table 6
Results from estimation of the theoretical model earnings equation for New Zealand in 1986, 1991 and 1996

Regressor	Specification		
	(1) *WLS*	(2) *WLS*	(3) *IV*
constant	1.957 [c]	2.008 [c]	2.234 [c]
	(0.018)	(0.025)	(0.156)
$\ln(1-u)$	1.550 [c]	1.578 [c]	2.119 [b]
	(0.112)	(0.112)	(0.832)
aged 26-40	0.353 [c]	0.353 [c]	0.353 [c]
	(0.003)	(0.003)	(0.003)
aged 41-60	0.393 [c]	0.393 [c]	0.393 [c]
	(0.003)	(0.003)	(0.003)
qualified	0.128 [c]	0.128 [c]	0.128 [c]
	(0.003)	(0.003)	(0.003)
non-Maori	0.057 [c]	0.058 [c]	0.059 [c]
	(0.004)	(0.004)	(0.004)
male	0.161 [c]	0.160 [c]	0.147 [c]
	(0.004)	(0.004)	(0.020)
full-time	-0.044 [c]	-0.044 [c]	-0.044 [c]
	(0.003)	(0.003)	(0.003)
occupation dummies	Yes	Yes	Yes
$\ln p$	0.005	-0.036 [a]	-0.038 [c]
	(0.017)	(0.022)	(0.077)
m^{-1}		-0.015 [c]	-0.062 [c]
		(0.005)	(0.021)
urban area dummies	Yes	Yes	Yes
time dummies	Yes	Yes	Yes
number of observations	20,302	20,302	20,302
$\overline{R}^2$	0.714	0.714	0.712

Table 7

Testing for differences in the earnings equation between genders or occupations for New Zealand in 1986, 1991 and 1996

Regressor	Specification				
	(1) *WLS*	(2) *WLS*	(3) *IV*	(4) *IV*	(5) *WLS*
constant	2.069 [c]	1.827 [c]	2.185 [c]	1.781 [c]	2.011 [c]
	(0.043)	(0.037)	(0.061)	(0.065)	(0.029)
ln(1-u)	0.201	-0.180	1.046 [a]	-0.595	1.612 [c]
	(0.269)	(0.216)	(0.635)	(0.441)	(0.162)
ln(1-u)$_{OC2}$					-0.551 [c]
					(0.156)
ln(1-u)$_{OC3}$					-0.335 [c]
					(0.130)
ln(1-u)$_{OC4}$					-0.336 [b]
					(0.152)
ln(1-u)$_{OC5}$					0.010
					(0.139)
ln(1-u)$_{OC6}$					0.832 [c]
					(0.151)
ln(1-u)$_{OC7}$					-0.278
					(0.210)
ln(1-u)$_{OC8}$					-0.244
					(0.180)
aged 26-40	0.461 [c]	0.289 [c]	0.461 [c]	0.289 [c]	0.353 [c]
	(0.005)	(0.004)	(0.005)	(0.004)	(0.003)
aged 41-60	0.528 [c]	0.320 [c]	0.528 [c]	0.320 [c]	0.393 [c]
	(0.005)	(0.004)	(0.006)	(0.004)	(0.003)
qualified	0.126 [c]	0.112 [c]	0.126 [c]	0.112 [c]	0.128 [c]
	(0.005)	(0.003)	(0.005)	(0.003)	(0.003)
non-Maori	0.087 [c]	0.032 [c]	0.088 [c]	0.032 [c]	0.058 [c]
	(0.006)	(0.005)	(0.006)	(0.005)	(0.004)
male					0.162 [c]
					(0.004)
full-time	-0.126 [c]	-0.045 [c]	-0.126 [c]	-0.045 [c]	-0.044 [c]
	(0.007)	(0.003)	(0.007)	(0.003)	(0.003)
occupation dummies	Yes	Yes	Yes	Yes	Yes
ln$p$	0.028	-0.011	0.152 [b]	0.010	-0.036 [a]
	(0.036)	(0.025)	(0.060)	(0.043)	(0.022)
m^{-1}	-0.018 [b]	-0.004	-0.025 [a]	-0.004	-0.015 [c]
	(0.009)	(0.006)	(0.014)	(0.008)	(0.005)
urban area dummies	Yes	Yes	Yes	Yes	Yes
time dummies	Yes	Yes	Yes	Yes	Yes
number of observations	9784	10,518	9784	10,518	20,302
$\bar{R}^2$	0.713	0.768	0.711	0.768	0.716

Appendix: Definitions of the variables used

Y	Median income from all sources for persons in the subgroup over the past year.
H	Mean weekly hours worked by persons in the subgroup. This refers to the previous week at the 1986 and 1991 Censuses, but refers to the weekly hours "normally" worked in 1996.
t	Census year subscript taking the values 1986, 1991 and 1996.
i	Urban area subscript taking the following values:

1	Whangarei
2	Auckland
3	Hamilton
4	Tauranga
5	Rotorua
6	Gisborne
7	Napier/Hastings
8	New Plymouth
9	Wanganui
10	Palmerston North
11	Wellington
12	Nelson
13	Christchurch
14	Dunedin
15	Invercargill
16	Pukekohe
17	Tokoroa
18	Taupo
19	Whakatane
20	Hawera
21	Feilding
22	Levin
23	Kapiti
24	Masterton
25	Blenheim
26	Greymouth
27	Ashburton
28	Timaru
29	Oamaru
30	Gore

aged 26-40	Dummy variable identifying individuals aged 26-40.
aged 41-60	Dummy variable identifying individuals aged 41-60.
qualified	Dummy variable identifying individuals with post secondary school qualifications.
non-Maori	Dummy variable identifying Europeans.
male	Dummy variable identifying males.
full-time	Dummy variable identifying individuals who work 30 hours or more per week.

OC	Occupation group dummy variable taking the following values: 1 Health professionals 2 Teaching professionals 3 Office clerks 4 Salespersons, demonstrators and models 5 Personal and productive service workers 6 Skilled and semi-skilled production workers 7 Construction workers 8 Labourers and related elementary service workers
P	National consumer price index value in a particular year (base of 1 in 1986).
w	Real hourly wage for persons in the subgroup, calculated as $w = \frac{Y}{(52H)P}$.
U_M	Number of males unemployed in the particular urban area and year.
U_F	Number of females unemployed in the particular urban area and year.
L_M	Number of males in the labour force in the particular urban area and year. This is the sum of all males who are unemployed, employed full-time or employed part-time.
L_F	Number of females in the labour force in the particular urban area and year. This is the sum of all females who are unemployed, employed full-time or employed part-time.
N	Number of people aged between 15 and 59 in the particular urban area and year.
M	Number of people in an urban area whose address five years ago was elsewhere in New Zealand or overseas.
m	Gross in-migration rate of an urban area, calculated as $\frac{M}{N}$.
p	Average sale price of houses on the freehold open market for the half year ending June of the particular year, relative to the national value.
r	Proportion of private dwellings in a particular urban area and year that are rented.
W_R	Annual rainfall in millimetres expressed as a percentage of the normal value for 1961-1990.
W_S	Annual sunshine in hours expressed as a percentage of the normal value for 1951-1980.
I	Weighted average of the national growth rates over the previous year of employment in eight industry groups, namely, forestry and mining, manufacturing, electricity, building, wholesale and retail trade, transport, business services and personal services. The weights were calculated as the share of an urban area's employment in each industry in the particular year.

Sources: All variables were obtained from the 1981, 1986, 1991 and 1996 Censuses of Population and Dwellings, except the following: p, which was from Quotable Value New Zealand's *Urban Property Sales Statistics*, W_R and W_S, which were supplied by the National Institute of Water and Atmospheric Research, and I, which was derived from both the census and Quarterly Employment Survey.

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